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Peter Sebastian Eppinger

aus Göppingen

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Dekan:	Prof. Dr. rer. soc. Josef Schmid
1. Gutachter:	Prof. Dr. rer. soc. oec. Wilhelm Kohler
2. Gutachter:	Prof. Dr. rer. pol. Bernhard Boockmann

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Chapter 1

Introduction

The economic research presented in this dissertation is motivated by two important empirical observations that have shaped international trade and investment in the last decades. The first observation is the great trade collapse following the global financial crisis of 2008, which caused a drop in the volume of international trade that was substantially larger than the decline in production in many countries. This event raises the natural question of how the financial aspect of the crisis, the credit crunch, has affected real economic activity in general, and cross-border trade in particular. Chapters 2 to 4 revolve around this question. The second observation is the fragmentation of production across the globe, which has gained importance over several decades. This fragmentation, also known as offshoring, involves vast amounts of international trade in intermediate goods, and increasingly also trade in services. Furthermore, this type of trade is frequently accompanied by foreign direct investment (FDI), since many of the largest firms decide to organize parts of global production chains within their firm boundaries. The rise of offshoring and multinational firms raises many intriguing questions concerning the optimal organization of global production and its implications for firm performance and employment. These issues are investigated in Chapters 5 to 7. In this introduction, I begin by describing the broader goal of my research, then I provide a brief overview of each chapter, before highlighting the common thread in my theoretical and empirical methodology. I conclude by outlining important implications for future work.

The overarching goal of my research agenda, initiated by this dissertation, is to contribute to a better understanding of the determinants, the organization, and the consequences of international trade and investment in the modern era. Due to its potentially adverse effects, e.g. on income inequality, social standards, and the environment, a number of anti-globalists, who regained popularity in the mid-2010s, are seeking to stop globalization. But turning back the wheel of time is neither a viable option nor a very promising one. The overall efficiency gains from trade are obvious to most economists and also to a large share of the world population. Therefore, the main open question is not how to stop globalization from happening, but how it can be organized in an efficient and inclusive way. Several obstacles to an efficient organization of global production and international trade remain important today, including financial market imperfections, contractual frictions, and trade costs (which are still particularly high for many services). This dissertation sheds new light on several important aspects of these obstacles in order to improve our understanding of how firms, workers, and policy makers can optimally respond to globalization.

Chapter 2, entitled *The Great Trade Collapse and the Spanish Export Miracle: Firm-level Evidence from the Crisis*, is based on joint work with Nicole Meythaler, Marc-Manuel Sindlinger, and Marcel Smolka, which has been accepted for publication in *The World Economy*. The chapter investigates the

micro-structure of international trade in Spain before, during, and after the global financial crisis. Using a panel dataset of Spanish manufacturing firms and descriptive as well as econometric analysis, it uncovers two striking empirical facts. First, Spanish exports were surprisingly resilient to the crisis. We find no adverse effects of the financial crisis on foreign market entry or exit, but a considerable increase in the export intensity of firms after the peak of the crisis. These findings constitute novel evidence of the so-called "Spanish export miracle" at the micro level. Second, exporting firms were significantly more likely to survive the crisis, and they substantially increased their performance advantage relative to other firms (in terms of employment, output, and productivity) over the crisis years. The chapter concludes by discussing the role of macroeconomic devaluation and microeconomic substitution of firm sales across different markets in rationalizing these observations.

Chapter 3, entitled *External Financial Dependence and Firms' Crisis Performance across Europe*, is based on joint work with Katja Neugebauer. It analyzes the impact of the credit crunch in the global financial crisis on several dimensions of firm performance, including employment, productivity, and exports. It starts from the standard approach to measuring external financial dependence (EFD) based on the seminal contribution by Rajan and Zingales (1998). Many studies following this approach have transferred an EFD index based on US data to other countries under the assumption of a stable ranking of industries in terms of EFD across countries. We find only weak correlations of the US-based index with a direct question on financial dependence from a comparable firm survey in seven European countries. Also, correlations of the survey-based measure of EFD across countries are weak, casting doubt on the assumption of a stable EFD ranking across countries. Both measures are then employed in an empirical application to assess the impact of the financial crisis on the real economy. We use a difference-in-differences regression model to test whether the financial crisis affected firms differently across industries with varying degrees of EFD. We find that only the country-specific, survey-based measure of EFD can identify a stronger negative crisis impact on firm performance in financially dependent industries.

Chapter 4, entitled *Firm Exports, Foreign Ownership, and the Global Financial Crisis*, is based on joint work with Marcel Smolka. It demonstrates how foreign ownership by multinational firms has stabilized the export performance of Spanish firms in the financial crisis. We begin by setting up a trade model of heterogeneous firms facing credit constraints, as in Manova (2013). Into the model, we introduce the possibility for foreign-owned firms to access foreign capital markets via their multinational parents. The model predicts that the foreign-owned firms exploit this financial advantage in order to export greater volumes and shares of their output. This difference in export shares across foreign-owned and domestically owned firms is predicted to increase in a credit crunch, particularly among financially vulnerable firms. We investigate these predictions, using the same panel data on Spanish firms as in Chapter 2, by combining a difference-in-differences model with firm fixed effects and a propensity score reweighting approach. We further exploit rich information on firms' debt situation at the onset of the crisis to pin down the credit channel. We find strong and robust evidence that the positive effect of foreign ownership on firms' export shares was magnified in the crisis, and more so among financially vulnerable firms. Our data further suggest that the credit channel is intertwined with a market access channel, which also gained importance in the crisis. Foreign-owned firms increasingly relied on their foreign parents' distribution networks for exporting in the crisis, which constituted an alternative way to save on trade costs and circumvent credit constraints. By focusing on the role of foreign ownership in shaping firms' export behavior, this chapter lies at the intersection between international trade and FDI, and it builds a bridge to the second part of my dissertation.

Chapter 5, entitled *Optimal Ownership and Firm Performance: Theory and Evidence from China's FDI Liberalization*, is based on joint work with Hong Ma. The chapter provides a first quantification of the gains from the optimal allocation of ownership rights within multinational firms. We identify these gains by exploiting a unique liberalization of China's restrictions on foreign ownership in 2002. To guide the empirical analysis, we develop a theoretical model building on the property-rights theory of the multinational firm by Antràs (2003) and Antràs and Helpman (2004). In the model, a foreign multinational firm and its Chinese affiliate need to contribute relationship-specific and non-contractible investments to a joint production process in China. We model the continuous ownership decision of the foreign firm in its affiliate and introduce the Chinese policy restrictions as an upper bound on this foreign ownership share. The model predicts that FDI liberalization triggers an increase in the foreign ownership share to the optimal level, which induces a shift in investment incentives between the two parties, resulting in efficiency gains and higher firm output. Our empirical analysis compares firms increasing their foreign ownership share after liberalization to other foreign-owned firms within the liberalized industries in a difference-in-differences model augmented by propensity score reweighting. Thereby, we can identify the performance effects of a move to optimal ownership from several other benefits of foreign ownership, such as the financial advantage identified in Chapter 4. We find that the move to optimal ownership caused a shift in input ratios in line with our model and substantial output gains of up to 34% between 2001 and 2003. These output gains are even twice as large after six years, and they are also accompanied by productivity gains over the medium term. Our findings fill a major gap in the literature identified by Antràs (2015): We provide direct evidence that multinational firms' organizational decisions matter substantially for firm performance in practice, and we deliver a first estimate of the size of the gains from optimal ownership.

Chapter 6, entitled *Contracting Institutions and Firm Boundaries*, is based on joint work with Bohdan Kukharskyy. It describes and rationalizes a novel stylized fact on firm boundaries: Firms tend to integrate their subsidiaries more deeply into their boundaries in countries with better contracting institutions. We establish this pattern in a global micro data set on ownership shares across firm pairs, which provides an exceptionally broad and detailed view of both domestic and multinational ownership links. While the observed pattern is seemingly at odds with a transaction cost view of the firm, which would suggest that firms rely more on markets where contracts can more easily be enforced, we show that this pattern can be rationalized by a model based on the seminal property-rights theory of the firm by Grossman and Hart (1986) and Hart and Moore (1990). We model continuous ownership decisions, akin to the theory in Chapter 5, but here our approach is different and focuses on the role of varying degrees of contracting institutions in the subsidiary's country and varying relationship-specificity of investments. Our model predicts that better contracting institutions in subsidiaries' countries favor higher ownership shares, in particular in relationship-specific industries. We proceed by conducting a large-scale investigation of these predictions in our global micro data set and find strong support for our model.

Chapter 7, entitled *Service Offshoring and Firm Employment*, is single-authored. It is concerned with the labor market impact of service offshoring, a form of fragmentation that has gained importance in recent years and triggered fears of job losses in many developed countries. However, direct causal evidence on the employment effects of service offshoring is scarce, partly due to the lack of appropriate data and due to obvious endogeneity concerns. I fill this void by combining a unique data set, covering almost the entire universe of German firms' service offshoring activities over the years 2002-2013, with a novel instrumental variables (IV) approach. Following Hummels et al. (2014), the firm-specific IV

exploits detailed export supply shocks by partner countries and service types over time, which affect firms differentially according to their specific import mix. I find strong evidence of positive employment effects from service offshoring at the firm level. These findings can be rationalized by cost savings from offshoring predicted by the seminal trade in tasks model by Grossman and Rossi-Hansberg (2008). In line with the logic of this model, the employment gains are larger in firms that have previously offshored more services and hence benefit more from further cost reductions.

In terms of methodology, this dissertation combines formal theoretical modeling with rigorous empirical analysis based on state-of-the-art microeconometric methods. Several chapters build on modern trade theory, which has stressed the importance of firm heterogeneity in international trade and investment since the pioneering work by Melitz (2003). Chapter 4 examines the exports of heterogeneous firms in the presence of financial frictions and extends the model by Manova (2013) by allowing foreign-owned firms to access foreign capital markets via their parent company. Chapter 5 models the ownership decisions of heterogeneous multinational firms by generalizing the framework of Antràs and Helpman (2004). Firm heterogeneity also plays an important role for rationalizing the empirical findings obtained in Chapters 2, 3, and 7, even though the theory is not explicitly formalized in these chapters. Only Chapter 6 abstracts from firm heterogeneity in the theoretical analysis to keep the model simple and focused on the core issue of contracting institutions, but the theoretical results in this chapter extend in a straightforward manner to a setting in which firms differ in terms of productivity.

The empirical methodology employed in this dissertation strongly reflects my focus on the microstructure of international trade and investment. A common theme of all chapters is that they make use of micro data at (or even below) the firm level, either from a single country (Spain in Chapters 2 and 4, China in Chapter 5, and Germany in Chapter 7) or from multiple countries (in Chapters 3 and 6), which provide a rich picture of world trade and multinational firm activity. The econometric toolkit used to establish causality in the different chapters represents the state of knowledge in the discipline. It encompasses panel data econometrics, including difference-in-differences regression (in Chapters 2 to 5), IV estimation (in Chapters 6 and 7) , as well as propensity score matching and reweighting techniques (in Chapters 4 to 7). These methods are applied in an innovative way to the various datasets in order to gain reliable insights into optimal firm behavior and into the consequences of trade, offshoring, and FDI.

This dissertation points to several, quantitatively important benefits of globalization. My findings indicate that integration into the world economy through exporting and FDI was beneficial for firm performance in the global financial crisis. And they further suggest that allowing for multinational firm activity may serve to reduce inefficiencies that arise due to frictions in financial markets or imperfect contracts, and can thereby improve real outcomes. These results need to be complemented by evidence from other countries and further structural and theoretical analysis in order to explore their aggregate implications and arrive at welfare conclusions. Beyond the contributions of this dissertation, many important and pressing questions revolving around international trade and global production remain to be answered in future research. Provided that the current trends of financial globalization and innovation, the growing importance of multinational firms, and the increasing tradability of services will continue, the issues that I have discussed are likely to remain important in the future. I hope that the novel insights, newly accessed datasets, and methodological advancements contributed by this dissertation will provide a good starting point for addressing these upcoming challenges.

Bibliography

Antràs, P., *Global Production: Firms, Contracts, and Trade Structure*, Princeton University Press, 2015.

Antràs, Pol, "Firms, Contracts, And Trade Structure," *The Quarterly Journal of Economics*, 2003, *118* (4), 1375–1418.

__ **and Elhanan Helpman**, "Global Sourcing," *Journal of Political Economy*, 2004, *112* (3), 552–580.

Grossman, Gene M. and Esteban Rossi-Hansberg, "Trading Tasks: A Simple Theory of Offshoring," *American Economic Review*, 2008, *98* (5), 1978–1997.

Grossman, Sanford J. and Oliver D. Hart, "The Costs and Benefits of Ownership: A Theory of Vertical and Lateral Integration," *Journal of Political Economy*, 1986, *94* (4), 691–719.

Hart, Oliver and John Moore, "Property Rights and the Nature of the Firm," *Journal of Political Economy*, 1990, *98* (6), 1119–1158.

Hummels, David, Rasmus Jørgensen, Jakob Munch, and Chong Xiang, "The Wage Effects of Offshoring: Evidence from Danish Matched Worker-Firm Data," *American Economic Review*, 2014, *104* (6), 1597–1629.

Manova, Kalina, "Credit Constraints, Heterogeneous Firms, and International Trade," *Review of Economic Studies*, 2013, *80* (2), 711–744.

Melitz, Marc J., "The Impact of Trade on Intra-Industry Reallocations and Aggregate Industry Productivity," *Econometrica*, 2003, *71* (6), 1695–1725.

Rajan, Raghuram G. and Luigi Zingales, "Financial Dependence and Growth," *American Economic Review*, 1998, *88* (3), 559–86.

Chapter 2

The Great Trade Collapse and the Spanish Export Miracle: Firm-level Evidence from the Crisis*

We provide novel evidence on the micro-structure of international trade during the 2008 financial crisis and subsequent global recession by exploring a rich firm-level data set from Spain. The focus of our analysis is on changes at the extensive and intensive firm-level margins of trade, as well as on performance differences (jobs, productivity, and firm survival) across firms that differ in their export status. We find no adverse effects of the financial crisis on foreign market entry or exit, but a considerable increase in the export intensity of firms after the financial crisis. Moreover, we find that exporters were more resilient to the crisis than non-exporters. Finally, while exporters showed a significantly more favorable development of total factor productivity after 2009 than non-exporters, aggregate productivity declined substantially in a large number of industries in Spanish manufacturing. We also briefly explore two factors that might help explain the surprisingly strong export performance of Spain in the aftermath of the great trade collapse: improved aggregate competitiveness due to internal and external devaluation and a substitutive relationship between domestic and foreign sales at the firm level.

*This chapter is based on joint work with Nicole Meythaler, Marc-Manuel Sindlinger, and Marcel Smolka, which has been accepted for publication in *The World Economy*.

2.1 Introduction

The global recession that followed the 2008 financial crisis continues to place a heavy burden on the world economy. One important aspect of the crisis that has caught a lot of attention among both policy-makers and economists was the sudden, synchronized, and more than proportional decline in global trade relative to global production – the so-called "great trade collapse" (Baldwin, 2009). While the causes and consequences of this event have been subject to extensive debate, the available evidence derives largely from aggregate data rather than from detailed firm-level data.[1] This is somewhat surprising, since the issue of firm heterogeneity and the fact that only a fraction of firms access foreign markets have become cornerstones of modern trade literature. To what extent have firms decided to leave foreign markets in response to the crisis? Are firms today relying less on imports and exports than before the crisis? And did firms perform better or worse during the crisis if they were active on foreign markets? A fine-grained analysis of the micro-structure of international trade in the years surrounding the financial crisis can provide answers to these questions by uncovering patterns in the data that would go unnoticed in an analysis based on aggregate data alone.

The objective of this paper is to provide such a fine-grained analysis using a representative sample of Spanish manufacturing firms over the period 2005-2012. Spain is a particularly interesting case to look into. On the one hand, the country was deeply affected by the financial crisis and subsequent recession. In the first half of 2009, real industrial production contracted by 21.4% relative to the first half of 2008.[2] Importantly, Spain went through very difficult times also after the financial crisis. Following zero growth in 2010, total production contracted again in 2011-2013, reflecting what is sometimes called a "double dip" recession. On the other hand, and perhaps surprisingly, the country showed a relatively strong export performance over the crisis period. Figure 2.1 demonstrates that the Spanish economy was able to improve its competitive position on international markets compared to other economies in Europe. For example, between 2007 and 2013 exports from Italy and France *decreased* by 10% and 7%, respectively. In contrast, exports from Spain *increased* by 13% over the same period. This development (celebrated by some as the "Spanish export miracle"[3]) put Spain ahead of not only other countries in economic turmoil, but also countries that quickly returned to economic growth after 2008, such as Germany and the UK.

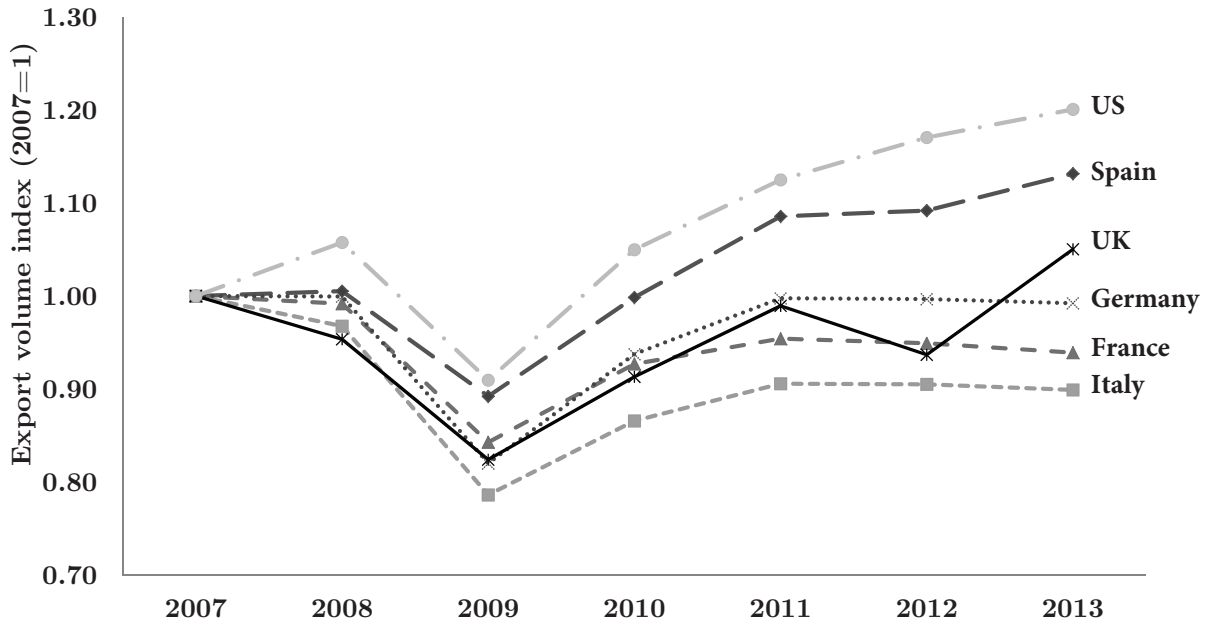
In this paper, we zoom in on the Spanish crisis experience. Adopting a micro-level perspective on Spanish firms allows us to investigate two important issues related to the crisis episode that remain obscure in aggregate data. First, we can disentangle the effects of the crisis at the extensive and the intensive firm-level margins of trade, i.e., we can separate a firm's decision to access foreign markets at all from the volume of a firm's exports and imports (as shares of its total sales and purchases, respectively). This distinction is crucial for the purpose of our analysis and allows us to address two interesting questions: Did aggregate trade decline because of firms exiting foreign markets, or because of a contraction in firm-level trade volumes? And was the subsequent recovery and export boom due to firms scaling up their exports, or due to new firms entering foreign markets? These questions are important because a destruction of cross-border trade linkages at the firm level can have long-lasting adverse effects on the economy (Monarch and Schmidt-Eisenlohr, 2016), and these effects are not expected for adjustments

[1] Important contributions using aggregate or sector-level trade data include Chor and Manova (2012) and Eaton et al. (2016). We discuss the existing micro-level evidence further below.

[2] Annual industrial production in 2009 declined by 16.2% relative to 2008. For real manufacturing exports, the same number is 21.2%. The data come from the Spanish Instituto Nacional de Estadística (INE).

[3] See, for example, the article "El *milagro* del sector exterior de España: admirable, pero con algunos claroscuros," published on May 10th, 2013, in the Spanish daily newspaper elEconomista.es.

Figure 2.1: *Export volumes, 2007-2013*



Note: The data are taken from the World Development Indicators (WDI) provided by the World Bank.

at the intensive margin. Conversely, the entry of new exporters might soften the adverse effects of the crisis by increasing the potential for future economic growth, because new exporters in Spain are more likely to engage in productivity-enhancing technology upgrading than non-exporters (Hanley and Pérez, 2012). In terms of methodology, we follow the literature estimating firm-level models of exporting and importing based on panel data (e.g. Bernard and Jensen, 1997, 1999). However, this literature is typically interested in the evolution of firm-specific characteristics (e.g. productivity, management, or labor force composition) and how these influence firms' export and import decisions, respectively. In contrast, our focus is on the direct effects of the financial crisis and subsequent recession, i.e., changes in macro conditions that are beyond the control of individual firms.

The second issue we investigate are differences in firm performance and crisis resilience between exporting and non-exporting firms. Since exporting firms are known to be larger and more productive, on average, than non-exporting firms, their behavior can be important for aggregate outcomes. It is thus crucial to understand the performance of these firms in times of exceptional economic distress. Does exporting to foreign markets make firms more immune to shocks, or does it make them more vulnerable? This is an interesting question that should be settled empirically, as there exist theoretical arguments supporting either view. While allocating sales across various markets, domestically and abroad, insures the firm against an adverse demand shock in one market, there is also a substantial risk involved in exporting (e.g. currency risk, non-payment risk, transport risk etc.), and relying on foreign markets in times of a globally synchronized crisis might prove particularly harmful to firm performance. To answer this question, we estimate differences between exporters and non-exporters in terms of size, productivity, and survival, so-called exporter premia (Bernard and Jensen, 1999), and we study the evolution of these premia over the crisis years. Importantly, increasing exporter premia during the crisis could be taken as an indication that economies become less vulnerable to economic shocks through exporting. Furthermore, if it is primarily non-exporting firms that are forced to exit the market due to the crisis, then this might (in the medium to long run) induce a reallocation of resources away from non-exporting firms towards

exporting firms, where they are put to more efficient use. The same logic applies if for non-exporters the evolution of productivity through the crisis and afterwards is less favorable than for exporters. Hence, differences in crisis performance of exporters vs. non-exporters are relevant also for the long-run growth perspective of the Spanish economy.

The main results of our empirical analysis can be summarized as follows. First, the sharp drop in international trade that the Spanish manufacturing sector experienced in 2009 took place at the intensive margin, not the extensive margin. This means that, while the financial crisis caused a strong reduction in firm-level imports and exports, it did not prompt firms to exit foreign markets altogether. In the years *after* the financial crisis, we do see changes at the extensive margin, but we see more, rather than less, firms starting to enter foreign markets. As a result, there is now a larger share of firms involved in international trade than before the crisis. Furthermore, firms have diversified their export portfolios to include more distant destinations outside the European Union.

Second, while firms active in the export market saw their export volumes plummet in the financial crisis, this decline was not limited to exports, but rather, it was visible to the same extent in their domestic sales. This observation might seem surprising in light of the discussion about the great trade collapse. Moreover, the decline in exports was fully made up for (and even overcompensated) already by 2011. Those firms that entered the financial crisis as exporters have in fact been allocating ever larger shares of their production to foreign markets over the past few years. It seems that these firms have effectively compensated for the lack of domestic demand by expanding their sales abroad. In this sense, firms in the Spanish manufacturing sector are on average more, not less, 'globalized' today than they were before the financial crisis.

Third, we find that it made a significant difference for key economic performance indicators (such as jobs, productivity, and survival) whether or not firms were active on export markets when the crisis hit the Spanish economy. While all firms strongly reduced their output and laid off large numbers of workers during and after the financial crisis, firms that entered the crisis as exporters (and continued to export throughout the crisis years) saved more jobs, stayed more productive, and were more likely to survive. One of the more alarming findings is that from 2007 to 2009 firms' average total factor productivity (TFP) deteriorated by around 15%. For non-exporters, TFP continued to decline by another 15% from 2009 to 2011. Exporters, in contrast, maintained about the same level of productivity in 2011 as they had in 2009. Our analysis also shows that *aggregate* TFP in the Spanish manufacturing sector declined as a result of the crisis.

After having documented these facts, we explore two possible explanations for the favorable development of Spanish exports after the financial crisis. The point of departure of both explanations is the fact that aggregate demand was hit much harder in Spain than in most other large economies, both within and outside the European Union. We argue (and provide evidence for the idea) that, as a result of this, the Spanish economy has become more competitive internationally through internal as well as external devaluation. This is the first explanation we discuss in our paper. The second explanation is that the more than proportional decline in domestic demand has prompted firms to substitute domestic with foreign sales. Importantly, such a response is not implied by the standard Melitz (2003) model, but has recently been rationalized in trade models in which short-run production costs are convex in total output (Blum et al., 2013; Soderbery, 2014; Vannoorenberghe, 2012). We present some evidence drawn from our data that is consistent with this idea. However, we should like to emphasize that a rigorous causal analysis or an investigation into the relative importance of the two explanations for the strong export performance

of the Spanish economy is beyond the scope of our paper. In a similar vein, we do not wish to claim that these two explanations are the most important, let alone the only, factors behind the strong export performance of the Spanish economy.

Our paper contributes to the small empirical literature that investigates firm behavior in response to the financial crisis with a focus on firms' trading activities. Two prominent studies in this literature using French and Belgian data, respectively, are by Bricongne et al. (2012) and Behrens et al. (2013), who carefully gauge the crisis-induced drop in international trade along the dimensions of firms, products, and trading partners.[4] Closely related to our paper are the studies by Giri et al. (2014) and Álvarez and Sáez (2014), which provide evidence on exports and firm performance during and after the crisis using Mexican and Chilean firm-level data, respectively. Studies with a particular focus on firm survival over the crisis years depending on firms' trading activities are Costa et al. (2014, for Italy) and Görg and Spaliara (2014, for the UK). There seems to be a consensus emerging from this literature that most of the crisis adjustment in firm exports took place at the intensive margin. Exporting firms are typically found to be more resilient to the crisis and there is some evidence that firms' financial conditions played a relevant role for their crisis performance.[5] We complement this literature with evidence on both importing and exporting firms in Spain during and after the financial crisis. More generally, our paper fits into the large empirical literature analyzing firm heterogeneity in international trade. Reviews of this literature can be found in Bernard et al. (2012) and Greenaway and Kneller (2007).

Our paper also relates to an ongoing discussion about the export performance of the Spanish economy in the period *before* the financial crisis. Antràs (2011) observes that the share of Spanish exports in world trade was stable throughout the period 2000-2010 despite rising unit labor costs relative to other Eurozone countries between 2000 and 2008. As trade models with homogeneous firms cannot account for this fact, he argues in favor of an explanation based on firm heterogeneity. Correa-López and Doménech (2012) suggest that a number of strategic actions taken at the firm level (e.g. technology and skill upgrading, product innovations, and financial optimization) contributed to the internationalization of Spanish firms over the period 1990-2010. In contrast to these studies, we focus explicitly on the years surrounding the financial crisis, which involved a number of particular challenges and changes in the competitive position of Spanish firms. In addition, we document and analyze differences in firm performance and crisis resilience in relation to firms' export activities.

The rest of the paper is organized as follows. In Section 2.2, we describe the data used in our empirical analysis. Section 2.3 presents the main analysis of our paper. We start with a decomposition of changes in total trade into extensive and intensive margins in Section 2.3.1. Section 2.3.2 analyzes the probability of firms to engage in exporting and importing before, during, and after the financial crisis. Section 2.3.3 proceeds by analyzing how firms allocated their sales across the foreign and the domestic market and which share of their purchases they chose to source from abroad (rather than domestically). In Section 2.3.4, we take up the issue of firm competitiveness and crisis resilience by investigating performance differences depending on firms' export status. Section 2.4 provides a discussion of two important factors that are likely to have contributed to the strong export performance of the Spanish economy. Section 2.5 concludes.

[4] Abreha et al. (2016) provide evidence for Denmark.

[5] For micro-level evidence regarding the credit shock on exports in the global financial crisis see also Görg and Spaliara (2013) and Paravisini et al. (2014).

2.2 Firm-level Data

The primary data source for our analysis is the Encuesta Sobre Estrategias Empresariales (ESEE, or Survey on Business Strategies). The ESEE is an annual survey of about 2,000 Spanish manufacturing firms with 10 or more employees. It includes rich information on strategic firm decisions (such as pricing, international trade and investment activities, or innovation strategies) along with key items of firms' balance sheets as well as profit and loss statements. Importantly, the ESEE is a panel data set representative for the Spanish manufacturing sector at large and covering the period 1990-2012. This data set allows us to provide a comprehensive, high-resolution perspective on the micro-structure of international trade, and to portray the evolution of Spanish manufacturing over the years before, during, and after the financial crisis. The initial sampling of the data in 1990 had a two-tier structure, combining exhaustive sampling for firms with more than 200 employees and stratified sampling for firms with 10-200 employees. In later years, special efforts have been devoted to minimizing the incidences of panel exit as well as to including new firms through refreshment samples aimed at preserving the representativeness of the data.[6]

The sample we use for our analysis covers the period 2005-2012 (unless indicated otherwise). It consists of an unbalanced panel of more than 3,100 firms, roughly 800 of which are observed throughout the entire period. The ESEE uses the main activity (industries at the 2-digit level of the NACE Rev. 2 classification)[7] and the size group of firms (in terms of the number of employees) as stratification variables. It distinguishes between 20 different industries and six different size groups defined by the average number of workers employed during the year: 10-20; 21-50; 51-100; 101-200; 201-500; >500. As far as the descriptive analysis is concerned, we employ sampling weights to account for the sampling scheme used to collect the data.[8] When applying regression methods in our analysis, we use fixed effects for the sampling strata (defined by pairs of industries and size groups) in order to obtain consistent estimation of the parameters of interest. Summary statistics of the most important variables used in our empirical analysis can be found in Table 2.A.1 in Appendix 2.A. Throughout the paper, we express all value variables in constant 2010 prices using industry-level price indexes from INE (similarly to Guadalupe et al., 2012).

A particularly important variable in our analysis is a firm's total factor productivity (TFP). We obtain TFP as a firm-specific and time-varying residual from industry-specific Cobb-Douglas production functions, which we estimate by the consistent three-step procedure proposed by Olley and Pakes (1996). This procedure derives from a dynamic model of firm behavior incorporating firm-specific productivity differences that exhibit idiosyncratic changes through time. The model tackles a potential endogeneity issue due to simultaneity between input choices and unobserved productivity shocks by using firm-specific capital investments as a proxy variable. In contrast to an alternative approach proposed by Levinsohn and Petrin (2003), the model by Olley and Pakes (1996) also takes into account the issue of sample selection due to firms entering and exiting the market. This is potentially important for the period of economic turbulence considered in our analysis. We estimate industry-specific production functions by using annual ESEE data on a firm's value added, investment, capital stock, labor employment, and mar-

[6] More detailed information on the design, management, and sampling properties of the survey are available from the Spanish Sociedad Estatal de Participaciones Industriales (SEPI) foundation at https://www.fundacionsepi.es/esee/en/epresentacion.asp.

[7] Until 2009, the survey defined industries according to the NACE Rev. 1 classification. We accommodate the two classifications based on concordance information provided by the SEPI foundation.

[8] Sampling weights are based on the composition of the population of Spanish firms in 2010, taken from INE: http://www.ine.es/en/inebmenu/mnu_empresas_en.htm.

ket exit decisions over the period 2000-2012. Value added is the sum of the total production value plus other operating income (i.e., income from rent and leasing, industrial property, commissions, and certain services), minus the total expenditure on intermediate inputs and external services. Investment is the total investment value in tangible fixed assets (land, buildings, and equipment). The capital stock is the value of tangible fixed assets. Labor employment is measured in effective working hours. As regards exit decisions of firms, our data allow us to distinguish firms shutting down production from those that stay in the market but exit the ESEE panel for other reasons.[9]

A brief comment on the measure of labor employment that we use in our analysis seems in order. In contrast to many other firm-level data sets used in the literature, the ESEE data include an almost ideal measure of labor employment, namely effective working hours. This reduces the possibility of measurement error and thus mitigates endogeneity concerns in the estimation of firm-level TFP. Since the ESEE data also include a more common measure of labor employment (the average number of workers a firm employs during a year), we can investigate different firm-level margins of labor adjustment, viz. the number of workers (or jobs) and the number of effective working hours. Interestingly, we find very small differences in the within-firm variation between these two variables. This implies that the reductions in effective working hours observed at the firm level are fully attributable to workers being laid off and jobs being lost.

While the focus of our analysis is on Spain, we occasionally also draw on firm-level survey data from the EFIGE project, which was designed to enhance the understanding of how European firms and economies are affected by the process of globalization (Altomonte and Aquilante, 2012).[10] This data set enables us to compare firms in Spain with firms in six other European countries in the year 2008: Austria, France, Germany, Hungary, Italy, and the UK. It includes 14,444 firms and, importantly, it is also representative of manufacturing firms with more than 10 employees in these countries. The focus of the EFIGE data is on the experience and competition of firms in foreign markets, as well as their responses to the challenges posed by the financial crisis. Among the firm-level information included in the data set are: sales, employment, innovation activities, international investment, and values of imports and exports.[11] The data set also includes some questions on the changes in sales, employment, and trade values in the year 2009.

2.3 Empirical Analysis

2.3.1 Decomposition of changes in total trade

We start by decomposing changes in both exports and imports in the Spanish manufacturing sector into extensive and intensive margins. We do this along the lines of Bernard et al. (2009) and Behrens et al. (2013) using the ESEE firm-level data set. Total exports in any given year t can be written as the product of the number of exporters (the extensive margin) in t and the average value of exports per exporting firm (the intensive margin) in t: $exports_t = number\ of\ exporters_t \times average\ exports_t$. Hence, we can

[9] Detailed results from these TFP estimations are available from the authors upon request.

[10] The EFIGE project is called "European firms in a global economy: Internal policies for external competitiveness."

[11] Altomonte et al. (2012) provide a full-fledged analysis of firms in different countries based on the EFIGE data set. Crespo et al. (2011) use the data to specifically compare firms in Spain with firms in other European countries.

decompose changes in total exports (and analogously for imports) as follows:

$$\frac{\Delta exports_t}{exports_{t-1}} \approx \frac{\Delta number\ of\ exporters_t}{number\ of\ exporters_{t-1}} + \frac{\Delta average\ exports_t}{average\ exports_{t-1}}, \tag{2.1}$$

where $\Delta exports_t \equiv exports_t - exports_{t-1}$ (and accordingly for the number of exporters and average exports).

Table 2.1 shows the decomposition of annual changes in total exports and imports according to Equation (2.1) over the period 2006-2012. Total trade in the Spanish manufacturing sector contracted in both years 2008 and 2009, with a drop of more than 15% in exports and more than 20% in imports in the main crisis year 2009. Importantly, these changes took place almost exclusively at the intensive margin of trade. More specifically, on the export side, average exports per firm decreased by 15% in 2009, which means that the intensive margin almost fully explains the overall drop in exports. On the import side, trade at the extensive margin even increased slightly in 2009, counteracting the drop at the intensive margin. It is interesting that exports quickly recovered in 2010 and 2011, in particular at the intensive margin, while imports had not recovered by the year 2012 (the last year of data we use in our analysis). In 2011, both exports and imports decreased slightly at the extensive margin, but increased again quite spectacularly in 2012.[12]

Table 2.1: *Decomposition of annual changes in total trade in Spanish manufacturing*

	Exports			Imports		
	Total	Extens. margin	Intens. margin	Total	Extens. margin	Intens. margin
2006	9.74	8.20	-1.09	19.26	6.84	11.39
2007	13.11	0.99	8.15	15.05	0.44	15.08
2008	-0.27	2.59	-2.69	-4.18	-0.86	-2.83
2009	-15.36	-0.47	-15.14	-20.26	0.12	-20.27
2010	6.86	3.26	4.02	5.28	1.91	5.48
2011	10.63	-4.15	15.16	3.74	-2.65	7.25
2012	3.62	13.90	-8.78	-5.32	9.48	-13.79

Note: This table shows annual percentage changes in total exports and imports in Spanish manufacturing, as well as a corresponding decomposition into extensive and intensive margins. Sampling weights apply. Source: Authors' calculations based on ESEE data.

Two comments on this decomposition exercise are in order. First, the changes at the intensive margin that we examine here are changes that took place at the level of the firm. Hence, they may include adjustments at several additional extensive margins that are only visible at a more disaggregated level: the number of products traded, the number of destination and source countries, and the number of buyers and sellers for each firm. While we cannot disentangle these margins for Spain, Behrens et al. (2013) find in more disaggregated data from Belgium that even within firm-country-product cells the intensive margin accounts for 97% of the overall drop in Belgian exports caused by the crisis.

Second, in the above decomposition we hold all prices constant, so that changes at the intensive margin of trade are due to changes in the quantities traded (rather than changes in the prices of traded products). To shed some light on the evolution of nominal trade values, we examine annual variations in both sales prices and input prices. We find that on average firms lowered their sales prices in 2009,

[12]The decrease in both exports and imports at the intensive margin of trade in 2012 stems in part from the comparatively low volume of trade by foreign market entrants in that year.

but only by 0.59%.[13] In contrast, the prices of inputs continued to rise in 2009, though at a lower rate (1.70%) than before or after 2009. Overall, we can thus say that nominal trade values dropped sharply in 2009, but that this drop is due to a reduction in the quantities traded rather than a decline in prices.

2.3.2 Foreign market entry and exit

In this section, we focus on the extensive margin of trade. What share of firms in the Spanish manufacturing sector is active on foreign markets? And how did this share develop over the recent period of financial and economic turmoil? When looking at the full sample of firms, we find that in the pre-crisis period 2005-2008 on average 46% of all firms were exporters, while 43% were importers; see Figure 2.2(a). We observe significant overlap between exporter and importer status, reflected in 30% of firms in 2005 being engaged in both exporting and importing at the same time (not depicted). This suggests that exporting and importing are complementary activities at the level of the firm, an issue that has been taken up in recent research and to which we will return below. Two observations stand out. First, there was only a very small decrease in the shares of exporting and importing firms in 2009, following the peak of the financial crisis. Second, both shares rose sharply in the subsequent years. By 2012, the shares of exporters and importers had both grown to all-time highs of 57% and 51%, respectively.[14]

While these numbers suggest a growing tendency among firms to serve foreign markets, they partly reflect firm entry into and exit from production, as well as changes in the sample composition over time (due to sample attrition caused by nonresponse of firms, as well as due to the inclusion of new firms through refreshment samples). For this reason, in Figure 2.2(b), we balance the sample on firms that are observed in each year from 2005 to 2012. This allows for a clean view on changes at the extensive margin of trade among incumbent and surviving firms.[15] The figure confirms that, whether we look at exporting or at importing, the financial crisis had a very small impact on the extensive margin of trade in 2009. Instead, we see constant or rising trade participation rates over time. The share of exporters has been subject to a slight upward trend that was only shortly interrupted in 2010, but accelerated thereafter and reached more than 51% in 2012 (up from less than 47% in 2005). Import participation, in contrast, has not changed much in the balanced sample. Before the financial crisis, the share of importers stood at about 47%. In 2009, the year following the peak of the financial crisis, this share decreased by one percentage point. Although it has been increasing in each year thereafter, import participation has not returned to its pre-crisis level by 2012. In any case, the figure shows that the overall changes that we find around the crisis years are rather small for incumbent and surviving firms. Importantly, the apparent differences in the evolution of trade participation rates across the full sample and the balanced sample (Figures 2.2(a) and 2.2(b)) can be reconciled by differential firm survival rates across trading and non-trading firms. We will take this issue up in Section 2.3.4, where we show that firms that entered the crisis as exporters had higher chances to survive the crisis than firms that were confined to the domestic market.

Do these numbers mask important variation across source and destination countries? A regional decomposition of trade available for 2006 and 2010 allows us to provide a preliminary answer to this

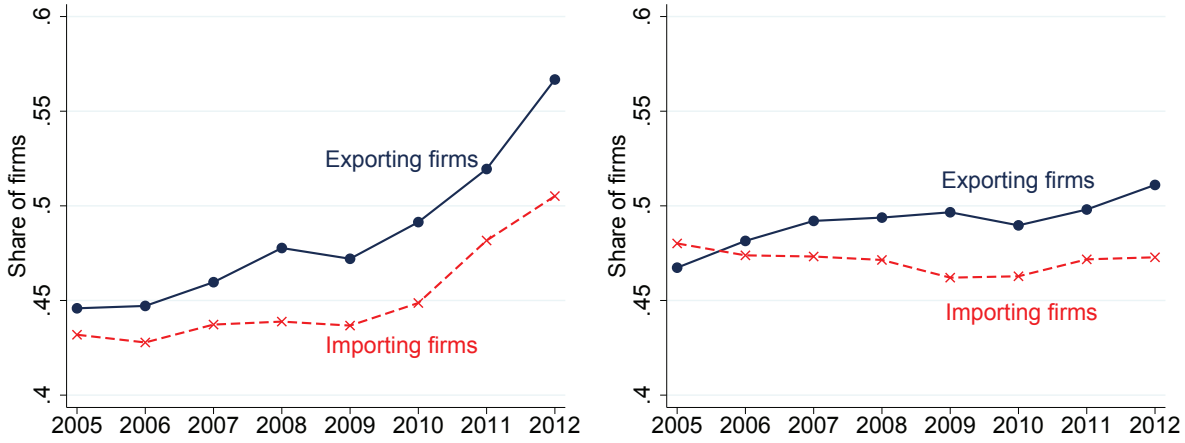
[13] This drop was only slightly larger for exporters (-0.71%) than for non-exporters (-0.48%).

[14] These are the highest trade participation shares observed over the period 1990-2012 (i.e., the period for which ESEE data were available at the time of writing this paper); see http://www.fundacionsepi.es/investigacion/esee/en/salgunos_resultados.asp.

[15] Note that the sample used in Figure 2.2(b) is thus not representative for the manufacturing sector at large. It is in fact biased towards larger firms (in terms of output and employment), as these were more likely to survive the crisis. This bias explains why trade participation rates in 2005 are higher in the balanced sample than in the full sample.

Figure 2.2: *Trade participation, 2005-2012*

(a) *Full sample* (b) *Balanced sample*



Note: In Figure 2.2(a) we use the full sample of firms, whereas in Figure 2.2(b) we balance the sample on 782 firms that are observed in each year from 2005 to 2012. Sampling weights apply. Source: Authors' calculations based on ESEE data.

question.[16] In Figure 2.3, where we balance the sample on firms that are observed in both years, we see very little time variation in export participation for most world regions that we can distinguish in our data: the European Union (EU), Latin America, the rest of the OECD, and the rest of the world (ROW, a residual category, including all of Africa, Eastern Europe, and Asia except Japan and South Korea). 43-44% of all firms exported to the EU, 12-13% exported to Latin America, and 17-18% exported to other OECD countries. This cross-sectional pattern is broadly consistent with a gravity model of trade in which distance and market size play important roles. The largest change over time can be observed for exports to the ROW, where export participation increased by more than one fifth (from 18% in 2006 to 22% in 2010). Hence, the average exporting firm started to penetrate new markets over the crisis years, and thus diversified its export portfolio. This observation is consistent with the behavior observed for Danish firms, which started to enter new markets (especially in Asia) during the recovery after the financial crisis (Abreha et al., 2016).[17] We find similar changes over time for imports as we do for exports. However, the overall share of firms importing from regions other than the EU is relatively small, especially compared to that of exporters: in 2010, it was 3% for Latin America, 8% for other OECD countries, and 13% for the ROW (up from 11% in 2006).

How do the numbers we find for Spain compare with those observed for other European countries? To answer this question, we exploit the EFIGE data set, which provides consistent trade participation shares for seven European countries in 2008. We find, perhaps surprisingly, that among these countries, export participation is lowest in Germany (41%) and France (45%), closely followed by Spain (48%), while Austria and the UK have much higher exporter shares (56%), exceeded only by Italy (63%). German firms also report the lowest importer share (25%), followed by Italy (35%) and Spain (40%).[18] The highest import participation is found for France, where more than half of all firms engage in importing.

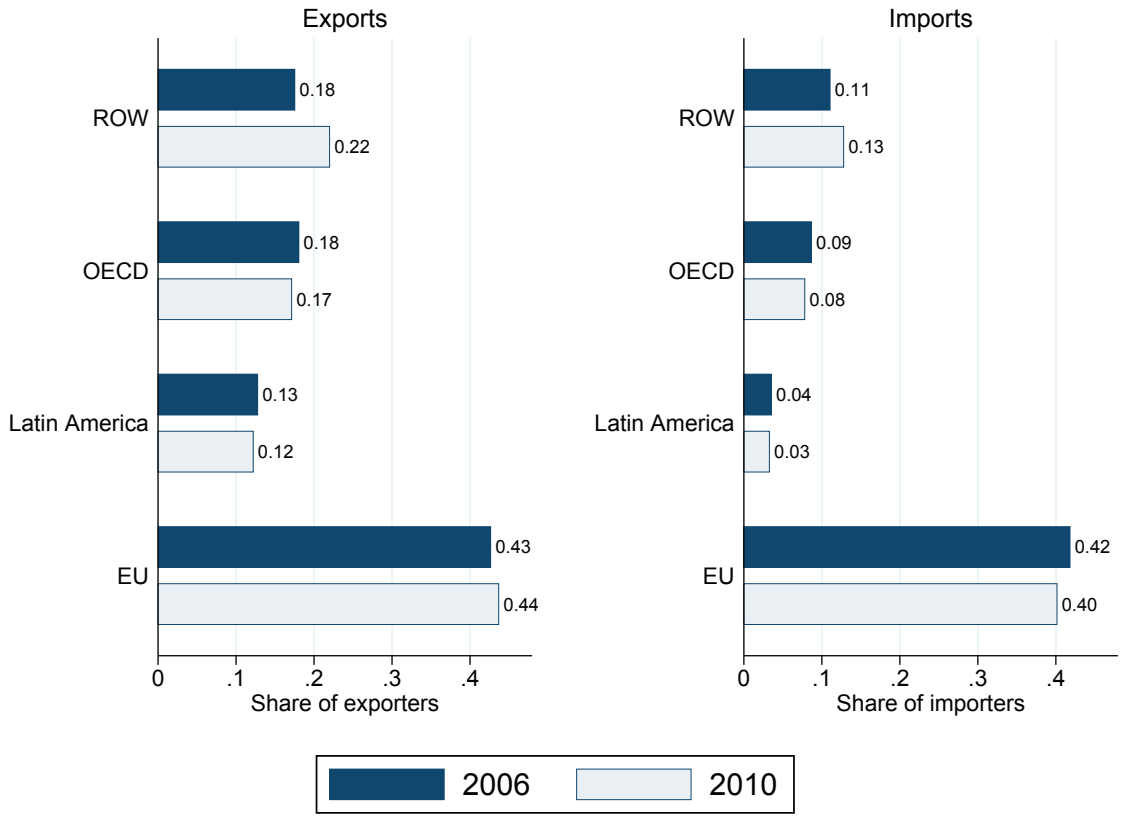
We next estimate a series of probability models for both exporting and importing. This allows us to narrow down the factors that were crucial for trade participation over the crisis years. To do so, we distin-

[16] Information on the composition of imports and exports by world region is available in ESEE every four years.

[17] Similar evidence is provided for Chile and Italy by Álvarez and Sáez (2014) and Costa et al. (2014), respectively.

[18] As far as imports are concerned, the EFIGE data tend to underestimate trade participation, as the questionnaire is limited to imports of goods and services that are used in the production process.

Figure 2.3: *Trade participation by region, 2006 and 2010*



Note: The sample is balanced on 1,247 firms observed in both years 2006 and 2010. In this sample, 780 (780) firms reported positive exports (imports) in 2006, and 797 (770) reported positive exports (imports) in 2010. Sampling weights apply. Source: Authors' calculations based on ESEE data.

guish between those factors that are directly related to the financial crisis and subsequent recession (i.e., macro-level changes taking place outside the firm and captured in our analysis by year dummies) and those related to the evolution of firm-specific characteristics (i.e., micro-level changes taking place inside the firm). The latter also capture indirect effects of the financial crisis (e.g. if some firms experienced a decline in their productivity over time). In order to account for (and exploit) possible complementarities between exporting and importing at the firm level, we estimate two equations simultaneously in a bivariate Probit framework. More specifically, we define two indicator variables, one for the export status of a firm, $Exporter_{it}$, and one for its import status, $Importer_{it}$. The variable $Exporter_{it}$ is equal to one if firm i reports positive exports at time t (and zero otherwise), and accordingly for $Importer_{it}$. We assume that a firm exports if current and expected revenues from exporting are greater than costs:

$$Exporter_{it} = \begin{cases} 1 & \text{if } \Pi^e_{it} > 0 \\ 0 & \text{otherwise,} \end{cases}$$

where Π^e_{it} is the unobserved (latent) net present value of current and expected profits from exporting. We assume that these can be linearly approximated as follows:

$$\Pi^e_{it} = \boldsymbol{\gamma}^e \cdot \mathbf{X}^e_{it} + \delta^e_t + \delta^e_i + \delta^e_{ks} + \varepsilon^e_{it}, \tag{2.2}$$

where $\mathbf{X}_{it}^{e}$ is a column vector collecting time-varying firm characteristics, $\boldsymbol{\gamma}^{e}$ is a vector of parameters to be estimated, δ_{t}^{e} is a year fixed effect, δ_{i}^{e} is a firm-specific effect,[19] δ_{ks}^{e} is a constant specific to the industry-size-group combination corresponding to firm i in year t (with industries being indexed by k and size groups by s), and ε_{it}^{e} is a firm-and-year-specific stochastic profit shock. An expression similar to Equation (2.2) is assumed for importing:

$$\Pi_{it}^{i} = \boldsymbol{\gamma}^{i} \cdot \mathbf{X}_{it}^{i} + \delta_{t}^{i} + \delta_{i}^{i} + \delta_{ks}^{i} + \varepsilon_{it}^{i}. \tag{2.3}$$

In contrast to much of the existing literature, we estimate the decisions of exporting and importing jointly. This strategy is motivated by recent evidence on fixed and sunk cost complementarity between the two activities (Kasahara and Lapham, 2013).[20] We thus assume that the stochastic profit shocks are drawn from a bivariate normal distribution:

$$\begin{bmatrix} \varepsilon_{it}^{e} \\ \varepsilon_{it}^{i} \end{bmatrix} \sim \mathcal{N}\left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} 1 & \chi \\ \chi & 1 \end{bmatrix} \right),$$

where χ is a parameter measuring the (residual) correlation between exporting and importing. Allowing (and testing) for $\chi > 0$ is important in our analysis, as it tells us whether a firm-specific negative effect of the crisis that directly affected one activity spilled over to the other activity (and thus entailed more harmful consequences than the direct effect alone).

In the model described above, we are mainly interested in the year fixed effects, $\delta_{05}^{\ell}, \ldots, \delta_{12}^{\ell}$, $\ell \in \{e, i\}$, as these indicate changes in the profitability of exporting and importing over time that cannot be explained by the firm-specific variables collected in $\mathbf{X}_{it}^{\ell}$. The year fixed effects thus pick up the (net) macro-level effects driven by changes in both demand-side and supply-side factors. The variables contained in $\mathbf{X}_{it}^{\ell}$ are: labor productivity (value added over effective working hours, in logs) to control for the firm's level of competitiveness;[21] capital intensity (tangible fixed assets over the number of workers, in logs); R&D intensity (R&D expenses over sales, in logs); skill intensity (number of graduate workers over total number of workers, in logs); foreign ownership (as dummy variables indicating the share of foreign capital in the firm's joint capital: 0%, >0% & <=50%, or >50%); multinational corporation (MNC) status (as a dummy variable indicating whether the firm has a foreign affiliate); the type of good produced (as dummy variables indicating final goods, intermediate goods, or not defined); and, in the case of exporting, internet presence (as a dummy variable indicating whether the firm is operating a website). Including a dummy for internet presence in the equation for exports, but not for imports, is based on the idea that a website is important as part of the firm's marketing and distribution strategy, but has no impact on the firm's purchasing and sourcing activities. Importantly, the fact that $\mathbf{X}_{it}^{e} \neq \mathbf{X}_{it}^{i}$ leads to efficiency gains in the estimation.

In our first estimation of the bivariate Probit model in Equations (2.2) and (2.3), we treat δ_{i}^{ℓ} as a random variable that is uncorrelated with the other covariates. We compute marginal effects evaluated at the sample means of all regressors. For the year dummies for 2006-2012, the effects can be interpreted as conditional differences in trade participation compared to the base year 2005. Statistical inference is based on robust standard errors clustered at the firm level, which allows for arbitrary forms of het-

[19] We impose different assumptions on the firm-specific effect δ_{i}^{ℓ}, as we shall detail below.

[20] Aristei et al. (2013) also investigate the two-way relationship between exporting and importing.

[21] In alternative specifications we use estimated TFP (rather than labor productivity) to control for the firm's competitiveness, to find that our main results do not change with this modification.

eroskedasticity and accounts for the autocorrelation implied by the firm-specific effect δ_i^ℓ.

Columns (1) and (2) of Table 2.2 report the estimation results. In line with the descriptive evidence presented in Table 2.1 and Figure 2.2, there is no indication of a significant decline in import or export participation in the years surrounding the financial crisis (2007-2009). On the contrary, our results suggest that macro-level developments in the aftermath of the financial crisis (those beyond the influence of individual firms) have prompted more firms to access foreign markets. We find that the probability of exporting is 4.1 percentage points (or 9.1%) higher in 2012 than it was in the base year 2005. The same number for importing is 3.2 percentage points (7.3%). Statistically significant differences between pre- and post-crisis export and import participation are first visible in 2012. These differences cannot be explained by the firm-level characteristics that the literature has consistently identified to influence both exports and imports at the extensive margin (such as productivity), as these are controlled for in the estimation. Regarding these firm-specific control variables, we find that the results accord well with known stylized facts. We find that those firms that are more productive as well as those more intensive in capital, R&D, and skills are more inclined to both exporting and importing. Moreover, we see large and significant differences (with a two-digit margin) between foreign-owned and domestically owned firms, as well as between MNCs and non-MNCs. Finally, the results demonstrate strong firm-level complementarities between exporting and importing ($\hat{\chi} = 0.525$, significant at the 1% level).

One important limitation of the bivariate Probit model is that identification is based on between-firm variation in the data, and that the model thus assumes firm-specific unobserved heterogeneity (denoted by δ_i^ℓ above) to be uncorrelated with the other covariates. However, it is likely that unobserved firm characteristics with strong serial correlation (such as managerial ability) do not only affect a firm's decision to access foreign markets, but that they are also correlated with the other covariates in the model (e.g. productivity). Addressing this issue by estimating firm fixed effects in the Probit framework suffers from the incidental parameters problem and would hence result in inconsistent estimation of all model parameters. We therefore estimate a system of seemingly unrelated regression equations with fixed effects (SUR FE), where each equation describes a linear probability model rather than a non-linear Probit model. On the one hand, this model may deliver implausible predictions for the trading probabilities outside the unit interval. On the other hand, it has the advantage of controlling for unobserved firm-specific heterogeneity through firm fixed effects. Identification of the parameters of interest then comes from within-firm variation in the data, i.e., changes in export and import participation over time.[22]

The estimation results for the SUR FE model are reported in columns (3) and (4) of Table 2.2. The main conclusions drawn from the bivariate Probit model are upheld in this model. In particular, there is no evidence that the financial crisis had any detrimental effect on trade participation. On the contrary, the probability of exporting increased by 3.2 percentage points over the period 2005-2012 due to macro-level effects (statistically significantly at the 1% level). On the import side, we find positive and significant effects in the years 2006 to 2008, but again the probability of importing was significantly higher in 2012 than in any pre-crisis year (by 4.3 percentage points compared to 2005). In contrast to the results obtained from the bivariate Probit model, the only firm-specific variable that consistently and significantly increases the probability of both exporting and importing is productivity. Hence, a firm

[22]Our data set includes information about unusual events that can change the scale and nature of the firm, such as mergers, acquisitions, and divestments. We exclude such firms from the sample whenever we exploit the within-firm variation in our analysis. This leads to a reduced sample size in the corresponding regressions.

Table 2.2: *Probability model for trade participation*

	Bivariate Probit Model		SUR Fixed Effects Model	
	Exporter	*Importer*	*Exporter*	*Importer*
	(1)	(2)	(3)	(4)
Year dummy 2006	-0.00773	0.00407	0.00769	0.0175***
	(0.00715)	(0.00785)	(0.00547)	(0.00670)
Year dummy 2007	-0.00925	-0.00499	0.00930	0.0157**
	(0.00829)	(0.00895)	(0.00616)	(0.00757)
Year dummy 2008	-0.00521	-0.00292	0.00951	0.0181**
	(0.00936)	(0.00966)	(0.00648)	(0.00787)
Year dummy 2009	-0.00716	-0.0135	0.00989	0.0116
	(0.0107)	(0.0107)	(0.00720)	(0.00851)
Year dummy 2010	0.000998	-0.0202*	0.00881	0.0114
	(0.0112)	(0.0116)	(0.00698)	(0.00876)
Year dummy 2011	0.0186	0.00836	0.0196***	0.0275***
	(0.0120)	(0.0121)	(0.00738)	(0.00898)
Year dummy 2012	0.0406***	0.0317**	0.0323***	0.0426***
	(0.0126)	(0.0128)	(0.00773)	(0.00977)
Labor productivity (in logs)	0.0486***	0.0657***	0.0144***	0.0210***
	(0.00817)	(0.00797)	(0.00490)	(0.00641)
Capital intensity (in logs)	0.0404***	0.0471***	-0.00330	-0.0275***
	(0.00684)	(0.00592)	(0.00792)	(0.00972)
R&D intensity (in logs)	1.218***	1.596***	-0.0604	-0.00165
	(0.364)	(0.318)	(0.160)	(0.244)
Skill intensity (in logs)	0.0604	0.127***	-0.00109	-0.0317
	(0.0394)	(0.0437)	(0.0204)	(0.0286)
Multinational dummy	0.261***	0.112***	0.00199	0.0591**
	(0.0313)	(0.0259)	(0.0170)	(0.0274)
Type of good: intermediate good	0.0587***	0.00696	0.0171	0.0276
	(0.0185)	(0.0173)	(0.0139)	(0.0171)
Type of good: not defined	-0.0228	-0.0591***	0.00808	0.0364**
	(0.0181)	(0.0164)	(0.0135)	(0.0179)
Foreign ownership: $> 0\%$ & ≤ 50 %	0.0490	0.00693	-0.0255	-0.00557
	(0.0588)	(0.0478)	(0.0273)	(0.0360)
Foreign ownership: $> 50\%$	0.212***	0.174***	0.00840	0.0179
	(0.0289)	(0.0264)	(0.0182)	(0.0184)
Internet dummy	0.124***		0.0309**	
	(0.0127)		(0.0124)	
Industry-size-group fixed effects	Yes		Yes	
Firm fixed effects	No		Yes	
Number of observations	14,887		13,209	
Number of firms	2,860		2,601	
Cross-equation correlation	0.525***		0.112***	
R2			0.0071	0.0127

Note: This table presents estimated marginal effects on both export and import probabilities obtained from fitting a bivariate Probit model (columns (1) and (2)), as well as a system of seemingly unrelated regression equations (SUR) with fixed effects (columns (3) and (4)). The dependent variables are dummy variables indicating positive exports or imports, respectively. For dummy variables as regressors we report the effects of a discrete change from zero to one. In the bivariate Probit model, marginal effects are evaluated at the sample means of all regressors. Robust standard errors (clustered at the firm level) are given in parentheses. *,**,*** denote significance at the 10%, 5%, 1% levels, respectively. Source: Authors' estimations based on ESEE data.

that experiences a productivity gain over time is more likely to enter foreign markets.[23] This finding adds to the overwhelming evidence emphasizing the importance of firm heterogeneity in the study of international trade, and it is consistent with the seminal work by Melitz (2003). Finally, the positive and significant (residual) correlation between exporting and importing is confirmed in the SUR FE model.

Thus far, we have assumed that any persistence in export status over time stems from possible autocorrelation in the independent variables (including firm fixed effects) and the errors. Similar to other firm-level data sets, persistence in export status is indeed a salient feature of our data. Balancing the panel on 1,037 firms that are observed in each year from 2005-2010, we find that 601 firms exported in each and every year, while 276 firms never exported. Hence, a vast majority of 84.6% of all firms maintained their export status throughout the six-year period considered.

There are at least two sources of persistence in export participation that we have not considered in the models described above and that are reviewed and modeled in Roberts and Tybout (1997) and Bernard and Jensen (1997). The first is learning by doing, which refers to the accumulation of knowledge (through production and exporting) that reduces future costs of production and exporting. The second are sunk costs for foreign market entry, for example in the form of information and distribution costs. Similar ideas apply to importing. While we cannot separately identify these two channels, we may hypothesize based on the above considerations that the firm's current and expected profits from exporting will depend positively on past export status:

$$\Pi^e_{it}(\textit{Exporter}_{it-1} = 1, \cdot) - \Pi^e_{it}(\textit{Exporter}_{it-1} = 0, \cdot) > 0.$$

In such a dynamic framework, a negative transitory shock to foreign demand due to the financial crisis would generate a negative effect on export participation that carries over to future time periods (implying gradual adjustment of the probability to export).

To allow for dynamics in trade participation, we specify the following model for exporting:

$$\textit{Exporter}_{it} = \rho^e \cdot \textit{Exporter}_{it-1} + \boldsymbol{\gamma}^e \cdot \mathbf{X}^e_{it} + \delta^e_t + \delta^e_i + \delta^e_{ks} + \varepsilon^e_{it}, \tag{2.4}$$

and accordingly for importing. Of course, the larger the autoregressive parameter (i.e., the coefficient of the lagged dependent variable), the stronger (i.e., long-lasting) is the dynamic effect. First, we estimate these models by the standard fixed effects approach with the right-hand side of the equation including the lagged dependent variable (LDV FE model). Second, we use the first-differenced general method of moments (diff-GMM) approach developed by Arellano and Bond (1991).

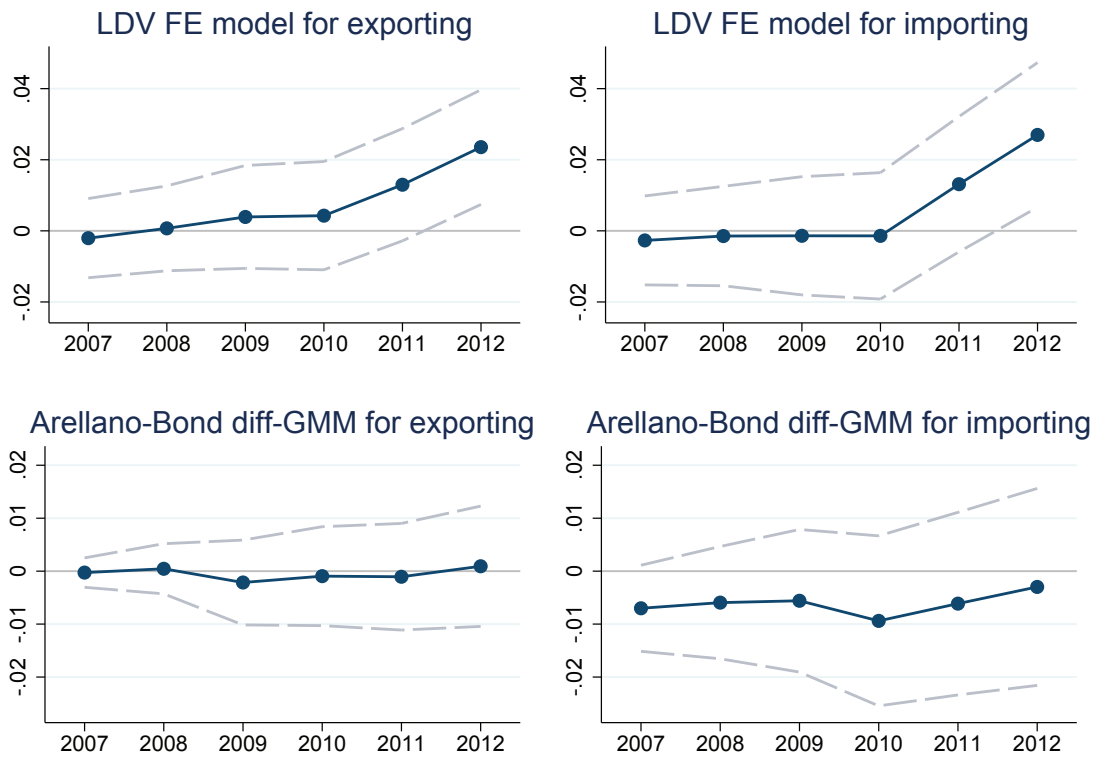
In the LDV FE models, we apply the within-transformation to the data in order to get rid of the firm fixed effects δ^e_i and δ^i_i, respectively. Estimates of ρ^ℓ in the LDV FE models serve as lower bounds for the true parameter values.[24] We find values of $\hat{\rho}^e \approx 0.237$ and $\hat{\rho}^i \approx 0.193$ (both significant at the 1% level) in the LDV FE models. In the diff-GMM approach, the model is estimated in first differences to cancel the firm fixed effects. In addition to the lagged dependent variable, we treat labor productivity, capital intensity and foreign ownership as endogenous variables, and R&D and skill intensity along with MNC

[23] There is strong evidence in the literature for self-selection of the more productive firms into exporting as well as importing; see e.g. Bernard and Jensen (1999), Smeets and Warzynski (2013), and Kohler and Smolka (2014). There is also some evidence for both exporting and importing to increase productivity; see e.g. De Loecker (2007), Halpern et al. (2015), and Feng et al. (2012).

[24] The Nickell bias for the autoregressive parameter, ρ^ℓ, is $\text{plim}_{N\to\infty}\left(\widehat{\rho^\ell} - \rho^\ell\right) \approx -(1+\rho^\ell)/T$, with $\ell \in \{e, i\}$ and $T = 7$ in our application.

status as pre-determined variables. Lagged levels of the dependent variable, the predetermined variables, and the endogenous variables are used as internal instruments. We allow for the maximum number of available lags for use as instruments. To accommodate heteroskedasticity, we use the two-step version of the diff-GMM estimator. We estimate values of $\hat{\rho}^e \approx 0.382$ and $\hat{\rho}^i \approx 0.373$ (both significant at the 1% level) in the diff-GMM models.[25]

Figure 2.4: *Dynamic probability models for trade participation*



Note: This figure shows estimated coefficients of year dummies in dynamic probability models as specified in Equation (2.4) for the exporter dummy (left-hand side) and importer dummy (right-hand side), respectively, alongside 90% confidence intervals. The effects are changes in the probability to export and import, respectively, relative to 2006. Source: Authors' estimations based on ESEE data.

Figure 2.4 summarizes the coefficients of the year dummies in these models estimated by both approaches. The LDV FE models confirm the increase in both probabilities for exporting and importing after the crisis (both significant at 5% for 2012). The diff-GMM estimations, in contrast, cannot identify any statistically significant effect of the financial crisis and subsequent recession on trade participation. Importantly, none of our dynamic estimation approaches provides any evidence of a detrimental crisis effect on the extensive margin of trade.

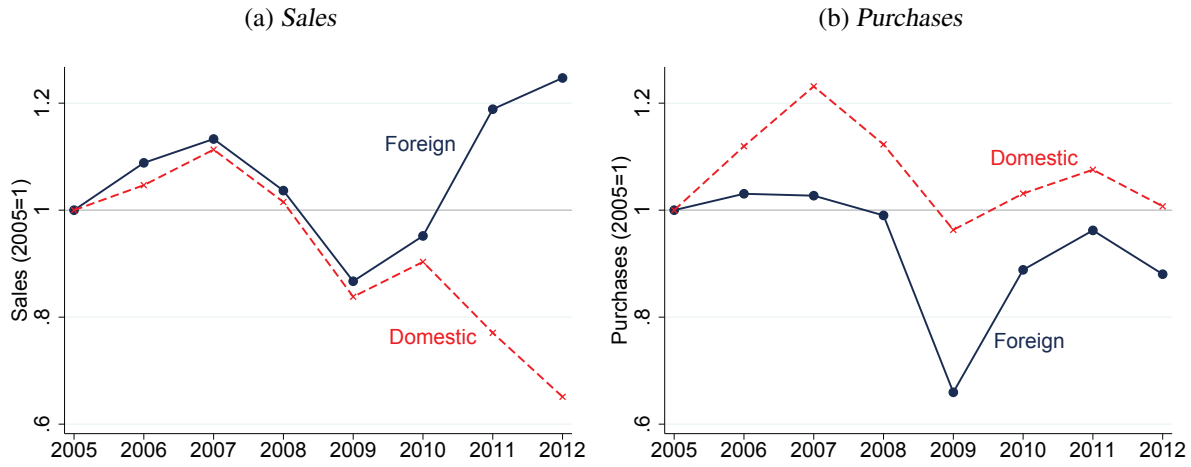
2.3.3 Export and import intensity

Next, we analyze the evolution of firms' trade *volumes* (i.e., the intensive margin of international trade). Over the pre-crisis period, the average exporter was shipping goods and services worth 11.2 million €

[25] Neither for exporting ($p = 0.8969$) nor for importing ($p = 0.5363$) can we reject the null hypothesis that the overidentifying restrictions are valid (Hansen specification test of the instrument condition). For both exporting and importing, the Arellano-Bond test for zero autocorrelation in first-differenced errors of order one is rejected ($p = 0.000$), while that of order two cannot be rejected ($p = 0.1733$ for exporting and $p = 0.5363$ for importing).

abroad (per year), and the average importer was purchasing goods and services worth 7.8 million € from abroad (per year). Figure 2.5 depicts real export values for those firms that were continuous exporters over the period 2005-2012 (Figure 2.5(a)), and real import values for those that were continuous importers (Figure 2.5(b)). The solid lines demonstrate that the financial crisis had a very strong negative effect at the intensive margin of international trade. Real trade values of both exporting and importing plummeted drastically from 2007 to 2009, but recovered partly in 2010, and further so in 2011. While exports had fully recovered by 2011 and increased further in 2012, imports had not returned to their pre-crisis level by 2012. These findings are in line with our insights from the decomposition exercise above.

Figure 2.5: *Sales and purchases, 2005-2012*



Note: In Figure 2.5(a) (Figure 2.5(b)), the sample is balanced on 456 (412) firms that are continuously exporting (importing) over the period 2005-2012. Foreign and domestic sales (Figure 2.5(a)) as well as foreign and domestic purchases (Figure 2.5(b)) are normalized to one in 2005. Sampling weights apply. Source: Authors' calculations based on ESEE data.

How strong were the adjustments in 2008 and 2009 compared to the concurrent drop in domestic activities? The dashed lines in Figure 2.5 show that while imports experienced a much stronger decline than domestic purchases (-36% versus -22% from 2007 to 2009), the drop in domestic sales was equally pronounced as the drop in exports (-24%). Thus, speaking of a great trade collapse in the Spanish case, while justified for firm-level imports, seems unwarranted when looking at firm-level exports. Most noteworthy, however, is the shift in sales that we observe in the years after 2010, away from the domestic market towards the foreign market. Within just two years, exports *increased* by 29%. Domestic sales, in contrast, *decreased* by 28%. Hence, it seems that firms were compensating for the collapse in aggregate demand in Spain (in the course of the events associated with the European sovereign debt crisis) by channeling their sales into the export market. This was possible because other countries like France and Germany, the two largest economies in Europe and the top export destinations for Spain, had positive economic growth in each year from 2010 to 2012. Spain, in contrast, had negative economic growth over the same period.[26] The finding that firms substituted domestic with foreign sales accords well with macro-level evidence by Belke et al. (2015). We will return to this issue in our discussion in Section 2.4. Importantly, the observation that the foreign market has gained in importance relative to the domestic market has no correspondence on the import side, where domestic and foreign purchases were largely moving in parallel to one another.

We now use regression analysis to investigate the factors influencing the *trade intensity* of firms,

[26] Source: Eurostat at http://ec.europa.eu/eurostat/.

defined as the share of exports in total sales or the share of imports in total purchases, over the period 2005-2012. We should like to emphasize the difference between these *shares* and the *levels* of firms' trade volumes; the latter we have used to isolate changes at the 'intensive margin' in the decomposition exercise in Section 2.3.1. There are three reasons for using the trade intensity rather than the log of exports or imports in our estimations. First, the trade intensity is an important measure of globalization at the firm level, which indicates how strongly firms are integrated into the global economy through international trade.[27] Second, the trade intensity is defined not only for exporters or importers, but also for firms that do not engage in international trade. This allows us to circumvent an obvious selection problem that arises when non-trading firms are excluded from the sample.[28] Third, we can use the full sample of firms to investigate how export and import intensity are intertwined by estimating the two equations (one for export intensity and one for import intensity) simultaneously. This also leads to efficiency gains in the estimation. For these three reasons it is both convenient and meaningful to use trade intensities as dependent variables in our regression analysis. However, for the interpretation of our results we must keep in mind that these regressions are not suitable for showing evidence on the great trade collapse as such, simply because domestic activity in Spain dropped substantially as well (as is clear from Figure 2.5).

We estimate a SUR model, where the first equation is specified as:

$$ExpInt_{it} = \boldsymbol{\gamma}^e \cdot \mathbf{X}^e_{it} + \delta^e_t + \delta^e_i + \delta^e_{ks} + \varepsilon^e_{it}, \tag{2.5}$$

with $ExpInt_{it}$ denoting the export intensity (exports over total sales) of firm i in year t, and accordingly for the second equation with $ImpInt_{it}$ (imports over total purchases) as the dependent variable.[29] As we did above in the model for trade participation, we assume $E[\varepsilon^e_{it}\ \varepsilon^i_{jt}|\cdot] = 0$ whenever $i \neq j$, whereas $E[\varepsilon^e_{it}\ \varepsilon^i_{it}|\cdot] = \chi$. In a first specification, we treat the firm effects, δ^e_i and δ^i_i, as random variables that are not correlated with the other covariates, and thus exploit between-firm variation in the data. In a second specification, we relax this assumption by treating the firm effects as fixed effects that are explicitly controlled for, and we identify the parameters of interest from within-firm variation in the data by using SUR FE estimation.

Overall, the regression results reported in Table 2.3 are consistent with the evolution of trade intensities shown in Figure 2.5. Importantly, we do not find evidence that firms decreased their trade intensities due to the financial crisis. Hence, while both foreign and domestic activity declined sharply, firms did not become less 'globalized' in the crisis. On the contrary, we find a steady increase in the export intensity of firms over the post-crisis period 2009-2012, as documented by $\hat{\delta}^e_t > \hat{\delta}^e_{t-1}$ for $t \geq 09$. In the SUR FE model, the year-to-year differences that we find are statistically significant at the 1% level for $t \geq 10$. The rise in the export intensity identified in the data is not accompanied by a contemporaneous rise in import intensity, where we find no significant differences across years.

To substantiate these results, and to address similar concerns as in the previous section, we also

[27] Behrens et al. (2013, p. 703) also examine ratios of international activity over total activity and point out that analyzing "the recent trade collapse using firm-level data on both trade and domestic operations [...] is necessary to gauge whether international activity has been disproportionately hit by the crisis."

[28] We explicitly model the process governing selection into exporting or importing in a robustness analysis discussed below.

[29] The parameters in these equations are of course different from the ones in Equations (2.2) and (2.3), but for convenience we use the same notation as before.

Table 2.3: *Model for trade intensity*

	SUR Model		SUR Fixed Effects Model	
	Export intensity	*Import intensity*	*Export intensity*	*Import intensity*
	(1)	(2)	(3)	(4)
Year dummy 2006	-0.00276	-0.00672	-0.00157	0.000339
	(0.00353)	(0.00871)	(0.00252)	(0.00889)
Year dummy 2007	-0.00501	-0.0132	-0.00151	-0.00423
	(0.00417)	(0.00897)	(0.00257)	(0.00893)
Year dummy 2008	-0.00521	-0.0149	-0.000320	-0.00471
	(0.00492)	(0.00945)	(0.00294)	(0.00850)
Year dummy 2009	0.00124	-0.0122	0.00566*	-0.00292
	(0.00586)	(0.00969)	(0.00302)	(0.00794)
Year dummy 2010	0.0106*	-0.0146	0.00871***	-0.00248
	(0.00632)	(0.00990)	(0.00319)	(0.00786)
Year dummy 2011	0.0260***	-0.00565	0.0198***	-0.000175
	(0.00688)	(0.0100)	(0.00345)	(0.00775)
Year dummy 2012	0.0453***	-0.00348	0.0358***	0.000513
	(0.00758)	(0.0103)	(0.00399)	(0.00785)
Labor productivity (in logs)	0.0111**	0.0304***	-0.000243	0.00654**
	(0.00541)	(0.00413)	(0.00265)	(0.00304)
Capital intensity (in logs)	0.0281***	0.0220***	0.00428	-0.00315
	(0.00401)	(0.00338)	(0.00304)	(0.00551)
R&D intensity (in logs)	0.527***	0.396**	0.0752	0.219
	(0.178)	(0.184)	(0.0755)	(0.299)
Skill intensity (in logs)	-0.0232	0.0399**	-0.00532	0.00647
	(0.0189)	(0.0176)	(0.0123)	(0.0124)
Multinational dummy	0.110***	0.0263*	0.00142	0.0194*
	(0.0182)	(0.0140)	(0.0125)	(0.0112)
Type of good: intermediate good	0.0403***	-0.0145	0.00210	-0.0233
	(0.0106)	(0.0114)	(0.00700)	(0.0173)
Type of good: not defined	0.0366***	-0.0282***	0.00584	-0.00172
	(0.00981)	(0.0107)	(0.00726)	(0.0122)
Foreign ownership: $> 0\%$ & ≤ 50 %	0.00137	0.0195	0.00899	0.00266
	(0.0343)	(0.0244)	(0.0181)	(0.0236)
Foreign ownership: $> 50\%$	0.104***	0.209***	0.00731	0.0436
	(0.0168)	(0.0154)	(0.0146)	(0.0294)
Internet dummy	0.0352***		-0.00253	
	(0.00829)		(0.00510)	
Industry-size-group fixed effects	Yes		Yes	
Firm fixed effects	No		Yes	
Number of observations	14,902		13,209	
Number of firms	2,861		2,601	
Cross-equation correlation	0.1148***		0.0146*	
R^2	0.370	0.289	0.0276	0.0042

Note: This table presents estimated coefficients from fitting a system of seemingly unrelated regression equations (SUR) for export and import intensities (both without and with firm fixed effects). The dependent variables are export and import intensities, respectively. Robust standard errors (clustered at the firm level) are given in parentheses. *,**,*** denote significance at the 10%, 5%, 1% levels, respectively. Source: Authors' estimations based on ESEE data.

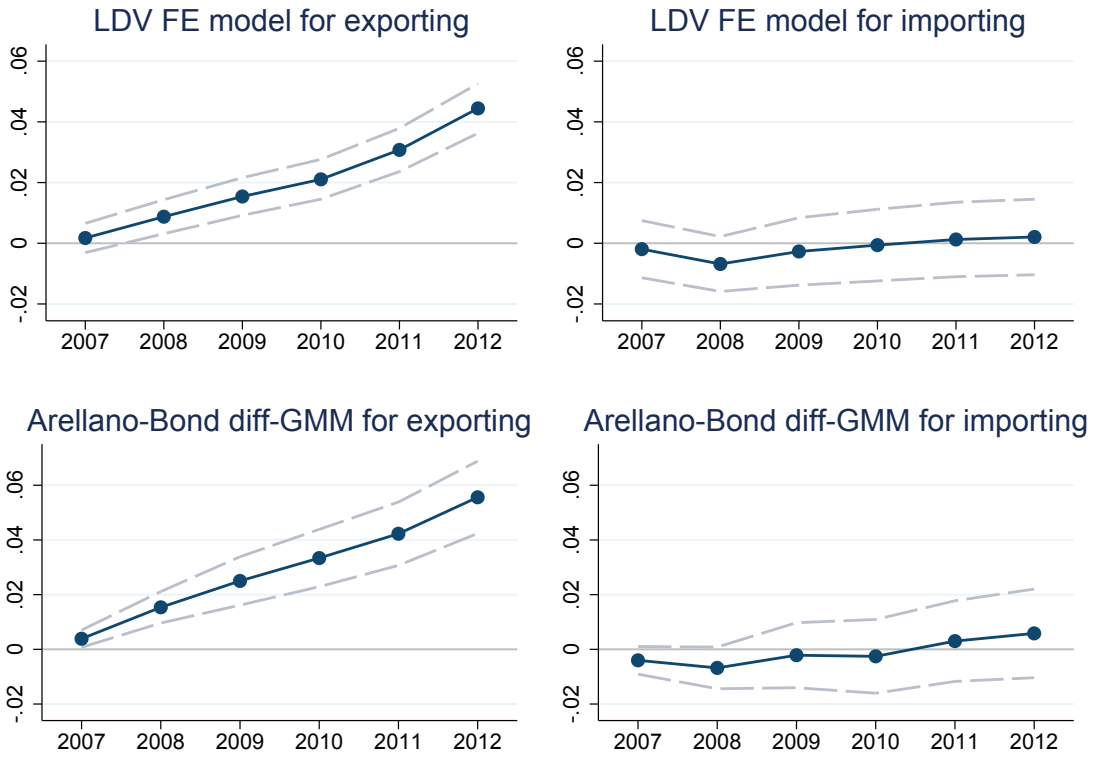
consider the following dynamic specification of the model for exporting:

$$ExpInt_{it} = \rho^e \cdot ExpInt_{it-1} + \boldsymbol{\gamma}^e \cdot \mathbf{X}^e_{it} + \delta^e_t + \delta^e_i + \delta^e_{ks} + \varepsilon^e_{it}, \tag{2.6}$$

and accordingly for importing. The models are estimated alternatively by LDV FE and diff-GMM. In the LDV FE models, we find values of $\hat{\rho}^e \approx 0.220$ (significant at the 1% level) and $\hat{\rho}^i \approx 0.0132$ (not statistically significant). In the two-step diff-GMM estimations, we choose the same endogenous and pre-determined variables as well as the same number of lags as in Section 2.3.2. We estimate values of $\hat{\rho}^e \approx 0.182$ and $\hat{\rho}^i \approx 0.038$ (both significant at the 1% level) in the diff-GMM models.[30]

Figure 2.6 shows the estimated coefficients of the year dummies in dynamic models for the export intensity (left part) and the import intensity (right part), respectively, alongside 90% confidence intervals. The effects are changes in the trade intensities relative to 2006. Irrespective of the estimator we use, we find that the export intensity of firms has been on the rise ever since 2007, and continuously throughout the years of the financial crisis and subsequent recession. In line with the results obtained from our static models, the dynamic models do not indicate any significant change in the import intensity over time.

Figure 2.6: *Dynamic models for trade intensity*



Note: This figure shows estimated coefficients of year dummies in dynamic models as specified in Equation (2.6) for the export intensity (left-hand side) and the import intensity (right-hand side), respectively, alongside 90% confidence intervals. The effects are changes in the export intensity and import intensity, respectively, relative to 2006. Source: Authors' estimations based on ESEE data.

[30] Neither for exporting ($p = 0.3758$) nor for importing ($p = 0.1364$) can we reject the null hypothesis that the overidentifying restrictions are valid (Hansen specification test of the instrument condition). For both exporting ($p = 0.000$) and importing ($p = 0.012$), the Arellano-Bond test for zero autocorrelation in first-differenced errors of order one is rejected, while that of order two cannot be rejected ($p = 0.7233$ for exporting and $p = 0.1752$ for importing).

Notice that we include both exporters (importers) and non-exporters (non-importers) in the above estimations. An implicit assumption underlying this approach is that the intensive margin of trade is governed by the same factors (and in the same way) as the extensive margin. However, it is not clear theoretically why this should be the case. For example, in the Melitz (2003) model, the workhorse model of international trade with heterogeneous firms, the foreign and domestic sales of a firm react proportionally to changes in the firm's productivity, conditional on exporting. Hence, while productivity gains are expected to increase the likelihood of a firm to export, they need not increase the export intensity of a firm that already exports.[31] To address this issue, we also model the selection into exporting and importing explicitly by using a two-stage Heckman selection model with skill intensity as an exclusion restriction in the first-stage equation.[32] The results (not reported) indicate a selection bias for the export intensity (by a significant coefficient of the inverse Mills ratio in the second-stage equation), but not for the import intensity. While the effects of a few control variables on the export intensity (e.g. productivity) change with the selection correction compared to the SUR model, the year fixed effects hardly change at all.

2.3.4 Firm competitiveness and crisis resilience

According to our data, in the pre-crisis period 2005-2008, exporting and importing firms alone were responsible for about 85% of total output, and about 74% of all jobs in Spanish manufacturing. These numbers are considerably higher in 2012 (92% for output and 82% for jobs), which attests to a growing importance of the global economy for the manufacturing sector in Spain. This development is partly explained by new firms entering foreign markets in recent years, as shown in Section 2.3.2, but it might also be the result of an exceptional degree of competitiveness and crisis resilience of those firms that had already been active on foreign markets before the financial crisis. Exploring this issue in greater detail is the purpose of this section.

Figure 2.7 depicts the evolution of various measures of firm performance over the period 2007-2011 depending on the firm's export status. We look at four different firm characteristics that are informative for the analysis of firm behavior in the financial crisis: real output, effective working hours[33], total factor productivity (TFP), and the (average) hourly wage paid by the firm. Moreover, we distinguish between four different groups of firms[34]: continuous exporters (henceforth called exporters), export market entrants, firms leaving the export market, and continuous non-exporters (henceforth called non-exporters). To abstract from the effects of entry into and exit from production (which is analyzed separately below) as well as changes in sample composition due to nonresponse of firms and refreshment samples, we focus on the pre-crisis cohort and balance the sample on firms that are observed in each year over the period considered. All values are normalized to one in 2007.

There are several insights to be gained from Figure 2.7. First, firm output and employment were under strong pressure during the financial crisis and contracted sharply for all groups of firms. Both exporters and non-exporters reduced their output by more than 25% from 2007 to 2009. Those firms

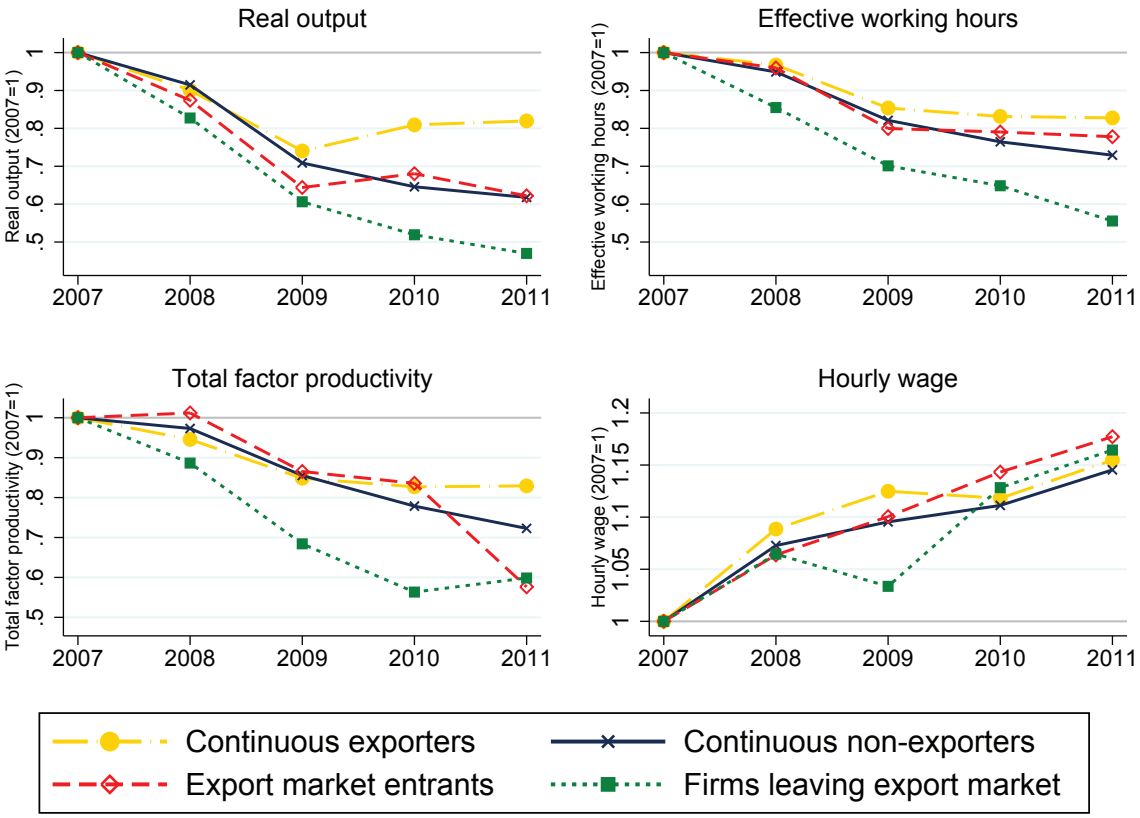
[31] This could help explain why productivity enters insignificantly in the fixed effects specification above, since changes in the export status over time are rare, while changes in the export intensity of exporters are frequent.

[32] The underlying assumption is that the skill intensity of the firm determines the firm's trade status (i.e., the decision to export or import), but does not have a partial effect on the intensity of trade.

[33] As explained in Section 2.2, effective working hours is our preferred measure of employment and should be interpreted as an input-based indicator of firm size, not as a measure of productivity. All results reported in this section look almost identical if we use the number of employees (head count) instead.

[34] The precise definition of each group is given in the note to Figure 2.7.

Figure 2.7: *Firm competitiveness and crisis resilience, 2007-2011*



Note: The sample is balanced on firms that are observed in each year from 2007 to 2011. Continuous exporters are 671 firms that export in each of the years from 2005 to 2012 (if observed); continuous non-exporters are 310 firms that do not export in any of the years from 2005 to 2012 (if observed); export market entrants start exporting in one of the years from 2009 to 2012 and stay in the export market after entry (43 firms); firms leaving the export market stop exporting in one of the years from 2009 to 2012 and do not re-enter after exit (21 firms). All variables are normalized to one in 2007. Sampling weights apply. Source: Authors' calculations based on ESEE data.

leaving or entering the export market reduced their output even more drastically (by more than 35%). The reductions in employment were smaller than those in output for all groups, ranging from 15% to 30%, but we observe the same ranking across firms with different export status.

Second, output and employment stabilized after 2009, but this development is fully attributable to exporters and export market entrants. Non-exporters and firms leaving the export market, in contrast, continued to shrink further. More generally, it turns out that a firm's export status is a good indicator for how well firms did both during and after the peak of the financial crisis, as exporters outperformed all other firms over the period 2007-2011. The cumulative differences that we find between exporters and non-exporters are remarkable. For example, we find that non-exporters produced 38% less in 2011 than in 2007, while for exporters the reduction was 18%. Similarly, non-exporters destroyed 27% of their jobs from 2007 to 2011, whereas for exporters the same number is 17%.[35]

Third, the evolution of TFP shows marked differences between exporters and non-exporters. While from 2007 to 2009 the TFP of both types of firms declined similarly by about 15%, the trajectory *after* 2009 correlates strongly with the firm's export status. For exporters, the level of TFP was about the same

[35] Not surprisingly, firms leaving the export market in one of the years 2009-2012 performed weakest in terms of real output and employment throughout the period from 2007 to 2011.

in 2011 as it was in 2009. For non-exporters, in contrast, the level of TFP had deteriorated by another 15% in 2011 relative to 2009. Hence, after the financial crisis, non-exporters have lost part of their technical and managerial efficiency in production (i.e. their ability to transform inputs into outputs).[36] This is a remarkable observation that is likely to shape the dynamics of the Spanish manufacturing sector over the next couple of years. To gauge the importance of this development for aggregate performance, we have used our firm-level estimates of TFP to compute changes in industry-level productivity. Aggregate productivity is influenced not only by firm-level TFP, but also by the allocation of factors across firms. Low-productivity firms exiting the market and freeing up resources to be used by high-productivity firms leads to aggregate productivity gains. We have computed industry-level productivity as the market-share weighted average of firm-level TFP and found strong heterogeneity in the evolution of aggregate TFP over the period 2005-2012, with some industries experiencing a drastic decline in TFP by more than 65% in response to the financial crisis (such as the industries "computer, electronic and optical products" or "ferrous and non-ferrous metals"), and very few showing a positive performance (such as "other transportation equipment"). The overall performance at the industry-level was poor: in 18 out of 20 industries TFP declined between 2007 and 2012. While beyond the scope of this paper, analyzing these issues in more detail might prove fruitful in future research (see Hospido and Moreno-Galbis, 2015, for a first study in this direction).

Fourth, the (nominal) hourly wage paid by firms increased on average by around 7.5% from 2007 to 2008 and by a compound annual growth rate of about 2.3% thereafter.[37] Importantly, although wage moderation efforts are visible during the financial crisis, *real* wages continued to increase even after 2007, given a compound annual inflation rate of 1.6% over the period 2008-2011.[38] Overall, the evolution of wages is very similar across the four different groups of firms. For exporters, real wages declined slightly after 2009, making Spanish exports more competitive internationally.

Putting these insights together, we may reflect that the Spanish labor market adjusted first and foremost through a contraction in labor demand causing a sharp increase in involuntary unemployment. Owing to the dual nature of the Spanish labor market (highly protected permanent vs. poorly protected temporary workers),[39] this took the form of massive lay-offs of low-skilled and medium-skilled workers with temporary contracts, rather than a reduction of the employment intensity of individual workers.[40] The observed development can entail negative effects for future economic growth, as the skills of unemployed workers erode substantially, especially for longer unemployment spells.[41]

A particularly important and fundamental dimension of firm performance is firm survival. Looking

[36] One might be tempted to argue that firm-specific input and output price changes are responsible for this observation. However, firm-level input and output pricing information available in the ESEE data allow us to demonstrate that this is not the case, as we find hardly any differences in the evolution of prices between exporters and non-exporters; see Figure 2.A.1 in Appendix 2.A.

[37] Due to data limitations we cannot distinguish wages by different types of workers (e.g. high-skilled vs. low-skilled workers). Hence, the observed wage changes at the firm level may be due to adjustments in both the wages of continuously employed individuals and the composition of employment (e.g. in terms of skills, types of contracts, or migration background).

[38] The inflation data are elicited from consumer price data provided by the OECD at http://stats.oecd.org/Index.aspx?DataSetCode=MEI_PRICES.

[39] This peculiarity of the Spanish labor market is heavily criticized by leading Spanish economists; see for instance chapter four in Garicano (2014).

[40] The opposite happened in Germany, where the unemployment rate hardly increased at all through the crisis years. Burda and Hunt (2011) discuss this issue as "the German Labor Market Miracle".

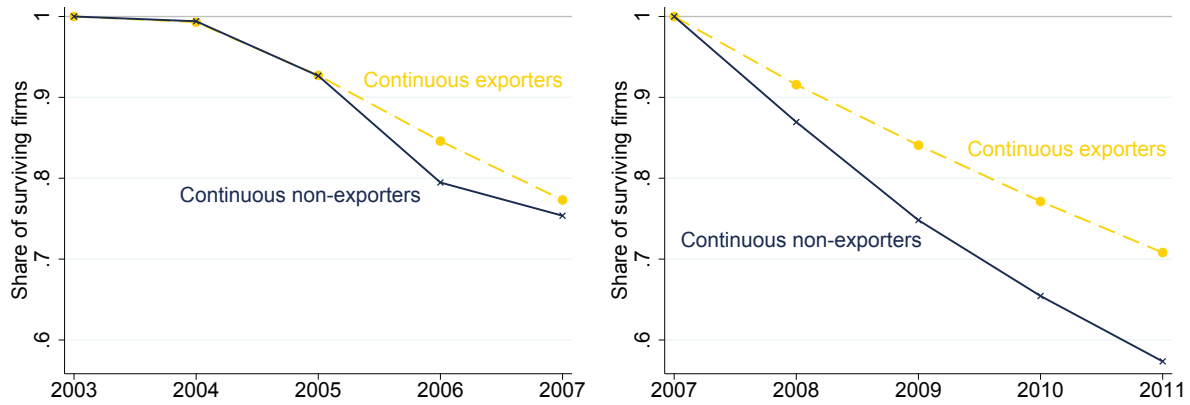
[41] Gregory and Jukes (2001) provide empirical evidence on this mechanism by estimating the effect of unemployment on earnings following re-employment for British men over the period 1984-1994. However, one of the conclusions that can be drawn from their analysis is that human capital depreciation is lowest for young and low-paid workers and highest for middle-aged and high-paid workers. This might ameliorate concerns about the future growth of the Spanish economy, as the crisis caused an increase in unemployment mainly among less experienced and unskilled workers.

into firm survival in the context of our paper is interesting, because the challenges posed by a crisis as severe as the one in 2008/09 are more demanding than the ones associated with the usual business cycle. Figure 2.8 serves to illustrate differences in firm survival depending on firms' export status. The figure follows two different firm cohorts over a five-year period, and depicts the share of surviving firms in each cohort (separately for exporters and non-exporters): the first cohort refers to firms observed in 2003 (Figure 2.8(a)) and the second cohort to firms observed in 2007 (Figure 2.8(b)). We define market exit (or, equivalently, firm death) as going out of business or terminating manufacturing activities (and we exclude firms that ceased to collaborate, did not respond to the questionnaire, or could not be localized). For simplicity, we examine only continuous exporters and continuous non-exporters.

Consider first Figure 2.8(b), which follows the 2007 cohort of firms through the crisis period 2007-2011. The figure shows that the share of surviving firms decreases significantly faster for non-exporters than for exporters throughout this period. Out of 100 firms that were producing and selling *only* in the domestic market in 2007, 43 firms had exited the market by 2011. In contrast, out of 100 exporters observed in 2007, only 29 had exited the market over the same period. Hence, those firms entering the crisis as exporters (and staying in the export market) had higher chances to survive the crisis than those starting out as non-exporters. For the sake of comparison, consider next Figure 2.8(a), which conducts the same exercise by following the 2003 cohort of firms through the pre-crisis period 2003-2007. We find that in this earlier period survival rates are higher and very similar across exporters and non-exporters. To conclude this part of our analysis, we may thus state that exporters were more resilient than non-exporters to the exceptional economic distress associated with the financial crisis and the subsequent recession.

Figure 2.8: *Firm survival, before (2003-2007) and during the crisis (2007-2011)*

(a) *2003 cohort of firms* (b) *2007 cohort of firms*



Note: In Figure 2.8(a), the sample is restricted to the 2003 cohort of firms; continuous exporters are 714 firms that export in each of the years from 2001 to 2008 (if observed); continuous non-exporters are 341 firms that do not export in any of the years from 2001 to 2008 (if observed). In Figure 2.8(b), the sample is restricted to the 2007 cohort of firms; continuous exporters are 949 firms that export in each of the years from 2005 to 2012 (if observed); continuous non-exporters are 544 firms that do not export in any of the years from 2005 to 2012 (if observed). Source: Authors' calculations based on ESEE data.

We next address the issue of performance differences between exporters and non-exporters in a more rigorous way by using econometric methods. It is a well-known fact that exporters have a competitive edge over non-exporters. Bernard and Jensen (1999) and others have shown that exporters are on average more productive than non-exporters, have higher sales, and employ more workers. These differences have been quantified in terms of so-called *exporter premia*. We estimate *time-varying* exporter premia for several measures of firm performance and document the evolution of these premia during the financial crisis and subsequent global recession. We also identify (and quantify) the advantage of exporters

regarding the likelihood to survive the crisis.

Building on the methodology established in the literature, we estimate variants of the following econometric model:

$$\textit{Performance}_{it} = \lambda_t \cdot \textit{Exporter}_{it} + \boldsymbol{\theta} \cdot \mathbf{Z}_{it} + \phi_t + \phi_{ks} + \varepsilon_{it}, \tag{2.7}$$

where $\textit{Performance}_{it}$ is one of the following four variables: real output (total production value, in logs), effective working hours (in logs), TFP (in logs), and survival (as a dummy variable indicating that the firm is still active and producing *in the following year*). As above, the variable $\textit{Exporter}_{it}$ is a dummy variable for positive exports, λ_t represents the coefficients of interest (with $t = 05, \ldots, 12$), ϕ_t is a year fixed effect, ϕ_{ks} is an industry-size-group fixed effect, and ε_{it} is the error term. The vector $\mathbf{Z}_{it}$ collects a number of firm-specific and time-varying control variables, and the vector $\boldsymbol{\theta}$ includes the corresponding parameters to be estimated. This setup allows us to estimate the evolution of *conditional* performance differences between exporters and non-exporters, as we control for the industry-size-group combination corresponding to the firm, as well as a common set of firm-level characteristics (identical to those used in the previous section, but excluding the performance variables themselves). We estimate Equation (2.7) by OLS without firm fixed effects, which allows us to identify the *levels* of different exporter premia as well as their evolution over time.

Figure 2.9 displays our estimates of the year-specific exporter premia $\hat{\lambda}_t$ for the different performance variables. In terms of output, employment, and productivity, our results demonstrate that exporters were outperforming non-exporters throughout the period of analysis. Furthermore, these differences have been increasing over time, in particular in the years 2011 and 2012, so exporters magnified their size and productivity advantages after the financial crisis. More precisely, in 2007, exporters were on average 22% larger in terms of output and 6% larger in terms of employment compared to non-exporters. These differences had widened to more than 50% for output and 15% for employment by 2012. Similarly, exporters increased their TFP premium from 5% to 13% between 2007 and 2012.

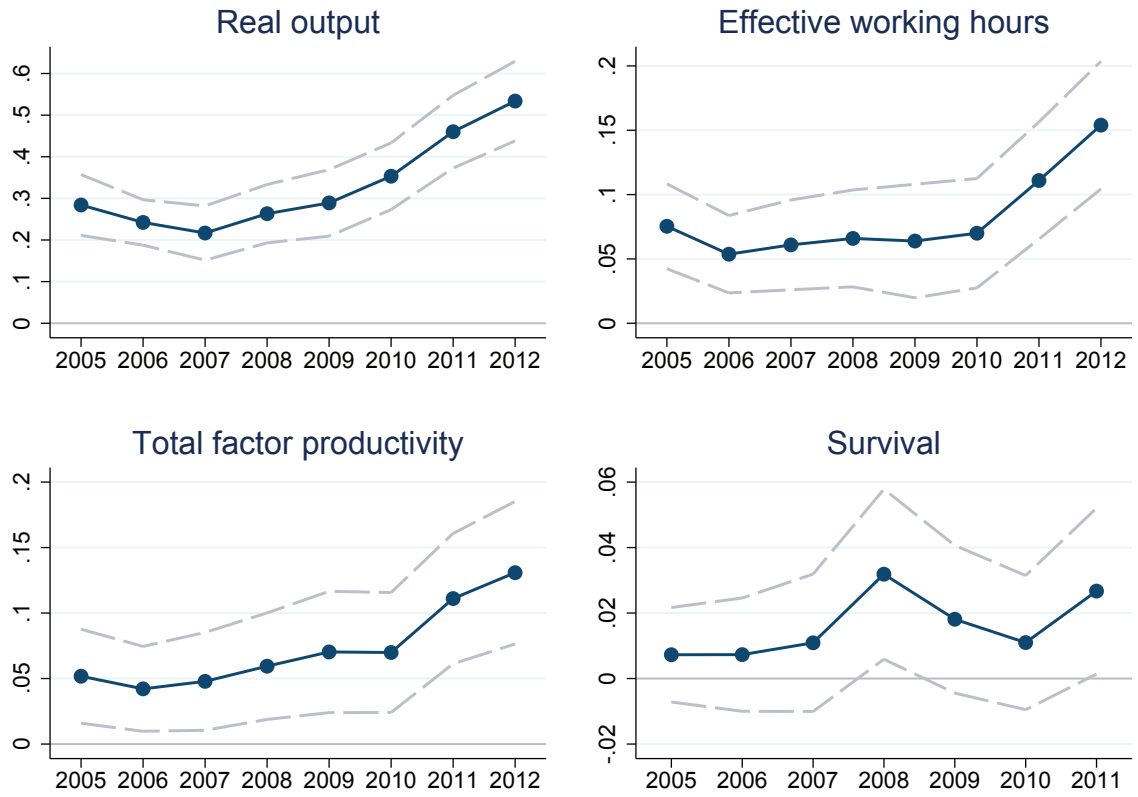
Two comments on these results are in order. First, the estimated exporter premia accord well with the evolution of real output, employment, and TFP depicted in Figure 2.7. As the performance of exporters was less detrimentally affected by the crisis than the performance of non-exporters, the performance premia of exporters consequently increased over the crisis period. Second, the levels of, and the changes in, the estimated exporter premia are arguably determined by a host of different factors.[42] An investigation into these factors is beyond the scope of our analysis. In a similar vein, while our estimates show significant and increasing performance differences between exporters and non-exporters and thus point to extra benefits of exporting in times of crisis, we must emphasize that our regressions do not allow for a *causal* interpretation. It seems at least conceivable that exporting firms fared better during the crisis than non-exporting firms due to factors that gained in importance in the crisis but are not directly related to exporting (such as superior management quality or a more flexible and efficient labor force). A rigorous analysis of these different factors and the question of whether exporting was *causally* associated with a stronger performance in the crisis is left for future research.

As for the estimated exporter premium for firm survival in Figure 2.9, we find that it is positive throughout, but small and insignificant in the pre-crisis period. It is significantly positive for the first

[42]The usual explanation for observing positive exporter premia is that success leads to exporting (based on the idea that only the 'good' and successful firms are able to cover the additional costs associated with exporting) or that exporting leads to success (due to technology or knowledge spillovers and learning by exporting). These explanations are of course not mutually exclusive.

time in 2008, where it reaches its peak of 3.2 percentage points. This estimate suggests that exporters were more likely than non-exporters to survive the peak of the crisis and still be producing in 2009. In the subsequent years, the survival premium remains above the pre-crisis level, but is only significantly different from zero in 2011. Again, this finding is in line with the analysis in Figure 2.8 and suggests that exporting was beneficial for firm survival in the crisis.

Figure 2.9: *Exporter premia, 2005-2012*



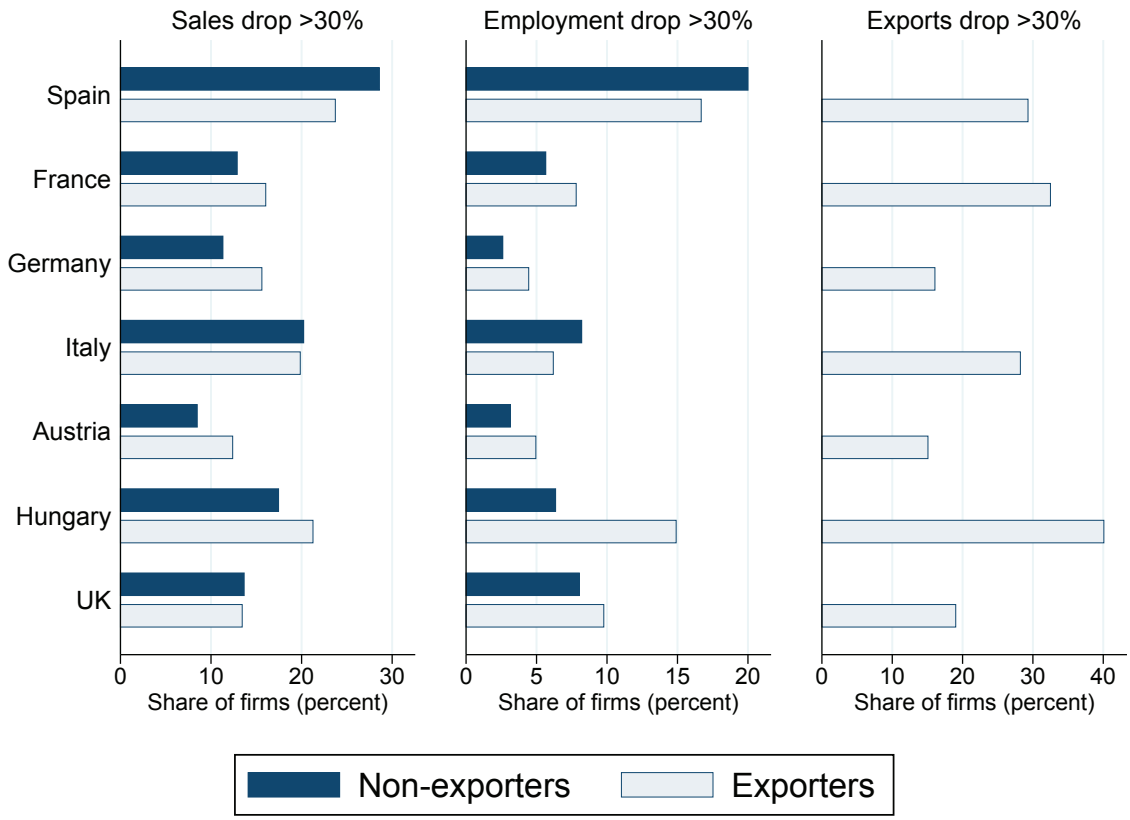
Note: The figure shows estimates of λ_t in Equation (2.7) for $t = 05, \ldots, 12$, alongside 95% confidence intervals. These estimates can be interpreted as year-specific exporter premia for the variables indicated in the subfigures. For output, effective working hours, and TFP, the exporter premia are given in percentages. For survival, the premium is given in percentage points. The estimated premia are conditional on the share of foreign capital in the firm's joint capital (0%; $> 0\%$ & $\leq 50\%$; $> 50\%$), the firm's capital intensity, R&D intensity, skill intensity, type of good produced (intermediate good; final good; not defined), multinational status, as well as the industry-and-size-group cluster to which the firm belongs. Confidence intervals derive from robust standard errors clustered at the firm level. Source: Authors' estimations based on ESEE data.

Is the fact that exporters were so much more resilient to the crisis than non-exporters special to the case of Spain, or do we observe a similar pattern in other European countries too? To answer this question, we exploit direct questions on firms' export status and their crisis performance in the EFIGE data set. As indicators of crisis performance, we compute the shares of firms (by country and export status) that reduced their sales, employment, and exports by at least 30% in 2009 compared to 2008.[43]

Figure 2.10 suggests that Spain is indeed a special case, at least in two dimensions. First, we see that, whether we look at exporters or non-exporters, firms in Spain were hit the hardest by the crisis in terms

[43]The results are qualitatively similar when examining the shares of firms that experienced any reduction at all instead of applying the 30% threshold.

Figure 2.10: *Firms' crisis performance across European countries*



Note: This figure reports the share of firms (by country and export status, in percent) that reduced their sales, employment, and exports by more than 30% in 2009 (compared to 2008). The number of firms reporting changes in sales are 2,832 for Spain, 2,973 for France, 2,929 for Germany, 3,021 for Italy, 439 for Austria, 488 for Hungary, and 2,046 for the UK. The number of firms for all other variables are slightly smaller. Sampling weights apply. Source: Authors' calculations based on EFIGE data.

of overall performance.[44] The share of firms that drastically reduced their sales and their employment was larger in Spain than in any other country covered by the EFIGE data. For example, one fifth of all non-exporting firms in Spain shed more than 30% of their labor in 2009, more than twice the share observed in any other country in our sample.

The second insight from Figure 2.10 is that the superior crisis performance of exporters over non-exporters seems to be the exception rather than the rule in Europe. Not only do the EFIGE data confirm our previous findings that in Spain it was the non-exporting firms that suffered the most in 2009; the data also show that it is *only* in Spain that a significantly larger share of non-exporters than of exporters saw their sales plummet. The same holds true for the strong reductions in employment (with Italy being the only exception and showing a pattern similar to Spain, although on a much lower level). It would be interesting to investigate these cross-country differences (and the underlying causes) in more detail in future work.

We finally notice that in terms of export reductions Spain ranks in the middle of the seven-country sample. This is perhaps surprising in light of the strong relative export performance described in the introduction and depicted in Figure 2.1. A possible explanation may be that in Spain it is very few

[44] Due to data limitations, the export status refers to the year 2008 instead of 2009, but it is highly persistent in the EFIGE sample. More than 97% of exporters, but virtually none of the non-exporters, had exported before 2008 (as reported by firms in the data set).

very successful firms that are responsible for the increase in exports over the last couple of years (as suggested by Antràs, 2011). Moreover, one should keep in mind that in Figure 2.10 we only look at changes between 2008 and 2009, whereas Figure 2.1 describes the evolution of exports over the entire period 2007-2013.

2.4 Discussion

2.4.1 Improved aggregate competitiveness

The analysis provided in this paper is a micro-level analysis emphasizing the role of firm heterogeneity in international trade during the crisis years. Yet, the economic and political discussions surrounding the recent performance of the Spanish economy are often couched in terms of what is vaguely referred to as "international competitiveness" or "aggregate competitiveness". Can recent improvements in competitiveness explain the exceptionally strong export performance of the Spanish economy? We now provide a brief look into relevant data that allow us to give a tentative answer to this question. At this point it proves useful to recall some insights of standard macroeconomic theory. The key assumption is some nominal imperfection in the short run (e.g. price stickiness), such that the real wage does not instantaneously adjust to fluctuations in aggregate demand. This implies that the short-run aggregate supply curve is not vertical, so that a negative shock to aggregate consumption or investment leads to lower output and involuntary unemployment. For a country like Spain, which is well integrated into the global economy, international trade can provide a remedy for such a negative demand shock. This can happen through a nominal depreciation of the Euro vis-à-vis other currencies (over the short and medium run), as well as through differential price changes that improve the Spanish terms of trade within the Eurozone (over the medium and long run). We find evidence that both channels are likely to have played a significant role in the favorable development of Spanish exports.

As for the first channel, nominal exchange rate depreciation, we examine the two most important non-Eurozone destinations for Spanish exports: the United Kingdom (the fifth largest importer of Spanish goods in 2013) and the United States (the sixth largest importer).[45] When we look at the relevant period from January 1, 2009, to December 31st, 2013, the Euro depreciated relative to both the Pound sterling and the US dollar (in each case by 5% when comparing yearly averages). As a result, Spanish products became less expensive in the UK and the US, whereas imports from these two countries became more expensive for Spain. This development seems to be reflected in the trade data: from 2009 to 2013, total Spanish exports to the UK and the US increased by a remarkable 55%, whereas total imports increased by just 9%. In contrast, when considering the Eurozone countries among the top six export destinations (France, Germany, Portugal, and Italy), for which nominal depreciation did not play any role, total Spanish exports increased by much less (viz. by 24%), while imports increased at a similar rate (viz. by 7%).

To offer a more systematic view of the issue, we examine the evolution of nominal *effective* exchange rates for Spain. Nominal effective exchange rates are weighted averages of the usual bilateral nominal exchange rates, where the weights reflect the importance of trading partners in terms of their volume of international trade with Spain (considering only trade in manufacturing goods). It turns out that, irrespective of the details of the weighting scheme applied, we observe a depreciation of Spain's nominal effective exchange rate (EER) over the period 2009-2013, and thus a corresponding improvement of its

[45] All data mentioned in this paragraph are readily accessible through the website of INE.

competitive position. For example, for the ECB EER-38 group of currencies plus the latest composition of the Eurozone, the index for Spain's nominal effective exchange rate decreased from 107.5 in 2009 to 104.8 in 2013 (having normalized the index to 100 in 1999).[46] This observation corroborates the presumption that nominal depreciation has contributed to boosting Spanish exports, while it has made imports from outside the Eurozone more expensive (on average).

As for the second channel, differential price changes may improve the competitive position of Spain even within the Eurozone, where nominal exchange rates are fixed at unity. Since the drop in aggregate demand was stronger in Spain than in most other Eurozone countries (see Figure 2.10 for supporting evidence), we should see a change in *real* exchange rates improving Spain's competitive position vis-à-vis other Eurozone countries (especially those that quickly returned to economic growth after the financial crisis). To investigate this possibility, we first examine real effective exchange rates (REER) based on the ECB's harmonized index of consumer prices (HICP). We find that the HICP-based index for Spain's REER vis-à-vis the latest composition of the Eurozone plus the ECB EER-12 group of currencies decreased only slightly from 114.8 in 2009 to 113.6 in 2013.[47] Moreover, a similar or even more pronounced decrease in this index is observed for many other Eurozone countries as well (e.g. Germany, France, Greece, and Italy). Hence, differences in the development of consumer prices within the Eurozone are not a relevant channel when it comes to explaining the recent export performance of Spain.

One issue with the HICP-based REER is that the evolution of consumer prices might not accurately reflect changes in a country's competitive position in international trade. The reason is that a significant share of consumption expenditure is on non-tradable goods and services, while many tradable goods are not included in the consumption basket (e.g. capital goods). Hence, an important alternative to looking at consumer prices is to focus on unit labor costs. It turns out that for Spain this distinction is extremely important. In particular, the index for Spain's REER based on unit labor costs (but otherwise defined as above) decreased from 119.5 in 2009 to 105.8 in 2013, while it basically stagnated for Germany, Italy, and France. Only Greece, arguably the Eurozone country that experienced the most adverse shock to aggregate demand, showed a development comparable to that of Spain. Hence, Spain has become more competitive internationally through internal devaluation (i.e., real wages growing less than productivity relative to other Eurozone countries).

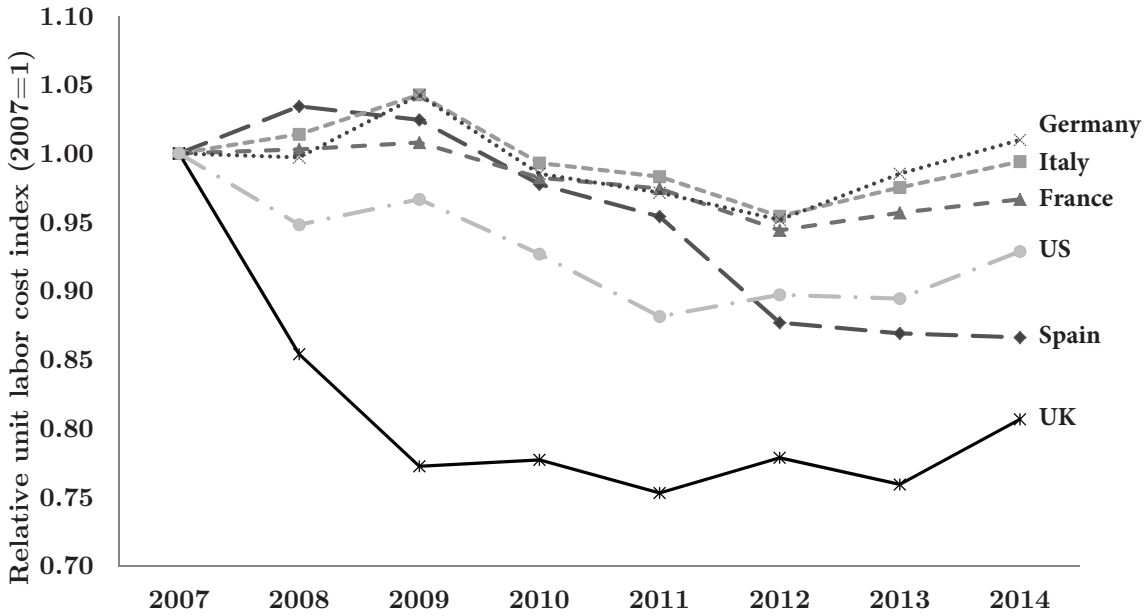
One way to summarize this discussion is to look at the evolution of *relative unit labor costs*, a broad competitiveness measure that reflects both nominal exchange rate depreciation and internal devaluation at the same time. Figure 2.11 shows that, unlike other countries such as Germany, Italy, France, the UK, or the US, Spain has experienced a *steady decline* in relative unit labor costs since 2008. The decrease in relative unit labor costs was particularly pronounced in 2012, the year for which we have estimated the strongest conditional difference in both export participation and export intensity compared to the pre-crisis period (see Tables 2.2 and 2.3). Interestingly, the development since 2008 is in stark contrast to the one before 2008. Antràs (2011) describes how exports from Spain fared surprisingly well in spite of *deteriorating* competitiveness observed prior to the crisis. Hence, the evolution of relative unit

[46] The depreciation is not monotonic over the period considered and has its minimum in 2012 at 102.6. The countries in the ECB EER-38 group of currencies are the non-Eurozone EU member states (Bulgaria, Czech Republic, Denmark, Hungary, Poland, Romania, Sweden, and the United Kingdom), plus Algeria, Argentina, Australia, Brazil, Canada, Chile, China, Croatia, Hong Kong, Iceland, India, Indonesia, Israel, Japan, Malaysia, Mexico, Morocco, New Zealand, the Philippines, Russia, Norway, Singapore, South Africa, South Korea, Switzerland, Taiwan, Thailand, Turkey, the United States, and Venezuela. All data on nominal and real exchange rates come from the Statistical Data Warehouse of the ECB: http://sdw.ecb.europa.eu/.

[47] The countries in the ECB EER-12 group of currencies are Australia, Canada, Denmark, Hong Kong, Japan, Norway, Singapore, South Korea, Sweden, Switzerland, the United Kingdom, and the United States.

labor costs may play a role in explaining the strong export performance of Spain in the years after 2008, while it is more difficult to reconcile with the export performance over the previous decade. It should be noted, however, that the largest firms in Spain experienced a more favorable development of unit labor costs than many of their European competitors already before 2008 (as pointed out by Antràs, 2011).[48] Because the large firms are responsible for the lion's share of a country's exports, it seems plausible that this development had a positive impact on aggregate exports also before 2008.

Figure 2.11: *Relative unit labor costs, 2007-2014*



Note: Competition-weighted relative unit labor costs for the overall economy in US dollar terms. Competition weights take into account the structure of competition in both export and import markets of the goods sector of 49 countries. An increase in the index indicates a real effective appreciation and a corresponding deterioration of the competitive position. The index accounts for annual shifts in the composition of trade flows. Source: OECD Economic Indicators. See also OECD Economic Outlook Sources and Methods at http://www.oecd.org/eco/sources-and-methods.htm.

It emerges from our micro-level analysis that the factors underlying the observed increase in aggregate competitiveness in Spain are unrelated to technological improvements or a more efficient resource allocation, as a large majority of manufacturing industries in Spain experienced a *decrease* in aggregate TFP (as demonstrated above). Rather, the enhanced competitiveness of Spain seems closely linked to the country's problematic labor market performance. Record unemployment rates around 25% in 2012 and 2013 put downward pressure on real wages.[49] In the face of downward nominal wage rigidity and overall low inflation rates, however, it is to a significant extent the increase in real wages in some of Spain's most important trading partners that boosted Spain's competitive position. For example, real wages in Germany increased by 3.6% (while unemployment decreased) over the period 2010-2013. This is by far the largest increase the country has experienced over any three-year period in the recent past.[50] Exports from Spain have thus clearly benefited from the cross-country differences in labor market adjustments within Europe.

An interesting implication of our analysis in Section 2.3.4 is that exporting firms contributed the most

[48] In a similar vein, Antràs (2011) argues that in terms of productivity the largest firms in Spain outperform not only their domestic peers, but also their European competitors.

[49] See Eurostat data available at http://ec.europa.eu/eurostat/web/lfs/data/database.

[50] See the (real) wage data available from the OECD at http://stats.oecd.org/.

to enhanced aggregate competitiveness in terms of relative unit labor costs. These firms increased their output after 2009, while at the same time reducing both their employment and their real wages. While our data show that this development is not driven by technological improvements within the firm,[51] it is in sharp contrast to non-exporting firms, which continued to reduce their output and saw their productivity decline even after the financial crisis.

2.4.2 Substituting domestic with foreign sales

Another explanation for the strong export performance of Spain that we want to discuss in some detail is directly tied to the fact that aggregate demand was more strongly adversely affected in Spain than in many other countries in the aftermath of the financial crisis and through the subsequent years.[52] As mentioned in the introduction, Spain went through a double-dip recession in 2009 and 2011-13. The findings of our micro-data analysis, in particular the observations that more firms entered the export market after the financial crisis, while existing exporters increased their share of exports in total sales, suggest that firms in Spain substituted domestic with foreign sales.

Under which conditions can we expect firms to treat domestic and foreign sales as substitutes? Not necessarily under the conditions of standard trade theory (e.g. Melitz, 2003), where firms maximize profits in the domestic and the export market independently of one another. Moreover, failure in one market might actually trigger failure in the other market, for example when liquidity constraints take hold (as sales in the domestic market might generate the extra liquidity necessary to generate sales in the foreign market). However, two recent papers construct theoretical models that help rationalize a substitutive rather than complementary relationship between domestic and foreign sales. Both papers are based on the assumption of production costs being convex in the short run; we briefly discuss both of these papers in turn and argue that they are likely to be relevant in the Spanish case.

A first explanation is advanced by Vannoorenberghe (2012). In his model with labor as the only factor of production, firms face increasing marginal costs of production in the short run, as it becomes increasingly expensive for the firm to employ more labor. This implies a negative correlation between foreign and domestic sales within the firm, and a negative domestic demand shock will induce the firm to sell larger quantities on the export market in the short run. The overall rigidity and dual nature of the Spanish labor market, granting high job protection to permanent employees, makes such cost convexities very likely, as they might rule out the first-best response of firms to demand shocks (in terms of adjusting the labor input). Hence, the theory by Vannoorenberghe (2012) might partly explain the significant increase in the export intensity of incumbent exporters over the crisis period that we have identified in Table 2.3.

Can a slump in domestic demand also trigger the entry of new firms into the export market that we observe in Spain after 2009? Focusing on the extensive margin, Blum et al. (2013) set up a Melitz (2003) model in which firms face capacity constraints due to investments in fixed capital ex ante (i.e., before demand shocks in the domestic and the export market are known to the firm).[53] This assumption again implies convex production costs in the short run. In response to an adverse shock to domestic demand, their model predicts not only a shift of sales to the foreign market within exporters (as in Vannoorenberghe, 2012), but also export market entry of firms that were previously confined to the

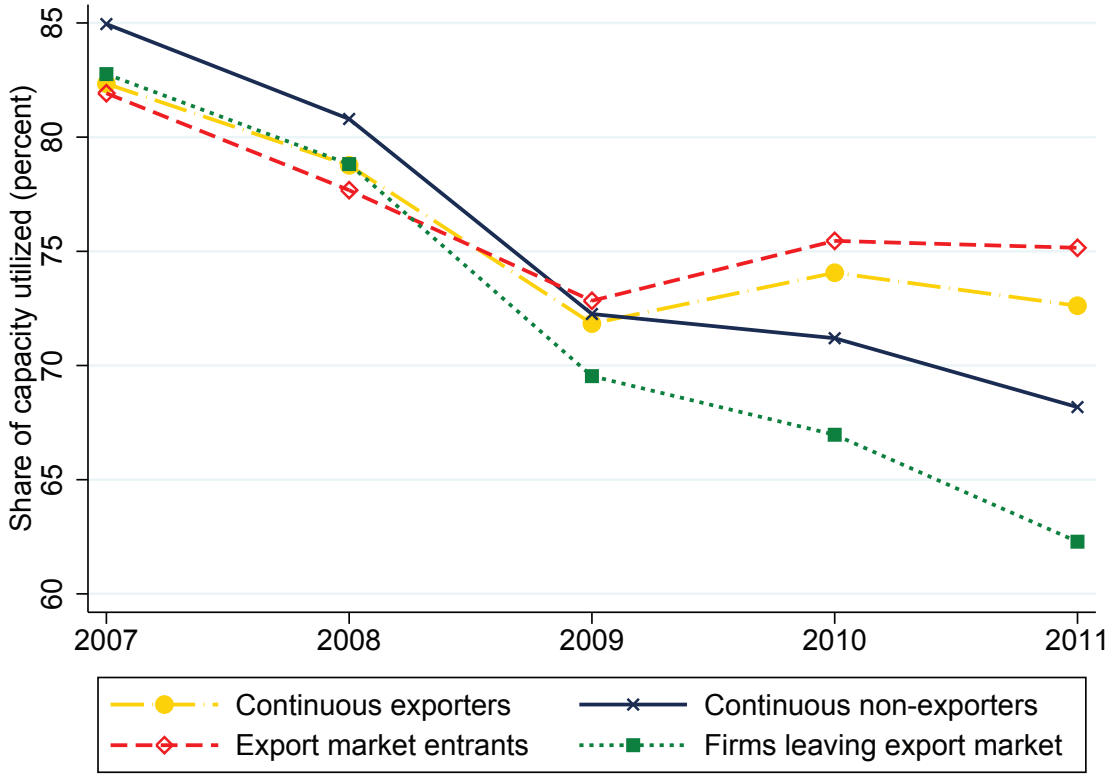
[51] As shown in Figure 2.7, exporting firms had about the same level of TFP in 2009 as in 2011.

[52] For example, GDP data from INE as well as the data depicted in Figure 2.10 suggest that the drop in aggregate demand in 2009-2012 was more dramatic in Spain than in most of its major trading partners.

[53] See Soderbery (2014) for a model that yields similar predictions based on linear demand and capacity constraints.

domestic market. The intuition for this prediction is that if capacity is utilized below a certain threshold, it is optimal for the firm to bear the cost of exporting. For some firms, exporting then essentially serves as a means to survive the shock. Extreme changes in domestic demand, like the deep recession in Spain, can thus push more firms into the export market in the short run. This intuition is also in line with the findings by Belke et al. (2015), who provide empirical support for a substitutive relationship between domestic and foreign sales in their analysis of macro data for several Eurozone countries. The evidence is particularly strong for the southern European countries Italy, Portugal, and Spain.

Figure 2.12: *Capacity utilization, 2007-2011*



Note: The sample is balanced on firms that are observed in each year from 2007 to 2011. Capacity utilization is measured in percentage points. For the definition of the different groups of firms, see the note to Figure 2.7. Sampling weights apply. Source: Authors' calculations based on ESEE data.

Figure 2.12 demonstrates that our micro data are consistent with an explanation based on capacity constraints and hence convex production costs in the short run. It depicts the evolution of average capacity utilization by groups of firms defined according to their export status (in analogy to Figure 2.7). From 2007 to 2009, capacity utilization dropped from above 82% to below 73% for all groups of firms. A reduction of this size seems large enough to trigger the above-mentioned adjustments. Of particular interest is the heterogeneity in this drop and the development during the later years. On the one hand, for continuous non-exporters, capacity utilization continued to decline after 2009, giving rise to an accumulated loss of 17 percentage points (20%) for the period from 2007 to 2011.[54] For continuous and new exporters, on the other hand, capacity utilization not only declined by less from 2007 to 2009, but it also increased again after 2009. Interestingly, export market entrants show the most favorable development over the observed period, which suggests that starting to export indeed helped these firms to

[54] Firms leaving the export market experienced an even larger decline of 20 percentage points (25%).

exploit otherwise unused (and costly) capacities. Taken together, the available evidence suggests that cost convexities might have contributed to the fact that firms rely more on exports today than before the financial crisis, as we have documented in our main analysis.

2.5 Conclusion

We have explored a rich firm-level data set from Spain to provide novel evidence on firm behavior during and after the 2008 financial crisis. We have investigated changes at the extensive and intensive firm-level margins of trade, as well as performance differences (jobs, productivity, and firm survival) between exporting and non-exporting firms. We find that the trade collapse in 2009 is almost fully explained by adjustments at the intensive margin. The number of firms that were forced to exit the export market due to the crisis is negligible, and firms allocated a larger fraction of their sales to foreign markets, especially in the years after 2010. Moreover, we find a growing performance gap between exporters and non-exporters, which shows that exporters proved to be more resilient to the economic challenges they were facing during the global financial crisis and subsequent recession. Exporters now account for a larger share of output and jobs, and they contribute more to aggregate productivity than before the financial crisis. However, we observe a decline in the aggregate productivity of the manufacturing sector as a whole over the crisis period. These findings are of direct relevance for the ongoing political debate about the current and future economic situation in Spain.

We conclude by pointing out some interesting similarities between the current situation in Spain and the situation observed more than a decade ago in Germany.[55] From the mid-1990s and into the 2000s, Germany suffered from high unemployment and poor economic growth. However, relative unit labor costs had set out to decrease in 1995, boosting German exports through a gradual improvement of the country's competitive position in the global economy. The same seems to be currently happening in Spain. At the time of the financial crisis, Germany had already been the world champion of exports for several years, economic growth had returned, and unemployment had been brought down. Germany had transformed itself from the "sick man of Europe" into an "economic superstar" (Dustmann et al., 2014).

A compelling narrative behind this development, advanced by Dustmann et al. (2014), is that German labor market institutions were flexible enough to allow for a significant decentralization of the wage-setting process, away from the industry level towards the firm level. This decentralization, largely triggered by the fall of the iron curtain and the pressures of economic globalization, translated into a considerable decline in real wages at the lower end of the wage distribution, and ultimately to a more competitive economy. It was one of the principal aims of the 2012 labor market reform in Spain to widen the scope of collective bargaining at the firm level (OECD, 2014). Future research will need to show whether this reform can contribute to a better-functioning labor market that can mimic the German success and further enhance the international competitiveness of Spanish exports.

[55] See also the article "Spain's economy: Not yet the new Germany," published on March 9th, 2013, by The Economist.

Bibliography

Abreha, Kaleb G., Valérie Smeets, and Frédéric Warzynski, "Coping with the Crisis and Export Diversification," Aarhus University, mimeo 2016.

Altomonte, Carlo and Tommaso Aquilante, "The EU-EFIGE/Bruegel-Unicredit Dataset," Technical Report 753, Bruegel 2012.

_ , Filippo Di Mauro, Gianmarco Ottaviano, Armando Rungi, and Vincent Vicard, "Global Value Chains During the Great Trade Collapse: A Bullwhip Effect?," Technical Report 1412, European Central Bank 2012.

Álvarez, Roberto and Camila Sáez, "Post Financial Crisis and Exports Expansion: Micro-evidence From Chilean Exporters," MPRA Paper 60637, University Library of Munich 2014.

Antràs, Pol, "El Comportamiento de las Exportaciones Españolas," competitividad 02, Apuntes FEDEA november 2011.

Arellano, Manuel and Stephen Bond, "Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations," *Review of Economic Studies*, 1991, *58* (2), 277–297.

Aristei, David, Davide Castellani, and Chiara Franco, "Firms' Exporting and Importing Activities: Is There a Two-way Relationship?," *Review of World Economics*, 2013, *149* (1), 55–84.

Baldwin, Richard, "Introduction: The Great Trade Collapse: What Caused It and What Does It Mean?," in Richard Baldwin, ed., *The Great Trade Collapse: Causes, Consequences and Prospects*, 2009.

Behrens, Kristian, Gregory Corcos, and Giordano Mion, "Trade Crisis? What Trade Crisis?," *Review of Economics and Statistics*, 2013, *95* (2), 702–709.

Belke, Ansgar, Anne Oeking, and Ralph Setzer, "Domestic Demand, Capacity Constraints and Exporting Dynamics: Empirical Evidence for Vulnerable Euro Area Countries," *Economic Modelling*, 2015, *48* (C), 315–325.

Bernard, Andrew B. and J. Bradford Jensen, "Why Some Firms Export," Working Paper Department of Economics 97-26, Massachusetts Institute of Technology 1997.

_ and _ , "Exceptional Exporter Performance: Cause, Effect or Both?," *Journal of International Economics*, 1999, *47* (1), 1–25.

_ , _ , Stephen J. Redding, and Peter K. Schott, "The Margins of US Trade," *American Economic Review*, 2009, *99* (2), 487–93.

_ , _ , _ , and _ , "The Empirics of Firm Heterogeneity and International Trade," *Annual Review of Economics*, 2012, *4*, 283–313.

Blum, Bernardo S., Sebastian Claro, and Ignatius J. Horstmann, "Occasional and Perennial Exporters," *Journal of International Economics*, 2013, *90* (1), 65 – 74.

Bricongne, Jean-Charles, Lionel Fontagné, Guillaume Gaulier, Daria Taglioni, and Vincent Vicard, "Firms and the Global Crisis: French Exports in the Turmoil," *Journal of International Economics*, 2012, *87* (1), 134–146.

Burda, Michael and Jennifer Hunt, "What Explains the German Labor Market Miracle in the Great Recession?," *Brookings Papers on Economic Activity*, 2011, *42* (1), 273–335.

Chor, Davin and Kalina Manova, "Off the Cliff and Back? Credit Conditions and International Trade during the Global Financial Crisis," *Journal of International Economics*, 2012, *87* (1), 117–133.

Correa-López, Mónica and Rafael Doménech, "The Internationalisation of Spanish Firms," Technical Report 12/30, BBVA 2012.

Costa, Stefano, Carmine Pappalardo, and Claudio Vicarelli, "Financial Crisis, Internationalization Choices and Italian Firm Survival," MPRA Paper 54107, University Library of Munich 2014.

Crespo, Aránzazu, Klaus Desmet, and Susanna Esteban, "Did Spanish Firms Perform Worse in the Wake of the 2008 Crisis?," EFIGE Country Report: Spain, Bruegel 2011.

Dustmann, Christian, Bernd Fitzenberger, Uta Schönberg, and Alexandra Spitz-Oener, "From Sick Man of Europe to Economic Superstar: Germany's Resurgent Economy," *Journal of Economic Perspectives*, 2014, *28* (1), 167–188.

Eaton, Jonathan, Samuel Kortum, Brent Neiman, and John Romalis, "Trade and the Global Recession," *American Economic Review*, 2016, *106* (11), 3401–38.

Feng, Ling, Zhiyuan Li, and Deborah L. Swenson, "The Connection Between Imported Intermediate Inputs and Exports: Evidence From Chinese Firms," NBER Working Paper 18260, NBER 2012.

Garicano, Luis, *El Dilema de España - Ser Más Productivos Para Vivir Mejor*, Península, 2014.

Giri, Rahul, Enrique Seira, and Kensuke Teshima, "Exporters During the Trade Collapse: The (Surprising) Resiliency of the Small Exporter," Working Papers 2014-06, Banco de México 2014.

Görg, Holger and Marina-Eliza Spaliara, "Export Market Exit, Financial Pressure and the Crisis," Kiel Working Papers 1859, Kiel Institute for the World Economy 2013.

_ and _ , "Exporters in the Financial Crisis," *National Institute Economic Review*, 2014, *228* (1), R49–R57.

Greenaway, David and Richard Kneller, "Firm Heterogeneity, Exporting and Foreign Direct Investment," *Economic Journal*, 2007, *117* (517), F134–F161.

Gregory, Mary and Robert Jukes, "Unemployment and Subsequent Earnings: Estimating Scarring Among British Men 1984-94," *Economic Journal*, 2001, *111* (475), 607–625.

Guadalupe, Maria, Olga Kuzmina, and Catherine Thomas, "Innovation and Foreign Ownership," *American Economic Review*, 2012.

Halpern, László, Miklós Koren, and Adam Szeidl, "Imported Inputs and Productivity," *American Economic Review*, 2015, *105* (12), 3660–3703.

Hanley, Aoife and Joaquín Monreal Pérez, "Are Newly Exporting Firms More Innovative? Findings from Matched Spanish Innovators," *Economics Letters*, 2012, *116* (2), 217–220.

Hospido, Laura and Eva Moreno-Galbis, "The Spanish Productivity Puzzle in the Great Recession," IZA Discussion Paper 8891, IZA 2015.

Kasahara, Hiroyuki and Beverly Lapham, "Productivity and the Decision to Import and Export: Theory and Evidence," *Journal of International Economics*, 2013, *89* (2), 297–316.

Kohler, Wilhelm and Marcel Smolka, "Global Sourcing and Firm Selection," *Economics Letters*, 2014, *124* (3), 411–415.

Levinsohn, James and Amil Petrin, "Estimating Production Functions Using Inputs to Control for Unobservables," *Review of Economic Studies*, 2003, *70* (2), 317–341.

Loecker, Jan De, "Do Exports Generate Higher Productivity? Evidence from Slovenia," *Journal of International Economics*, 2007, *73* (1), 69–98.

Melitz, Marc J., "The Impact of Trade on Intra-Industry Reallocations and Aggregate Industry Productivity," *Econometrica*, 2003, *71* (6), 1695–1725.

Monarch, Ryan and Tim Schmidt-Eisenlohr, "Learning and the Value of Relationships in International Trade," CESifo Working Paper Series 5724, CESifo Group Munich 2016.

OECD, "The 2012 Labour Market Reform in Spain: A Preliminary Assessment," OECD Publishing, OECD, Paris 2014.

Olley, G Steven and Ariel Pakes, "The Dynamics of Productivity in the Telecommunications Equipment Industry," *Econometrica*, 1996, *64* (6), 1263–1297.

Paravisini, Daniel, Veronica Rappoport, Philipp Schnabl, and Daniel Wolfenzon, "Dissecting the Effect of Credit Supply on Trade: Evidence from Matched Credit-Export Data," *The Review of Economic Studies*, 2014.

Roberts, Marc J. and James R. Tybout, "The Decision to Export in Colombia: An Empirical Model of Entry with Sunk Costs," *American Economic Review*, 1997, *87* (4), 545–564.

Smeets, Valérie and Frédéric Warzynski, "Estimating Productivity with Multi-Product Firms, Pricing Heterogeneity and the Role of International Trade," *Journal of International Economics*, 2013, *90* (2), 237–244.

Soderbery, Anson, "Market Size, Structure, and Access: Trade with Capacity Constraints," *European Economic Review*, 2014, *70*, 276 – 298.

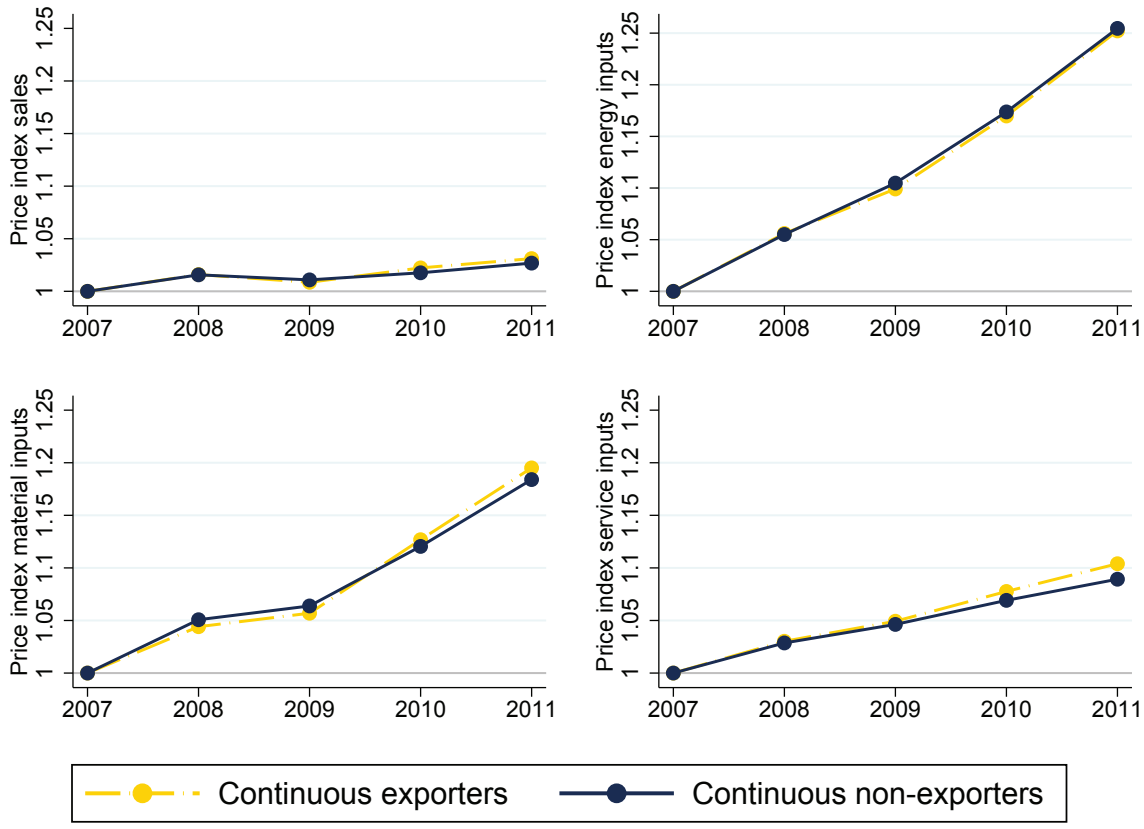
Vannoorenberghe, Gonzaghue, "Firm-level Volatility and Exports," *Journal of International Economics*, 2012, *86* (1), 57 – 67.

2.A Appendix

Table 2.A.1: *Summary statistics*

Variable	Obs.	Mean	Std. Dev.	Min.	Max.
Exporter dummy	15,074	0.641	0.48	0.000	1.000
Importer dummy	15,051	0.63	0.483	0.000	1.000
Export value (1,000 EUR in prices of 2010)	15,074	26,270	205,000	0	7,731,165
Domestic sales (1,000 EUR in prices of 2010)	15,074	39,065	167,334	0	6,058,877
Import value (1,000 EUR in prices of 2010)	15,051	16,417	124,840	0	3,780,449
Domestic purchases (1,000 EUR in prices of 2010)	15,048	35,744	179,316	0	5,463,883
Market exit	12,433	0.033	0.179	0.000	1.000
Multinational dummy	15,074	0.090	0.286	0.000	1.000
Foreign ownership: =0%	15,060	0.846	0.361	0.000	1.000
Foreign ownership: >0% & <=50%	15,060	0.021	0.142	0.000	1.000
Foreign ownership: >50%	15,060	0.133	0.340	0.000	1.000
Labor productivity (in logs)	14,935	12.268	1.696	2.785	19.16
Capital intensity (in logs)	15,043	4.357	1.140	-2.303	9.013
R&D intensity (in logs)	15,042	0.008	0.032	0.000	2.636
Skill intensity (in logs)	14,629	0.192	0.253	0.000	2.760
Type of good: final good	14,743	0.162	0.368	0.000	1.000
Type of good: intermediate good	14,743	0.264	0.441	0.000	1.000
Type of good: not defined	14,743	0.574	0.494	0.000	1.000
TFP (in logs)	14,914	-0.105	1.371	-4.947	5.555
Real output (in logs)	15,074	15.887	1.969	9.921	22.765
Effective working hours (in logs)	15,051	11.563	1.416	7.496	17.029
Hourly wage (in logs)	14,996	2.853	0.406	0.397	4.825
Internet dummy	15,060	0.774	0.418	0.000	1.000

Figure 2.A.1: *Evolution of firm-level output and input prices, 2007-2011*



Note: The sample is balanced on firms that are observed in each year from 2007 to 2011. Continuous exporters are 671 firms that export in each of the years 2005 to 2012 (if observed); continuous non-exporters are 310 firms that do not export in any of the years 2005 to 2012 (if observed). All variables are normalized to one in 2007. Sampling weights apply. Source: Authors' calculations based on ESEE data.

Chapter 3

External Financial Dependence and Firms' Crisis Performance across Europe*

How do financial market conditions affect real economic performance? Empirical investigations of this question have often relied on measures of external financial dependence (EFD) that are constructed using U.S. data and applied to other countries under the assumption of a stable industry ranking across countries. This paper exploits unique, comparable survey data from seven European countries to show that correlations of EFD across countries are weak, casting some doubt on this assumption. We then use the novel survey-based EFD index to show that the global financial crisis had a disproportionately negative impact on the real performance of financially dependent firms.

*This chapter is based on joint work with Katja Neugebauer.

3.1 Introduction

How do financial market conditions impact on real economic performance? This question has been examined at least since Schumpeter (1911) and regained particular relevance after the global financial crisis that started in 2008. Rajan and Zingales (1998, henceforth RZ) achieved significant progress towards establishing a causal effect of financial development on real growth by exploiting differences in external financial dependence (EFD) across industries. In their seminal contribution, RZ measure industry-level EFD as the share of investment not financed by internal cash flow in the median listed U.S. firm (from the Compustat database) over the 1980s. Their approach rests on two main assumptions: First, if the U.S. capital market is close to perfect, credit demand by listed U.S. firms should not be contaminated by supply-side imperfections, but instead reflect technological fundamentals.[1] Second, in applying the EFD index of U.S. industries to other countries, RZ assume that the industry ranking is constant across countries. It is the second assumption that we investigate more closely in this paper.

Since the seminal contribution by RZ, their EFD index and updated versions of it have been used in many applications to different research questions and countries.[2] For instance, Manova (2013) uses the index to identify the role of credit constraints for international trade, Dell'Ariccia et al. (2008) and Kroszner et al. (2007) examine how the real effects of banking crises vary by EFD, and Chor and Manova (2012) analyze the differential impact of the global financial crisis on exports.

In this paper, we exploit a unique survey question in the EFIGE dataset[3] to obtain a novel, country-specific index of EFD for seven European countries. We first correlate industry rankings of EFD across this EFIGE index and an updated version of the Compustat index, computed from U.S. data following RZ. To shed some light on the stable-ranking assumption, we then examine correlations of the EFIGE index across countries. Finally, we apply both the Compustat index and the country-specific EFIGE index in firm-level regression analysis to examine the differential effect of the global financial crisis on the real performance of manufacturing firms across industries with varying EFD.

We find that industry rankings of EFD are weakly correlated across the two measures and also across European countries, which casts some doubt on the stable-ranking assumption. The subsequent regression analysis reveals that the crisis had a disproportionately negative impact on firm performance in financially dependent industries according to the EFIGE index of EFD. By contrast, the Compustat index delivers insignificant or counter-intuitive estimates.

The paper is structured as follows. In Section 3.2, we characterize the data used and introduce the novel index of EFD based on the EFIGE survey. Section 3.3 shows how this measure correlates with the updated Compustat index and across countries. Section 3.4 exploits both EFD measures to examine how EFD affected firms' performance in the financial crisis. The last section draws conclusions for future research.

[1] While this conjecture is not the main subject of this paper, the global financial crisis of 2008, which originated in the U.S., has revealed that U.S. capital markets are still far from frictionless even in the 21^{st} century.

[2] At the time of this writing, the paper by RZ ranks among the top 100 most cited research papers in economics (https://ideas.repec.org/top) with more than 7,500 citations registered on Google scholar (https://scholar.google.com, both accessed on February 17, 2017).

[3] The data were collected in the project "European Firms in a Global Economy" (Altomonte and Aquilante, 2012, see http://bruegel.org/efige/).

3.2 Data

To obtain the EFIGE index of EFD, we exploit the following question in the survey:

> *In the industry your firm works, how dependant [sic] are companies on external financing? To give your answer please use a score from 1 (not dependent [at] all) to 5 (Extremely dependent).*

Our novel index of EFD is the arithmetic mean of firms' responses by industry j and country c. This measure has three key advantages: First, it is directly comparable across seven European countries, as the identical question was posed to 14,364 (representatively sampled) manufacturing firms at the same time in 2010.[4] Second, it mitigates reporting bias by addressing general conditions in the firm's industry rather than the firm's own financial situation. Third, its general formulation is designed to cover all relevant aspects of EFD.

Since the original Compustat index by RZ is not available for the European industry classification used in EFIGE (NACE Rev. 1.1), we follow RZ in computing the index from Compustat data on U.S. firms. Each Compustat firm is assigned the NACE code corresponding to its SIC code. We select data from the more recent, pre-crisis period 1990-2005 and apply the measure to all countries for our subsequent analysis, following the RZ assumption that EFD as a fundamental industry characteristic should be stable across countries and over time.[5]

For the analysis of firm performance, we merge the industry-level EFD measures to the Orbis firm dataset provided by Bureau van Dijk (BvD). Our panel includes 190,418 manufacturing firms from the seven EU countries under study over the period 2005-2010. It covers only firms (i) whose core activity is classified as manufacturing by their NACE code and (ii) which belong to the size classes medium, large, and very large, as defined by BvD. We compute real growth rates of performance variables (see Section 4), using producer price indices at the most disaggregate industry level that is available from Eurostat for each country (typically 4-digit NACE).

3.3 Comparing Industry Rankings of EFD

Industry rankings of EFD for the updated Compustat index, computed from U.S. data following RZ, and the survey-based EFIGE index by country are listed in Table 3.A.1 in Appendix 3.A. Table 3.1 shows Spearman rank correlation coefficients for all pairwise comparisons between these rankings.

Two observations stand out. First, the ranking of U.S. industries based on the Compustat index is not positively correlated with the rankings of EFD as perceived by European firms and reported in the EFIGE survey. Instead, the correlation coefficients reported in the first row of Table 3.1 are even negative for most countries except Italy and Spain, for which they are small and insignificant.

Second, when comparing the EFIGE index across countries, the industry ranking is highly unstable. Only for six out of 21 pairwise comparisons does the correlation coefficient exceed 0.3 and it is only significant at the 5% level (based on a two-sided t-test) in three of these cases. The correlation is close

[4] The data contain representative samples for manufacturing firms with 10 or more employees in Austria (AUT), France (FRA), Germany (DEU), Hungary (HUN), Italy (ITA), Spain (ESP), and the U.K. (GBR).

[5] See Appendix 3.A for details of this procedure. In a related paper, Ferrando et al. (2008) compute the Compustat index for listed European firms.

to zero for most country pairs and even negative in eight cases.[6]

Provided that the EFD score reported by firms for their industry is systematically related to the fundamental EFD, these observations indicate that (i) the Compustat index based on U.S. data is uncorrelated with EFD in European industries and (ii) even within Europe, there are substantial differences in the industry rankings of EFD across countries. These findings cast some doubt on the stable-ranking assumption and suggest that it might be advisable to take country-specific factors into account when investigating the role of EFD.

Table 3.1: *Correlations of EFD rankings across countries*

	AUT	DEU	ESP	FRA	GBR	HUN	ITA
U.S. (Compustat)	-0.2707	-0.0200	0.0889	-0.1680	-0.2087	-0.0652	0.1104
AUT (EFIGE)		-0.0767	**0.5609****	0.2887	-0.3699	-0.3263	-0.5414**
DEU (EFIGE)			0.2739	**0.4279****	-0.1174	0.2925	**0.3600***
ESP (EFIGE)				**0.5178****	-0.1196	0.2105	-0.1937
FRA (EFIGE)					0.0761	-0.0446	0.0247
GBR (EFIGE)						0.2826	**0.3391**
HUN (EFIGE)							**0.3982***

The table shows Spearman rank correlation coefficients for pairwise comparisons between the rankings of EFD across countries listed in Table 3.A.1. The EFD index for U.S. firms is computed from Compustat for 1990-2005, following RZ. The remaining measures are based on average values of reported EFD by industry and country from the EFIGE survey. Correlation coefficients exceeding 0.3 are marked in bold. Asterisks indicate significance levels based on a two-sided t-test: * p<0.10, ** p<0.05.

3.4 Firm Performance in the Global Financial Crisis

We now use the Orbis panel dataset for 2005-2010 to analyze the differential impact of the global financial crisis on firms' real performance depending on EFD. This exercise fulfills the double purpose of (i) assessing the detrimental impact of the crisis on firm performance through the credit channel, and (ii) evaluating the usefulness of the alternative EFD measures for this purpose.[7]

We estimate the following econometric model:

$$\Delta \ln Y_{cijt} = \beta \, \mathit{Crisis}_{ct} \times \mathit{EFD}_{cj} + \delta_{ct} + \delta_{cij} + \varepsilon_{cijt}, \tag{3.1}$$

where $\Delta \ln Y_{cijt} \equiv \ln Y_{cijt} - \ln Y_{cij,t-1}$ measures real growth in the performance of firm i, which is active in country c and industry j in year t. We examine the following dimensions of firm performance Y_{cijt}: real turnover (operating revenues), employment (number of workers), real labor productivity (value added per worker), and real exports (only available for AUT, GBR, and HUN). The key explanatory variable is the interaction term $\mathit{Crisis}_{ct} \times \mathit{EFD}_{cj}$ between the EFD measure (either from Compustat or EFIGE) and the dummy variable Crisis_{ct}, which equals one in the years of the banking crisis, as classified by the Worldbank's Global Financial Development Database (GFDD, Cihak et al., 2012).[8] In theory,

[6] Some firms in the EFIGE dataset were surveyed in a pilot study, some months before the main survey. Also, some industries host few firms. In unreported robustness checks, we confirm that the general picture of weak correlations in Table 3.1 remains unchanged after excluding firms from the pilot study or restricting the sample to countries and industries with at least ten observations.

[7] Our seven-country sample does not offer sufficient cross-country variation in financial development to fully horse-race our new EFD measure against the original RZ specification. We therefore exploit the shock to credit conditions in the crisis for identification, which also offers the advantage of controlling for firm-specific effects.

[8] The GFDD indicate that the banking crisis started already in 2007 in GBR, but only in 2008 in the other six countries, and it did not end until 2010.

we would expect that a negative credit supply shock in the crisis tightens existing credit constraints and thereby reduces the quantities of inputs employed and output produced by constrained firms (captured by Y_{cijt}).[9] Furthermore, the effect of credit constraints should be stronger in industries that depend more on external finance, as shown theoretically by Manova (2013) for exports. Based on this hypothesis, we expect $\beta < 0$.

Importantly, the firm fixed effect δ_{cij} in Equation (3.1) accounts for any time-invariant characteristics of countries, industries, and firms (such as firm size, productivity, and the level of EFD). The country-year fixed effect δ_{ct} controls for the overall crisis impact in each country and any other country-specific shocks. Equation (3.1) is essentially a firm-level variant of the main specification by Dell'Ariccia et al. (2008), who assess the effects of banking crises on real performance in a panel of countries and industries. Compared to their specification, our approach cannot include industry-year fixed effects, because we look at a single crisis, but it has the significant advantage of exploiting within-firm variation.

Table 3.2 summarizes our results of estimating Equation (3.1) for different performance variables and the two alternative EFD measures. When measuring EFD based on the EFIGE survey, our hypothesis is confirmed: all dimensions of firm performance were more negatively affected by the crisis in financially dependent industries compared to industries with low EFD. The estimated interaction effect is always negative and significant at conventional levels (with p-values in the range of 1-9%). In contrast, the interaction effect with the Compustat index is zero for employment and exports, and it suggests a counter-intuitive positive correlation for turnover and labor productivity. These results indicate that *if* the credit crunch had a differential effect on firm performance in line with our hypothesis and the existing literature, then the EFIGE measure is able to identify this effect for European firms, while the Compustat index, which relies on U.S. data, is not.

Table 3.2: *Differential crisis impact on firm performance by EFD*

	Turnover		Employment		Labor productivity		Exports	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Crisis × *EFD* (EFIGE)	-0.075**		-0.041*		-0.080**		-0.083**	
	(0.036)		(0.024)		(0.036)		(0.032)	
Crisis × *EFD* (Compustat)		0.010**		0.001		0.011**		0.003
		(0.005)		(0.002)		(0.004)		(0.012)
Observations	707,039	707,039	505,612	505,612	305,026	305,026	91,791	91,791
Firms	190,418	190,418	167,537	167,537	105,219	105,219	27,177	27,177
Clusters	163	163	163	163	139	139	70	70
R^2 (within firm)	0.114	0.114	0.013	0.013	0.048	0.048	0.029	0.029

The table shows OLS estimates of Equation (3.1). The dependent variable for each column is the annual growth rate (in logs) of the respective variable indicated in the header. All regressions control for firm fixed effects and country-year fixed effects. Standard errors clustered by industry-country cell are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05.

One might suspect that the estimations using the EFIGE measure suffer from an endogeneity issue due to reverse causality. If firms rated their industry lower in terms of EFD because they were hit harder by the crisis, this effect might bias our estimates of β downward. Even though we cannot fully rule out such a bias, we have three reasons to believe that it is not driving our results. First, the survey question is not concerned with the firm's own current circumstances, but targets general conditions in the industry.

[9] Empirical studies using linked firm-bank data have established a causal effect of the credit supply shock in the crisis on firm employment (Chodorow-Reich, 2013) and exports (Amiti and Weinstein, 2011).

Table 3.3: *Estimated interaction effects in robustness checks*

	Turnover		Employment		Labor productivity		Exports	
	EFIGE	Compustat	EFIGE	Compustat	EFIGE	Compustat	EFIGE	Compustat
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A. Measuring EFD by the median share of investments not financed internally (from EFIGE)								
Crisis × *EFD*	-0.045**		-0.010		-0.040**		-0.042*	
	(0.021)		(0.008)		(0.017)		(0.024)	
N	707,039		505,612		305,026		91,791	
R^2	0.114		0.013		0.048		0.029	
Panel B. Controlling for interaction terms of year dummies with industry characteristics								
Crisis × *EFD*	-0.074**	0.011**	-0.046*	0.000	-0.079**	0.007*	-0.058	-0.006
	(0.037)	(0.005)	(0.024)	(0.003)	(0.039)	(0.004)	(0.035)	(0.010)
N	707,039	707,039	505,612	505,612	305,026	305,026	91,791	91,791
R^2	0.115	0.115	0.013	0.013	0.049	0.049	0.030	0.030
Panel C. Including 2011-13 as a non-crisis period								
Crisis × *EFD*	-0.052**	0.010**	-0.011	0.005***	-0.042**	0.004	-0.070***	0.009
	(0.020)	(0.004)	(0.009)	(0.002)	(0.018)	(0.002)	(0.025)	(0.009)
N	1,168,899	1,168,899	886,490	886,490	541,994	541,994	144,475	144,475
R^2	0.068	0.068	0.009	0.009	0.034	0.034	0.018	0.018
Panel D. Excluding small countries and industries								
Crisis × *EFD*	-0.087*	0.011**	-0.046*	0.001	-0.082**	0.011***	-0.109**	0.003
	(0.047)	(0.005)	(0.026)	(0.002)	(0.040)	(0.004)	(0.046)	(0.013)
N	667,582	667,582	482,072	482,072	297,204	297,204	81,120	81,120
R^2	0.118	0.118	0.013	0.013	0.049	0.049	0.026	0.026
Panel E. Excluding potential outliers								
Crisis × *EFD*	-0.066**	0.009**	-0.029	0.001	-0.062*	0.011***	-0.078**	0.003
	(0.031)	(0.004)	(0.020)	(0.002)	(0.033)	(0.004)	(0.032)	(0.007)
N	692,636	692,636	494,696	494,696	298,926	298,926	89,939	89,939
R^2	0.230	0.230	0.026	0.025	0.085	0.085	0.043	0.043
Panel F. Excluding EFIGE firms from the pilot study								
Crisis × *EFD*	-0.061*		-0.037*		-0.052		-0.055**	
	(0.031)		(0.022)		(0.035)		(0.024)	
N	707,039		505,612		305,026		91,791	
R^2	0.114		0.013		0.048		0.029	

The table shows the results of estimating (variations of) Equation (3.1). The dependent variable for each column is the annual growth rate (in logs) of the respective variable indicated in the header. Odd columns report interaction effects for the EFIGE EFD measure, even columns report interaction effects for the Compustat EFD measure. All regressions control for firm fixed effects and country-year fixed effects. In panel A, we measure EFD by one minus the industry median response to the question "How were these investments in plants, machines, equipment and ICT financed on average in the last three years (2007-2009)? – Self-financing (use of internal sources)" (in percent/100) in EFIGE. In panel B, we add interaction terms of year dummies with the following industry characteristics: the logs of industry-level averages of capital intensity, the share of tangible assets, and operating revenues (as a proxy for firm size), as well as the Herfindahl index of operating revenues (all observed in 2005). In panel C, we extend the sample to the years 2011-13, in which we set the crisis dummy to zero. In panel D, we exclude the following small countries and industries for which the EFIGE EFD measure is based on less than ten firms: Austria, Hungary, "Tobacco" (NACE code 16), "Coke and refined petroleum products" (23), and "Office machinery and computers" (30). In panel E, we exclude potential outliers, defined as the observations with the highest and lowest 1% values of the respective dependent variable. In panel F, we exclude firms from the EFIGE pilot study when computing the EFIGE EFD measure. Standard errors clustered by industry-country cell are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

Second, for our results to be unbiased, we do not require that the reported EFD is entirely unaffected by the crisis. In particular, a uniform increase in the reported EFD of all firms in a given country would be absorbed by country-year fixed effects. Since the firms were surveyed simultaneously and since the crisis was highly synchronized across countries, as pointed out by Baldwin (2009) and confirmed in

industry-level data,[10] we would not expect the EFD ranking in 2010 to differ systematically from the fundamental ranking. Third, in an important robustness check, we construct an alternative EFD measure based on questions in the EFIGE survey, which inquire about how the firm has financed its investments over the years 2007-2009. We compute the share of investments not financed internally for the median firm by industry, reflecting the idea of RZ. This alternative (country-industry specific) EFD measure is based on the firm's financial accounts and hence not prone to subjective judgment. The regressions using this measure, reported in Panel A of Table 3.3, confirm the differential crisis effects found in our main analysis. This finding further suggests that the differences across EFD indexes documented in Table 3.1 are not merely driven by differences in measurement, but reflect inherent differences in EFD rankings across countries.

We conduct a series of additional robustness checks, which are detailed in Table 3.3. The pattern that we find in our main regressions is insensitive to (i) controlling for additional interaction terms of year dummies with industry characteristics (capital intensity, share of tangible assets, average firm size, and the Herfindahl index of turnover in 2005), (ii) including the non-crisis period 2011-13, (iii) considering only the countries and industries for which the EFIGE EFD measure is based on at least ten firms, (iv) excluding potential outliers with extreme growth rates (top and bottom 1% of our dependent variables), and (v) excluding firms from the EFIGE pilot study when computing the EFIGE index. In these robustness checks, the interaction term of the crisis dummy with the EFIGE index of EFD is always estimated to be negative and remains statistically significant with few exceptions, while the interaction effect is never negative and significant for the Compustat index.

3.5 Concluding Remarks

This paper has documented that industry rankings of a novel, survey-based index of external financial dependence (EFD) (i) are virtually uncorrelated with the standard index computed based on Compustat data for U.S. firms following Rajan and Zingales (1998), and (ii) differ substantially across seven European countries. These findings suggest that an industry which is highly financially dependent in one country may rank low on EFD in another country. Investigating the fundamental determinants of these international differences seems an interesting area for future research.

Our results indicate that it might not be adequate to apply an EFD index computed from U.S. data to other countries, as is currently standard practice in the literature. In a related paper, Ciccone and Papaioannou (2016) argue that this approach will cause a "benchmarking bias" if the U.S. index is a less noisy proxy (i.e., a better benchmark) for some countries than for others. In light of these insights, we suggest that future research on financial dependence should not rely exclusively on the U.S. index, but consider country-specific measures as complementary whenever possible.

Finally, we have contributed to the literature investigating the real effects of the global financial crisis by providing comparable firm-level evidence across seven European countries. Our investigation using the survey-based EFD index suggests that the credit channel significantly contributed to reducing real firm performance in the crisis. While this finding is in line with most of the existing literature, the relative importance of the credit shock compared to other adjustments in the crisis remains a controversial topic.

[10] EFIGE questionnaires were completed by all firms (except those from the pilot study) between January and May 2010 (Altomonte and Aquilante, 2012), while the majority of country-industry pairs in our sample experienced the steepest drop in monthly output between October 2008 and March 2009 (based on seasonally adjusted volume indices of production for 2-digit NACE industries from Eurostat).

Bibliography

Altomonte, Carlo and Tommaso Aquilante, "The EU-EFIGE/Bruegel-Unicredit Dataset," Technical Report 753, Bruegel 2012.

Amiti, Mary and David E. Weinstein, "Exports and Financial Shocks," *The Quarterly Journal of Economics*, 2011, *126* (4), 1841–1877.

Baldwin, Richard, "Introduction: The Great Trade Collapse: What Caused It and What Does It Mean?," in Richard Baldwin, ed., *The Great Trade Collapse: Causes, Consequences and Prospects*, 2009.

Chodorow-Reich, Gabriel, "The Employment Effects of Credit Market Disruptions: Firm-level Evidence from the 2008-9 Financial Crisis," *The Quarterly Journal of Economics*, 2013, *129* (1), 1–59.

Chor, Davin and Kalina Manova, "Off the Cliff and Back? Credit Conditions and International Trade during the Global Financial Crisis," *Journal of International Economics*, 2012, *87* (1), 117–133.

Ciccone, Antonio and Elias Papaioannou, "Estimating Cross-industry Cross-country Interaction Models Using Benchmark Industry Characteristics," Technical Report 22368, National Bureau of Economic Research 2016.

Cihak, Martin, Asli Demirgüç-Kunt, Erik Feyen, and Ross Levine, "Benchmarking financial systems around the world," Policy Research Working Paper 6175, The World Bank 2012.

Dell'Ariccia, Giovanni, Enrica Detragiache, and Raghuram Rajan, "The real effect of banking crises," *Journal of Financial Intermediation*, 2008, *17* (1), 89–112.

Ferrando, Annalisa, Petra Köhler-Ulbrich, and Rozália Pál, "Is the Growth of Euro Area Small and Medium-sized Enterprises Constrained by Financing Barriers?," Industrial Policy and Economic Reforms Paper 6, Enterprise and Industry Directorate-General European Commission 2008.

Kroszner, Randall S., Luc Laeven, and Daniela Klingebiel, "Banking Crises, Financial Dependence, and Growth," *Journal of Financial Economics*, 2007, *84* (1), 187–228.

Manova, Kalina, "Credit Constraints, Heterogeneous Firms, and International Trade," *Review of Economic Studies*, 2013, *80* (2), 711–744.

Rajan, Raghuram G. and Luigi Zingales, "Financial Dependence and Growth," *American Economic Review*, 1998, *88* (3), 559–86.

Schumpeter, Joseph A., *A Theory of Economic Development*, Harvard University Press, Cambridge, MA, 1911.

3.A Data Appendix

We closely follow RZ when calculating their index of external financial dependence (EFD). We use the North America Segment of the Compustat database to calculate the EFD measure over the time horizon 1990-2005.[11] We only keep entries with cash flow statements (SCF) codes 1, 2, 3, and 7, and those with Industrial Format (INDFT = INDL). Furthermore, we drop all firms for which the country of incorporation (FIC) is not the U.S. or for which the financial year (FYEAR) is missing.

As RZ, we construct EFD as the share of investment that cannot be financed through internal cash flows, i.e., capital expenditures minus cashflow from operations divided by capital expenditures, for the median firm by industry. Capital expenditures (CAPX) are readily available in Compustat. Following RZ, we define cash flow from operations as the sum of funds from operations and changes in working capital. For SCF codes 1, 2, and 3, we construct the cash-flow measure as the sum of total funds from operations (FOPT) plus increases in accounts payable ($\text{AP}_t - \text{AP}_{t-1}$), decreases in inventories ($\text{INVT}_{t-1} - \text{INVT}_t$), and decreases in receivables ($\text{RECT}_{t-1} - \text{RECT}_t$) for each financial year t. For SCF code 7, total funds from operations is not available and therefore calculated as the sum of income before extraordinary items (IBC), depreciation and amortization (DPC), deferred taxes (TXDC), equity in net loss/earnings (ESUBC), sale of property, plant and equipment and investments (SPPIV), and other funds from operations (FOPO). We use the levels of the working capital variables and calculate the changes manually instead of using the reported changes provided by Compustat. This is because there are fewer missing values for the levels than for the changes in the dataset. Furthermore, we treat cash flow from operations as missing if any of its components is missing.

We calculate capital expenditures minus cash flow from operations for each firm over the period 1990-2005 and divide by the sum of capital expenditures over the respective time period, provided that both variables are non-missing. We then assign to each firm the 2-digit or 3-digit NACE Rev. 1.1 code (as reported in Table 3.A.1) corresponding to its SIC code.[12] Finally, we use the median value by industry as our Compustat EFD index.

[11] Note that the current data items in Compustat that we use are labeled differently from the ones used in RZ. However, Compustat provides conversion tables between the old and new variable names.

[12] Since official correspondence tables between these classifications are unavailable to the best of our knowledge, and since chains of correspondences involve numerous ambiguities, we match industry codes manually. The resulting correspondence is available on request.

Table 3.A.1: *EFD rankings across countries*

NACE Rev 1.1 industry		U.S. (Compustat)			AUT (EFIGE)			FRA (EFIGE)			DEU (EFIGE)			ITA (EFIGE)			ESP (EFIGE)			HUN (EFIGE)			GBR (EFIGE)		
code	name	EFD	rank	firms	EFD	rank	firms	EFD	rank	firms	EFD	rank	firms	EFD	rank	firms	EFD	rank	firms	EFD	rank	firms	EFD	rank	firms
16	Tobacco	-3.4462	1	13							4.0000	25	3				3.6432	24	4						
19	Leather and footwear	-1.3422	2	27				2.3369	5	32	2.6854	2	13	2.3101	10	115	2.8816	1	47	2.9467	6	4	3.7160	24	10
361	Furniture	-0.5680	3	66							3.1252	17	67	2.5925	21	141	3.1542	8	206	3.3559	15	14	3.0595	6	90
22	Publishing and printing	-0.4268	4	166	3.3747	14	34	3.0589	23	148	3.1547	19	215	2.4744	19	105	3.3323	18	100	2.4970	2	27	3.2662	17	208
28	Fabricated metal products	-0.3272	5	173	3.2828	13	70	2.6043	17	839	3.1429	18	510	2.3607	13	611	3.2016	10	580	3.1897	11	101	3.1667	12	301
35	Other transport equipment	-0.3057	6	92	5.0000	20	2	2.6581	19	16	3.4073	23	20	2.2043	2	33	3.6122	23	42	4.3720	22	3	2.9638	5	21
150	Food (excl. beverages)	-0.1454	7	197	3.0028	5	28	2.4341	10	194	3.0753	14	285	2.4027	15	200	3.0976	6	402	3.2721	13	52	3.2786	19	137
21	Pulp, paper and paper products	-0.1343	8	89	3.1800	8	10	2.4652	11	83	2.9631	10	62	2.3601	12	71	2.8854	2	27	3.1055	9	16	3.1883	14	47
23	Coke and refined petroleum products	-0.1114	9	63				2.3333	4	3	2.5695	1	4	2.1253	1	8							2.6991	1	6
26	Non-metallic mineral products	-0.0884	10	74	3.4615	17	18	2.3615	6	153	2.8727	7	94	2.3268	11	167	3.5672	22	163	3.5639	18	30	3.1494	11	56
20	Wood products, except furniture	-0.0627	11	62	3.2144	11	21	2.6703	20	93	3.4386	24	103	2.5789	20	88	3.2587	15	212	3.4507	16	17	3.2313	15	89
17	Textiles	-0.0427	12	75	3.8595	19	8	2.5975	16	118	2.8296	4	77	2.2046	3	196	3.4058	21	46	3.0567	8	7	3.1762	13	52
240	Chemicals (excl. pharamaceuticals)	0.0047	13	300	2.5833	2	5	2.4253	8	79	2.8719	6	76	2.2595	7	88	3.0600	5	104	3.2483	12	18	3.6003	23	94
34	Motor vehicles	0.0759	14	140	2.8214	4	6	2.4947	12	73	3.1681	20	41	2.2219	6	47	3.2527	14	64	3.2851	14	11	3.1297	10	33
27	Basic metals	0.0870	15	144	3.0087	6	13	2.9574	22	68	3.2563	21	58	2.4630	18	76	3.2279	12	68	4.2548	21	7	3.2682	18	54
18	Wearing apparel and fur	0.1021	16	124	3.6062	18	5	2.6242	18	55	2.7240	3	17	2.2743	8	109	3.3879	20	50	3.1564	10	17	3.3609	20	42
25	Rubber and plastic products	0.1205	17	150	3.2621	12	22	2.5890	14	226	2.9824	11	192	2.3638	14	167	3.2049	11	148	3.4907	17	40	3.2613	16	122
29	Machinery and equipment	0.1255	18	397	3.2055	10	48	2.4290	9	249	3.0743	13	503	2.4175	16	381	3.1288	7	305	3.0420	7	68	3.0931	8	208
31	Electrical machinery and apparatus	0.3269	19	615	3.4449	16	13	2.2367	3	113	2.9376	8	106	2.2066	4	143	2.9768	3	60	2.4296	1	18	3.0822	7	116
360	Other manufacturing (excl. furniture)	0.3719	20	159	2.1752	1	5	2.1247	2	16	2.9420	9	105	2.6424	23	70	3.0394	4	52	3.8399	20	4	3.1111	9	168
159	Beverages	0.3992	21	74	2.7172	3	4	2.5896	15	18	2.8582	5	62	2.6245	22	38	3.3667	19	57	3.7313	19	10	3.3663	21	10
30	Office machinery and computers	0.6565	22	365	3.1667	7	7	1.8791	1	8	3.3431	22	28	2.6667	24	9	3.2489	13	6	5.0000	23	1	3.3976	22	8
33	Medical/ precision/ optical instruments	1.0336	23	818	3.3855	15	15	2.3879	7	58	3.0637	12	192	2.2967	9	71	3.2590	16	25	2.7671	4	6	2.7771	2	80
32	Radio/ TV/ communication equipment	1.1559	24	296	3.1834	9	5	2.5131	13	94	3.1101	15	56	2.4211	17	49	3.1744	9	25	2.9181	5	9	2.7901	3	101
244	Pharmaceuticals	8.6029	25	731				2.7505	21	23	3.1121	16	19	2.2075	5	20	3.2990	17	17	2.5000	3	2	2.9621	4	10

The table shows rankings of industries in terms of external financial dependence (EFD) by country. The first three columns report the EFD measure for U.S. firms computed from Compustat data over the period 1990-2005, following RZ. The remaining columns report averages of EFD by industry reported in the EFIGE survey in 2010.

Chapter 4

Firm Exports, Foreign Ownership, and the Global Financial Crisis*

The global financial crisis of 2008/09 caused a sharp decline in production, and an even greater collapse in international trade. This paper demonstrates that foreign ownership substantially reduced the adverse effect of the crisis on firm exports in Spain. We provide theory and firm-level evidence suggesting that internal capital markets within multinational firms can explain this pattern. In our model, heterogeneous exporting firms face credit constraints, which can be alleviated by foreign ownership. The model predicts that a credit crunch hits the exports of foreign-owned firms less severely, and this difference is magnified by financial vulnerability. We test these predictions in a representative panel of Spanish manufacturing firms over the years 2005-2012. We examine firm exports relative to sales in a difference-in-differences model and control for unobservables via firm fixed effects and propensity score reweighting. To pin down the credit channel in a triple difference model, we further distinguish firms by their debt service-to-sales ratio at the onset of the crisis. We find that foreign-owned firms significantly increased their competitive advantage on export markets in the crisis, in particular among highly indebted firms. This pattern is robust to controlling for asymmetric demand shocks and various other benefits of foreign ownership. Our findings highlight important firm-level complementarities between international trade and foreign direct investment.

*This chapter is based on joint work with Marcel Smolka.

4.1 Introduction

Foreign-owned firms are much more actively engaged in exporting than other firms. In Spain, for instance, the share of exports in sales is more than twice as large for foreign-owned compared to domestically owned firms. What can explain this pattern? Manova et al. (2015) show that the export strength of foreign-owned firms in China is greater in financially dependent industries and propose that internal capital markets within multinational firms play an important role. Since exports are known to depend more heavily on financing than domestic sales (Minetti and Zhu, 2011), and since foreign-owned firms can alleviate credit constraints by accessing internal capital markets (Desai et al., 2004; Egger et al., 2014), this financial advantage of foreign ownership may be responsible for the observed pattern. However, the credit channel is hard to disentangle from other benefits of foreign ownership, which may include the transfer of superior management practices and technologies, or market access through the foreign parent's distribution network.

In this paper, we exploit the global financial crisis of 2008/09 as a major shock to credit conditions in order to identify how important the financial advantage of foreign ownership is for firm exports. The financial crisis caused the sharpest decline in real economic activity around the globe since the end of World War II. International activities were even more severely affected by this shock than domestic production and sales; a fact that has become known as the 'great trade collapse' (Baldwin, 2009). Several studies have established that the credit crunch played an important role in the trade collapse (see e.g. Chor and Manova, 2012; Paravisini et al., 2014). Our identification strategy combines this credit shock with information on ownership and the debt situation of Spanish firms at the onset of the crisis. We uncover a strong pattern: Foreign-owned firms significantly magnified their competitive advantage on export markets in the crisis, and this effect was concentrated in small and financially vulnerable firms. The evidence suggests that multinational firms indeed provide their affiliates with a substantial financial advantage, which helped to stabilize firm exports in the crisis.

We formalize this argument in a simple theoretical model of trade finance and foreign ownership to guide our empirical analysis. In the model, heterogeneous firms have limited funds to finance the costs of exporting, so they need to obtain credit in an imperfect capital market (as in Manova, 2013). Into this framework, we introduce the possibility for foreign-owned firms to access foreign capital markets through their multinational parent. The internal capital market is not frictionless, but it may offer lower interest rates. The financial crisis can be understood in terms of the model either as a deterioration of the credit market imperfections or as a shock to the interest rate at which all firms can borrow externally. Both of these scenarios cause a drop in exports relative to domestic sales among credit constrained exporters. Importantly, this effect is mitigated in foreign-owned firms because they can rely on the internal capital market. The model further predicts that this differential crisis impact should be more pronounced among financially vulnerable firms, which need to finance a larger share of their exporting costs and hence benefit even more from foreign ownership in the crisis.

We empirically investigate these predictions using a rich, high-quality panel data set, which covers a representative sample of manufacturing firms in Spain over the years 2005-2012. Crucially for the purpose of this paper, the data contain annual information on firms' foreign ownership status, export activities, and debt structure. In the first step of our empirical analysis, we test the impact of the financial crisis on the exports of foreign-owned compared to domestically owned firms in a difference-in-differences model. To identify the crisis impact on exports beyond its effect on production, we focus

on the export share, defined as exports over total sales, as our main outcome variable. Throughout the analysis, we control for firm-specific fixed effects and for arbitrary industry-specific shocks over time. To account for observable and unobservable firm characteristics that may determine foreign ownership and exports, we further augment the model by a propensity score reweighting approach.

The estimations show that, as the credit crunch hit Spain in 2009, foreign-owned firms significantly increased their export shares compared to domestically owned firms. This difference prevailed over the subsequent crisis years, while we cannot reject a common trend in the export shares of both groups over the pre-crisis years. Interestingly, the differential effect is concentrated in small firms, which are more likely to be credit constrained (see, e.g. Gertler and Gilchrist, 1994), and absent in large firms. These results constitute indicative evidence of the credit channel described by our model. However, it is conceivable that, in addition to internal capital markets, other benefits of foreign ownership also gained importance in the crisis for small firms. We further exploit our exceptionally rich firm data set to pin down the role of credit constraints in two steps: We first provide direct evidence on the credit channel, and then explicitly account for and examine other benefits of foreign ownership.

Our main finding is that financial frictions have played a decisive role for the superior export performance of foreign-owned firms in the crisis. Recall from our theoretical model that, if the credit channel is important, the effect of foreign ownership on exports in the crisis should be particularly pronounced among financially vulnerable firms. We test this prediction by estimating a triple difference model. In addition to the comparison before vs. after the credit crunch and between foreign vs. domestic ownership, we add as a third margin the firm's financial vulnerability. To obtain a firm-specific measure of financial vulnerability, we exploit detailed information in our data on the firms' debt structure (credit volumes and interest rates on short-term and long-term debt). Arguably, firms that had a particularly high debt service-to-sales ratio at the onset of the crisis in 2008 were more likely to face binding credit constraints in the subsequent years. Our estimates confirm that among these financially vulnerable firms, the positive effect of foreign ownership ownership on the export shares was substantially greater in the crisis. Since our measure of financial vulnerability is unlikely to be correlated with other confounding factors in the full-fledged and reweighted estimation model, we interpret these findings as strong evidence for the fact that foreign ownership promotes exports through the credit channel.

Scrutinizing our main finding, we distinguish the effect of foreign ownership in the crisis on export market entry and exit, the extensive margin of exports, and on changes in the volume of export sales, the intensive margin. We find no significant difference in the crisis impact on the extensive margin of exports across firms with a different ownership status or degree of financial vulnerability. Instead, the effects are mainly driven by the intensive margin. This result is in line with micro evidence from several countries showing that the global financial crisis reduced firm exports predominantly at the intensive margin (see Behrens et al., 2013; Bricongne et al., 2012; Eppinger et al., forthcoming; Paravisini et al., 2014). This is the reason for the focus of our theoretical and empirical analysis on the intensity of exporting. In several robustness checks, we then show that our main result is insensitive to various modifications of the estimation approach, it is confirmed in fractional probit estimations, and it proves robust to controlling for world-region specific demand shocks. All of these analyses strengthen our conclusion that foreign ownership promoted the exports of financially vulnerable firms in the crisis.

In the last step of our empirical analysis, we explicitly consider other channels through which foreign ownership can boost export performance, but which are not directly related to credit constraints. It is a unique feature of our data that we can approximate or directly observe several important benefits

that foreign ownership may bring to the firm, namely innovative management practices, a diversified export portfolio, or access to the export market via the parent's distribution network. This allows us to investigate the role of these non-financial benefits of foreign-ownership and how they interact with the credit channel. We find that a more innovative management and higher export market diversification were indeed associated with a higher export share in the crisis, but their effect does not systematically vary by financial vulnerability, and our main finding is upheld after controlling for these other channels.

More interestingly, we find direct evidence for the fact that an increasing number of foreign-owned firms relied on their parent's distribution network for exporting in the crisis. Since this decision was associated with a significantly higher export share, distribution networks constitute another important channel through which foreign ownership promoted firm exports in the crisis on top of the financial advantage. Furthermore, our estimation results suggest that these two benefits of belonging to a multinational firm are intertwined. More precisely, the financial advantage matters less in foreign-owned firms that exported via their parent's distribution network before the crisis. It seems plausible that this way of accessing export markets reduces the firm's cost of exporting that require financing, and hence, internal credit markets are less important for these firms. In line with this idea, we find that starting to use the parents distribution network was particularly conducive to exporting among financially vulnerable firms. Essentially, this constituted an alternative way of circumventing credit constraints in the crisis.

Our paper contributes to two strands of the international economics literature that recently emerged at the intersection of corporate finance and international trade and investment (see Foley and Manova, 2015, for an overview). The first is the literature on international trade and financial frictions. This literature has argued that export sales are inherently more dependent on external finance than domestic sales due to additional costs of exporting, longer shipping times, and greater risk involved in international transactions.[1] Therefore, financial development can affect the pattern of comparative advantage.[2] To study the impact of financial frictions across heterogeneous exporting firms, theoretical work by Chaney (2016), Feenstra et al. (2014), and Manova (2013) has extended the Melitz (2003) model with capital market imperfections due to asymmetric information or incomplete credit contracts. These models predict that financial frictions reduce exports via intensified selection of firms into production and exporting, and they may potentially also lower the level of firm exports. Several empirical studies have confirmed the patterns predicted by these theories, most prominently Manova (2013).[3]

A focus of this literature has been on financial crises in general and the global financial crisis in particular. The question of how much financial frictions have contributed to the great trade collapse remains controversial.[4] Auboin (2009) and Ahn et al. (2011) argue that trade finance played an important role and Chor and Manova (2012) provide evidence supporting this notion from monthly US import data by country and industry. By contrast, Levchenko et al. (2009) find no evidence of the credit channel in the US. Further studies at the micro level by Behrens et al. (2013, for Belgium) and Bricongne et al. (2012,

[1] Manova (2013) models the higher costs, Ahn (2011) models the risk, and Feenstra et al. (2014) model the longer shipping times of exports. Antràs and Foley (2015) and Schmidt-Eisenlohr (2013) explicitly study different trade finance regimes.

[2] This hypothesis was first proposed by Kletzer and Bardhan (1987), confirmed empirically at the country level by Beck (2002), and further refined by Egger and Keuschnigg (2009).

[3] These studies conclude that improved financial market conditions due to financial development (Beck, 2002; Berman and Héricourt, 2010) or equity market liberalizations (Manova, 2008) boost exports. As an exception, Greenaway et al. (2007) do not find that financially healthy firms are more likely to start exporting, but show that conversely, exporting improves financial health.

[4] Most empirical studies attribute a large part of the trade collapse to changes in the volume and composition of demand (Bems et al., 2013; Eaton et al., 2016; Levchenko et al., 2009). Further explanations include inventory adjustments (Alessandria et al., 2010) and a rise in protectionism (Baldwin and Evenett, 2009).

for France) show that financial frictions played a relevant but subordinate role for the aggregate decline. Recent contributions by Amiti and Weinstein (2011, for Japan) and Paravisini et al. (2014, for Peru) have made further progress towards identifying a causal effect of financial crises on trade by linking micro-level trade data with bank data, and they confirm that financial frictions impede trade.[5] Our findings contribute to this evidence and suggest that the tightening of credit constraints did affect Spanish firms' exports in the crisis. However, since this effect was concentrated in small and domestically owned firms, which account only for a small share of total trade, it seems likely that the contribution of the credit channel to the aggregate trade collapse was rather small in Spain.

The second strand of the literature closely related to our work focuses on multinational firms and corporate finance. We build on the work by Desai et al. (2004), who find that foreign affiliates of US multinational firms access internal capital markets to circumvent financial frictions in external capital markets. In line with this pattern, foreign affiliates have been shown to intensify their activities in currency crises and outperform domestically owned firms in terms of sales growth (Desai et al., 2008). The recent crisis is investigated by Alfaro and Chen (2012), who apply matching techniques to a worldwide panel of establishments to show that foreign-owned plants fared significantly better in terms of sales growth at the time of the credit crunch. This effect is larger in sectors with stronger financial linkages or higher financial dependence and it is very weak in the pre-crisis years. Garicano and Steinwender (2016) investigate investment behavior using the same dataset as in this study and show that *only* domestically owned firms in Spain decreased their investments during the crisis, in particular long-term ones. All of these findings support the view that foreign ownership improves the crisis resilience of firms' real activities by alleviating financial constraints.[6] We complement this literature with a detailed analysis of multinational firms' role in mitigating the impact of the global financial crisis on firm exports, a dimension of firm performance that is known to depend in particular on financial conditions.

Most intimately linked to our work is the recent contribution by Manova et al. (2015) connecting these two strands of the literature. Manova et al. (2015) find that foreign-owned firms in China account for disproportionate shares of exports in industries that are characterized by greater financial vulnerability. This specialization pattern is in line with the idea that internal capital markets provide foreign-owned firms with a comparative advantage in financially vulnerable industries.[7] While Manova et al. (2015) identify the effect of credit constraints from variation in financial vulnerability across industries within multi-product firms, our identification strategy exploits the shock to credit market conditions in the crisis to identify the effect within firms over time.[8] We show that foreign ownership promotes exports through access to finance not only in China, but also in an advanced economy like Spain. In addition, since we have data on firms' domestic sales, we can identify the effect of foreign ownership and the crisis on exports *relative* to domestic activity.[9] Our analysis also sheds new light on the mechanisms behind the financial advantage of foreign ownership: In addition to relying on internal capital markets, financially vulnerable firms increasingly used their parent's distribution network in the crisis, which constitutes an

[5] In a similar vein, Muûls (2015) exploits Belgian data on credit ratings.

[6] As a notable exception, Alvarez and Görg (2007) find no stabilizing effect of foreign ownership for the Chilean crisis in the late 1990s.

[7] This is an endogenous outcome in Antràs et al. (2009), who show how financial frictions shapes the pattern of foreign direct investment (FDI) in theory and in firm-level data.

[8] In related work, Wang and Wang (2015) exploit time variation in foreign ownership, by comparing foreign acquisitions to domestic acquisitions in China, to show that foreign ownership improved firm measures of financial health and export performance, in line with our main finding.

[9] Hence, our findings are also in line with the pattern found by Minetti and Zhu (2011) for a cross-section of Italian firms: Exports depend more heavily on external finance than domestic sales.

alternative way to reduce trade costs and the need for external funds.

What are the policy implications of our work? The fact that foreign ownership can alleviate credit constraints points to important complementarities between FDI and international trade at the firm level. Our results suggest that in the presence of financial frictions, multinational firms can bring important benefits to their affiliates. The substantial effects of foreign ownership identified in Spain might even be magnified in emerging economies, where financial development is weak. Inward FDI may serve as a vehicle for these countries to alleviate the local firms' credit constraints and promote export-led growth. Finally, the stabilizing effects of foreign ownership on exports in the global financial crisis gives rise to the hope that increasing multinational firm activity may also moderate the impact of future financial crises on international trade.

The remainder of this paper is organized as follows. Section 6.2 develops our theoretical model of trade finance and foreign ownership and derives predictions regarding the impact of the financial crisis on firm exports. Section 4.3 presents the panel data on Spanish manufacturing firms used to test these predictions and provides some descriptive evidence. In Section 4.4, we describe our empirical strategy and our estimation results. Section 4.5 offers some conclusions.

4.2 A Simple Theory of Trade Finance and Foreign Ownership

In this section, we present a partial equilibrium model of exporting under credit constraints to motivate and guide the empirical analysis. Our theoretical framework adapts the trade model with heterogeneous firms and credit constraints developed by Manova (2013), to which we add the possibility for foreign-owned firms to access foreign capital markets through their parent.

4.2.1 General setup

Consider a world of two countries.[10] The preferences of the representative consumer in each country constitute a Cobb-Douglas aggregate over industries j, each of which consists of a set of varieties indexed by ω that are substitutable with constant elasticity $\sigma > 1$:

$$U = \prod_j \left[\int_{\omega \in \Omega_j} q_j(\omega)^{\frac{\sigma-1}{\sigma}} \mathrm{d}\omega \right]^{\theta_j \frac{\sigma}{\sigma-1}},$$

where the set of varieties is denoted by Ω_j, and $\theta_j \in (0,1)$ represents the consumer's share of expenditure on each industry. Demand for each variety follows as

$$q_j(\omega) = p_j(\omega)^{-\sigma} \theta_j Y P_j^{\sigma-1}, \tag{4.1}$$

with aggregate income Y, price $p_j(\omega)$, and CES price index $P_j = \left[\int_{\omega \in \Omega_j} p_j(\omega)^{1-\sigma} \mathrm{d}\omega \right]^{\frac{1}{1-\sigma}}$, which firms take as given.

Each industry is characterized by a continuum of heterogeneous firms that compete monopolistically. As in Melitz (2003), firms are identical ex ante and may choose to pay a sunk cost $b_j f_E$ in order to enter the market. Upon entry, each firm draws its productivity $1/a$ from a known distribution $G(a)$ with

[10] In anticipation of our empirical analysis, these countries may be thought of as Spain and the rest of the world. We neglect the country index where it is not essential to keep the notation simple.

support $[\underline{a}, \overline{a}]$, $\overline{a} > \underline{a} > 0$. If $1/a$ turns out to be high enough to break even, the firm will start producing; otherwise it exits immediately. Since a is specific to the firm and each firms produces a distinct variety, we will henceforth use a to index product ω. The cost-minimizing input bundle is denoted by b_j, resulting in variable production costs $b_j a$ per unit of output.

To service a specific market, the firm has to incur fixed market-access costs and variable trade costs. More precisely, fixed costs are $b_j f_X$ for exporting and $b_j f_H$ for servicing the home market. We make the standard assumption that the fixed costs of market access, marketing, and distribution are higher in the export market: $f_X > f_H > 0$. Variable costs of exporting take the usual iceberg form, such that $\tau > 1$ units of the good need to be shipped in order for one unit to arrive in the other country. These assumptions imply the standard Melitz-type sorting pattern, according to which only the most productive firms can export profitably.

4.2.2 Trade finance

As in Manova (2013), we relax the assumption of perfect capital markets that is implicit in Melitz (2003). We assume that firms possess only a limited amount of liquidity, which they use to finance domestic operations and a share $1 - d(a)$ of exporting costs.[11] The remaining share $d(a) \in (0, 1)$ of all costs associated with exporting (production costs as well as fixed and variable trade costs) need to be financed by an investor. The investor will be repaid with exogenous probability $\lambda \in (0, 1)$ due to frictions in the capital market.[12]

We first consider a domestically owned firm, which needs to borrow the required funds from a domestic bank since it has no other financing option by assumption. We further assume that there is a large number of domestic banks in perfect competition and that all parties are risk neutral. As a result, the firm maximizes its expected profits from exporting:

$$\Pi_X(a) = p_j(a)q_j(a) - [1 - d(a)] [q_j(a)\tau b_j a + f_X b_j] - \lambda R(a) \tag{4.2}$$

$$\text{subject to} \quad R(a) \leq p_j(a)q_j(a) - [1 - d(a)] [q_j(a)\tau b_j a + f_X b_j], \tag{FC}$$

$$\text{and} \quad \lambda R(a) \geq (1 + r_D) [d(a) (q_j(a)\tau b_j a + f_X b_j)], \tag{PC-D}$$

and subject to demand from equation (4.1).[13] The financial constraint (FC), states that the firm cannot repay more than its total export revenue. The participation constraint of the domestic investor (PC-D), states that her net return (expected repayment minus credit) must exceed her outside option. The investor's outside option reflects the returns from investing the amount of credit (the term in brackets) into an alternative project at the real interest rate $r_D > 0$ that prevails in the perfectly-competitive domestic capital market.

Since the capital market is perfectly competitive, investors are paid their outside option and equation (PC-D) holds with equality in equilibrium. We can plug this condition and demand from equation (4.1) into equation (4.2). Solving this maximization problem for highly productive firms, which face

[11] As the notation suggests, we allow internal funds to differ potentially across firms.

[12] This notion of credit constraints is agnostic about the underlying source of the financial friction, but simply invokes that credit will not be repaid with certainty if capital markets function imperfectly.

[13] This is the program formulated by Manova (2013, Web Appendix, equation 2), in which firms need to finance part of *both* fixed and variable trade costs, with two adaptations. First, we abstract from collateral since it plays no role in our empirical analysis. The model easily can be extended to allow for a non-zero share of fixed entry costs as collateral, which does not alter our main theoretical predictions. Second, we allow for a non-zero outside option of the investor, such that the interest rate can play a role in the model.

no binding financial constraint, yields the optimal prices $p_j^*(a)$ and quantities $q_j^*(a)$:

$$p_j^*(a) = \frac{\sigma}{\sigma-1}\left[1 - d(a) + (1+r_D)d(a)\right]\tau b_j a \quad \text{and} \tag{4.3}$$

$$q_j^*(a) = \left(\frac{\sigma}{\sigma-1}\left[1 - d(a) + (1+r_D)d(a)\right]\tau b_j a\right)^{-\sigma}\theta_j Y P_j^{\sigma-1}. \tag{4.4}$$

The productivity threshold $(1/a_H^D)$ below which domestically owned firms cannot obtain external finance to export at first-best levels is obtained by plugging the optimal price and quantity along with the binding participation constraint into the binding financial constraint:

$$\frac{1}{a_H^D} = \left[\frac{\left[1 - d(a) + (1+r_D)d(a)/\lambda\right]\left(\frac{\sigma}{\sigma-1}\left[1 + r_D d(a)\right]\right)^{\sigma}}{\frac{\sigma}{\sigma-1}\left[1 + r_D d(a)\right] - \left[1 - d(a)\right] - (1+r_D)d(a)/\lambda}\right]^{\frac{1}{\sigma-1}}\left(\frac{f_X b_j^{\sigma}}{\theta_j Y}\right)^{\frac{1}{\sigma-1}}\frac{\tau}{P_j}. \tag{4.5}$$

Only firms above this threshold can export at first-best levels.

Firms with productivity levels just below this threshold will also export, but a only a smaller quantity (at a higher price) than in the first-best case because they are credit constrained. These firms reduce their export scale in order to lower the repayment required by the investor. The optimal prices $p_j^{**}(a)$ for these constrained exporters are determined by the binding financial constraint. They can be obtained by plugging equation (4.1) and equation (PC-D) into equation (FC), all holding with equality:

$$p_j^{**}(a)^{1-\sigma} - p_j^{**}(a)^{-\sigma}\tau b_j a\left[1 - d(a) + (1+r_D)d(a)/\lambda\right] = \frac{f_X b_j\left[1 - d(a) + (1+r_D)d(a)/\lambda\right]}{\theta_j Y P_j^{\sigma-1}} \tag{4.6}$$

This equation implicitly determines $p_j^{**}(a)$. Appendix 4.A.1 shows that the left-hand side of equation (4.6) is increasing in the optimal price, while the right-hand side is constant. Hence, equation (4.6) implicitly defines the optimal prices for constrained exporters, which is negatively related to the quantity sold (via the demand schedule in equation (4.1)) and the associated export revenues.

We can further determine the threshold $1/a_L^D$, below which domestically owned firms cannot export at all. The highest price that constrained exporters may want to set is

$$p_L(a) = \frac{\sigma}{\sigma-1}\tau b_j a\left[1 - d(a) + (1+r_D)d(a)/\lambda\right], \tag{4.7}$$

which maximizes the left-hand side of equation (4.6). Firms with productivity levels below $1/a_L^D$ cannot compensate the investor even if they set this price and offer all revenues as a repayment to the investor. Formally, we obtain $1/a_L^D$ by plugging $p_L(a)$ from equation (4.7) back into equation (4.6):

$$\frac{1}{a_L^D} = \frac{\sigma}{\sigma-1}\left[1 - d(a) + (1+r_D)d(a)/\lambda\right]^{\frac{\sigma}{\sigma-1}}\left(\frac{\sigma f_X b_j^{\sigma}}{\theta_j Y}\right)^{\frac{1}{\sigma-1}}\frac{\tau}{P_j} \tag{4.8}$$

Our assumptions imply the same sorting pattern for domestically owned firms as in Manova (2013), which is illustrated (above the line) in Figure 4.1. While the least productive firms below the entry cutoff $1/a_E$ exit the market immediately, those with productivity levels $1/a \in [1/a_E, 1/a_L^D)$ produce, but serve only the domestic market. Constrained exporters with productivity levels $1/a \in [1/a_L^D, 1/a_H^D)$ sell both to the domestic and to the export market, but they export lower quantities at higher prices than in the first-best case. Only the most productive firms export at first-best levels.

4.2.3 Foreign ownership and internal capital markets

We proceed by considering foreign-owned firms in the next step. We model foreign ownership as facilitating access to the foreign capital market. While belonging to a multinational firm may have benefit exports through several additional channels, our theory focuses exclusively on this financial advantage.[14] To be precise, we assume that accessing the foreign capital market is costless for foreign-owned firms via their parent firm. This assumption is motivated by ample evidence demonstrating how multinational firms use internal capital markets to finance their subsidiaries' activities (see e.g. Desai et al., 2004; Egger et al., 2014). For domestically owned firms, in contrast, we assume that the cost of finding and contracting a suitable lender abroad is prohibitively high, so they cannot access the foreign capital market at all. Note that this assumption is made for simplicity and stronger than necessary to arrive at our main theoretical predictions.

To keep matters simple, all differences between the domestic and foreign capital market are summarized by the real interest rate. We denote the real interest rate on the foreign (or world) capital market by r_F. In the context of this model, it can be interpreted as the lowest interest rate at which the multinational parent firm can borrow abroad and lend to its affiliate.

The maximization problem of the foreign-owned firm is identical to that of the domestically owned firm, but it includes one additional constraint, namely the participation constraint of the foreign investor:

$$\lambda R(a) \geq (1 + r_F) \left[d(a) \left(q_j(a) \tau b_j a + f_X b_j \right) \right]. \tag{PC-F}$$

Also, the foreign-owned firm makes one more decision: it chooses between domestic and foreign financing. Since equation (PC-F) is formulated in analogy to equation (PC-D), this additional option may provide a financial advantage to the foreign-owned firm. While equation (FC) need not be binding for highly productive firms, either equation (PC-D) or equation (PC-F) will always be binding in equilibrium because any firm needs to raise finance for part of its exporting costs, and both the domestic and foreign capital market are perfectly competitive.[15]

It is obvious from comparing the two participation constraints (PC-D) and (PC-F) that foreign-owned firms will choose foreign financing if and only if $r_F < r_D$. If this inequality is reversed, they will opt for domestic financing.[16] Note that both cases are possible in principle. We may expect $r_F < r_D$ as long as the multinational parent firm can obtain cheaper funds in some other country and extend a load at this reduce rate to its affiliate. This case seems plausible due to the greater liquidity of the world capital market and the size of multinational firms, which allows them to raise funds directly in the credit market by issuing corporate bonds and to diversify financial activities across countries. We will focus below on this interesting case in which foreign-owned firms enjoy a financial advantage. However, a situation where $r_D \leq r_F$ may arise in case the MNE itself is liquidity constrained and the outside option of investing the available funds in another affiliate promises higher returns than r_D. Note that even in this case, foreign-owned firms will never be worse off because they can always choose to borrow from

[14] This view of foreign ownership and internal capital markets is admittedly simplistic. For a formal treatment of the choice between internal and external capital markets, see Gertner et al. (1994). We deliberately abstract from other benefits of foreign ownership to show that this financial advantage can explain the observed differential impact of the credit shock on firms' export volumes and intensities. In the empirical analysis, we explicitly account for and study some of these other benefits of foreign ownership (see Section 4.4.3).

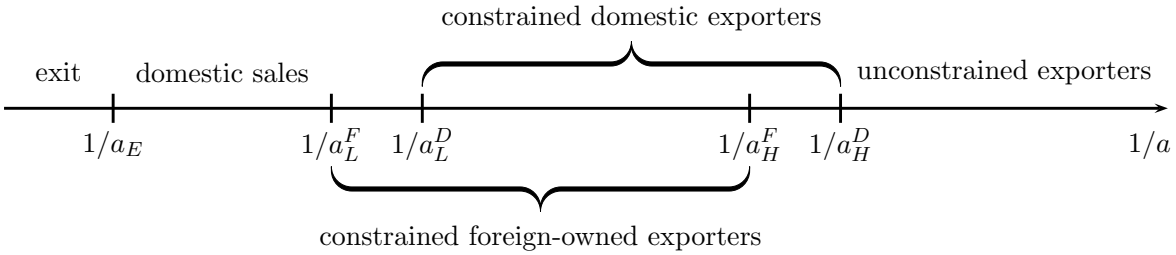
[15] Note that the foreign parent plays no role in the financial transaction as such, but only passes on the credit (without any surcharge) to its affiliate.

[16] Except for the knife-edge case of indifference, firms will never choose mixed financing.

local banks at r_D.

In the first case, where $r_F < r_D$, the world-market interest rate r_F replaces r_D in the (otherwise unchanged) equilibrium conditions (4.5)-(4.8) for the foreign-onwed firm. These conditions pin down two cut-offs above which foreign-owned firms can export at all ($1/a_L^F$) and above which they can export at first-best levels ($1/a_H^F$). These cut-offs will lie to the left of the corresponding cut-offs for domestically owned firms, as illustrated in Figure 4.1, because both cut-offs are increasing in the interest rate. Since foreign-owned firms have access to cheaper credit, more of them can raise the funds necessary to sell on the export market at all and at first-best levels in equilibrium.

Figure 4.1: *Sorting pattern of firms by productivity for* $r_F < r_D$



4.2.4 Comparative statics regarding the impact of the financial crisis

We are interested in how the credit crunch in the financial crisis affected the export intensities (*exports/sales*) of domestically owned relative to foreign-owned firms. In our framework, the credit crunch may have an impact on firm behavior via two channels. It may be modeled as (i) a negative shock to the efficiency of the capital market, reflected in the repayment probability λ, or (ii) a positive shock to the interest rates r_D and r_F (or the interest rate differential $r_D - r_F$). We choose to focus on the first case in the main text and consider the second case in Appendix 4.A.2. This choice is motivated by the idea that the financial crisis substantially increased the uncertainty of loan repayments, beyond the uncertainty associated with firms' fundamental characteristics. Around the peak of the financial crisis, marked by the bankruptcy of Lehman Brothers in September 2008, even loan repayments by major financial institutions were perceived as uncertain, which brought the interbank lending market to the verge of collapse. We view the deterioration of capital market efficiency reflected in the drop in λ as a temporary but global shock, which is the same for domestic and foreign financing.

We already know from the above analysis that firms of ownership type $f \in \{D, F\}$ with productivity levels below $1/a_L^f$ will not export, and those with very high productivity levels above the cut-off $1/a_H^f$ are not credit constrained and hence export at first-best levels. Note that our model predicts that the drop in λ will shift these cut-offs up and induces some firms to change their export status. However, it turns out that these effects of the financial crisis on export market entry and exit, which may be called the extensive margin of exports, turn out to be small and insignificant empirically in Spain (see Section 4.4.2 and Eppinger et al., forthcoming). Hence, this discussion concentrates on analyzing the export intensities of the most interesting firms with productivity levels $1/a$ in the critical range $\left[1/a_L^D(\lambda_{crisis}), 1/a_H^F(\lambda_{initial})\right)$, which are constrained exporters both before and after the shock. While not studied in detail here, the effects at the extensive margin generally work in the same direction as those on the intensive margin. They can be thought of as reinforcing the adverse effect of the financial crisis on total exports and the differential effect on foreign-owned compared to domestically owned firms.

In the interesting case of $r_F < r_D$, all foreign-owned exporting firms choose foreign financing both before and after the shock. equation (4.6) delivers an implicit solution for $p_j^{**}(a)$ for constrained exporters, depending on their type $f \in \{D, F\}$, which determines the interest rate they are facing depending on their internal funds $d(a)$. It allows us to derive the effect of the change in λ on prices $p_j(a)$ conditional on these characteristics. Comparative statics for export quantities $q_j^{**}(a)$ and export revenues $p_j^{**}(a)q_j^{**}(a)$ follow from this, as they are inversely related to $p_j^{**}(a)$ via the demand schedule from equation (4.1).

We can use this comparative statics exercise to show that a deterioration in credit market efficiency λ will reduce the export revenues of all constrained exporters. Since domestic sales are financed internally and remain unaffected by the crisis, this reduction translates directly into a reduction of the export intensity. Crucially, the detrimental crisis effect will be relatively larger for domestically owned firms, which face a higher interest rate r_D and have no other source of external funding. Intuitively, as the credit market friction gets worse, the financial advantage of foreign ownership becomes more important. By a similar logic, the differential effect will be larger among more financially vulnerable firms that have less internal funds (a high $d(a)$). These predictions are summarized in:

Proposition 4.1. *Among constrained exporters, a decrease in capital market efficiency in the financial crisis (i) weakly decreases the export intensity of domestically owned firms relative to foreign-owned firms, and (ii) this differential effect is larger for more financially vulnerable firms.*
Proof. *See Appendix 4.A.1.*

The qualification 'weakly' in Proposition 4.1 acknowledges that there is no differential effect for $r_F \geq r_D$ because in this case, there is no financial advantage of foreign ownership and both types of firms are identical. Note that unconstrained exporters are unaffected by the credit shock since their optimal prices and revenues are independent of λ.

We have held the real interest rates r_D and r_F constant throughout this comparative static exercise. However, one may argue that the credit crunch in the financial crisis has reduced credit supply and increased the real interest rate. Appendix 4.A.2 studies an alternative scenario, which models the financial crisis as a positive shock to the domestic real interest rate r_D, while assuming that the costs of foreign financing r_F (and also λ) remain unaffected. This modeling approach is motivated by the fact that multinational firms have access to finance in many countries and can choose the source of finance for which the interest rate remained lowest in the crisis. The predictions for constrained exporters in this scenario parallel those stated in Proposition 4.1. A positive shock to the domestic interest rate in the financial crisis tends to decrease the export intensity of domestically owned firms relative to foreign-owned firms, and this effect is stronger for more financially vulnerable firms. Interestingly, such an interest rate shock would additionally affect unconstrained domestically owned exporters (but not unconstrained foreign-owned exporters) because their export revenues are tied to the credit cost r_D via the participation constraint of the domestic investor.

4.3 Data

Our data come from the Encuesta Sobre Estrategias Empresariales (ESEE), an annual survey of about 2,000 manufacturing firms in Spain. The ESEE is a panel data set available starting in 1990 and managed by the Sociedad Estatal de Participaciones Industriales (SEPI). The sample we use for our analysis covers the period 2005-2012, which allows us to track firms through the years before, during, and after the

financial crisis. The initial sampling of the data in 1990 followed a two-tier structure designed to guarantee representativeness of the data for the manufacturing sector at large. Survey questionnaires were sent out to all 'large' firms (those with more than 200 employees), and to a subset of 'small' firms (those with 10 to 200 employees). Small firms were selected through stratified, proportional, and systematic sampling with a random seed. Industry affiliation and size class (defined by the number of employees) serve as stratification variables. Industries are defined by product category at the 2-digit level of the NACE Rev. 2 industry classification.[17] To mitigate the problem of sample attrition, SEPI includes refreshment samples, i.e., new firms are added to the survey as other firms exit. This is done in such a way that the representativeness of the data is preserved over time.

A crucial feature of our data set is that it includes not only information about the ownership structure of firms (foreign vs. domestic), but also about their export sales and their financial and debt situation. We define a firm as foreign owned if a foreign company owns more than 50% of its capital, and as domestically owned otherwise.[18] Around 14% of the firms in our sample are foreign owned over the period 2005-2012. Similar to firm-level data sets from other countries, this number is considerably higher for large firms (around 38%), and lower for small firms (around 6%). In terms of employment, sales, and exports, foreign-owned firms are quite important in our data set. Of the total number of effective working hours reported in 2012, 42% can be attributed to foreign-owned firms. For sales and exports the numbers are even higher, standing at 56% and 66%, respectively.[19]

Table 4.1: *Descriptive statistics*

	Domestically owned	Foreign owned	Observations (1)/(2)
	(1)	(2)	(3)
Export share (*exports/sales*)	0.173 (0.259)	0.412 (0.304)	13,441/2,187
Export dummy	0.597 (0.491)	0.960 (0.197)	13,441/2,187
Export sales (in logs)	14.107 (2.652)	16.738 (2.048)	8,022/2,099
Total sales (in logs)	15.484 (1.811)	18.067 (1.517)	13,452/2,194

Notes: This table shows means and standard deviations (in parentheses) of export and sales variables by ownership (domestic vs. foreign). The numbers of observations reported in the final column correspond to the firm-year observations in columns (1) and (2). The sample spans the years 2005-2012. Export sales and total sales are expressed in constant 2005 prices. *Source:* Authors' computations based on ESEE data.

Table 4.1 reveals a pronounced difference in the export orientation between foreign-owned and do-

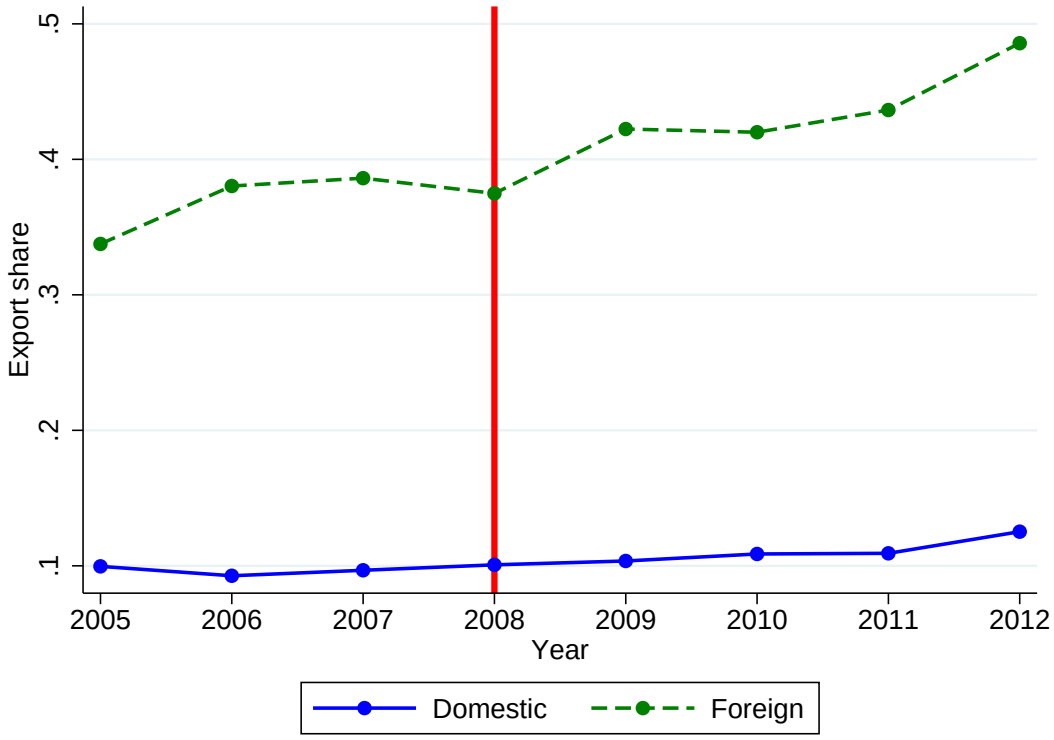
[17] Until 2009, the survey defined industries according to the NACE Rev. 1 classification. We accommodate the two classifications based on concordance information provided by SEPI. More information about the content of the survey along with its sampling properties is available at http://www.fundacionsepi.es/esee/en/epresentacion.asp.

[18] A small number of firms is foreign owned by more than zero, but not more than 50%, and thus labeled as domestically owned. We have checked that these firms are not driving our results by excluding them from the sample or by relabeling them as foreign owned. Our data set also includes information about unusual events such as mergers, acquisitions, and splitting up of firms. Such events can contaminate the analysis, as they often imply a drastic change not only in the ownership structure of the firm, but also in the type and scale of its output and exports (e.g. because the acquired entity is a large exporter). Hence, whenever a firm experiences one of the above-mentioned events, we treat it as a different firm afterwards. This applies to 30 firms in our sample.

[19] These numbers refer to the firms included in our raw data. When we apply sampling weights to correct for the oversampling of large firms in our data set we get somewhat smaller numbers that more closely match the statistics for the manufacturing sector at large. In terms of employment, sales, and export shares foreign-owned firms account, respectively, for 26%, 42%, and 55% in 2012.

mestically owned firms. We pool all observations across the years 2005-2012 and then sort them into groups of firms depending on their ownership structure. Virtually all firms in foreign ownership—96%—are exporters, while the same number is less than 60% among domestically owned firms. Among exporting firms, foreign-owned firms also export larger volumes, on average, than domestically owned firms.[20] These differences in both export status and export volumes translate into a considerably higher export share (i.e., *exports*/*sales*) for foreign-owned firms compared to domestically owned firms (41% vs. 17%). The firm-level export share as a measure of export orientation and export performance will be the focus of our empirical analysis in the next section, but rather than looking at mean differences between firms with different ownership structures, we will look into *changes* in this variable over time.

Figure 4.2: *Export share by ownership status*



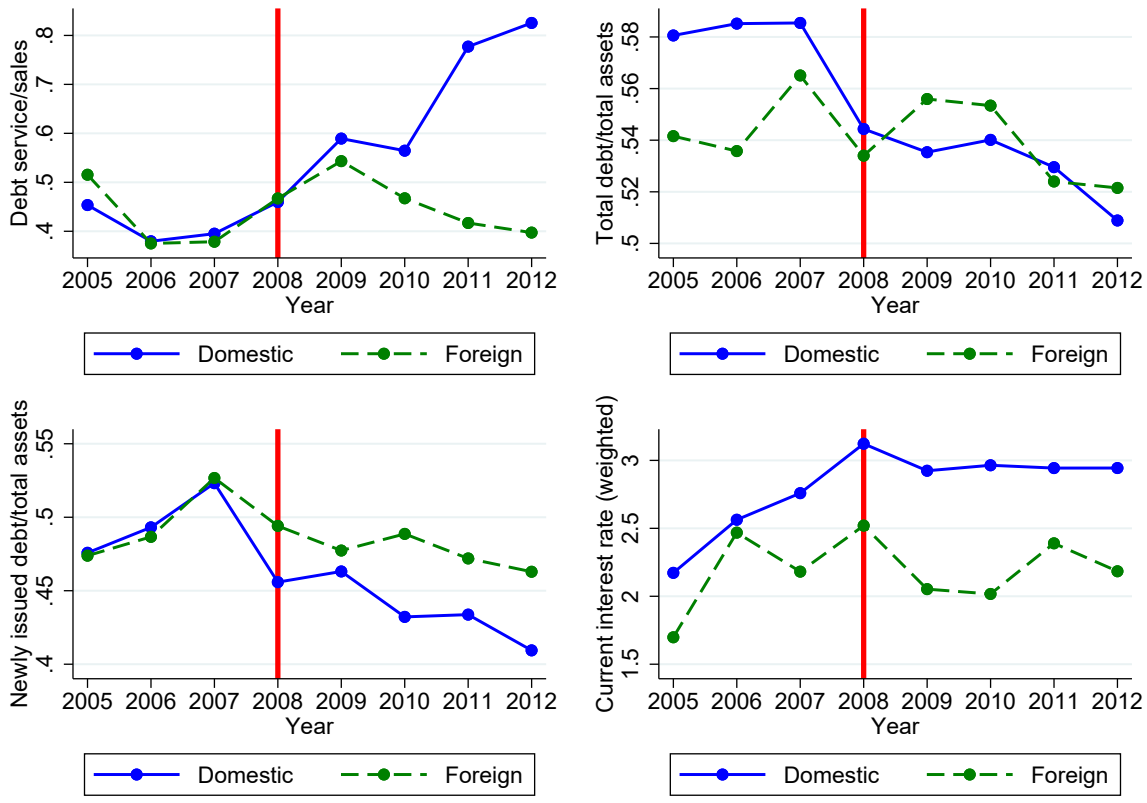
Notes: This figure depicts the average firm-level export shares by ownership (domestic vs. foreign). The panel used to construct the figure is balanced on firms with positive sales over the period 2007-2010. Sampling weights apply. *Source:* Authors' tabulations based on ESEE data.

Figure 4.2 provides a first glance at the evolution of export shares over time by ownership of firms (foreign vs. domestic). In contrast to Table 4.1 we now apply sampling weights to provide a picture representative for all firms in the Spanish manufacturing sector. Moreover, we balance the panel on firms with positive output in each and every year over the period 2007-2010. This allows us to abstract from the effects of market entry and exit in the years of the financial crisis. Two observations stand out. First, in line with the evidence in Table 4.1, foreign-owned firms have a consistently higher export share than domestically owned firms. Secondly, while the line for domestically owned firms is rather flat around the years of the financial crisis, indicating an almost constant export share at around 10%, we see a

[20] Throughout the paper, we express export sales and total sales in constant 2005 prices using firm-level output price indexes from ESEE, which are complemented by industry-level price indexes from the Spanish Instituto Nacional de Estadística (INE) when missing.

considerable increase in the export share among foreign-owned firms in 2009 (from 37% to more than 42%).[21] Why did the export share increase in the financial crisis? And why do we observe this increase only among foreign-owned firms, but not among domestically owned firms? In the next section, we will conduct a systematic econometric analysis of these questions, and we will focus on the possibility that foreign ownership was especially helpful in facilitating exports for those firms that were likely to run into credit constraints due to the credit supply shock of the financial crisis, as predicted by our theoretical model in Section 4.2.

Figure 4.3: *Financial and debt situation by ownership status (2005-2012)*



Notes: This figure depicts the average debt service-to-sales ratio; the average debt ratio; the average ratio of newly issued debt to total assets; and the average interest rate on newly issued debt (computed as a weighted average taking into account the amounts of different types of debt and their corresponding interest rates) by ownership (domestic vs. foreign). Sampling weights apply. *Source:* Authors' tabulations based on ESEE data.

A descriptive look into our data supports the view that the financial crisis had a differential effect on firms' debt situation depending on their ownership structure. Figure 4.3 depicts the evolution of important financial variables from 2005-2012 separately for firms in domestic and foreign ownership, respectively. We begin with our measure of financial vulnerability, shown in the top left panel, and defined as the ratio of debt service (i.e., the sum of debt and interest payments) to total sales. Intuitively, the variable measures what share of a firm's sales needs to be used for servicing the firm's debt, and

[21] Notably, balancing the panel on firms *exporting* over the period 2007-2010, rather than on firms *producing*, yields a very similar picture (available on request). This means that the extensive margin of exporting is not driving the differential change in export shares visible in 2009.

the higher the debt service-to-sales ratio, the more financially vulnerable is the firm.[22] We find average debt-to-service ratios around 0.45 over the period 2005-2008, and an extremely close comovement of this ratio for domestically owned and foreign-owned firms. After 2008, however, the two series start to diverge. Domestically owned firms experienced a striking increase in their debt service-to-sales ratio from 0.47 in 2008 to more than 0.8 in 2012, while foreign-owned firms had about the same degree of financial vulnerability in 2012 as in 2007 (viz. 0.4), in spite of a peak at 0.55 in 2009. In our econometric analysis in the next section, we will use this measure of financial vulnerability to predict how strongly firms reacted to the crisis in terms of their export intensity. However, rather than using a time-varying measure, which is strongly influenced by the crisis itself, we look at a firm's financial situation in 2008 (i.e., at the onset of the financial crisis), and show that it is a strong predictor of how a firm's export intensity evolved through the crisis.

The top right panel of Figure 4.3 suggests that domestically owned firms reduced their debt ratios (i.e., total debt over total assets) more strongly in the crisis than foreign-owned firms, in line with the idea of a credit supply shock that restricts access to credit differently depending on firms' ownership structures. The bottom left panel also supports this view by showing that, relative to their total assets, domestically owned firms issued not only less new debt in the crisis than before the crisis, but also considerably less than foreign-owned firms. This divergence is in line with our conjecture that domestically owned firms are more credit constrained. Finally, the bottom right panel shows that in 2009 and 2010 credit costs (measured by the firm's current interest rate) evolved more favorably for foreign-owned firms than for domestically owned firms. Notably, foreign-owned firms benefited from lower average interest rates throughout the sample period, which squares well with our assumption that they can access internal or external credit markets abroad at a lower cost (see Section 4.2). Taken together, our data provide remarkable evidence for the idea that the credit supply shock of the financial crisis had a much stronger negative effect on the debt situation of firms in domestic ownership as opposed to firms in foreign ownership.

4.4 Empirical Analysis

We now turn to our main empirical analysis. We proceed in three steps. First, we estimate the effect of foreign ownership on firms' export shares in the financial crisis using a difference-in-differences (DiD) framework. We find clear evidence that firms in foreign ownership (the treatment group) increased their export shares during the credit supply shock in 2009 relative to firms in domestic ownership (the control group), and we find that the effect is strongly concentrated among small firms, i.e., firms more likely to run into credit constraints in 2009. Secondly, we focus on the credit channel by exploring heterogeneity in the effect of foreign ownership in relation to the financial situation of the firm at the onset of the crisis in 2008 using a triple difference (DiDiD) framework. We find that those firms that had a higher debt service-to-sales ratio and were thus more financially vulnerable at the time the crisis hit the economy enjoyed much greater benefits from foreign ownership than firms with a less risky debt profile. In the final

[22] Debt service in any given year is determined through both short-term and long-term debt contracts signed in previous years. Debt service arising from short-term debt (i.e., with maturity of less than one year) is precisely reported in our data. Debt service arising from long-term debt can be approximated using the yearly stock of long-term debt reported by the firm as well as the average interest rate the firm reports to pay for its long-term debt (both the stock of long-term debt and the average interest rates are reported separately for financial and non-financial institutions, respectively). To keep things simple, and since we have no information on the number and maturity of the different long-term credit contracts the firm signs, we assume that each year the firm pays back 10% of its stock of long-term debt reported in the previous year plus the interest payment.

part of our empirical analysis, we investigate whether other benefits of foreign ownership (i.e., benefits not directly related to improved access to credit through the foreign parent) have also contributed to a relatively better firm-level export performance in the financial crisis. Our results indicate that superior management quality, stronger export market diversification, and better access to foreign markets via the parent's distribution network also play a role in explaining the higher export shares that we observe among foreign-owned firms.

4.4.1 The effect of foreign ownership

We now lay out a suitable econometric framework for our analysis and begin with the following linear model in order to estimate the relationship between foreign ownership and a firm's export share:

$$(exports/sales)_{it} = \gamma \cdot F_{it} + \sum_{r=2006}^{2012} \phi_r \cdot (F_{it} \times Y_r) + D_i + D_{st} + \varepsilon_{it}, \tag{4.10}$$

where the dependent variable is the export share of firm i in year t, F_{it} is a dummy variable for foreign ownership with a corresponding coefficient γ, and Y_r is a dummy variable equal to one if $r = t$, so that the sum collects interaction terms between the foreign ownership dummy and a comprehensive set of year dummies with corresponding coefficients ϕ_r (indexed in a self-explanatory way by $r = 2006, \ldots, 2012$).[23] The parameter D_i is a firm fixed effect, while D_{st} is an industry-year fixed effect. Finally, ε_{it} is an error term with zero conditional mean. Statistical inference will be based on robust standard errors clustered by firm, allowing for arbitrary forms of both serial correlation and heteroskedasticity.

The purpose of this model is to study in a systematic and flexible way the evolution of firm-level export shares through the financial crisis, and to exploit the credit crunch in 2008/09 as a source of exogenous variation in credit supply. In particular, we want to see whether and to what extent the response to the crisis was a function of the ownership structure of the firm. Against the backdrop of Proposition 4.1, we hypothesize that foreign-owned firms perform better on export markets in the years of the financial crisis than their domestically owned peers. Importantly, rather than restricting the estimation by defining the crisis period to start in 2008 or 2009, we keep the estimation flexible and interact the foreign ownership dummy with the full set of year dummies. Hence, we let the data speak as to when the crisis left measurable evidence in terms of its effect on firm-specific export shares. Adopting this flexible approach has two more advantages: First, it allows us to investigate common pretrends between foreign-owned and domestically owned firms before the crisis, and thus test whether our identifying assumption (i.e., common trends in export shares across domestically owned and foreign-owned firms) is likely to hold. And secondly, it allows us to investigate whether the expected effect on the export share is persistent, or whether it fades out.

Crucial for the interpretation of our estimates is the variation in the data we use for identification. Notice that the set of fixed effects included in the model absorb a wide variety of different effects. In particular, the firm fixed effects capture any time-invariant heterogeneity in observable or unobservable firm characteristics, such as firm-specific productivity, management quality, or the fixed cost of exporting. The industry-year fixed effects flexibly account for the industry-specific evolution of comparative advantage and arbitrary domestic or foreign demand shocks across industries in the crisis. The interac-

[23] Thus, the baseline category is the year 2005.

tion effects ϕ_r in equation (4.10) are hence identified from differences in within-firm variation over time across the two ownership groups, after controlling for industry-specific shocks.

The descriptive analysis of our data Section 4.3 revealed a strong competitive advantage of foreign-owned relative to domestically owned firms on export markets. As pointed out by Manova et al. (2015), there are many potential reasons for this advantage. Apart from enjoying better financing conditions, foreign-owned firms may have access to their parents' superior distribution network, employ better management practices, or have a more diversified export portfolio relative to domestically owned firms. The use of firm fixed effects in the estimation of (4.10) allows us to see whether foreign-owned firms *magnified* their competitive advantage on export markets during the crisis. Such an effect is less likely to be driven by omitted variables than a level difference in export shares, as firm characteristics that are constant through time and correlated with foreign ownership are controlled for.

In order to go one step further in tackling unobserved heterogeneity associated with both foreign ownership and firm-specific export shares, we combine the fixed effects model with a propensity score reweighting estimator.[24] Specifically, we construct propensity scores and reweight each observation in order to generate a similar distribution of key observable characteristics across domestically owned and foreign-owned firms. By matching along observable firm characteristics, we hope to also match the distribution of important unobservable characteristics. To estimate the propensity scores, we first consider the years 2007, 2008, and 2009 in our panel and sort those firms that are foreign owned in all three years into the treatment group and those that are domestically owned into the control group. Hence, we exclude those firms that switch their ownership status over time, as well as those firms that enter or leave the sample in one of the years 2007, 2008, and 2009. We then obtain the propensity scores for the remaining firms by running cross-sectional probit regressions of foreign ownership in 2007 (the treatment) on firm-specific sales, sales growth, capital intensity (all in logs), export share, and a full set of industry dummies. The firm-specific variables are all lagged by one year, i.e., they are observed in 2006. Each treated firm is reweighted by $1/\hat{p}$ and each control firm by $1/(1-\hat{p})$, where $\hat{p}$ is the estimated propensity score.[25] Importantly, $\hat{p}$ indicates the estimated probability that the firm is in foreign ownership shortly before the crisis materialized (i.e., in 2007), and is thus orthogonal to the crisis itself.

Columns (1) and (2) of Table 4.2 present estimates of γ and $\phi_r, r = 2006, \ldots, 2012$, based on the fixed effects estimator and the propensity score reweighting estimator, respectively. The results indicate a significant increase in the export share of foreign-owned relative to domestically owned firms in the crisis, as predicted by part (i) of Proposition 4.1. The fixed effects estimator identifies a differential increase by 2.3 percentage points in 2009 (significant at the 10% level) and a further increase in the two subsequent years, peaking at 3.9 percentage points in 2011 (significant at the 1% level) relative to the base year. The propensity score reweighting estimator confirms this significant divergence in export shares across the two groups, although the number of observations is considerably smaller (two thirds of the initial sample size). The point estimates of the effects are slightly larger, 3.7 percentage points in 2009 and 5.2 percentage points in 2011, but the coefficients are less precisely estimated. The differential

[24] The reweighting approach goes back to DiNardo et al. (1996) and has been applied by Garicano and Steinwender (2016) to the same Spanish firm data set used in this study.

[25] We only keep those observations in the analysis that are in the region of common support, and we have checked that the balancing property is supported in the data, i.e., after stratifying the population into six 'blocks' according to their propensity scores, all observed characteristics of domestically owned and foreign-owned firms are balanced in each block of the population. This means that within each block we cannot reject the null hypothesis (at any conventional significance level) that the means of the variables are identical across treatment and control group. We also winsorize the propensity scores at the 99th percentile. Detailed results of the propensity score estimation are available from the authors on request.

Table 4.2: *The effect of foreign ownership on the export share (DiD)*

	Dependent variable: Export share (*exports*/*sales*)					
	Full sample		Small firms		Large firms	
	(1)	(2)	(3)	(4)	(5)	(6)
Foreign	-0.00603	-0.00613	-0.0216	-0.0115	-0.000952	0.00516
	(0.0154)	(0.0188)	(0.0291)	(0.0285)	(0.0202)	(0.0229)
Foreign $\times$ Y_{2006}	0.00281	-0.000747	0.0242	0.0204	-0.00782	-0.0194
	(0.00769)	(0.00827)	(0.0162)	(0.0128)	(0.0128)	(0.0123)
Foreign $\times$ Y_{2007}	-0.000234	-0.0144	0.0244	-0.0124	-0.0132	-0.0187
	(0.00937)	(0.0106)	(0.0215)	(0.0200)	(0.0132)	(0.0118)
Foreign $\times$ Y_{2008}	-0.000703	-0.00266	0.0185	0.0128	-0.00217	-0.0119
	(0.0104)	(0.0108)	(0.0220)	(0.0164)	(0.0145)	(0.0148)
Foreign $\times$ Y_{2009}	0.0227*	0.0365*	0.0522*	0.0611	0.0260	0.0223
	(0.0123)	(0.0189)	(0.0281)	(0.0375)	(0.0174)	(0.0201)
Foreign $\times$ Y_{2010}	0.0282**	0.0493**	0.0731**	0.0956**	0.0148	0.0171
	(0.0133)	(0.0221)	(0.0306)	(0.0447)	(0.0179)	(0.0187)
Foreign $\times$ Y_{2011}	0.0385***	0.0521**	0.0689**	0.0952**	0.0240	0.0173
	(0.0141)	(0.0220)	(0.0329)	(0.0468)	(0.0200)	(0.0199)
Foreign $\times$ Y_{2012}	0.0311*	0.0438*	0.0724*	0.0923*	0.00698	0.00609
	(0.0161)	(0.0240)	(0.0371)	(0.0502)	(0.0218)	(0.0212)
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Industry-year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Propensity score		Yes		Yes		Yes
Observations	15628	10325	11542	6978	4086	3347
R-squared	0.055	0.088	0.058	0.121	0.092	0.105

Notes: The dependent variable in all columns is the firm-specific export share. Foreign is a dummy variable for foreign ownership (equal to one if the firm is foreign owned by more than 50% and zero otherwise). For details on the propensity score reweighting estimator see the text. Small firms are those with 10 to 200 employees, while large firms are those with more than 200 employees. Robust standard errors (in parentheses) are clustered at the firm level. *,**,*** denote significance at the 10%, 5%, 1% levels, respectively.

crisis impact on the export shares is sizable: It constitutes an increase in the foreign ownership premium of around 9-15% in 2009 compared to 2005.[26] Importantly, the estimation results support the idea that domestically owned and foreign-owned firms share a common trend in export shares *before* the crisis, as we find no significant differences in the evolution of export shares between the two types of firms over the period 2005-2008.[27]

The estimation results presented in columns (1) and (2) show that foreign-owned firms widened their competitive edge on export markets in the wake of the credit crunch. While suggestive of the relevance of the credit channel, these results could be caused by factors unrelated to credit constraints. For example, the quality of the management could have been more important during the financial crisis when firms were facing greater economic challenges and risks for their businesses. Moreover, the international distribution network maintained by MNEs might have helped foreign-owned firms to channel their sales into foreign markets that were faring better during the crisis than the domestic market (e.g. emerging-

[26]These numbers are computed by dividing the estimates of ϕ_{2009} from, respectively, Columns (1) and (2) of Table 4.2 by the unconditional difference in weighted averages for the two groups in 2005 of around 24 percentage points (see Figure 4.2).

[27]Notice that γ is identified only through a small number of firms switching into and out of foreign ownership. We have verified that our results are not driven by these firms, as they remain virtually unchanged when we exclude switchers from the sample altogether.

market economies). We therefore need an additional margin of variation in the data in order to clearly disentangle the credit channel from other potentially confounding factors.

One such margin could be firm size. There is by now an extensive body of evidence suggesting that small firms are more likely to be credit constrained than large firms (see Gertler and Gilchrist, 1994; Guiso et al., 2004; Beck et al., 2008). This fact can be rationalized by our theoretical model of heterogeneous firms under credit constraints. As we have seen in Section 4.2, only the less productive, smaller exporting firms are credit constrained, while the most productive firms can always export at first-best levels (as in Manova, 2013). Hence, in order to further scrutinize the potential export effect of credit constraints, we estimate the model in (4.10) separately for small and large firms; see columns (3) to (6) in Table 4.2. We find that the above-described pattern of estimated coefficients is borne out by the data in the sample of small firms, but not in the sample of large firms. We also find larger coefficients when using the sample of small firms than before when using the full sample, suggesting a more important role foreign ownership among small firms. While we believe that these results serve as an indication for the credit channel to be at work, we cannot rule out the possibility that other benefits of foreign ownership unrelated to credit constraints also gained importance for small firms' exports in the crisis. The above-mentioned differences in management quality between foreign-owned and domestically owned firms, for instance, could be more pronounced in small firms than in large firms. An alternative and more promising route to pin down the credit channel is suggested by our model: The effect of foreign ownership on exports in the credit crunch should be more pronounced for financially vulnerable firms, which depend more heavily on credit to finance their exports. Therefore, we directly focus on firms' financial vulnerability in the next section.

4.4.2 The credit channel of foreign ownership

We now focus explicitly on the credit channel of foreign ownership. To do so and identify the importance of credit constraints for exports in the financial crisis, we use a triple difference identification strategy that exploits variation in the ownership structure of firms (foreign versus domestic), variation in financial vulnerability across firms at the onset of the financial crisis, as well as the negative credit supply shock caused by the financial crisis. We first present our main results using the export share of firms as the dependent variable. We then investigate the effects at the extensive and the intensive margins of exports separately, distinguishing between the discrete choice of whether or not to export on the one hand, and the continuous choice of how much to export on the other hand. In the final part, we offer a variety of robustness checks.

Main results

We derive our main results from estimating the following equation:

$$\begin{aligned}(exports/sales)_{it} \;=\; & \gamma \cdot F_{it} + \rho \cdot (F_{it} \times V_i) + \sum_{r=2006}^{2012} \phi_r \cdot (F_{it} \times Y_r) + \sum_{r=2006}^{2012} \delta_r \cdot (V_i \times Y_r) + \\ & \sum_{r=2006}^{2012} \theta_r \cdot (F_{it} \times V_i \times Y_r) + D_i + D_{st} + \varepsilon_{it}, \end{aligned} \tag{4.11}$$

where V_i is our firm-specific measure of financial vulnerability (i.e., a firm's debt service-to-sales ratio in 2008, as described in Section 4.3), and the parameters of main interest are the coefficients of the triple

interaction terms: $\theta_r,\ r = 2006, \ldots, 2012$.[28] Based on part (ii) of Proposition 4.1, we expect to find $\theta_r > 0$ in the crisis years. The basic idea behind this model is to investigate whether the effect of foreign ownership is stronger for firms that are more likely to face binding credit constraints when hit by a credit supply shock. The key assumption is that firms with a higher debt service-to-sales ratio in 2008 will have found it more difficult to finance their business activities, and in particular their exporting activities, when liquidity dried out in the financial crisis. Importantly, we measure financial vulnerability in 2008, based on credit contracts signed in 2007 for newly issued short-term and long-term debt, as well as credit contracts signed before 2007 for older long-term debt. Hence, the firm's debt service in 2008 was unlikely to be influenced by the credit crunch, which came unexpected and the severity of which had not been anticipated. Moreover, and as we have discussed in Section 4.3, the series of domestically owned and foreign-owned firm of our measure of financial vulnerability moved in parallel over the period 2005-2008. Hence, we have reason to believe that the selection of firms into more or less risky debt profiles before the crisis was independent of the ownership structure of firms.[29]

Table 4.3 reports our main estimation results: We find that foreign ownership increased firms' export shares in the crisis more strongly among financially vulnerable firms, and this effect was concentrated in small firms. The table reports estimates of ϕ_r and $\theta_r,\ r = 2006, \ldots, 2012$ in equation (4.11) first for the full sample (columns (1) to (2)), and then separately for the samples of small and large firms (columns (3) to (4) and (5) to (6), respectively).[30] In each case, we first use the standard fixed effects estimator and then the propensity score reweighting estimator described in the previous section. We find economically and statistically significant triple interaction effects in 2009 and the subsequent years in the full sample. These estimates show that the effect of foreign ownership on exports was greater the more financially vulnerable the firm was in 2008. As in the case of our DiD results, this DiDiD effect is even larger among small firms, while it is not statistically significant in the sample of large firms, which were less likely to face binding credit constraints. These result provide strong support for part (ii) of our Proposition 4.1 and the financial advantage of foreign ownership.

The economic magnitude of the effects estimated in Table 4.3 is substantial. To make sense of our results in a quantitative way, we evaluate the crisis impact of foreign ownership at the 75th and the 25th percentile of financial vulnerability. We also focus on the sample of small firms, where we have identified the main differential effect of foreign ownership. While the effect of foreign ownership on export shares in 2009 was close to zero evaluated at the 25th percentile of financial vulnerability, it was large among highly financially vulnerable firms. The estimated effect of foreign ownership in the crisis was greater by around 5.3 percentage points for firms at the 75th compared to the 25th percentile of financial vulnerability, independent of the estimator used.[31] This differential effect is more than twice as large, reaching up to 11.6 percentage points, if we compare firms at the 90th to the 10th percentile of financial vulnerability. Note that these effects are estimated relative to the base year 2005, but since we find zero effects before 2009, they are also approximately equal to the effect in 2009 relative to 2008. We conclude that foreign ownership had a large positive impact on the export performance of highly

[28] Further parameters to be estimated in equation (4.11) are ρ, ϕ_r and $\delta_r,\ r = 2006, \ldots, 2012$, i.e., the coefficients of the interaction terms, as well as γ, the coefficient of the foreign ownership dummy. The fixed effects D_i and D_{st} absorb all other individual elements of the triple interaction, viz. V_i and $Y_r,\ r = 2006, \ldots, 2012$.

[29] In a similar vein, Garicano and Steinwender (2016) use the ratio of short-term debt with financial institutions to total debt in 2007 as a measure of financial vulnerability at the onset of the crisis.

[30] The remaining interaction terms between Foreign, FinVul, and the year dummies are not reported to save space.

[31] The 25th and 75th percentiles of the debt-service-to-sales ratio in our sample are 0.231 and 0.529, respectively. We compute the these effects as $\hat{\theta}_{2009} \cdot 0.529 - \hat{\theta}_{2009} \cdot 0.231$, using the estimates from columns (3) and (4).

Table 4.3: *The credit channel of foreign ownership (DiDiD)*

	Dependent variable: Export share (*exports/sales*)					
	Full sample		Small firms		Large firms	
	(1)	(2)	(3)	(4)	(5)	(6)
Foreign $\times$ Y_{2006}	0.0101	0.00990	0.0407**	0.0344*	-0.0426	-0.0296
	(0.0124)	(0.0130)	(0.0208)	(0.0181)	(0.0280)	(0.0246)
Foreign $\times$ Y_{2007}	0.000290	-0.0147	0.0200	-0.0222	-0.0247	-0.0186
	(0.0144)	(0.0150)	(0.0280)	(0.0272)	(0.0266)	(0.0228)
Foreign $\times$ Y_{2008}	-0.00613	-0.00471	0.0107	0.0104	-0.0337	-0.0359
	(0.0141)	(0.0134)	(0.0255)	(0.0184)	(0.0269)	(0.0246)
Foreign $\times$ Y_{2009}	-0.0178	-0.0417**	-0.0343	-0.0457**	-0.0131	-0.0250
	(0.0219)	(0.0188)	(0.0218)	(0.0178)	(0.0309)	(0.0335)
Foreign $\times$ Y_{2010}	-0.0137	-0.0279	-0.0175	-0.00978	-0.0139	-0.0236
	(0.0256)	(0.0255)	(0.0285)	(0.0333)	(0.0320)	(0.0290)
Foreign $\times$ Y_{2011}	0.00402	-0.0210	-0.0219	-0.0145	0.00409	-0.00154
	(0.0237)	(0.0212)	(0.0296)	(0.0232)	(0.0374)	(0.0321)
Foreign $\times$ Y_{2012}	-0.00537	-0.0258	-0.0254	-0.0150	0.00237	0.0109
	(0.0254)	(0.0212)	(0.0383)	(0.0289)	(0.0384)	(0.0345)
Foreign $\times$ FinVul $\times$ Y_{2006}	-0.0185	-0.0296	-0.0190	-0.0368	0.0650	0.0186
	(0.0226)	(0.0220)	(0.0323)	(0.0241)	(0.0601)	(0.0508)
Foreign $\times$ FinVul $\times$ Y_{2007}	-0.00497	-0.00954	0.00707	-0.00405	0.0240	-0.00344
	(0.0223)	(0.0194)	(0.0308)	(0.0223)	(0.0551)	(0.0461)
Foreign $\times$ FinVul $\times$ Y_{2008}	0.00925	-0.000840	0.0148	-0.00978	0.0653	0.0526
	(0.0207)	(0.0173)	(0.0352)	(0.0173)	(0.0554)	(0.0487)
Foreign $\times$ FinVul $\times$ Y_{2009}	0.0793*	0.151***	0.178***	0.177***	0.0715	0.0957
	(0.0451)	(0.0395)	(0.0199)	(0.0111)	(0.0625)	(0.0758)
Foreign $\times$ FinVul $\times$ Y_{2010}	0.0821	0.145***	0.195***	0.176***	0.0596	0.0810
	(0.0532)	(0.0460)	(0.0234)	(0.0163)	(0.0623)	(0.0562)
Foreign $\times$ FinVul $\times$ Y_{2011}	0.0776*	0.134***	0.193***	0.173***	0.0467	0.0354
	(0.0471)	(0.0397)	(0.0253)	(0.0137)	(0.0728)	(0.0572)
Foreign $\times$ FinVul $\times$ Y_{2012}	0.0761	0.128***	0.189***	0.168***	0.0222	-0.0122
	(0.0480)	(0.0349)	(0.0306)	(0.0140)	(0.0696)	(0.0623)
Full set of interactions	Yes	Yes	Yes	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Industry-year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Propensity score		Yes		Yes		Yes
Observations	12488	9937	8873	6720	3615	3217
R-squared	0.069	0.166	0.094	0.290	0.106	0.115

Notes: The dependent variable in all columns is the firm-specific export share. Foreign is a dummy variable for foreign ownership (equal to one if the firm is foreign owned by more than 50% and zero otherwise). FinVul measures the financial vulnerability of the firm in 2008; it is defined as the debt service-to-sales ratio (see the text for details). All columns include a full set of interaction terms between Foreign, FinVul, and the year dummies. For details on the propensity score reweighting estimator see the text. Small firms are those with 10 to 200 employees, while large firms are those with more than 200 employees. Robust standard errors (in parentheses) are clustered at the firm level. *,**,*** denote significance at the 10%, 5%, 1% levels, respectively.

financially vulnerable firms, but much less on other firms. Our findings support the view that internal capital markets were a key advantage for these firms' exports in the crisis.

Note that in the sample of small firms, the results are very similar across the different estimators both in terms of magnitudes and significance levels. Also, it is apparent from our estimates that we cannot reject the null hypothesis of $\theta_r = \theta_{s \neq r} \; \forall \; r, s \geq 2009$, i.e., the coefficients of the triple interaction terms are remarkably stable over the period 2009-2012. For the remaining part of our analysis, we therefore define a crisis dummy variable which takes on the value zero for the period 2005-2008 and the value one for the period 2009-2012.

Extensive versus intensive margin of exports

We now distinguish between the extensive and the intensive margin of exports, in order to see whether foreign-owned firms were more likely to enter the export market in the crisis (or, for that matter, less likely to *exit* the export market), or whether they exported relatively more goods and services in the crisis, or both. In principle, this distinction would allow us to investigate whether it is the fixed or the variable cost of exporting that matter more for exports in the presence of capital market imperfections; a question that has been debated in the literature.[32] However, there are two caveats concerning this interpretation in the context of our analysis. First, the share of foreign-owned firms that are exporters is very high already before the crisis, leaving little scope for further export market entry in this group. Secondly, we define the extensive margin of exports at the firm level, while we do not have data destination countries or different products exported. Hence, we cannot investigate potential effects on these below-firm-level extensive margins, although there might be market-specific fixed costs of exporting. Hence, we cannot draw strong conclusions regarding the question of whether variable or fixed costs of exporting played a more important role in the financial crisis.

Before we estimate the model separately for the extensive and intensive margins of exports, we consider the crisis impact on total sales of foreign-owned versus domestically owned firms. We use the same triple difference model as before, but now we use the firm's total sales (deflated and in logs) as the dependent variable. The results reported in columns (1) and (2) of Table 4.4 indicate a quantitatively important average treatment effect of foreign ownership equal to 9.7 or 6.6 percentage points, depending on the estimator used. Interestingly, we do not find a stronger effect for financially more vulnerable firms. In light of our main result that foreign ownership was especially helpful in increasing the export shares of financially vulnerable firms (see above in Table 4.3), the evidence seems to suggest that domestic sales increased by less in these same firms. This finding may point to a substitutive relationship between domestic and export sales within Spanish firms in response to the crisis.[33]

In columns (3) to (6) of Table 4.4, we distinguish between the two different margins of exports. For the extensive margin, we define an export dummy which is equal to one if the firm is an exporter, and zero otherwise, and we use this export dummy in our triple difference estimation. We do the same for the intensive margin, where we use the volume of exports (deflated and in logs) as the dependent variable, which implies that we only include exporting firms in this estimation. The results show compelling evidence that there is no differential effect at the extensive margin of exports, but a large effect at the intensive margin of exports, where the coefficient of the triple interaction term is estimated to be

[32] See e.g. the discussions in Minetti and Zhu (2011) and Muûls (2015) of the setups in Chaney (2016) and Manova (2013).

[33] It is beyond the scope of our paper to explore this relationship further, but we believe that shedding more light on this pattern is an interesting question for future research. See Almunia et al. (2017) and Eppinger et al. (forthcoming) for studies that go into this direction.

positive and highly significant regardless of the estimator used. While the average treatment effect of foreign ownership on the volume of exports is close to zero evaluated at the average degree of financial vulnerability, we find large positive effects for highly financially vulnerable firms. The triple difference estimator suggests that foreign ownership increased exports in the crisis by 13.2% to 20.3% more for firms at the 75th compared to the 25th percentile of financial vulnerability, depending on the estimator used. Perhaps surprisingly, our estimates imply negative treatment effects for firms with very low degrees of financial vulnerability (i.e., firms with a debt service-to-sales ratio close to zero, meaning no or almost no debt). However, this is true only for a small fraction of firms.

Table 4.4: *Extensive versus intensive margin of exports (DiDiD)*

	Total sales		Export dummy		Export sales	
	(1)	(2)	(3)	(4)	(5)	(6)
Foreign × Crisis	0.097**	0.066*	-0.016	0.006	-0.174**	-0.217***
	(0.038)	(0.040)	(0.013)	(0.017)	(0.083)	(0.079)
Foreign × FinVul × Crisis	-0.000	-0.014	0.006	-0.001	0.444***	0.681***
	(0.058)	(0.064)	(0.014)	(0.015)	(0.153)	(0.127)
Full set of interactions	Yes	Yes	Yes	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Industry-year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Propensity score		Yes		Yes		Yes
Observations	12502	9948	12488	9937	8056	6585
R-squared	0.293	0.284	0.017	0.023	0.052	0.092

Notes: The dependent variable in columns (1) and (2) is the log of total sales; in columns (3) and (4) an export dummy; and in columns (5) and (6) the log of export sales. Export sales and total sales are expressed in constant 2005 prices. Foreign is a dummy variable for foreign ownership (equal to one if the firm is foreign owned by more than 50% and zero otherwise). FinVul measures the financial vulnerability of the firm in 2008; it is defined as the debt service-to-sales ratio (see the text for details). Crisis is a dummy variable equal to one in years 2009 to 2012, and zero otherwise. All columns include a full set of interaction terms between Foreign, FinVul, and the crisis dummy. Robust standard errors (in parentheses) are clustered at the firm level. *,**,*** denote significance at the 10%, 5%, 1% levels, respectively.

Robustness analysis

In this section, we offer an extensive robustness analysis. We consider in turn (i) alternative specifications of our triple difference model, (ii) a fractional probit model for export shares, and (iii) an analysis at the more disaggregate level of firm-by-export destinations to control for the possibility of heterogeneous demand shocks across different world regions. The bottom line is that our main result is robust to all of these modifications. The analysis thus supports our basic argument that foreign-owned firms were less credit-constrained in the crisis than their domestically owned peers, and that this fact translated into relatively higher export shares.

We begin by modifying our main specification in various ways. We first estimate equation (4.11), but now we use interactions with the crisis dummy instead of the year dummies. Both the fixed effects estimator and the propensity score reweighting estimator show significant treatment effects of foreign ownership; see columns (1) and (2) of Table 4.5, respectively. Clearly, our preferred estimator is the propensity score reweighting estimator, since it makes an additional effort in ensuring that the group of domestically owned firms is a suitable control group for the group of foreign-owned firms. This estimator implies an average treatment effect of foreign ownership for firms with an average degree of financial

vulnerability equal to 4.2 percentage points. The effect is by 4.5 percentage points higher at the 75th compared to the 25th percentile of financial vulnerability.

Foley and Manova (2015) emphasize that the relationship among firm productivity, firm size, and financial vulnerability is complex. Nevertheless, we want to make sure that the difference in the evolution of export shares that we find between foreign-owned and domestically owned firms is not purely driven by differences in firm size or productivity. In column (3), we therefore augment the model to include size-year fixed effects, where we distinguish between six different size groups of firms in terms of the number of employees ($\leq$20;21-50;51-100;101-200;201-500;$>$500). In column (4), we include firm-specific total factor productivity (TFP) (lagged by one year) as well as its interaction with the crisis dummy. We find that a higher TFP is associated with a higher export share, but we do not find that this effect is magnified or reduced in the crisis.[34]

We also examine the robustness of our main finding to including proxies for financial health in the propensity score estimation. This is done to account for the possibility that multinational firms may seek to acquire financially vulnerable exporting firms (Manova et al., 2015). In column (5), we provide estimates based on an enhanced propensity score reweighting procedure, where we also consider one-year lags of the firm's amount of debt service (in logs) and its debt ratio (total debt over total assets) as well as the size group of the firm (as a fixed effect) when estimating the propensity scores. In the last column, we combine all three modifications in one estimation, i.e., we include size-year fixed effects and firm-specific TFP as well as its interaction with the crisis dummy, and we estimate the model with our enhanced propensity score reweighting estimator. The bottom line is that the estimated coefficient of the triple interaction term proves to be very robust to all the modifications we introduce into the model in Table 4.5.

Next, we account for the fact that our dependent variable, the export share, is a fractional response variable, i.e., it is naturally bounded between zero and one. For this reason, our linear fixed effects estimator may deliver inconsistent estimates of the treatment effect. To evaluate whether this is a problem in our application, we have estimated a fractional probit model. A complication is that the fractional probit model does not allow controlling for firm fixed effects without introducing an incidental parameters problem into the estimation. Hence, we have estimated a pooled fractional probit model with a triple interaction along the lines of equation (4.11), but since we cannot control for firm fixed effects, we have included fixed effects for firm size groups as well as our index measure of TFP (lagged by one year). To the extent that differences in firm size and productivity, whether time-varying or time-constant, account for selection into foreign ownership in our model, thus rendering the selection of firms into the treatment and the control group random, this model allows for valid identification of the treatment effect of foreign ownership in the financial crisis.

The top panels in Figure 4.4 visualize the key results from the fractional probit estimation. They show the average discrete effect of foreign ownership for each of the years 2007 to 2012. The base year is always 2006, i.e., all effects are shown relative to 2006.[35] We illustrate the effects depending on the firm's degree of financial vulnerability, and we focus on small firms, since the effects are strongest within this group of firms (as in Table 4.3). The panels from left to right are for firms with low, medium, and high degrees of financial vulnerability (corresponding to the 25th percentile, the average value, and the

[34] Our measure of TFP is the same index used by Delgado et al. (2002) whose analysis is based on the same Spanish firm-level data source as ours. As in their paper, the index is constructed as the log of the firm's output minus a cost-share weighted sum of the log of the firm's inputs. This approach goes back to Caves et al. (1982).

[35] The year 2005 is lost since we control for lagged TFP.

Table 4.5: *Robustness — Various modifications*

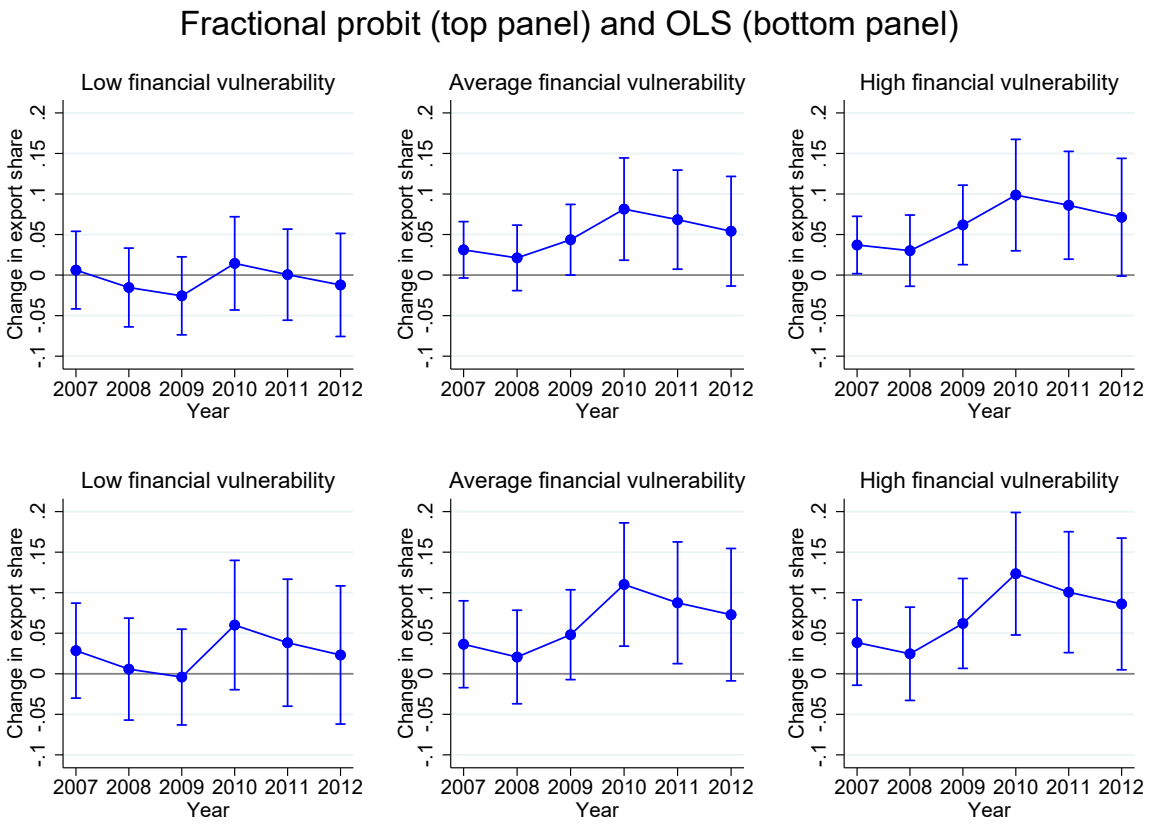
	Dependent variable: Export share (*exports/sales*)					
	FE	PS	Size-year FE	TFP	Enhanced PS	Combined
	(1)	(2)	(3)	(4)	(5)	(6)
Foreign × Crisis	-0.0105	-0.0279	-0.0316*	-0.0249	-0.0257	-0.0249
	(0.0217)	(0.0184)	(0.0186)	(0.0186)	(0.0222)	(0.0221)
Foreign × FinVul × Crisis	0.0833*	0.151***	0.150***	0.154***	0.128***	0.133***
	(0.0484)	(0.0395)	(0.0388)	(0.0408)	(0.0482)	(0.0488)
TFP				0.0392**		0.0245
				(0.0155)		(0.0156)
TFP × Crisis				-0.0210		-0.0167
				(0.0166)		(0.0183)
Full set of interactions	Yes	Yes	Yes	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Industry-year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Propensity score		Yes	Yes	Yes	Enhanced	Enhanced
Size-year fixed effects			Yes			Yes
Observations	12488	9937	9937	8487	9638	8245
R-squared	0.065	0.162	0.170	0.167	0.157	0.167

Notes: The table presents the results of various robustness checks. In columns (1) and (2) we use, respectively, the standard fixed effects estimator and the standard propensity score reweighting estimator as in Table 4.3. In column (3) we include size-year fixed effects, where we distinguish between six different size groups according to the firm's number of employees ($\leq$20;21-50;51-100;101-200;201-500;$>$500). In column (4) we include the lag of total factor productivity (in logs) as well as its interaction with the crisis dummy (see the text for details on our TFP measure). In column (5) we additionally interact TFP, FinVul, and Crisis and include a comprehensive set of corresponding interaction terms. In column (6) we use an enhanced propensity score reweighting estimator (see the text for details). In column (7) we combine the changes of columns (3) to (6) in one estimation. The dependent variable in all columns is the firm-specific export share. Foreign is a dummy variable for foreign ownership (equal to one if the firm is foreign owned by more than 50% and zero otherwise). FinVul measures the financial vulnerability of the firm in 2008; it is defined as the debt service-to-sales ratio (see the text for details). Crisis is a dummy variable equal to one in years 2009 to 2012, and zero otherwise. All columns include a full set of interaction terms between Foreign, FinVul, and the crisis dummy. Robust standard errors (in parentheses) are clustered at the firm level. *,**,*** denote significance at the 10%, 5%, 1% levels, respectively.

75th percentile of financial vulnerability). The estimates show that our main results from Section 4.4.2 prove to be robust to using the fractional probit model rather than the linear fixed effects estimator. We find significant effects of foreign ownership on export shares in the crisis, but these effects are more pronounced for firms with high degrees of financial vulnerability, and they are muted for firms with low degrees of financial vulnerability. To compare these estimates to a standard linear model, we have also estimated a pooled OLS model using the same specification as for the fractional probit model. To facilitate comparison, we show the OLS results in the bottom panels of Figure 4.4, and we get essentially the same picture as with the fractional probit model.

In a final robustness check concerning our main result, we want to control for demand shocks at the industry-destination level. These can be important omitted factors which may bias our estimates since they might occur concurrently with the credit shock, as argued by Paravisini et al. (2014). The export data we have available at the firm level are not ideal to tackle this issue, as they are not disaggregated by country of destination. However, we can exploit information on the regional composition of exports in 2006 and 2010 included in our data. In particular, firm-level exports in these two years are reported by destination at the level of four world regions. These regions are the European Union (EU), Latin

Figure 4.4: *Robustness — Average discrete effects of foreign ownership relative to 2006*



Notes: This figure depicts the average discrete effects of foreign ownership on the export share for the years 2007 to 2012, respectively (relative to the base year 2006). Low and high financial vulnerability refer to the 25th and the 75th percentile of the debt service-to-sales ratio, respectively. The top and bottom panel show estimates based on the fractional probit and the OLS estimator, respectively (see the text for details). The bars indicate 95% confidence intervals based on robust standard errors clustered at the firm level. The estimation sample is restricted to small firms, i.e., those with 10 to 200 employees. *Source:* Authors' tabulations based on ESEE data.

America, other OECD countries, and the rest of the world. While admittedly broad, this aggregation groups countries at similar income levels and geographic and cultural distance to Spain together. Also, the financial crisis had a very strong effect on aggregate demand in the EU and other OECD countries, since the financial crisis originated in the US and spilled over quickly to Western European countries. Latin America and the rest of the world (Africa, most of Asia, and Eastern Europe) were less severely affected. This is reflected in our data, where the growth rates of firm-level export shares from 2006 to 2010 were lower in the EU and other OECD countries than in Latin America and the rest of the world, respectively.

In 2006 and 2010, we have four observations (one per world region) for each firm. At this level, we can control for demand shocks by industries and world region by estimating the following equation:

$$\begin{aligned}(exports/sales)_{ijt} \;=\; & \gamma \cdot F_{it} + \rho \cdot (F_{it} \times V_i) + \phi_{2010} \cdot (F_{it} \times Y_{2010}) + \delta_{2010} \cdot (V_i \times Y_{2010}) + \\ & \theta_{2010} \cdot (F_{it} \times V_i \times Y_{2010}) + D_{ij} + D_{sjt} + \varepsilon_{ijt}, \end{aligned} \tag{4.12}$$

where $(exports/sales)_{ijt}$ is the share of exports by firm i to region j at time t (with $t = 2006, 2010$)in the firm's total sales, D_{ij} is a firm-region fixed effect, D_{sjt} is an industry-region-time fixed effect, and

ε_{ijt} is the error term. Importantly, demand shocks that are specific to an industry-region-year triplet are absorbed by our fixed effects in this model. The results are reported in Table 4.6 and demonstrate a high degree of robustness of our main result no matter whether we use the fixed effects estimator (column (1)) or the propensity score reweighting estimator (column (2)).

Table 4.6: *Robustness — Controlling for demand shocks*

	Dependent variable: Export share (*exports/sales*)	
	(1)	(2)
Foreign $\times$ Y_{2010}	-0.00921	-0.0129**
	(0.00671)	(0.00593)
Foreign $\times$ FinVul $\times$ Y_{2010}	0.0312**	0.0491***
	(0.0140)	(0.0115)
Full set of interactions	Yes	Yes
Firm-destination fixed effects	Yes	Yes
Industry-year-destination fixed effects	Yes	Yes
Propensity score		Yes
Observations	12726	10084
R-squared	0.027	0.071

Notes: The dependent variable in all columns is the firm-specific export share. Foreign is a dummy variable for foreign ownership (equal to one if the firm is foreign owned by more than 50% and zero otherwise). FinVul measures the financial vulnerability of the firm in 2008; it is defined as the debt service-to-sales ratio (see the text for details). Crisis is a dummy variable equal to one in years 2009 to 2012, and zero otherwise. All columns include a full set of interaction terms between Foreign, FinVul, and the crisis dummy. Robust standard errors (in parentheses) are clustered at the firm level. *,**,*** denote significance at the 10%, 5%, 1% levels, respectively.

4.4.3 Other channels

We now investigate other channels which could play an important role in our context and explain our main result. We focus on three channels that have received attention in the literature on foreign ownership and multinational activity. These channels have to do with management quality, export diversification, and the distribution network of multinational firms.

In a first step, we want to consider the possibility that management quality is an important determinant of the crisis performance of firms and thus the evolution of firm-level export shares in our sample. Since foreign-owned firms are often believed to be better managed than domestically owned firms (e.g. because the foreign parent transfers better management practices to its affiliated firm or installs a better manager), we want to make sure that our results in Table 4.11 are not driven by differential crisis effects of management quality. To measure management quality, we compute the average of all management innovations the firm reports across the years 2005-2012. Management innovations include innovations regarding a firm's sales channel, product design, pricing strategies, promotion activities, external relations management, and labor organization. In line with our expectations, foreign-owned firms appear to be better managed, on average, than domestically owned firms according to our measure. If we pool all observations across all years, as we did in Table 4.1, then the average management score for foreign-owned firms is 0.85 innovations per year, while the same number for domestically owned firms is 0.61. When we include our measure of management quality in our regression analysis and allow for differ-

ential importance in crisis versus non-crisis times, then we find that it is positively related to export performance in the crisis; see column (1) in Panel A of Table 4.7. We do not find, however, that more financially vulnerable firms enjoy larger benefits from higher management quality than less financially vulnerable firms; see column (2). Neither do we find that taking management quality into account affects our main result that foreign ownership leads to a strong relative increase in the export share of financially vulnerable firms in the crisis; see columns (3) and (4), where we employ the fixed effects and the propensity score reweighting estimator, respectively.

In a second step, we investigate whether a stronger export market diversification led to a more favorable export performance in the crisis. To do so, we use the number of international markets a firm reports to serve in 2007 (i.e., before the crisis) as a measure of export diversification. We find that foreign-owned firms served a considerably higher number of international markets than domestically owned firms in 2007 (the average number is 1.31 versus 0.66, respectively). The results of our regression analysis show that an ex ante more diversified export portfolio led to a significantly stronger export performance in the crisis; see column (1) in Panel B of Table 4.7. From column (2) one could get the impression that the number of international markets does not interact with the firm's degree of financial vulnerability, since the coefficient of the triple interaction turns out to be insignificant. The picture changes, however, when we bring in our main channel, viz. the triple interaction between foreign ownership, financial vulnerability, and the crisis dummy, which is important since foreign ownership and export market diversification are highly correlated as shown above. Using our preferred estimator in column (4), we find that the effect of export market diversification is less positive in more financially vulnerable firms. This is a plausible result since a more diversified export portfolio implies higher cumulative fixed cost of exporting. If firms are more likely to be credit constrained in the financial crisis, then they will find it more difficult to finance the fixed cost of exporting for a larger number of export destinations (ceteris paribus). Importantly, our main result regarding the role of foreign ownership in the crisis and how it interacts with financial vulnerability does not change by taking the channel of export market diversification into account.

Finally, we consider the distribution network of multinational firms that foreign-owned firms can access. Fortunately, in 2006 and 2010, our data set allows us to observe this channel, as it includes a dummy variable for whether or not the firm relies on the foreign parent's distribution network to access export markets. We start by using information in 2006 to see whether it made a difference for firms' crisis performance if they used this distribution network *before* the crisis. Column (1) in Panel C of Table 4.7 seems to suggest that this was not the case, while column (2) indicates a positive effect in more financially vulnerable firms. This is an important new result that accords well with the idea that the fixed cost of exporting can be significantly lower when firms use the distribution network of their foreign parent, rather than using their own distribution means. To shed more light on how the market access channel is related to the effect of foreign ownership working through financial vulnerability, we include in columns (3) and (4) the two relevant triple interaction terms simultaneously. Our preferred estimator confirms the stronger foreign ownership effect in the crisis for more financially vulnerable firms. However, we also find a negative and highly significant coefficient of the triple interaction term between the 'export via foreign parent' dummy, financial vulnerability, and the crisis. To understand this pattern, note that, by definition, the firm's foreign ownership dummy is always equal to one when the firm's export via foreign parent dummy is one, but not vice versa. Hence, this triple interaction effect should be interpreted as a comparison *within* the group of financially vulnerable foreign-owned firms between those firms that use their parent's distribution network and those that do not. If the market

Table 4.7: *Other channels*

	Dependent variable: Export share (*exports*/*sales*)			
Panel A: Management quality	(1)	(2)	(3)	(4)
Management quality × Crisis	0.00597**	0.00306	0.00702*	0.00584
	(0.00283)	(0.00459)	(0.00365)	(0.00716)
Management quality × FinVul × Crisis		0.00180	-0.00912	-0.0177
		(0.00861)	(0.00590)	(0.0158)
Foreign × Crisis			-0.0140	-0.0317*
			(0.0215)	(0.0184)
Foreign × FinVul × Crisis			0.0899*	0.160***
			(0.0477)	(0.0384)
Observations	15185	12488	12488	9937
R-squared	0.053	0.055	0.066	0.163
Panel B: Export market diversification				
No. international markets × Crisis	0.00574**	0.00659	0.00895*	0.0201***
	(0.00285)	(0.00481)	(0.00492)	(0.00569)
No. international markets × FinVul × Crisis		-0.00259	-0.00976	-0.0331***
		(0.00942)	(0.00999)	(0.0114)
Foreign × Crisis			-0.0124	-0.0216
			(0.0205)	(0.0151)
Foreign × FinVul × Crisis			0.0813*	0.127***
			(0.0448)	(0.0306)
Observations	12972	12488	12488	9937
R-squared	0.055	0.055	0.066	0.168
Panel C: Export via foreign parent				
Export via foreign parent × Crisis	0.00851	-0.00420	0.00710	0.0423**
	(0.0105)	(0.0129)	(0.0267)	(0.0213)
Export via foreign parent × FinVul × Crisis		0.0296**	-0.0489	-0.117***
		(0.0130)	(0.0513)	(0.0347)
Foreign × Crisis			-0.0104	-0.0324*
			(0.0262)	(0.0172)
Foreign × FinVul × Crisis			0.0973*	0.165***
			(0.0558)	(0.0302)
Observations	15628	12488	12488	9937
R-squared	0.052	0.055	0.066	0.169
Full set of interactions		Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes	Yes
Industry-year fixed effects	Yes	Yes	Yes	Yes
Propensity score				Yes

Notes: The dependent variable in all columns is the firm-specific export share. Foreign is a dummy variable for foreign ownership (equal to one if the firm is foreign owned by more than 50% and zero otherwise). FinVul measures the financial vulnerability of the firm in 2008; it is defined as the debt service-to-sales ratio (see the text for details). Crisis is a dummy variable equal to one in years 2009 to 2012, and zero otherwise. All columns include a full set of interaction terms between Foreign, FinVul, and the crisis dummy. Columns (2) to (4) include a full set of interaction terms corresponding to the various explanatory variables shown in the left-hand side of the table. Robust standard errors (in parentheses) are clustered at the firm level. *,**,*** denote significance at the 10%, 5%, 1% levels, respectively.

access channel reduces trade costs and financing needs, then the general financial advantage of foreign ownership working through the credit channel should be less important for these firms. The estimates

are in line with this conclusion.

One limitation of the analysis so far is that we have only used information on the distribution channels reported in 2006 (and hold this constant over the subsequent three years) in the above regressions. The reason is that we do not observe annual changes in the number of firms that rely on the distribution network of their foreign parent, and thus assume implicitly that those firms that used other means to export in 2006 continued to do so in the crisis years. This is obviously a strong assumption. It seems quite plausible that the use of the parent's distribution network gained importance during the crisis. As a matter of fact, our data show that the share of foreign-owned firms that used the foreign parent's distribution network to access the export market increased considerably between 2006 and 2010 from 43.1 to 54.9%. In our final set of regressions, we exploit these changes within firms from 2006 to 2010 in order to investigate whether those firms that *started* using their parent's distribution network improved their export performance in the crisis.

Table 4.8: *Export market access via foreign parent*

	Dependent variable: Export share (*exports/sales*)					
	(1)	(2)	(3)	(4)	(5)	(6)
Foreign × Crisis	-0.0355			-0.0601**	-0.0424**	-0.0295
	(0.0267)			(0.0271)	(0.0200)	(0.0223)
Foreign × FinVul × Crisis	0.122**			0.163***	0.120***	0.127***
	(0.0559)			(0.0545)	(0.0339)	(0.0425)
Export via foreign parent		0.0537**	0.00986	0.0445**	-0.00200	-0.0177
		(0.0232)	(0.0374)	(0.0221)	(0.0292)	(0.0299)
Export via foreign parent × FinVul			0.0879		0.105**	0.134***
			(0.0759)		(0.0518)	(0.0402)
Full set of interactions	Yes	Yes	Yes	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Industry-year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Propensity score						Yes
Observations	3198	2541	2021	2021	2021	1662
R-squared	0.082	0.076	0.086	0.142	0.150	0.296

Notes: The dependent variable in all columns is the firm-specific export share. Foreign is a dummy variable for foreign ownership (equal to one if the firm is foreign owned by more than 50% and zero otherwise). FinVul measures the financial vulnerability of the firm in different years as indicated at the top of each column; it is defined as the debt service-to-sales ratio (see the text for details). All columns include a full set of interaction terms between Foreign, FinVul, and the crisis dummy. For details on the propensity score reweighting estimator see the text. Robust standard errors (in parentheses) are clustered at the firm level. *,**,*** denote significance at the 10%, 5%, 1% levels, respectively.

Table 4.8 investigates the implications of increased reliance on foreign parents' distribution networks in the crisis. Column (1) shows estimates of the treatment effect of foreign ownership in the sample including only the two years 2006 and 2010, in which we directly observe how firms have accessed export markets. We find a robust positive effect of foreign ownership that is increasing in the firm's degree of financial vulnerability also in this drastically reduced sample. Column (2) shows that firms which started exporting via their parent's distribution network in the crisis had a significantly higher export share than firms that exported through other channels. However, column (3) seems to suggest that this effect is not stronger among financially vulnerable firms. Importantly, the samples in these and the subsequent regressions are restricted to exporting firms, and the coefficients of the export via foreign parent dummy are identified only through within-firm variation between 2006 and 2010 (i.e., through

firms changing their distribution channels).

We provide an integrated assessment of both the credit channel and the market access channel in Columns (4) to (6) of Table 4.8. Column (4) confirms our main result that foreign ownership promoted exports more strongly in financially vulnerable firms, and on top of this, there is a positive and significant effect stemming from the market access channel. Finally, we show that there is a strong interaction between the market access channel and the firm's financial vulnerability in columns (6) and (7), using the fixed effects and the propensity score reweighting estimator, respectively. After controlling for the differential effect of foreign ownership by firms' financial vulnerability, we find that using the parent's distribution network for exporting was also particularly conducive to exporting among these same, financially vulnerable firms. This is intuitive because exporting through the foreign parent's distribution network may be associated with lower trade costs than exporting by own distribution means. Therefore, more financially vulnerable firms benefited relatively more from doing so in the credit crunch. Our data strongly suggest that foreign-owned firms had two options to alleviate credit constraints in the crisis: They could borrow on the internal or external capital market abroad, or they could use the parent's distribution network in order to reduce trade costs and thereby their financing needs. Both of these channels seem to have been important for foreign-owned firms in Spain during the global financial crisis.[36]

4.5 Conclusions

Firms belonging to multinational corporations are generally much more active in exporting. This paper has demonstrated that foreign-owned firms in Spain further magnified their competitive edge on export markets when the global financial crisis hit the economy. We have extended the heterogeneous firms model by Manova (2013) to show that a financial advantage of foreign ownership due to access to internal and external capital markets abroad can rationalize this pattern. Our rich panel data on Spanish firms allows us to pin down this financial advantage after controlling for unobservable firm-specific effects, various confounding factors, and several other benefits of foreign ownership. In line with the model's predictions, we find that foreign ownership improved export performance in the crisis especially among small and financially vulnerable firms, which were highly indebted at the onset of the crisis.

Our contribution formalizes the hypothesis put forward by Manova et al. (2015) that internal capital markets within multinational firms promote firm exports, and it provides strong support for this hypothesis by exploiting the credit crunch in the financial crisis. We further contribute to the literature novel evidence on a second dimension of the financial advantage of foreign ownership. Our analysis suggests that multinational firms can help their affiliates to circumvent credit constraints not only through internal capital markets, but also by providing access to their distribution networks and thereby potentially reducing trade costs that need to be financed. We provide first direct evidence on this second channel and show how the effects of market access and internal capital markets interact within multinational firms.

Our findings point to important firm-level complementarities between FDI and trade, which gained particular importance in times of crisis. It remains an important task for future research to investigate whether foreign ownership can have similar stabilizing effects on exports in other countries and crises. Multinational firm activity clearly has the potential to mitigate the impact of future financial crises and may help to prevent another trade collapse.

[36]Note that this conclusion is in line with our previous analysis of this channel in Panel C of Table 4.7, although we used a different type of variation in the data there. Firms that *already* relied on the parent's distribution network before the crisis (used for identification in Table 4.7) benefited less from internal capital markets, because they had lower financing needs. Those that *started* using the foreign parent's means of distribution (in Table 4.8) benefited in particular if they were financially vulnerable.

Bibliography

Ahn, JaeBin, "A Theory of Domestic and International Trade Finance," IMF Working Papers 11/262, International Monetary Fund 2011.

_ , Mary Amiti, and David E. Weinstein, "Trade Finance and the Great Trade Collapse," *American Economic Review*, 2011, *101* (3), 298–302.

Alessandria, George, Joseph P. Kaboski, and Virgiliu Midrigan, "The Great Trade Collapse of 2008-09: An Inventory Adjustment?," Technical Report 16059, National Bureau of Economic Research 2010.

Alfaro, Laura and Maggie Xiaoyang Chen, "Surviving the Global Financial Crisis: Foreign Ownership and Establishment Performance," *American Economic Journal: Economic Policy*, 2012, *4* (3), 30–55.

Almunia, Miguel, Pol Antràs, David Lopez-Rodriguez, and Eduardo Morales, "Venting Out: Exports during a Domestic Slump," 2017. Harvard University, mimeo.

Alvarez, Roberto and Holger Görg, "Multinationals as Stabilizers? Economic Crisis and Plant Employment Growth," IZA Discussion Papers 2692, Institute for the Study of Labor (IZA) March 2007.

Amiti, Mary and David E. Weinstein, "Exports and Financial Shocks," *The Quarterly Journal of Economics*, 2011, *126* (4), 1841–1877.

Antràs, Pol and C. Fritz Foley, "Poultry in Motion: A Study of International Trade Finance Practices," *Journal of Political Economy*, 2015, *123* (4), 853–901.

_ , Mihir A. Desai, and C. Fritz Foley, "Multinational Firms, FDI Flows, and Imperfect Capital Markets," *The Quarterly Journal of Economics*, 2009, *124* (3), 1171–1219.

Auboin, Marc, "Boosting the Availability of Trade Finance in the Current Crisis: Background Analysis for a Substantial G20 Package," CEPR Policy Insight 35, CEPR 2009.

Baldwin, Richard, "Introduction: The Great Trade Collapse: What Caused It and What Does It Mean?," in Richard Baldwin, ed., *The Great Trade Collapse: Causes, Consequences and Prospects*, 2009.

_ and Simon Evenett, "Introduction and Recommendations for the G20," in Richard Baldwin and Simon Evenett, ed., *The Collapse of Global Trade, Murky Protectionism, and the Crisis: Recommendations for the G20*, London: CEPR, 2009, chapter 1, pp. 1–9.

Beck, Thorsten, "Financial Development and International Trade: Is There a Link?," *Journal of International Economics*, 2002, *57* (1), 107–131.

_ , Asli Demirgüç-Kunt, Luc Laeven, and Ross Levine, "Finance, Firm Size, and Growth," *Journal of Money, Credit and Banking*, 2008, *40* (7), 1379–1405.

Behrens, Kristian, Gregory Corcos, and Giordano Mion, "Trade Crisis? What Trade Crisis?," *Review of Economics and Statistics*, 2013, *95* (2), 702–709.

Bems, Rudolfs, Robert C. Johnson, and Kei-Mu Yi, "The Great Trade Collapse," *Annual Review of Economics*, 2013, *5* (1), 375–400.

Berman, Nicolas and Jérôme Héricourt, "Financial factors and the margins of trade: Evidence from cross-country firm-level data," *Journal of Development Economics*, 2010, *93* (2), 206–217.

Bricongne, Jean-Charles, Lionel Fontagné, Guillaume Gaulier, Daria Taglioni, and Vincent Vicard, "Firms and the Global Crisis: French Exports in the Turmoil," *Journal of International Economics*, 2012, *87* (1), 134–146.

Caves, Douglas W, Laurits R Christensen, and Walter Diewert, "The Economic Theory of Index Numbers and the Measurement of Input, Output, and Productivity," *Econometrica*, 1982, *50* (6), 1393–1414.

Chaney, Thomas, "Liquidity Constrained Exporters," *Journal of Economic Dynamics and Control*, 2016, *72* (C), 141–154.

Chor, Davin and Kalina Manova, "Off the Cliff and Back? Credit Conditions and International Trade during the Global Financial Crisis," *Journal of International Economics*, 2012, *87* (1), 117–133.

Delgado, Miguel A., Jose C. Fari nas, and Sonia Ruano, "Firm Productivity and Export Markets: A Non-parametric Approach," *Journal of International Economics*, 2002, *57* (2), 397–422.

Desai, Mihir A, C Fritz Foley, and James R Hines, "A Multinational Perspective on Capital Structure Choice and Internal Capital Markets," *The Journal of Finance*, 2004, *59* (6), 2451–2487.

Desai, Mihir A., C. Fritz Foley, and Kristin J. Forbes, "Financial Constraints and Growth: Multinational and Local Firm Responses to Currency Depreciations," *Review of Financial Studies*, November 2008, *21* (6), 2857–2888.

DiNardo, John, Nicole M. Fortin, and Thomas Lemieux, "Labor Market Institutions and the Distribution of Wages, 1973-1992: A Semiparametric Approach," *Econometrica*, 1996, *64* (5), 1001–1044.

Eaton, Jonathan, Samuel Kortum, Brent Neiman, and John Romalis, "Trade and the Global Recession," *American Economic Review*, 2016, *106* (11), 3401–38.

Egger, Peter and Christian Keuschnigg, "Corporate Finance and Comparative Advantage," Technical Report 2009-04, Department of Economics, University of St. Gallen 2009.

_ , _ , **Valeria Merlo, and Georg Wamser**, "Corporate Taxes and Internal Borrowing within Multinational Firms," *American Economic Journal: Economic Policy*, 2014, *6* (2), 54–93.

Eppinger, Peter, Nicole Meythaler, Marc-Manuel Sindlinger, and Marcel Smolka, "The Great Trade Collapse and the Spanish Export Miracle: Firm-level Evidence from the Crisis," *The World Economy*, forthcoming.

Feenstra, Robert C., Zhiyuan Li, and Miaojie Yu, "Exports and Credit Constraints under Incomplete Information: Theory and Evidence from China," *The Review of Economics and Statistics*, 2014, *96* (3), 729–744.

Foley, C. Fritz and Kalina Manova, "International Trade, Multinational Activity, and Corporate Finance," *Annual Review of Economics*, 2015, *7* (1), 119–146.

Garicano, Luis and Claudia Steinwender, "Survive Another Day: Using Changes in the Composition of Investments to Measure the Cost of Credit Constraints," *The Review of Economics and Statistics*, 2016, *98* (5), 913–924.

Gertler, M. and S. Gilchrist, "Monetary Policy, Business Cycles, and the Behavior of Small Manufacturing Firms," *Quarterly Journal of Economics*, 1994, *109* (2), 309–340.

Gertner, Robert H, David S Scharfstein, and Jeremy C Stein, "Internal versus External Capital Markets," *The Quarterly Journal of Economics*, 1994, *109* (4), 1211–30.

Greenaway, David, Alessandra Guariglia, and Richard Kneller, "Financial Factors and Exporting Decisions," *Journal of International Economics*, 2007, *73* (2), 377–395.

Guiso, L., P. Sapienza, and L. Zingales, "Does Local Financial Development Matter?," *Quarterly Journal of Economics*, 2004, *119* (3), 929–969.

Kletzer, Kenneth and Pranab Bardhan, "Credit Markets and Patterns of International Trade," *Journal of Development Economics*, 1987, *27* (1-2), 57–70.

Levchenko, Andrei A., Logan Lewis, and Linda L. Tesar, "The Collapse of International Trade During the 2008-2009 Crisis: In Search of the Smoking Gun," Technical Report 592, Research Seminar in International Economics, University of Michigan 2009.

Manova, Kalina, "Credit Constraints, Equity Market Liberalizations and International Trade," *Journal of International Economics*, 2008, *76* (1), 33–47.

_ , "Credit Constraints, Heterogeneous Firms, and International Trade," *Review of Economic Studies*, 2013, *80* (2), 711–744.

_ **, Shang-Jin Wei, and Zhiwei Zhang**, "Firm Exports and Multinational Activity under Credit Constraints," *Review of Economics and Statistics*, 2015, *97* (3), 574–588.

Melitz, Marc J., "The Impact of Trade on Intra-Industry Reallocations and Aggregate Industry Productivity," *Econometrica*, 2003, *71* (6), 1695–1725.

Minetti, Raoul and Susan Chun Zhu, "Credit Constraints and Firm Export: Microeconomic Evidence from Italy," *Journal of International Economics*, 2011, *83* (2), 109–125.

Muûls, Mirabelle, "Exporters, Importers and Credit Constraints," *Journal of International Economics*, 2015, *95* (2), 333–343.

Paravisini, Daniel, Veronica Rappoport, Philipp Schnabl, and Daniel Wolfenzon, "Dissecting the Effect of Credit Supply on Trade: Evidence from Matched Credit-Export Data," *The Review of Economic Studies*, 2014.

Schmidt-Eisenlohr, Tim, "Towards a Theory of Trade Finance," *Journal of International Economics*, 2013, *91* (1), 96–112.

Wang, Jian and Xiao Wang, "Benefits of Foreign Ownership: Evidence from Foreign Direct Investment in China," *Journal of International Economics*, 2015, *97* (2), 325–338.

4.A Mathematical Appendix

4.A.1 Proof of Proposition 4.1

The left-hand side (LHS) of equation (4.6) is increasing in the price in the relevant range $p_j^{**}(a) \in [p_j^*(a), p_L(a)]$:

$$\begin{aligned}\frac{\partial LHS}{\partial p_j^{**}(a)} &= (1-\sigma)p_j^{**}(a)^{-\sigma} + \sigma p_j^{**}(a)^{-1-\sigma}\tau b_j a\left[1-d(a)+(1+r_D)d(a)/\lambda\right] \\ &\geq \left((1-\sigma)\frac{\sigma}{\sigma-1}\left[1-d(a)+(1+r_D)d(a)\right]\tau b_j a + \sigma\tau b_j a\left[1-d(a)+(1+r_D)d(a)/\lambda\right]\right) p_j^{**}(a)^{-1-\sigma} \\ &= (1+r_D)d(a)\left(\frac{1-\lambda}{\lambda}\right)\sigma\tau b_j a p_j^{**}(a)^{-1-\sigma} > 0\end{aligned}$$

where the first inequality follows from plugging in the lowest optimal price $p_j^*(a)$ from equation (4.3) for $p_j^{**}(a)$ and the second follows from the fact that all parameters and prices are non-negative and $\lambda \in (0,1)$. Trivially, the right-hand side (RHS) of equation (4.6) is independent of the price.

To determine the impact of a deterioration in capital market efficiency on the optimal price of constrained exporters, we take the derivatives of LHS and RHS with respect to the capital market efficiency parameter λ:

$$\frac{\partial LHS}{\partial \lambda} = p_j^{**}(a)^{-\sigma}\tau b_j a(1+r)d(a)/\lambda^2 > 0, \quad \text{and} \quad \frac{\partial RHS}{\partial \lambda} = -\frac{f_X b_j(1+r)d(a)}{\theta_j Y P_j^{\sigma-1}\lambda^2} < 0,$$

$$\frac{\partial^2 LHS}{\partial \lambda \partial r} = p_j^{**}(a)^{-\sigma}\tau b_j a d(a)/\lambda^2 > 0, \quad \text{and} \quad \frac{\partial^2 RHS}{\partial \lambda \partial r} = -\frac{f_X b_j d(a)}{\theta_j Y P_j^{\sigma-1}\lambda^2} < 0.$$

Hence, a decrease in λ increases the optimal price for constrained exporters and this effect is stronger for higher interest rates. Due to $r_F < r_D$, the effect will be stronger for domestically owned firms compared to foreign-owned firms (while both types of firms are equally affected for $r_F \geq r_D$). Also, the effect is stronger for more financially vulnerable firms that need to finance a larger share $d(a)$ of exporting costs:

$$\frac{\partial^3 LHS}{\partial \lambda \partial r \partial d(a)} = p_j^{**}(a)^{-\sigma}\tau b_j a/\lambda^2 > 0, \quad \text{and} \quad \frac{\partial^3 RHS}{\partial \lambda \partial r \partial d(a)} = -\frac{f_X b_j}{\theta_j Y P_j^{\sigma-1}\lambda^2} < 0.$$

The comparative statics for quantities $q_j^{**}(a)$ and export revenues $p_j^{**}(a)q_j^{**}(a)$ follow from these results. They are the opposite of the price effects because quantities sold are inversely related to the price $p_j^{**}(a)$ via the demand schedule in equation (4.1). Thus, both export quantities and export revenues (i) decrease in response to a decrease in λ, (ii) they decrease more for domestically owned firms facing the higher interest rate, and (iii) this differential effect is stronger for more financially vulnerable firms.

Since firms finance all domestic activities internally and the foreign-owned firms borrow at the unaffected foreign interest rate $r_F < r_D$, the comparative statics regarding the export revenues translate directly into the export intensity *exports/sales*= *exports*/(*exports* + *domestic sales*) because domestic sales are unaffected by credit market conditions. This completes the proof of Proposition 4.1.

4.A.2 Modeling the financial crisis as an interest rate shock

In this appendix, we pursue an alternative modeling of the financial crisis as an increase in the domestic interest rate r_D relative to the foreign interest rate r_F. To keep the analysis simple, we model the financial crisis in this scenario as an increase in r_D, while r_F remains constant. The smaller drop in the foreign interest rate, at which foreign-owned firms can borrow, may be rationalized by a combination of three facts: the sheer size of most multinational firms, the large liquidity of the world capital market, and the geographical asymmetry of the crisis. In general, large multinational firms have multiple sources of financing, could offer substantial collateral also in the crisis, and continued to have access to thicker foreign capital markets. Most importantly, the financial crisis had an asymmetric effect across world regions, leaving banks in some countries (notably in East Asia and Eastern Europe) hardly affected at all. Arguably, multinational firms, and hence their affiliates, are more likely to have access to credit from such banks. They could tap the source of finance for which the interest rate remained lowest in the crisis.

For $r_F < r_D$, all foreign-owned firms access the foreign capital market both before and after the shock. In this case, we only need to consider how an increase in the real interest rate r_D affects the export intensity of domestically owned firms: $\partial(\textit{exports}/\textit{sales})/\partial r_D$. If the sign of this derivative is negative, foreign-owned firms will *ceteris paribus* maintain a higher export share in the crisis.

To determine the impact of an increase in r_D on the optimal price of constrained exporters, we take the derivatives of LHS and RHS with respect to r_D. It is easy to see from equation (4.6), that LHS is decreasing in r_D, while RHS is increasing in r_D:

$$\frac{\partial LHS}{\partial r_D} = -p_j^{**}(a)^{-\sigma}\tau b_j a d(a)/\lambda < 0, \quad \text{and} \quad \frac{\partial RHS}{\partial r_D} = \frac{f_X b_j d(a)}{\theta_j Y P_j^{\sigma-1}\lambda} > 0.$$

Constrained exporters will increase their optimal price in response to an increase in r_D and this differential effect is stronger for more financially vulnerable firms:

$$\frac{\partial^2 LHS}{\partial r_D \partial d(a)} = -p_j^{**}(a)^{-\sigma}\tau b_j a/\lambda < 0, \quad \text{and} \quad \frac{\partial^2 RHS}{\partial r_D \partial d(a)} = \frac{f_X b_j}{\theta_j Y P_j^{\sigma-1}\lambda} > 0.$$

It follows by a similar argument as in Section 4.A.1 that export quantities, export revenues, and export intensities for constrained exporters decrease in response to an increase in r_D. Foreign-owned firm's exports, in contrast, are unaffected by this shock as long as they can borrow at $r_F < r_D$. The resulting, differential effect on export revenues is larger for more financially vulnerable firms, in analogy to Proposition 4.1.

Interestingly, the optimal prices and quantities for unconstrained exporters also depend on the interest rate via the investor's participation constraint. This can be easily seen from the optimality conditions (4.3) and (4.4). Thus, if the r_D increases in the crisis, also *un*constrained *domestic* exporters have to reduce their export revenues and intensities. This prediction differs from the scenario where λ changes, and it implies a differential effect of the financial crisis across domestic and foreign-owned firms' export intensities also among the more productive firms.

Chapter 5

Optimal Ownership and Firm Performance: Theory and Evidence from China's FDI Liberalization*

Multinational firms organize the global production of highly specialized inputs in the face of inefficiencies arising from contractual frictions. An open question is by how much optimal firm organization can reduce these inefficiencies. This paper provides a first quantification of the performance gains from optimal ownership within multinationals. We exploit a unique policy change in China, which liberalized its severe restrictions on foreign ownership after its WTO accession. To guide our empirical analysis, we incorporate ownership restrictions into the property-rights theory of the multinational firm. Using a large panel of Chinese manufacturing firms, we show that increased foreign ownership after the liberalization induced changes in firms' input ratios and output in line with the theory. To quantify the gains from optimal ownership restructuring, we analyze the performance of foreign-acquired firms in liberalized industries in a difference-in-differences model. By choosing suitable control groups and applying propensity score reweighting, we carefully exclude other possible explanations. We find that optimal ownership restructuring led to firm-level output gains of up to 34% after two years, which increased further over the medium term.

*This chapter is based on joint work with Hong Ma.

5.1 Introduction

Multinational enterprises (MNEs) organize the production of highly specialized inputs around the globe in the face of contractual frictions. Through the optimal choice of their ownership structure, MNEs minimize the inefficiencies arising from these contractual frictions (Antràs, 2015). An open question is by how much firms' organizational decisions can reduce these inefficiencies. The goal of this paper is to quantify the gains in firm performance that can be achieved by the optimal allocation of ownership rights. The key challenge for such a quantification exercise lies in identifying an exogenous change from a constrained ownership structure towards the optimum. We overcome this challenge by exploiting a unique policy change in China, which can be considered exogenous to the individual firm. China has historically maintained severe restrictions on foreign direct investment (FDI), which prevented foreign MNEs from choosing the optimal ownership share in their Chinese suppliers. While FDI was entirely prohibited in some industries, others faced explicit upper bounds on foreign ownership shares. These restrictions were abolished to a large extent in 2002, after China's WTO accession. The massive FDI liberalization allowed MNEs to reoptimize their organizations and hence provides a unique opportunity for quantifying the efficiency gains from optimal ownership.

To guide our empirical analysis, we provide a theoretical model of MNEs that delivers sharp predictions for the effects of FDI liberalization on ownership and performance. The model extends the property-rights theory (PRT) of the multinational firm due to Antràs (2003) and Antràs and Helpman (2004) by allowing for continuous ownership shares, which enables us to introduce the Chinese FDI policy as an upper bound on the foreign ownership share.[1] In the model, a foreign firm and its Chinese supplier both contribute relationship-specific inputs to a joint production process in China. The firm provides capital, while the supplier is responsible for hiring labor. Due to contractual frictions, both parties face a risk of hold-up and hence underinvest into these inputs. By choosing a higher ownership share in its supplier, the firm improves its bargaining position and hence increases its own capital investments, but it also aggravates the supplier's underinvestment into labor. The optimal choice of ownership solves this trade-off, and it depends crucially on the importance of capital relative to labor in production. A policy restriction on the foreign ownership share distorts the investment incentives and worsens the overall inefficiency. FDI liberalization then induces a restructuring process towards higher foreign ownership shares among the previously constrained firm pairs, and it may additionally trigger entry of new foreign-owned suppliers. The model predicts that this optimal ownership restructuring raises the capital-labor ratio in production, because the foreign firm's investment incentives improve, while the Chinese supplier's incentives for optimal labor provision diminish. These adjustments generate overall efficiency gains and can result in higher firm output.

Our empirical analysis uses panel data on more than half a million manufacturing firms in China over the period from 1998 to 2007, combined with information on the FDI policy from official sources. The analysis proceeds in three steps. First, we confirm that general patterns of foreign ownership in unregulated industries are consistent with key predictions of our PRT model. Second, we verify that the Chinese FDI restrictions in general, and the major liberalization in 2002 in particular, had the expected direct effect on foreign ownership shares. Third, we exploit the inflow of FDI after liberalization in our main econometric analysis to estimate the performance effects of optimal ownership restructuring in a difference-in-differences fixed effects model for the years 2001 and 2003. We define the 'treated' (or

[1] The PRT goes back to the seminal contributions by Grossman and Hart (1986) and Hart and Moore (1990).

'foreign-acquired') firms in this model as firms increasing their foreign ownership share (across different thresholds) after the FDI policy in their industry was liberalized. To account for selection of firms by foreign investors and selection of liberalized industries by the Chinese government, we choose foreign-owned firms within the same industries as a preferred control group,[2] and we further augment the model with propensity score reweighting based on initial characteristics.

The main findings of our analysis are summarized as follows. The increase in foreign ownership after liberalization induced a significant rise in the capital-labor ratio and output in the foreign-acquired firms between 2001 and 2003. These findings are in line with the predictions from our theoretical model: Optimal ownership restructuring towards higher foreign ownership increases the investment incentives of the foreign firm, but it reduces the Chinese firm's incentives to contribute to the production process. This mechanism can rationalize the increase in the capital-labor ratio, provided that foreign owners contribute to their affiliates' capital stock, while hiring labor is typically the local firm's responsibility (see, for instance, the evidence discussed in Antràs, 2003). Since the positive incentive effect dominates for a move towards optimal ownership, overall efficiency and firm output increases. Our findings suggest that the performance gains from optimal ownership can be substantial. We estimate output gains of up to 34% after two years, with the exact magnitude depending on the ownership threshold defining foreign acquisitions. Furthermore, the output effect tends to increase over the medium term and can be more than twice as large after five years. Over the medium term, the ownership restructuring also leads to a significant increase in total factor productivity.

Our conclusions are strengthened by an investigation of several alternative control groups in the difference-in-differences model. We first verify that our estimates are not confounded by potential negative spillovers from FDI (e.g. through increased competition) by choosing two control groups of firms that are unlikely to compete directly with the treated firms: (i) foreign-owned firms in unregulated industries, which produce very different products, and (ii) firms in liberalized industries that export the majority of their sales and hence mainly compete on world markets. We also compare foreign-acquired firms to domestic-acquired firms to show that the effects cannot be explained by dynamic adjustments that might occur for any change in ownership. Finally, in the most ambitious identification strategy, we choose foreign-acquired firms in unregulated industries as a control group. These firms constitute an exceptionally suitable control group because they are contemporaneously selected by foreign investors, and should therefore share relevant unobservable characteristics with the treated firms and at the same time experience all the benefits related to foreign acquisitions per se. The only salient difference between the treated firms and this control group is the fact that foreign ownership in the treated firms was previously restricted, aggravating the distortions arising from contractual frictions, so these firms moved from the restricted to the optimal ownership share.[3] Our main findings are strongly confirmed in this analysis, supporting the view that optimal ownership restructuring is responsible for the positive effects on the capital-labor ratio and firm output.

We build on and contribute to the literature investigating MNEs' organizational decisions in the presence of contractual frictions. In his seminal work, Antràs (2003) incorporated the PRT into a model of international trade, which has since been extended along several dimensions to describe MNEs' activities

[2] Some firms within the liberalized industries are already foreign-owned before the reform because they produce different products that are not specifically regulated by the FDI policy (see Section 5.3.2).

[3] Note that we also account for initial observable differences in firm and industry characteristics through propensity score reweighting in all of these regressions.

(see Antràs, 2015, for a synopsis).[4] We generalize this theoretical framework by allowing for continuous ownership shares, in order to account for the existence of international joint ventures (JVs) and to investigate the effects of restrictions on foreign ownership shares in China. A similar model extension is developed by Bircan (2014), though with a very different focus on uncertainty and dynamic ownership patterns, and without policy restrictions on ownership. Eppinger and Kukharskyy (2017) adopt an alternative way of modeling JVs to study the relationship between contracting institutions and firm boundaries.

The predictions of the PRT have been investigated empirically by a growing literature using US intra-firm trade data at the industry level and, more recently, also using firm-level data.[5] As summarized in Antràs and Yeaple (2014), these studies have typically examined the empirical relationship between the prevalence of intra-firm trade and proxies for the relative importance of the inputs provided by both parties (the so-called headquarter intensity) and firm productivity (or productivity dispersion), and generally found patterns consistent with the PRT. Note that the existing firm-level evidence derives almost exclusively from data on firms' headquarters in developed countries.[6] This may be problematic if the joint production processes actually take place in the sourcing countries, as suggested by the high prevalence of processing trade in the case of China (Feenstra and Hanson, 2005). In this case, the characteristics of these production processes (headquarter intensity and productivity), which are the key determinants of the organizational decision in the model, cannot be well approximated using only data on the investor. Our paper complements this literature in an important way by providing firm-level evidence on the supplier side of the global sourcing relationship.

In his recent book, Antràs (2015) concludes that a major gap in the existing literature on the organization of MNEs is the lack of a quantitative assessment of the performance effects of MNEs' organizational choices. He suggests three possible avenues for future research towards closing this gap: (i) making better use of the existing firm-level data, (ii) exploiting exogenous variation due to technological or policy changes, and (iii) moving the empirical analysis closer to the theoretical structure. This paper makes progress on all three fronts. First, we exploit a firm-level dataset from China that includes precise information on the foreign ownership share and has not been used in this literature before. Second, the key novelty of our empirical approach is that we exploit China's FDI liberalization as an exogenous change in the set of ownership forms available to multinational firms. Third, we bring the theory closer to the data by incorporating the empirically important feature of partial ownership into the model and deriving clear predictions on the policy impact. In combination, these features allow us to answer a key question that has remained open to this date: "How do the organizational decisions of multinational firms shape firm-level performance?" (Antràs, 2015, p. 261)

Numerous studies have previously investigated the effects of foreign acquisitions on various dimensions of firm performance.[7] In this literature, the combination of difference-in-differences estimation

[4] Important extensions considering firm heterogeneity and partial contractibility are developed by Antràs and Helpman (2004, 2008). Antràs and Chor (2013), Alfaro et al. (2015), and Schwarz and Suedekum (2014) consider the organization of multiple input suppliers. Contractual frictions are also crucial in transaction-cost theories of MNEs, as developed by Grossman and Helpman (2002, 2003, 2005).

[5] Studies using US industry-level data include Antràs (2003), Bernard et al. (2010), Nunn and Trefler (2008, 2013), and Yeaple (2006). Feenstra and Hanson (2005) use highly disaggregated customs data to investigate the role of ownership and control over input purchases in Chinese processing trade. Firm-level studies on the organization of global sourcing include: Tomiura (2007) for Japan, Corcos et al. (2013) and Defever and Toubal (2013) for France, Federico (2010, 2012) for Italy, and Kohler and Smolka (2011, 2014, 2015) for Spain.

[6] Recent exceptions using data on subsidiaries are Fernandes and Tang (2012) on export processing by Chinese firms, Bircan (2014) on foreign-owned firms in Turkey, and Kukharskyy (2016) as well as Eppinger and Kukharskyy (2017) examining firm pairs around the globe.

[7] Early contributions in this vein include Harris and Ravenscraft (1991) and Swenson (1993).

with firm fixed effects and propensity score matching or reweighting, as applied in this paper, has become standard practice for estimating average treatment effects.[8] Several of these studies have identified positive effects of foreign acquisitions on firm productivity or output, though the evidence remains inconclusive and some papers find no significant performance gains (notably Wang and Wang, 2015, for China). Note that the findings in this literature may be explained by a variety of benefits which foreign ownership can bring to the firm, such as technology transfers, improved financial conditions, or foreign market access. We would like to emphasize that our contribution differs substantially from these studies in terms of its fundamental objective, the empirical setup, and the main conclusions. While the existing studies aim to identify general benefits of the 'foreignness' of the owner, our goal is to estimate the particular effects of ownership restructuring towards an optimal allocation of property rights.[9] We merely exploit the fact that foreign owners were constrained by the FDI policy, but we are not interested in the effect of their foreignness per se. By comparing foreign-acquired firms to *other foreign-owned firms* (within liberalized industries) or even to *other foreign-acquired firms* (in unregulated industries), we control for the various other benefits of foreign ownership and can clearly attribute our findings to optimal ownership restructuring.

Our analysis contributes to evaluating the Chinese policy towards inward FDI. This is an important policy field, since China has become the world's factory (Zhang, 2006) and one of the top destinations for FDI (UNCTAD, 2013) over the past decades. But interestingly, China also remains one of the countries with the most restrictive policies towards inward FDI in the 2010s, despite recent steps towards liberalization (Kalinova et al., 2010). Two studies have previously examined the major reform of the Chinese FDI policy in 2002, though with different objectives than our paper. Sheng and Yang (2016) show that the liberalization increased product variety in Chinese exports. Lu et al. (2017) examine possible spillovers from FDI induced by the liberalization and find some evidence for negative competition and positive agglomeration effects in domestic firms. However, their analysis excludes the foreign-acquired firms, which are naturally the main beneficiaries of the reform. Our analysis complements this assessment of the policy and arrives at more favorable conclusions, which suggest significant potential for efficiency gains from further liberalization of FDI.

The efficiency gains from optimal ownership identified in this paper are relevant well beyond MNEs' activities in China. Our results are likely to be applicable to other developing and emerging economies, which receive the bulk of (cost-seeking) vertical FDI and where contracting institutions are often weak. Also, foreign ownership restrictions are still commonplace in the twenty-first century, even among developed countries, as data from the OECD reveals.[10] Since MNEs account for a quarter of world GDP and participate in 80% of world trade in 2010 (UNCTAD, 2011, 2013), liberalizing foreign ownership restrictions can have sizable effects on global welfare. Our findings have potentially also more general implications for other types of restrictions on ownership within countries. A case in point is state involvement in firm ownership, which might deter necessary private investments.[11] While such existing

[8] Examples include Arnold and Javorcik (2009), Girma and Görg (2007a,b), Guadalupe et al. (2012), and Wang and Wang (2015).

[9] Chen (2011) shows explicitly how the performance of acquired US firms depends on the country of the investor, and Wang and Wang (2015) compare foreign to domestic acquisitions in China to identify the effects of foreign ownership.

[10] The 'foreign equity restrictions' component of the OECD's FDI restrictiveness index provides data for 62 countries over the years 1997 to 2016 (see Kalinova et al., 2010; OECD, 2016). Over this period, we observe a decrease in the index on average, but all countries included in the database maintain at least some restrictions on foreign equity in 2016. China's FDI liberalization caused a decrease in the index from around 0.38 to 0.22 (where 0 reflects no equity restrictions and 1 represents an economy closed to FDI), but it still ranks third in 2016 after the Philippines and Indonesia.

[11] See Dollar and Wei (2007) and Berkowitz et al. (2017) for evidence on the inferior performance of China's state-owned enterprises.

restrictions are often motivated by anti-trust or security considerations, our analysis suggests that these benefits should be carefully weighted against the costs of aggravated underinvestment in the presence of contractual frictions. We view our contribution as a first step towards quantifying the relative importance of different sources of inefficiencies that depress the productivity of developing countries. Our findings suggest that contractual frictions and distorted ownership decisions are likely to play an important role in explaining the greater misallocation of resources in China relative to developed countries identified by Hsieh and Klenow (2009).

The remaining paper is structured as follows. Section 5.2 presents our property-rights theory of the multinational firm with partial ownership and derives testable predictions for the effects of liberalizing foreign ownership restrictions. Section 5.3 reviews the policy background and presents our firm data. The empirical analysis proceeds in three steps: Section 5.4 shows that the patterns of foreign ownership prevailing in China are generally in line with the theory. In Section 5.5, we verify the effectiveness of the Chinese FDI policy in terms of its direct impact on foreign ownership. Section 5.6 presents the core part of the empirical analysis; it introduces the empirical methodology and discusses the estimated effects of optimal ownership restructuring on firm performance in detail. Section 5.7 concludes.

5.2 A Property-Rights Theory of the Multinational Firm under Foreign Ownership Restrictions

To motivate and guide our empirical analysis, this section presents a theoretical model of the multinational firm based on the seminal property-rights theory (PRT) due to Grossman and Hart (1986) and Hart and Moore (1990). Our approach generalizes the framework developed by Antràs (2003) and Antràs and Helpman (2004) to allow for partial ownership. This setup enables us to incorporate the restrictions on foreign ownership imposed by the Chinese FDI policy explicitly into the model, and to study the consequences of FDI liberalization for ownership and performance.

5.2.1 General setup

We examine a world of two countries: China and Foreign. For simplicity, we assume that all consumers and firms F producing final goods reside in Foreign, while all intermediate input suppliers (manufacturers) M reside in China. Each firm pair, consisting of one firm F and one supplier M, is indexed by i, and industries are indexed by j. All consumers share identical constant elasticity of substitution (CES) preferences over varieties of the final good. These preferences imply that the monopolistically competitive firms face iso-elastic demand for each variety of the final good:

$$p_j(i) = A_j \, y_j(i)^{\alpha-1}, \tag{5.1}$$

where $1/(1-\alpha)$ (with $0 < \alpha < 1$) is the elasticity of substitution between varieties and A_j is an industry-specific demand shifter (depending on aggregate expenditure and the industry's share in aggregate expenditure).

Each firm F requires one particular intermediate input $x_j(i)$ that is sourced from a supplier M in China.[12] As in Antràs (2003), we assume that the input is costlessly shipped to Foreign and used there

[12] For simplicity, we abstract from the location choice, which Antràs and Helpman (2004) analyze explicitly, by assuming that China has already been selected as the preferable production location, e.g. due to an exogenous cost advantage.

to produce the final good at zero additional cost: $y_j(i) = x_j(i)$.[13] The intermediate input is produced by M in China using the following Cobb-Douglas technology:

$$x_j(i) = \theta(i) \left(\frac{K_j(i)}{\eta} \right)^{\eta} \left(\frac{L_j(i)}{1-\eta} \right)^{1-\eta}, \tag{5.2}$$

where the inputs are capital $K_j(i)$, provided by the foreign firm, and labor $L_j(i)$, provided by the Chinese supplier, and the headquarter intensity is denoted by $\eta \in (0,1)$. These assumptions on the joint production process follow Antràs (2003), who cites evidence from several countries that cost sharing with suppliers is much more prevalent in capital compared to labor investments.[14] We further assume that the two parties' investments into the inputs $K_j(i)$ and $L_j(i)$ are fully relationship-specific because they need to be customized to the production process, and the investments are non-contractible because they cannot be verified by a third party.[15] The price of capital r and the wage rate w are determined in competitive factor markets in China and hence exogenous to F and M. The parameter $\theta(i)$ denotes the total factor productivity (TFP) of the firm pair.

Following Antràs and Helpman (2004), we assume that production relationships differ in their productivity $\theta(i)$. To enter the market, potential firms have to incur fixed entry costs in order to draw $\theta(i)$ from a known distribution $G(\theta(i))$, as in Melitz (2003). Upon entry, F obtains knowledge of $\theta(i)$ and decides, in the first stage of the game, whether to exit the market or start producing. Production requires a supplier in China, so if it starts producing, F must simultaneously decide on the organizational form of the production relationship with M, which is summarized by its ownership share $s \in [0,1]$ in M. When choosing s, the firm faces a trade-off involving the underinvestment problems arising from incomplete contracts, which are explained below, and the fixed organizational costs associated with this choice. To obtain a homothetic cost function (as in Antràs, 2003), we assume that in order to start production, the firm has to pay $r^{\eta} w^{1-\eta} f(s)$, where we leave the functional form of the fixed organizational costs $f(s)$ open for now. F then offers a production relationship under the chosen organizational form to potential suppliers, which are assumed to be large in number and competing in a perfect market. The firm concludes with one of the suppliers a contract specifying the ownership share s and a transfer payment. In the second stage, both parties non-cooperatively choose their investment levels into the inputs. Since these investments are non-contractible, both parties bargain about the surplus of the relationship ex post, after their investments are sunk. Bargaining takes place in the last stage of the game, then the final good is produced and sold, and revenues are shared as agreed in bargaining. Using equations (5.1) and (5.2) as well as $y_j(i) = x_j(i)$, the implied total revenues are

$$R_j(i) = A_j \theta(i)^{\alpha} \left(\frac{K_j(i)}{\eta} \right)^{\alpha\eta} \left(\frac{L_j(i)}{1-\eta} \right)^{\alpha(1-\eta)}. \tag{5.3}$$

Due to the relationship-specificity of their investments into $K_j(i)$ and $L_j(i)$, both parties are locked

[13] Note that this final production step is absent in Antràs and Helpman (2004). In their model, $x_j(i)$ denotes the final good and production takes place in the foreign firm's country. In our model, joint production takes place in China, which reflects the case that seems to be more relevant in our application, given the high prevalence of processing activities among Chinese manufacturing firms (see Feenstra and Hanson, 2005).

[14] In assuming that the foreign firm's capital input is necessary in the joint production process we follow Antràs and Helpman (2004). Even without imposing this assumption, the foreign firm will find it optimal to contribute capital to the joint production process as long as its bargaining power is large enough ($\beta > 0.5$), as shown by Antràs (2003).

[15] These assumptions are crucial to generate a meaningful trade-off between different organizational forms, but they are more restrictive than necessary to derive our main predictions. See Antràs (2015) and Eppinger and Kukharskyy (2017) for more general setups that allow for partial relationship specificity and partial contractibility.

in a bilateral monopoly after they have formed a relationship. Also, since courts cannot verify any ex-ante agreements regarding investments (or revenues), the parties bargain over the distribution of the surplus from the relationship after input production has taken place. This setup leads to a two-sided hold-up problem and underinvestment by both F and M relative to the first-best input levels. Importantly, we follow the PRT in assuming that contractual incompleteness and the resulting hold-up problem cannot be eliminated through integration, so ex-post bargaining takes place under any organizational form s. We adopt the standard approach by modeling this process as a generalized Nash bargaining, in which each party obtains their outside option plus a share of the surplus from the relationship, the so-called quasi-rent. We set M's outside option to zero and denote F's bargaining power by $\beta \in (0,1)$, while the supplier's bargaining power is $1-\beta$. The key assumption in this setup is that by choosing a higher ownership share in M ex ante, the firm F increases its outside option in the bargaining process ex post, and thereby increases its own share of the surplus from the relationship. Formally, we assume that if bargaining breaks down, the firm can produce a fraction $\delta(s) \in (0,1)$ of $x_j(i)$, where $\delta'(s) > 0, \forall s$. Intuitively, the output loss $1-\delta(s)$ derives from the fact that the firm only has residual property rights over a share s of M's input, and hence cannot produce the same amount of $x_j(i)$ as within the production relationship. The loss decreases in the ownership share, but it is always non-zero, because F can never produce the intermediate good as efficiently on its own as in a joint production process.[16] The resulting revenues $\delta(s)^\alpha R_j(i)$ that the firm can generate if the relationship breaks up constitute F's outside option. This modeling of continuous ownership shares generalizes the approach by Antràs (2003) and Antràs and Helpman (2004), which restricts the firm's decision to a binary choice between full integration and outsourcing. Our approach is motivated by the fact that we observe many JVs empirically, and it will prove useful when we introduce China's FDI regulation policy below.

5.2.2 Equilibrium choice of inputs and optimal ownership

We solve the model by backward induction for the optimal ownership share s^*. In stage two of the game, both parties choose their investments into inputs non-cooperatively, knowing the ownership structure and anticipating the Nash bargaining solution. The firm F chooses $K_j(i)$ to maximize $\pi_F = \beta_F R_j(i) - rK_j(i)$, and the supplier M chooses $L_j(i)$ to maximize $\pi_M = [1-\beta_F] R_j(i) - wL_j(i)$, where

$$\beta_F \equiv \delta(s^*)^\alpha + \beta\left(1-\delta(s^*)^\alpha\right) \tag{5.4}$$

denotes F's revenue share, which is uniquely determined by the chosen ownership structure s^* (and the parameters α and β).

The first-order conditions of these programs deliver the input choices $K_j(i)$ and $L_j(i)$ as functions of revenues, which can be plugged back into the revenue function (5.3) to solve for:

$$R_j^*(i) = A_j^{\frac{1}{1-\alpha}} \Theta(i) \alpha^{\frac{\alpha}{1-\alpha}} \left(\frac{\beta_F}{r}\right)^{\frac{\alpha\eta}{1-\alpha}} \left(\frac{1-\beta_F}{w}\right)^{\frac{\alpha(1-\eta)}{1-\alpha}}, \tag{5.5}$$

where we have transformed productivity as $\Theta(i) \equiv \theta(i)^{\alpha/(1-\alpha)}$. Using these results, we can solve for both parties' optimal investments into the inputs in the presence of contractual frictions (and mark-up

[16] Since an alternative modeling approach, which would allow the firm to obtain a share $\delta(s)$ of the supplier's *input* greatly complicates the analysis, we adopt the approach taken by Antràs (2003) and Antràs and Helpman (2004) and model the efficiency loss $\delta(s)$ in terms of *output*.

distortions), which we call second-best input quantities:

$$L_j^*(i) = \Theta(i)(1-\eta)A_j^{\frac{1}{1-\alpha}}\alpha^{\frac{1}{1-\alpha}}\left(\frac{\beta_F}{r}\right)^{\frac{\alpha\eta}{1-\alpha}}\left(\frac{1-\beta_F}{w}\right)^{\frac{1-\alpha\eta}{1-\alpha}}, \tag{5.6}$$

$$K_j^*(i) = \Theta(i)\eta A_j^{\frac{1}{1-\alpha}}\alpha^{\frac{1}{1-\alpha}}\left(\frac{\beta_F}{r}\right)^{\frac{1-\alpha+\alpha\eta}{1-\alpha}}\left(\frac{1-\beta_F}{w}\right)^{\frac{\alpha(1-\eta)}{1-\alpha}}. \tag{5.7}$$

These quantities imply the following capital-labor ratio:

$$\frac{K_j^*(i)}{L_j^*(i)} = \frac{\eta}{1-\eta}\frac{w}{r}\frac{\beta_F}{1-\beta_F}. \tag{5.8}$$

Note that the term $\beta_F/1-\beta_F$ reflects the relative input distortion as in Hsieh and Klenow (2009), which derives from the contracting inefficiency in this setup. For the knife-edge case of $\beta_F = 0.5$, the relative distortion vanishes, but the absolute input quantities remain inefficient. After plugging the input quantities from equations (5.6) and (5.7) into the production function (5.2), we obtain the second-best output level:

$$x_j^*(i) = \theta(i)^{\frac{1}{1-\alpha}}A_j^{\frac{1}{1-\alpha}}\alpha^{\frac{1}{1-\alpha}}\left(\frac{\beta_F}{r}\right)^{\frac{\eta}{1-\alpha}}\left(\frac{1-\beta_F}{w}\right)^{\frac{1-\eta}{1-\alpha}}. \tag{5.9}$$

As in Antràs and Helpman (2004), the solution of the model is simplified by assuming that the contract between F and M includes a transfer payment. Since there are many potential suppliers with zero outside options, F chooses the size of the transfer that makes M just indifferent between accepting and declining the contract, and thereby, F extracts the entire surplus from the relationship.

In the first stage of the game, the firm therefore chooses the ownership share s^* that maximizes the total operating profits $\Pi_j(i)$ in anticipation of the second-best input quantities from equations (5.6) and (5.7) and the resulting revenues from equation (5.5):

$$\max_{s^*}\Pi_j(i) = \Theta(i)\alpha^{\frac{\alpha}{1-\alpha}}A_j^{\frac{1}{1-\alpha}}\Psi(\beta_F(s^*)) - r^\eta w^{1-\eta}f(s^*), \tag{5.10}$$

$$\text{where}\quad \Psi(\beta_F(s^*)) \equiv \left(\frac{\beta_F(s^*)}{r}\right)^{\frac{\alpha\eta}{1-\alpha}}\left(\frac{1-\beta_F(s^*)}{w}\right)^{\frac{\alpha(1-\eta)}{1-\alpha}} \times\Big(1-\alpha\eta\beta_F(s^*)-\alpha(1-\eta)\left[1-\beta_F(s^*)\right]\Big), \tag{5.11}$$

and $\beta_F(s^*)$ is given by equation (5.4).

The ownership share s^* can generally affect profits via two channels: First, the term $\Psi(\beta_F(s^*))$ reflects the core trade-off created by contractual frictions and the resulting two-sided hold-up problem: One the one hand, a higher ownership share increases the investment incentives and the share of the surplus for F, but on the other hand, it decreases the investment incentives by M, which tends to reduce the overall size of the surplus. Second, the choice of s^* can affect profits via the fixed organizational costs reflected in the term $f(s^*)$. However, these differences in organizational costs are not at the heart of the model, and any assumption on the fixed cost ranking seems ad hoc, as it is not explained within the theory of the firm. While Antràs and Helpman (2004) assume that the fixed costs of integration exceed those incurred under outsourcing, which is in line with the pattern of firm selection in Spain (Kohler and Smolka, 2011, 2014), the opposite ranking seems to be consistent with firm data for

France (Defever and Toubal, 2013). Since none of these alternative ranking assumptions seems preferable *a priori*, and since this assumption is not essential for the core trade-off generated by contractual frictions, we simplify the analysis for the moment by abstracting from differences in fixed organizational costs across ownership forms (as e.g. in Kukharskyy and Pflüger, 2010). Assuming $f(s) = f$ allows us to focus on the core trade-off reflected in $\Psi(\beta_F(s^*))$. We return to a discussion of alternative fixed cost rankings below when considering the role of firm heterogeneity.

Which ownership share maximizes the firm's variable operating profits via $\Psi(\beta_F(s^*))$ in equation (5.10)?[17] Since the firm's revenue share $\beta_F(s^*)$ is implicitly determined by the choice of s^* and strictly increasing in s^* (via $\partial\beta_F(\cdot)/\partial\delta > 0$ and $\delta'(s^*) > 0$), this problem can be solved in two steps (see Appendix 5.A.1): We first solve $\partial\Psi(\beta_F(s^*))/\partial\beta_F(s^*) = 0$ for $\beta_F(s^*)$ and second compute the ownership share s^* required to implement $\beta_F(s^*)$. Using this result, equation (5.4), and the invertibility of $\delta(s)$, we can express the optimal ownership share s^* as:

$$s^* = \delta^{-1}\left[\left(\frac{\beta_F^* - \beta}{1-\beta}\right)^{\frac{1}{\alpha}}\right], \tag{5.12}$$

$$\text{where} \quad \beta_F^* = \frac{\eta(\alpha\eta + 1 - \alpha) - \sqrt{\eta(1-\eta)(1-\alpha\eta)(\alpha\eta+1-\alpha)}}{2\eta - 1}. \tag{5.13}$$

Equation (5.12) describes a unique mapping of the optimal ownership share into the optimal revenue share (5.13), which is the same expression as in Antràs and Helpman (2004). In contrast to their analysis, which restricts the ownership choice at this point to a discrete set of two options (outsourcing or full vertical integration), we proceed with the general case and allow the firm to choose any ownership share in the unit interval $s \in [0, 1]$.

The optimal choice is illustrated in Figure 5.1(a) in terms of the revenue share β_F^*. Since neither one of the two parties receives the full return on its investment, they both underinvest in equilibrium. As a consequence, the firm does not necessarily choose the highest possible ownership share in the optimum, but instead trades off its own share of the surplus and the underinvestment by both parties, which individually decrease the size of the overall surplus. This trade-off varies systematically with the importance of the investments by the two parties, reflected in the headquarter intensity parameter η. As shown graphically in Figure 5.1(a) and derived formally in Appendix 5.A.1, β_F^* (and hence s^*) is monotonically increasing in η. This result reflects the key intuition of the PRT (Grossman and Hart, 1986): The party undertaking the more important investment should optimally be assigned more ownership rights. Thus, a higher headquarter intensity η in the joint production process optimally requires a higher ownership share s^*. Consequently, firms optimally choose outsourcing ($s^* = 0$) in industries with a very low headquarter intensity, they organize in JVs in the intermediate range of η, where s^* is strictly increasing in η, and full integration ($s^* = 1$) is optimal for very high values of η, as shown in Figure 5.1(a). This discussion generalizes an important result from Antràs and Helpman (2004) to the case of continuous ownership shares:

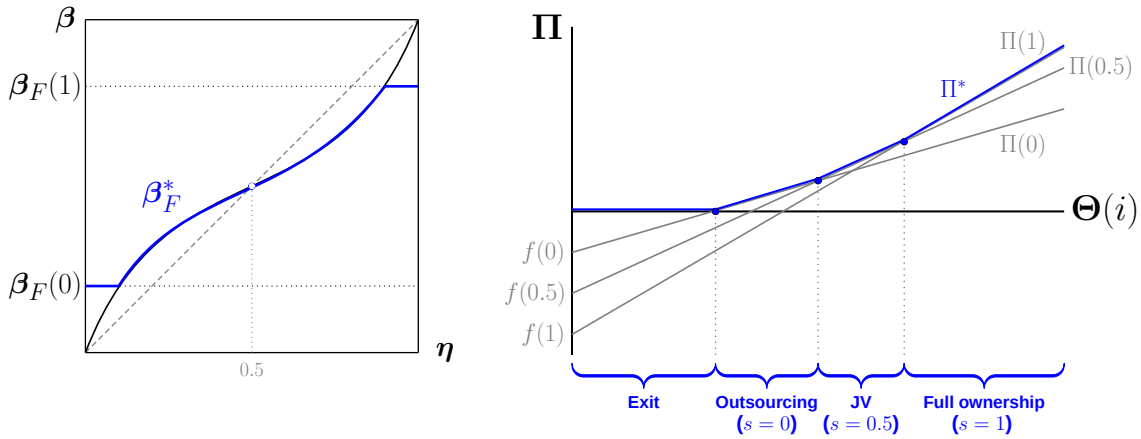
Proposition 5.1 (Headquarter intensity)**.** *The optimal ownership share s^* increases (weakly) in headquarter intensity η.*

Proof. *See Appendix 5.A.1.*

[17] Note that the firm treats market conditions reflected in $\alpha^{\alpha/(1-\alpha)}A_j^{1/(1-\alpha)}$, factor prices, and the productivity term $\Theta(i)$ as exogenous when solving this problem.

Figure 5.1: *Optimal revenue share, headquarter intensity, and productivity*

(a) *Optimal revenue share* (b) *Sorting pattern by productivity*



Note: Figure 5.1(a) illustrates the optimal revenue share β_F^* as a function of headquarter intensity η. Figure 5.1(b) illustrates maximum profits Π as a function of (transformed) productivity $\Theta(i)$ in an industry with a very high η, such that $s^* = 1$, for the example of an increasing fixed cost ranking $f(s)$ and $s = \{0, 0.5, 1\}$. It illustrates how firm pairs sort into organizational forms based on their productivity.

We now turn to a discussion of firm heterogeneity in terms of productivity $\theta(i)$ and alternative fixed cost rankings. Provided that the parameters α, β, and η are shared by all firms in the same industry, the model with constant fixed organizational costs (assumed above) predicts that all firm pairs active in a given industry adopt the same organizational form. However, overwhelming empirical evidence suggests that firms within narrowly defined industries choose different organizational forms (see e.g. Kohler and Smolka, 2011, 2014). To rationalize this fact, we follow Antràs and Helpman (2004) in assuming heterogeneity of firms in terms of their productivity in combination with fixed costs differences across organizational forms. We distinguish two types of rankings of fixed organizational costs, since their ordering is a priori not clear. First, we consider the case in which fixed organizational costs are strictly increasing in the degree of integration $f'(s) > 0 \, \forall s$, e.g. because the management of both firms' activities creates managerial overload, which dominates the associated managerial economies of scope (as in Antràs and Helpman, 2004). Second, we consider the opposite case in which fixed organizational costs are strictly decreasing $f'(s) < 0 \, \forall s$ (as in Defever and Toubal, 2013).

This set of assumptions can rationalize self-selection of firms into different organizational forms according to their productivity. The predicted sorting pattern is such that the most productive firms always choose the ownership share s^* that is optimal in their industry, as described by equations (5.12) and (5.13), because they can always cover the fixed costs of any organizational form.[18] Yet, for less productive firms, the optimal choice s^* may yield negative profits if the associated organizational fixed costs are too high. Hence, these firm pairs will either operate under the best feasible organizational form that yields non-negative profits, or they cannot produce at all. This best feasible organizational form depends on the headquarter intensity of the industry and the fixed cost ranking.

Figure 5.1(b) illustrates the sorting pattern of firms into organizational forms for one exemplary case of an industry characterized by a very high headquarter intensity η and an increasing fixed cost ranking ($f'(s) > 0$). It shows the maximum operating profits $\Pi_j(i)$ from equation (5.10) as a function

[18] This argument assumes that $f(s)$ is not too steep (such that the fixed organizational costs do not increase faster in s than $\beta_F(s)$) and that the upper bound to the productivity distribution is not too low (such that at least one firm can cover the cost $f(s^*)$).

of (transformed) productivity $\Theta(i)$ for three possible ownership forms $s = \{0, 0.5, 1\}$. In the industry depicted here, we have $\beta_F^* \geq \beta_F(1)$, so the most productive firm pairs choose full integration ($s^* = 1$). The maximum profit line is flatter for JVs, such as $s = 0.5$, reflecting the less efficient incentive structure. However, given the lower fixed costs, some firms in an intermediate productivity range choose $s = 0.5$. By the same argument, the less productive firms choose outsourcing ($s = 0$), and the least productive ones cannot operate under any organizational form and exit the market. These insights from Figure 5.1(b) generalize in a straightforward way to the case of continuous ownership shares in the unit interval.

More generally, we can distinguish three types of possible sorting patterns in industries with different degrees of headquarter intensity. First, in industries with a very *high* η, such that $\beta_F^* > \beta_F(1)$, we have that either more productive firm pairs choose *higher* ownership shares (for $f'(s) > 0$, the case in Figure 5.1(b)), or all operate under full integration (for $f'(s) < 0$) and there is no self-selection based on productivity. Second, in industries with a very *low* η, such that $\beta_F^* < \beta_F(0)$, we find that either all firm pairs conduct outsourcing (for $f'(s) > 0$), or more productive ones choose *lower* ownership shares (for $f'(s) < 0$). Third, in the range of η for which JVs are optimal, i.e., those implying $\beta_F^* \in (\beta_F(0), \beta_F(1))$, we may observe either type of selection pattern, depending on the fixed cost ranking: More productive firm pairs choose higher ownership shares, but never higher than s^* for $f'(s) > 0$, whereas more productive firms choose lower ownership shares, but never less than s^* for $f'(s) < 0$.

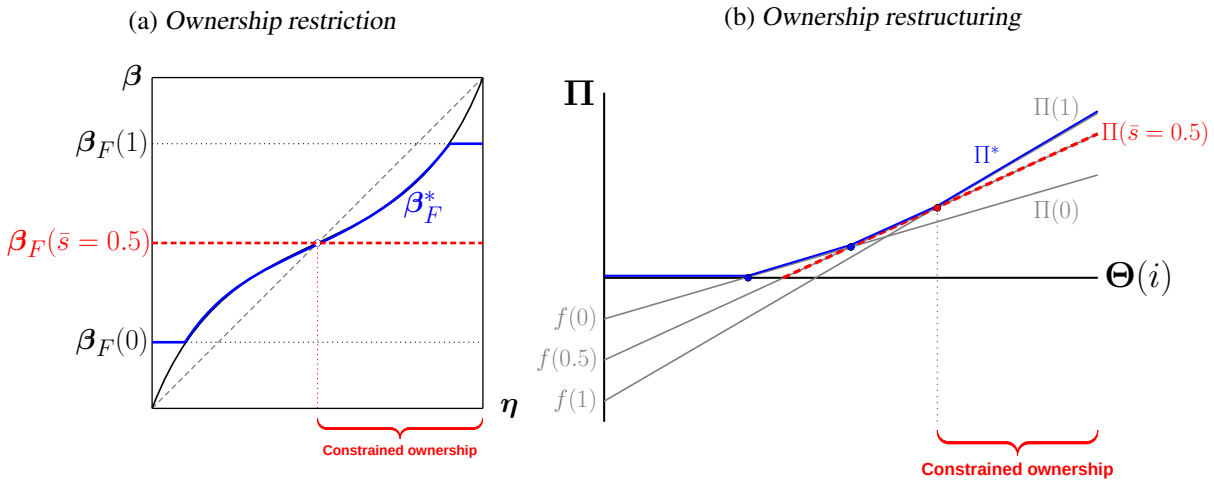
These results are reminiscent of the predictions derived by Kohler and Smolka (2015) for the case of a discrete choice between full integration and outsourcing. Formally, they follow from a modularity property of the maximum operating profit function: β_F^* maximizes $\Psi(\beta_F)$, which is the 'slope parameter' for the productivity term $\Theta(i)$ in equation (5.10). Intuitively, the incentive effect due to optimal ownership has more leverage for more productive firms. In the words of Kohler and Smolka (2015, p. 2): "the firm's productivity magnifies any per-unit production cost advantage". Hence, whichever degree of integration is preferable in a given industry on the basis of per-unit costs, the most productive firms are (weakly) more likely to choose this optimal degree of integration because they have a higher capacity to cover the fixed costs of the preferred organizational form. Our model extends these results to the case of partial ownership.

5.2.3 Modeling foreign ownership restrictions and liberalization

How do restrictions on foreign ownership affect firm performance in this model? As discussed in the introduction, the Chinese FDI policy has maintained various restrictions on foreign ownership across several industries. FDI was entirely prohibited in some industries and severely restricted in other industries, some of which faced explicit limits on foreign equity shares. In terms of the model, these policy restrictions can be understood as introducing an upper bound $\bar{s}$ on the foreign ownership share of the foreign firm F in the Chinese supplier M. This upper bound varies from zero, in case FDI is prohibited entirely, to just below one, in case the foreign investor is required to at least have a Chinese partner holding some shares in a Sino-foreign JV. The foreign ownership restriction $\bar{s}$ implies an upper bound on the firm's revenue share $\beta_F(\bar{s})$, as illustrated by the dashed line in Figure 5.2(a) for the example of $\bar{s} = 0.5$. To focus on the relevant cases, we consider only industries in which at least one firm pair is constrained by the policy, i.e., $\bar{s} < s^*$. It follows from our analysis in the previous section that in industries with a low headquarter intensity such a policy is less likely to be binding because low ownership shares are preferable on efficiency grounds. Constrained firm pairs will be those active in industries with a high headquarter intensity, as evident in Figure 5.2(a). Hence, we can focus our attention on industries with

a high headquarter intensity of $\eta > 0.5$. Furthermore, in these industries, more productive firm pairs are more likely to be constrained if the fixed cost ranking is increasing ($f'(s) > 0$). In the constrained production relationships, the policy aggravates existing distortions due to contractual incompleteness, which reduces the output below the second-best level. Therefore, abolishing $\bar{s}$ induces restructuring towards the optimal ownership share, reoptimized input quantities, and higher output as a consequence. We now develop these predictions.

Figure 5.2: *Ownership restriction and restructuring*



Note: Figure 5.2(a) illustrates the optimal revenue share β_F^* as a function of headquarter intensity η in the presence of an ownership restriction $\bar{s} = 0.5$. Figure 5.2(b) illustrates maximum profits Π as a function of (transformed) productivity $\Theta(i)$ in an industry with a very high η, such that $s^* = 1$, for the example of an increasing fixed cost ranking $f(s)$ and $s = \{0, 0.5, 1\}$. It shows that in such an industry the most productive firm pairs are constrained by the ownership restriction $\bar{s} = 0.5$.

Profit-maximizing firm behavior implies that a policy liberalization, which abolishes the foreign ownership restriction $\bar{s}$, induces the constrained firm pairs to increase s towards the optimal level s^*. The most productive firm pairs then move to the optimal ownership share s^*, corresponding to β_F^* in Figure 5.2(a). Some less productive firm pairs may increase the ownership less if they are constrained by fixed organizational costs.[19] Figure 5.2(b) provides an illustration for the same example as in the previous section (an industry with a very high η, such that $s^* = 1$, and three exemplary ownership shares with $f(1) > f(0.5) > f(0)$), where we assume that foreign ownership is restricted initially through $\bar{s} = 0.5$. Abolishing this threshold then induces ownership restructuring among all previously constrained firm pairs to $s^* = 1$. Note that the figure depicts a situation in which all organizational forms coexist in the equilibrium without the policy, though this need not be the case. There may also be industries in which production relationships cannot operate profitably except under foreign ownership, e.g. because the fixed costs $f(0)$ are higher than in the depicted situation (such that the intersection point $\Pi(0) = \Pi(1)$ lies in the loss domain). These firm pairs would be kept out of the market by the policy, but enter the industry with wholly foreign-owned supplier after liberalization. These predictions are summarized in:

Proposition 5.2 (Optimal ownership restructuring)**.** *Abolishing foreign ownership restrictions leads to (i) an increase in foreign ownership shares among previously constrained firm pairs and (ii) entry of new (partially) foreign-owned suppliers.*

Proof. *In the text.*

[19] This discussion focuses on the interesting case of a non-trivial sorting pattern among constrained firms. As noted earlier, there is no sorting if $f'(s) < 0$ in industries with a high η, and firms are unlikely to be affected by the policy at all in industries with a low η.

How are both parties' input choices and their joint performance affected by the ownership restructuring induced by the liberalization? The increase in s (and hence $\beta_F(s)$) induces the previously constrained firm pairs to adjust their input quantities towards the second-best levels from equations (5.6) and (5.7). From an examination of equation (5.8), it can be seen that this implies an increase in the capital-labor ratio in previously constrained production relationships (as we formally show in Appendix 5.A.2). This result is intuitive, since allowing the firms to optimally choose a higher ownership share ameliorates F's underinvestment into capital, while it aggravates M's underinvestment into labor. As a result, the capital-labor ratio increases in general for any increase in the ownership share, and in particular for the increase from $\bar{s}$ towards s^* induced by the policy liberalization.

The reoptimized investments lead to an increase in output from the policy-constrained level towards the second-best output level from equation (5.9) under plausible conditions. We show in Appendix 5.A.2 that output is increasing for an increase in the revenue share $\beta_F(s)$, and for a move from $\bar{s}$ to s^*, if $\eta > \beta_F(s)$. This condition is satisfied for any industry with a high headquarter intensity of $\eta > 0.5$, as is illustrated in Figure 5.1(a) and can be verified easily using equation (5.13). Recall that these are precisely the industries in which the policy is most likely to impose a binding constraint initially. It is further satisfied also in other industries if the policy restrictions are severe and imply a low $\beta_F(\bar{s})$. Intuitively, the firm pairs move closer to the second-best output level because the aggravated underinvestment by the Chinese supplier is more than compensated by the improved investment incentives of the foreign firm through its increased ownership share. This is how the efficiency gains from an optimal allocation of ownership rights can materialize in terms of higher output:

Proposition 5.3 (Firm performance)**.** *Optimal ownership restructuring after liberalizing ownership restrictions (i) increases the capital-labor ratio and (ii) results in overall higher output in previously constrained firms.*

Proof. *See Appendix 5.A.2.*

It should be noted that this core prediction is independent of the different rankings of fixed organizational costs discussed in Section 5.2.2.

At this point, two comments are in order regarding the effects on inputs and alternative measures of firm performance. First, the question arises: How are both input quantities K and L affected by the optimal ownership restructuring after liberalization? In principle, there are two types of effects at work for each input: a direct incentive effect, working in opposite directions for both parties, and an indirect size effect, which may increase the use of both inputs due to the increase in output. The effect on capital is clearly positive, as both the incentive effect for F and the overall efficiency gains lead to increased capital investments. Hence, the use of capital increases both relative to labor and in absolute terms. By contrast, the net effect on the quantity of labor employed is ambiguous in general, since the two effects work in opposite directions. The direct incentive effect on L is negative, since M retains less residual property rights, which aggravates the hold-up problem from the viewpoint of the supplier and reduces its incentives to contribute labor to the joint production process. This negative effect dominates in particular for a marginal increase in s towards the optimal ownership share s^*. However, provided that the optimal ownership restructuring increases output, this tends to raise the demand for labor. For a non-marginal increase from $\bar{s}$ to s^*, this size effect may overcompensate the incentive effect, resulting in a positive employment effect. Appendix 5.A.2 formally derives the marginal effects of an increase in $\beta_F(s)$ on K and L and illustrates by use of a numerical example how the effect of optimal ownership restructuring on L may turn positive.

Second, we have chosen output as our preferred measure of firm performance in anticipation of our empirical analysis. This choice is motivated by the fact that our theory delivers a clear prediction for the output effect of optimal ownership restructuring, and output is observable in the data. One may alternatively consider measuring firm performance in terms of profitability or productivity, but both of these measures have their drawbacks. Profits almost trivially increase for any move to optimal ownership, which makes them a natural candidate to look at. However, the economic profits from the relationship, corresponding to Π in our model, are unobservable in our data on Chinese firms, since (i) accounting profits are generally a poor proxy for economic profits due to tax-optimizing reporting behavior, and (ii) the profits from the relationship are likely to be captured by the foreign firm, whose profits we do not observe. Hence, we do not focus on profitability as a performance measure. Given our focus on efficiency gains, productivity would seem to be another suitable performance measure. Note that TFP θ in our model is exogenous, drawn in an entry lottery à la Melitz (2003), and does not respond to ownership changes. Hence, we would not expect any effect of optimal ownership restructuring on TFP, provided that it can be measured in a theory-consistent manner. Instead, the efficiency gains in the model arise purely through reoptimized input choices. Only to the extent that TFP captures aspects of the production process that can be subject to underinvestment, one may expect an impact of ownership restructuring on TFP, but this would not be an empirical analogue of θ. We will return to this question in the empirical analysis of firm performance in Section 5.6.

5.3 Policy Background and Data

To bring the theoretical predictions to the data, we take advantage of China's large-scale FDI liberalization following its WTO accession. In this section, we briefly introduce the institutional background on China's FDI regulation policy, describe the reform liberalizing this policy in many industries, and provide an overview of our firm panel data.

5.3.1 China's FDI regulation policy

China was historically closed to FDI and started opening its door to foreign capital in 1978, as part of the 'Reforming and Opening' program. The regulations on FDI, however, remained tight until the mid-1990s.[20] In 1995, the National Development and Reform Commission (NDRC) and the Ministry of Commerce (MOFCOM) co-published the first 'Catalogue for the Guidance of Foreign Investment Industries' (henceforth referred to as the Catalogue) as a guideline for regulations on FDI. In the Catalogue, products are sorted into four categories: 'restricted', 'encouraged', 'prohibited', and by implication 'permitted' (the last is the residual category of products not included in the Catalogue).[21] Different rules and restrictions apply to products in different categories. For products in the 'encouraged' category, foreign investors are granted freedom in choosing their ownership structures and often enjoy other privileges, such as preferential tax rates, subsidized land, or duty-free imports of inputs (Sheng and Yang, 2016). By contrast, no FDI is allowed in 'prohibited' products. Foreign firms that intend to invest in the 'restricted' products are subject to intense scrutiny and have to satisfy stringent requirements.

[20] For instance, before 1990, the 'Equity Joint Venture Law' required that the chairman of the board of an equity JV must be appointed by Chinese investors.

[21] To give some examples, based on the 2002 Catalogue, 'wool and cotton textiles' and 'container production' were both restricted categories for foreign investment, while investment in 'automobile and motor cycles' was encouraged.

Equity requirements are one particular type of restrictions that directly constrain foreign ownership structures, namely the *Sino-Foreign joint venture* requirement and the *majority Chinese ownership* requirement. Under the *Sino-Foreign joint venture* requirement, a foreign investor has to find a Chinese partner and form a joint venture (JV) to start a business. Hence, this regulation excludes full foreign ownership, corresponding to a foreign equity share of 100%. Under the *majority Chinese ownership* requirement, a foreign investor can only own 50% or less of the firm's equity shares.[22] Hence, it excludes majority foreign ownership. Other restrictions include local content requirements, technology transfer for market access, or mandatory export requirements, among others.[23]

From the information provided in the Catalogue, we construct two indicator variables for the policy categories 'restricted or prohibited' and 'encouraged'. These are obtained by mapping the product descriptions from the Catalogue to our firm data (described in Section 5.3.3) via the Chinese Industry Classification (CIC) at the 4-digit level if possible, and at the 3-digit or 2-digit level in some instances. Note that the different policy categories are not mutually exclusive at the 4-digit industry level, so FDI in a given industry may for instance be both encouraged (for some products) and restricted (for some other products), or an encouraged product may have explicit equity requirements. Furthermore, some foreign firms that export their manufactured products have also received preferential treatment.[24]

5.3.2 China's WTO accession and FDI policy liberalization

After its first publication in 1995, the Catalogue was revised in 1997, 2002, 2004, 2007, 2011, and 2015.[25] Through these revisions, the Catalogue has generally become more liberal and investor-friendly over time. Our main focus is on the 2002 revision of the Catalogue, which manifests a major liberalization of inward FDI as a direct consequence of China's accession to the WTO in 2001.

In accessing the WTO, China made exceptional commitments to liberalize its FDI policy. In compliance with the WTO Agreement on Trade-Related Investment Measures (TRIMs), China had to abolish any restrictions or distortions on foreign investment related to international trade. In particular, China had to abolish its performance requirements (such as local content or technology transfer) in restricted industries upon accession, and it was not granted a transition period to implement these changes.[26] What is more, China accepted several additional obligations that substantially exceed the standard WTO rules. These additional obligations are unique in WTO history (Qin, 2007). They further extended China's commitments to liberalize foreign investment by granting national treatment to foreign-invested enter-

[22] In some cases, the requirement is that the Chinese parties shall hold a higher equity share than any foreign parties in the joint project, which is known as a 'relative majority Chinese ownership' requirement.

[23] For example, a foreign firm may be required to export the majority if its manufactured products to get approval in a restricted product category.

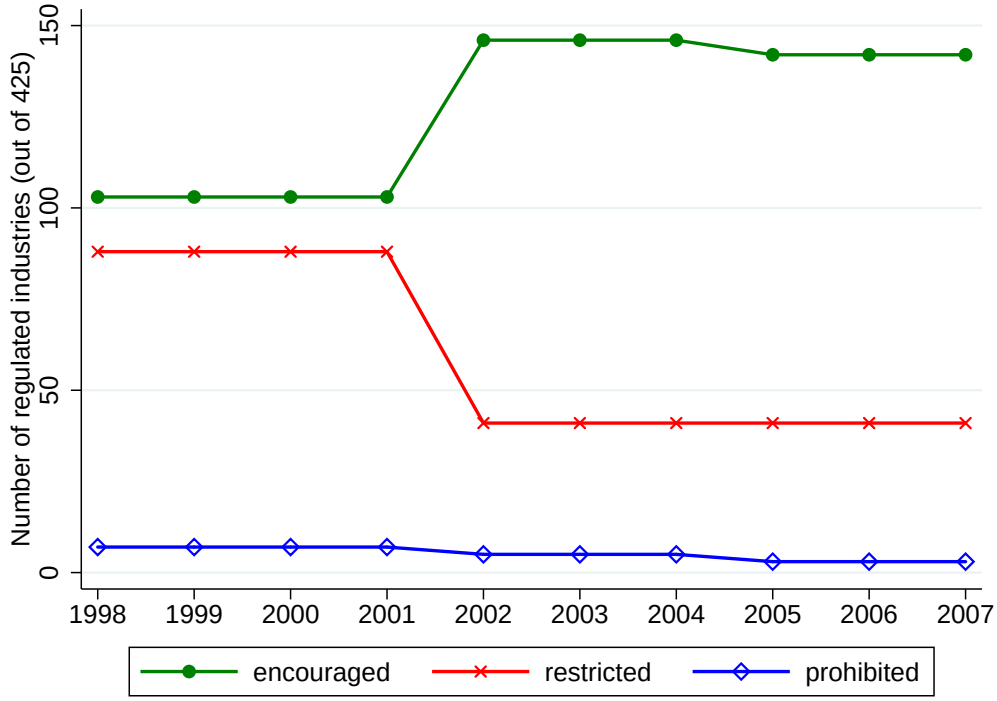
[24] To be concrete, foreign investment in 'permitted' product categories can be treated as 'encouraged' if all products are made for export. And foreign investment in 'restricted' product categories can potentially be treated as 'permitted' if the export share over total sales is above 70%, upon approval.

[25] For our empirical investigation, we code the policy changes by the dates when they were effectively implemented. The 2002 revision became effective on April 1, 2002, so we code the new policy as effective in 2002-2004. The 2004 revision became effective on January 1, 2005, so we code the new policy as effective in 2005-2007, as the 2007 revision became effective only on December 1, 2007.

[26] The Accession Protocol states that "China shall, upon accession, comply with the TRIMs Agreement, without recourse to the provisions of Article 5 of the TRIMs Agreement" (Section I.7.3). Article 5 grants a transition period to developing countries.

prises and abolishing equity restrictions in particular industries.[27] To fulfill these obligations, China undertook a major regulatory reform regarding foreign trade and investment in 2002, which involved a substantial liberalization of the Catalogue.

Figure 5.3: *FDI liberalization in Chinese manufacturing, 1997-2007*



Note: Authors' calculations based on the 'Catalogues of Industries for Foreign Direct Investment' (NDRC, various issues).

Figure 5.3 illustrates the major FDI liberalization in 2002 after China's WTO accession. It counts the number of industries in which products are regulated by the Catalogue in each year over the period 1997-2007. Within the manufacturing sector, 45 industries (out of 425) became newly 'encouraged', while 50 industries were removed from the 'restricted' list and 2 industries were no longer 'prohibited'. The net changes depicted show a clear liberalization process. Note that the changes in the Catalogue in 2004 were very small in comparison: Within the manufacturing sector, only three industries were liberalized. Hence, we focus on the major reform of 2002.

As briefly discussed in the introduction, our main difference-in-differences analysis in Section 5.6 compares the performance of firms increasing their foreign ownership share in liberalized industries to other firms in the same 4-digit industries, which were already foreign-owned before the reform. This approach exploits the fact that the products listed in the FDI Catalogue are generally defined at a more disaggregate level than the 4-digit industry codes by which we can map them to the firms. Hence, foreign-owned firms exist in the industries classified as 'liberalized' already before the reform because they produce unregulated products. We have made spot checks using information on firms' specific products from their websites to validate this assumption. Two illustrative examples come from the 4-digit industries 'jewelry' and 'cable and wire' manufacturing. First, the firm ChengDu TianHeYinLou

[27] Article 207 of the "Report of the Working Party on the Accession of China" states explicitly: "With respect to the manufacture of motor vehicle engines, the representative of China also confirmed that China agreed to remove the 50 per cent foreign equity limit for joint-ventures upon accession." This article was incorporated into the accession protocol pursuant section 1.2 of the same protocol and article 342 of the working party report.

LLC increased its foreign ownership share from 0 to 100% in 2003, after the FDI restriction on its main product 'golden or silver jewelry' was abolished. By contrast, ShangHai TianQi LLC, which was foreign-owned throughout the years 2001-2003, produces 'pearl jewelry', a product that was not listed in the 1997 Catalogue. Second, the main product of Xiamen City HuaLe LLC is a 'cable for high voltage', for which FDI was restricted (with a *Sino-Foreign joint venture* requirement) before the reform, thus it could only be acquired to 100% by a foreign owner in 2003. The comparable firm Beijing LuoBoSenGe Electronic LLC produces other 'electronic cable', which was not listed in the FDI Catalogue, allowing this firm to be wholly foreign-owned already before 2002. This setup allows us to use highly comparable firms that are similarly selected by foreign investors and active in the same narrowly defined industries as a control group in our main empirical analysis of firm performance.

5.3.3 Overview of firm data and key variables

To investigate foreign ownership and firm performance in China, we use data from the Chinese Annual Surveys of Industrial Production (ASIP), conducted by the National Bureau of Statistics of China (NBSC). This dataset is the most comprehensive survey data for industrial firms in China.[28] It forms the basis for major statistics published in China Statistical Yearbooks and has been widely used in economic research (mostly notably by Hsieh and Klenow, 2009; Brandt et al., 2012). The dataset covers all state-owned enterprises (SOEs) and all other firms with total sales exceeding 5 million RMB (around US$ 620,000 at the running exchange rate in 2005) in the industrial sector (including manufacturing, mining, and utilities). It is a panel for the period of 1998-2007, covering several years before and after the major FDI policy reform in 2002. For our purpose, we limit the analysis to manufacturing firms, resulting in a sample of close to half a million firms (on average around 180,000 firms per year). We follow Cai and Liu (2009) in taking necessary steps to clean the data. Brandt et al. (2012) suggest that one sixth of all firms that are observed for more than one year change their official ID over the sample period. Hence, we follow their procedures to track each firm over time using their ID, name, industry, address, and other information.

From the ASIP, we collect detailed firm-level production and balance-sheet information (gross output, value added, employment, wage bill, fixed assets, material input purchases, etc.). Most importantly, the data contain information on the firms' ownership structures. Our measure of foreign ownership is the share of foreign investors, including investors from Hong Kong, Macao, or Taiwan, in the firm's paid-in capital. This variable has the key advantage that it allows us to observe the degree (or depth) of foreign ownership instead of just a binary choice.

We also use the ASIP data to compute proxies for the key determinants of ownership shares according to our PRT model, namely headquarter intensity η and productivity θ. The crucial headquarter intensity parameter has typically been approximated by industry-level capital intensity or skill intensity since the pioneering work by Antràs (2003). This approach is based on the premise that the foreign owner is more likely to contribute to suppliers' relationship-specific investments in capital or human capital compared to investments in labor, as discussed in Section 5.2. This premise seems plausible also in the Chinese context, as foreign ownership has long played an important role for Chinese firms in gaining access to capital (Huang, 2008). Bache and Laugesen (2013), Corcos et al. (2013), and Defever and Toubal (2013) have taken this approach to the firm level, arguing that factor intensities vary even within narrowly defined industries. We follow the literature and compute capital intensity and skill intensity as proxies

[28] It accounts for over 90% of industrial output and over 70% of industrial employment in 2004 (Brandt et al., 2012).

for headquarter intensity, alternatively at the industry level and at the firm level. The firm-level capital intensity K/L is obtained by dividing the real capital stock K by the number of workers L in a given year. We compute the real capital stock by applying the perpetual inventory method to our firm panel. The firm-level skill intensity S/L is the share of workers with college education S in the total number of workers, as observed in the Chinese firm census in 2004. Industry-level measures are arithmetic means of these ratios by industry and year.

To obtain a measure for the total factor productivity (TFP) θ of the firm pairs, we estimate production functions at the 2-digit industry level based on the ASIP data. Since TFP is observable to the firm and affects its input choices and output, but it is unobserved by the econometrician, estimation of these production functions suffers from well-known endogeneity issues. We address these issues by adopting the method proposed by Ackerberg et al. (2015), which builds on the seminal insights from Olley and Pakes (1996) and Levinsohn and Petrin (2003). To adopt their control function approach, we assume that gross output is Leontief in material input $M_j(i)$ and the Cobb-Douglas production function $x_j(i)$, combining $K_j(i)$ and $L_j(i)$ as specified in equation (5.2), which has the interpretation of value added in this context.[29] We further assume that material input monotonically increases in TFP and can hence be inverted to control for TFP. For this procedure, material input M is deflated by the industry-specific input deflator and value added is deflated the industry-specific output deflator provided by Brandt et al. (2012).[30] From the estimated production functions, we obtain predicted TFP as the time-varying firm-specific residual.

5.4 Foreign Ownership Patterns in Chinese Manufacturing

5.4.1 Econometric model of foreign ownership

As a first step in our empirical analysis, we examine how the factors predicted to shape optimal ownership according to our PRT model correlate with foreign ownership shares in Chinese firms in *unrestricted* industries. This exercise serves to validate the model of Section 5.2, which is used to interpret our main results in Section 5.6, and it relates our paper to the existing literature on the organization of MNEs. To this end, we specify the following econometric model:

$$FO\ share_{it} = \boldsymbol{\chi} \cdot \mathbf{X}_{it} + \boldsymbol{\zeta} \cdot \mathbf{Z}_{it} + \rho_j + \rho_r + \rho_t + \upsilon_{it}\,, \tag{5.14}$$

where the dependent variable $FO\ share_{it}$ denotes the share of equity in firm i and year t owned by foreigners. The vector $\mathbf{X}_{it}$ (with associated coefficient vector $\boldsymbol{\chi}$) contains different sets of variables that determine ownership shares according to our PRT model of the MNE. First, to proxy for the key headquarter intensity parameter η, we follow the literature in using the industry-level capital intensity. To allow for the possibility that headquarter intensity varies within industries, we also use as alternative proxies for η the one-year lags of the firm-level capital intensity $\ln K/L$ and the skill intensity $\ln S/L$ (see Section 5.3.3 for data and measurement). Second, we interact these headquarter intensity proxies with the one-year lag of the firm's estimated TFP. Since the effect of TFP might be non-linear in headquarter

[29] This assumption avoids the identification issue raised by Gandhi et al. (2016). We have also experimented with estimating gross output production functions following their approach. Note that we relax the assumption of constant returns to scale of $x_j(i)$ when estimating production functions, as is customary to obtain a better empirical fit.

[30] For the subsequent empirical analysis, exports are deflated by the same output deflator, where appropriate, and the wage (bill) is deflated by the consumer price index from the World Bank's World Development Indicators.

intensity (see Section 5.2.2 and Kohler and Smolka, 2015), we also adopt a more flexible specification that interacts TFP with quintiles ($Q2 - Q5$) of firm-level capital intensity (similar to the industry-level analysis by Nunn and Trefler, 2014).

As additional firm-level control variables $\mathbf{Z}_{it}$ (with coefficient vector $\boldsymbol{\zeta}$), we include: firm age, which may be important for optimal ownership shares in joint ventures (see Bircan, 2014), the lagged export share (*exports/ sales*), since ownership may depend on whether the firm supplies the domestic market, other markets, or its foreign owner, and the lagged log of the value added ratio (*VA/ sales*) as a proxy for the importance of the Chinese firm in the value chain (following Antràs, 2003). We always include fixed effects ρ_j, ρ_r, and ρ_t by 4-digit industry j, province r, and year t, respectively. The error term is denoted by υ_{it}.

We estimate all variants of equation (5.14) by Ordinary Least Squares (OLS). Standard errors are clustered by firm or by industry, depending on the level at which the regressors are varying.[31] We use all the available years in our panel, but restrict the sample to industry-years in which FDI was not restricted or prohibited, in order to abstract from the effects of the policy for now.

We conduct this exercise to verify whether the patterns of foreign ownership in Chinese firms are generally in line with the PRT model developed in Section 5.2. More specifically, Proposition 5.1 predicts that ownership shares are higher in industries and firms with a high headquarter intensity. Furthermore, based on the discussion in Section 5.2.2, we would expect that ownership shares increase in the productivity of the production process for a high headquarter intensity, while they tend to decrease in productivity for a low headquarter intensity. These theoretical considerations suggest a positive interaction effect of headquarter intensity and productivity. Note, however, that these estimations should not be interpreted as a strict test of Proposition 5.1 due to an important endogeneity issue suggested by the same theory. Since increased foreign ownership raises the investment incentives of the foreign owner, but decreases those of the Chinese firm, the factor intensities used to proxy for headquarter intensity are endogenous to ownership (see Proposition 5.3). This concern is ameliorated here by using industry-level and lagged firm-level proxies for headquarter intensity and productivity, but it cannot be fully eliminated.

5.4.2 Empirical evidence on foreign ownership patterns in China

Table 5.1 summarizes the results of estimating equation (5.14). It demonstrates that foreign ownership patterns in the Chinese firm data are generally consistent with our PRT model of the MNE. In column 1, as a first proxy for headquarter intensity, we examine the log of the industry-level capital intensity. The estimated correlation is positive and highly statistically significant, as predicted by Proposition 5.1. Similarly, in column 2, we find that foreign ownership shares are positively correlated with the firm-level capital intensity (lagged by one year) and skill intensity (as observed in 2004).[32] These findings are consistent with the PRT, provided that these variables reflect the fundamental importance of physical and human capital in production and that the foreign owner is more likely to contribute these inputs to production, whereas the local firm is responsible for hiring (low-skilled) labor. The evidence suggests that firms with a higher headquarter intensity are foreign-owned to a larger degree. The coefficient estimates for the control variables in these and all subsequent regressions suggest that younger, more

[31] Since TFP is a generated regressor, we also experiment with cluster-bootstrapped standard errors (based on 500 replications). However, since the bootstrapped standard errors for the main regressors turn out to be smaller for the main regressors in almost all specifications, we prefer the more conservative cluster-robust standard errors.

[32] Since skill intensity can only be measured for firms observed in 2004, the sample is reduced in this regression. For the same reason, we use capital intensity as our preferred firm-level proxy for headquarter intensity in the remaining regressions.

export-oriented firms with lower value added ratios tend to have higher foreign ownership shares.

Table 5.1: *Foreign ownership patterns*

Dependent variable: *FO share*	Headquarter intensity		Productivity × headquarter intensity		
	Industry	Firm	Industry	Firm	Firm quintiles
	(1)	(2)	(3)	(4)	(5)
ln K/L (industry)	0.0222***		0.00398		
	(0.00801)		(0.00841)		
ln K/L (firm, lag)		0.0515***		0.0585***	
		(0.000719)		(0.000943)	
ln S/L (firm, 2004)		0.0204***			
		(0.000838)			
TFP (firm, lag)			-0.0453***	-0.0220***	-0.0000685
			(0.0111)	(0.00221)	(0.00147)
TFP × ln K/L (industry)			0.0135***		
			(0.00256)		
TFP × ln K/L (firm, lag)				0.00890***	
				(0.000567)	
Q2 K/L (firm, lag)					0.0355***
					(0.00229)
Q3 K/L (firm, lag)					0.0594***
					(0.00248)
Q4 K/L (firm, lag)					0.0887***
					(0.00262)
Q5 K/L (firm, lag)					0.180***
					(0.00310)
TFP × Q2 K/L (firm, lag)					0.00152
					(0.00154)
TFP × Q3 K/L (firm, lag)					0.00580***
					(0.00165)
TFP × Q4 K/L (firm, lag)					0.0107***
					(0.00172)
TFP × Q5 K/L (firm, lag)					0.0307***
					(0.00195)
Age (firm)	-0.00285***	-0.00352***	-0.00331***	-0.00342***	-0.00336***
	(0.000198)	(0.0000533)	(0.000245)	(0.0000530)	(0.0000527)
Exports/ sales (firm, lag)	0.340***	0.370***	0.335***	0.342***	0.340***
	(0.0195)	(0.00277)	(0.0221)	(0.00358)	(0.00358)
ln (*VA/ sales*) (firm, lag)	-0.00125	-0.00111	-0.00544***	-0.00427***	-0.00425***
	(0.00155)	(0.000988)	(0.00177)	(0.00110)	(0.00110)
Observations	1,068,379	672,193	500,818	499,163	500,818
R^2	0.289	0.335	0.359	0.387	0.383

The table reports OLS estimates of equation (5.14) in the sample of unrestricted industries, controlling for 4-digit industry, province, and year fixed effects. TFP (in logs) is the firm-level residual from industry-level production functions, estimated as described in Section 5.3. K/L is the capital-labor ratio and S/L is the skill intensity of the industry (in columns 1 and 3) or firm (in columns 2, 4, and 5), respectively. $Q2 - Q5$ denote quintile dummies. Standard errors clustered by industry in columns 1 and 3 and clustered by firm in in columns 2, 4, and 5 are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

Second, we allow for a heterogeneous effect of productivity on ownership across different values of headquarter intensity. This is captured by the interaction term between TFP and headquarter intensity in columns 3-5 of Table 5.1. Again, we start by considering the industry-level capital intensity as a proxy for headquarter intensity in column 3, which shows a direct positive (but insignificant) correlation with the

FO share. The coefficient of the interaction term is positive and highly statistically significant. It suggests that in capital-intensive industries, more productive firms have higher foreign ownership shares. This pattern is in line with our model for a ranking of fixed organizational costs that is increasing in the degree of integration (as discussed in Section 5.2.2). Also, the negative coefficient of TFP can be rationalized by our framework for a decreasing fixed cost ranking: In industries with a very low headquarter intensity, lower ownership shares are optimal, and these are more likely to be chosen by more productive firms if outsourcing is associated with relatively high fixed organizational costs. Hence, a higher productivity tends to favor lower ownership shares in industries with a low headquarter intensity. The final column adopts a more flexible specification by considering quintiles of the distribution of capital intensity across all firms. It reveals that the positive association of the *FO share* with our proxy for headquarter intensity is monotonic. Furthermore, the positive correlation of foreign ownership with TFP is concentrated in the highest three quintiles of K/L and increases monotonically in this variable. Again, this findings is in line with our model for an increasing fixed cost ranking. Overall, we find that the patterns of foreign ownership observed in Chinese manufacturing are fully in line with key predictions of our PRT model.

It is interesting to compare our results to the existing literature which has empirically examined ownership patterns in global sourcing. Both industry-level and firm-level studies have so far mainly focused on firms' headquarters and generally found evidence in line with the theory (see e.g. Nunn and Trefler, 2013; Kohler and Smolka, 2015). These papers implicitly assume that factor intensities and productivity of the headquarters are suitable proxies for the parameters characterizing the joint production process. We complement these studies with evidence on the supplier side of the global sourcing relationship and continuous ownership shares. Since we are interested in production processes taking place in China, we derive proxies for headquarter intensity and productivity using data on Chinese firms. Similar to the previous literature, our findings confirm the ownership patterns predicted by the PRT.

5.5 Direct Effect of the FDI Policy on Foreign Ownership

5.5.1 Descriptive evidence of the direct policy impact

Figure 5.4 provides indicative evidence that the FDI policy reform in 2002 had the expected positive effect on foreign ownership. Figure 5.4(a) focuses on new entrants into the panel. It depicts the share of equity JVs (*FO share* of 25-99.99%) and wholly foreign owned firms (*FO share* of 100%) among all entrants in each year. While there is slightly more entry of wholly foreign-owned firms compared to JVs in the years 1999-2001, both lines seem to evolve in parallel over this pre-reform period. With the 2002 FDI policy reform, there is a clear divergence, with full ownership becoming more prevalent among entrants. This observation is in line with the second part of Proposition 5.2. In 2002 and every later year, 9-12% of all entrants were wholly foreign owned, compared to only 3-5% JV entrants.

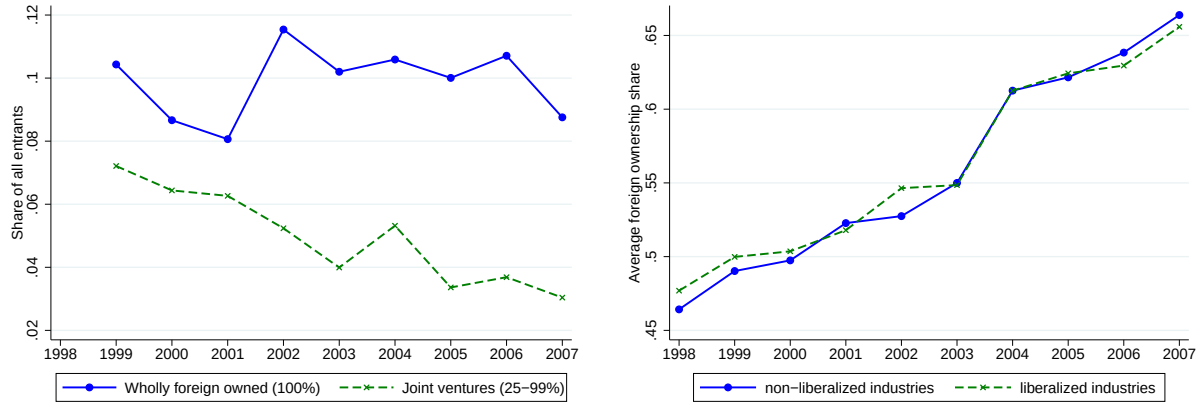
Figure 5.4(b) moves on to examine the average foreign equity share among incumbent firms, excluding all entrants in a given year. It shows a clear trend towards higher foreign ownership shares throughout the period of 1998-2007, both for non-liberalized and liberalized industries, defined as those for which a 'restricted' or 'permitted' FDI policy was abolished in 2002.[33] It is remarkable that foreign ownership in these two types of industries seems to follow very closely a common trend both before 2001 and after 2003. The only notable exception is the reform year 2002, in which the trend towards foreign owner-

[33]This is also true if we define incumbents more narrowly by considering only firms that are observed throughout the entire period.

ship accelerated in liberalized industries (increasing from around 52% to 55% average foreign ownership shares), while it slowed down in the other industries. These observations are in line with the first part of Proposition 5.2. To sum up, a clear general trend towards higher degrees of foreign ownership is visible both for entrants and incumbents for the period of 1998-2007, as is a temporary acceleration of this trend with the major FDI policy reform in 2002.

Figure 5.4: *Foreign ownership over time*

(a) *Shares of foreign-owned firms among entrants* (b) *Foreign ownership share among incumbents*



Note: Figure 5.4(a) illustrates the share of Sino-foreign equity JVs (*FO share* of 25-99.99%) and wholly foreign owned firms (*FO share* of 100%) among all entrants in a given year. Figure 5.4(b) illustrates the average *FO share* among all incumbent firms, excluding entrants in a given year. It distinguishes industries by the FDI policy change in 2002. Industries for which a 'restricted' or 'permitted' FDI policy was abolished in 2002 are labeled 'liberalized' industries, all others are non-liberalized.

5.5.2 Econometric specification

To assess the direct effect of the FDI regulation policy on ownership structures more rigorously, we regress measures of foreign ownership FO_{it} on a set of policy regulation variables $\mathbf{R}_{jt}$, indicating whether FDI in industry j of firm i in year t is 'restricted or prohibited' or 'encouraged':

$$FO_{it} = \boldsymbol{\gamma} \cdot \mathbf{R}_{jt} + \boldsymbol{\psi} \cdot \mathbf{X}_{it-1} + \boldsymbol{\xi} \cdot \mathbf{Z}_{it} + \boldsymbol{\kappa} \cdot \mathbf{T}_{jt} + \delta_j + \delta_r + \delta_t + \nu_{it}\,. \tag{5.15}$$

As the dependent variable FO_{it}, we consider alternatively the continuous foreign ownership share (*FO share*$_{it}$) or dummy variables indicating either (i) a foreign ownership share at or above the standard legal threshold of 25% (*FO dummy*$_{it}$), or (ii) majority foreign ownership (*MFO dummy*$_{it}$), or (iii) whole foreign ownership (*WFO dummy*$_{it}$).[34] We control for the preferred set of the observed, theory-based determinants of the integration decision contained in $\mathbf{X}_{it}$ (with coefficient vector $\boldsymbol{\psi}$) from column 4 of Table 5.1 and the same firm-level control variables $\mathbf{Z}_{it}$ (with coefficient vector $\boldsymbol{\xi}$). Since China's WTO accession involved trade policy changes that might affect the attractiveness of different industries for FDI, we also control for Chinese import tariffs on the firm's output good and its inputs (inferred from

[34] A firm with a *FO share* $\geq$25% is officially classified as a foreign-invested enterprise (FIE) in China.

Chinese input-output tables), summarized in $\mathbf{T}_{jt}$ (with coefficient vector $\boldsymbol{\kappa}$).[35] Finally, we include fixed effects by 4-digit industry (δ_j), province (δ_r), and year (δ_t) to account for unobservables. We assume that the error ν_{it} has mean zero and is i.i.d.. Naturally, we expect the estimated parameters $\boldsymbol{\gamma}$ to reflect the policy's intention that foreign ownership is reduced by FDI restrictions and increased by the encouraged policy. In particular, Proposition 5.2 predicts that liberalizing a 'restricted or prohibited' policy increases the foreign ownership share.

We first estimate equation (5.15) for the full panel of 1998-2007 by OLS. This approach provides a general picture of the policy impact on foreign ownership within industries over time. Notably, it captures the combined effect of the policy on foreign ownership along three different margins of adjustment: (i) existing firms being acquired (to a higher degree) or divested by foreigners, (ii) firms switching into and out of regulated industries, and (iii) firms entering or exiting the panel. In a second step, we focus on the effect of the major FDI policy reform in 2002 on the first of these margins. To this end, we estimate equation (5.15) in first differences using only data for one year before and one year after the policy change, i.e., 2001 vs. 2003. We exclude all firms changing their 4-digit industry code over the respective period, and the setup implies that firms entering or exiting the panel between these years are also excluded. Note that first differencing not only controls for heterogeneity specific to industries or regions, but it also accounts for all time-invariant firm characteristics. We continue to include the lagged theory-based variables $\mathbf{X}_{it-1}$ and firm age, and we also add the first-differenced export share and log value added ratio. To account for other contemporaneous changes in policy and market conditions, we further control for the first differences in output and input tariffs as well as 3-digit industry fixed effects in the first-differenced model. Since the policy variables are industry-specific, standard errors are clustered at the 4-digit industry-level in all of these regressions.

5.5.3 Estimation results for the direct policy impact

Table 5.2 presents the results of estimating equation (5.15) and shows that the FDI restrictions had the expected direct effect on foreign ownership. Panel A presents OLS estimation results in the full panel for the alternative dependent variables. All four measures of foreign ownership are negatively affected by the 'restricted or prohibited' FDI policy, as expected, though the effect on the *WFO dummy* is insignificant. The estimate in column 1 suggests that abolishing FDI restrictions has increased the *FO share* on average by 1.2 percentage points (or around 8% relative to the lagged average ownership share). Note that this number reflects only the average treatment effect across all firms, but it involves many large increases in foreign ownership for the subset of treated firms. Similarly, dropping an FDI restriction has increased the probability of being foreign owned above the 25% or 50% thresholds by 1.1-1.5 percentage points or 7-8%.[36] By contrast, the effects of the 'encouraged' policy are estimated to be small, negative, and insignificant, suggesting that this policy did not have the intended effect of fostering FDI. This finding may not be too surprising, at least for columns 1, 3, and 4, since exceeding the legal foreign

[35] See Brandt et al. (2017) for evidence on the impact of the tariff cuts due to WTO accession. Data on Chinese import tariffs come from the Worldbank's World Integrated Trade Solutions (WITS) database at the 6-digit level of the harmonized system (HS6) and are matched to Chinese 4-digit industries via a correspondence table. Output tariffs correspond to the firm's industry, and in the case of multiple HS6 codes per industry, the tariffs are weighted by trade volumes, excluding processing trade (which is typically exempt from tariffs). Input tariffs are obtained using input coefficients of the firm's industry according to Chinese input-output tables for 2002, which mandates a higher level of aggregation, and they are similarly weighted by non-processing trade volumes.

[36] These estimates are based on linear probability models with their well-known limitations. Using a fractional logit model for the *FO share* and logit models for the *FO*, *MFO*, and *WFO dummies* yields estimates for the policy's average partial effects of a similar magnitude. These results are available from the authors on request.

ownership threshold of 25% may in practice be sufficient to obtain the extra benefits associated with the 'encouraged' policy. Therefore, we would not necessarily expect that it provides incentives to increase the foreign ownership share at the margin or to 100%.[37] Note that these estimates reflect ownership changes as well as industry switching, entry, and exit.

Table 5.2: *Effect of the FDI regulation policy on foreign ownership*

Panel A. Pooled OLS – Full panel: 1998-2007	(1)	(2)	(3)	(4)
Dependent variable in header:	*FO share*	*FO dummy*	*MFO dummy*	*WFO dummy*
Restricted or prohibited policy	-1.171**	-0.0148***	-0.0111**	-0.00823
	(0.509)	(0.00532)	(0.00514)	(0.00530)
Encouraged policy	-0.367	-0.00279	-0.00482	-0.00254
	(0.432)	(0.00484)	(0.00467)	(0.00475)
Observations	1,192,196	1,192,196	1,192,196	1,192,196
R^2	0.317	0.291	0.281	0.245

Panel B. First Differences (FD) – 2001 vs. 2003				
Dependent variable in header:	*FO share*	*FO dummy*	*MFO dummy*	*WFO dummy*
Restricted or prohibited policy	-0.639***	-0.00800**	-0.00443	-0.00843***
	(0.246)	(0.00383)	(0.00288)	(0.00303)
Encouraged policy	0.0613	0.00207	-0.000218	-0.00235
	(0.276)	(0.00364)	(0.00350)	(0.00362)
Observations	79,013	79,013	79,013	79,013
R^2	0.003	0.003	0.002	0.003

The table reports estimates of equation (5.15). The measure of foreign ownership FO used as a dependent variable is indicated in the header for each column. Panel A reports pooled OLS estimates for the full panel, including all covariates from column 4 of Table 5.1 and trade-weighted averages of input and output tariffs. Panel B reports FD estimates for the years 2001 and 2003, excluding firms that switch their 4-digit industry code over this period, and controlling for lagged TFP, $\ln K/L$, and TFP $\times \ln K/L$ (as observed in 2001), firm age, the first differences in the remaining covariates from column 4 of Table 5.1 and input and output tariffs, as well as 3-digit industry fixed effects. Standard errors clustered by 4-digit industry are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

Panel B of Table 5.2 focuses on the major reform of the FDI policy in 2002. In these regressions, the policy effects are identified from within-firm variation in a first-differenced (FD) version of equation (5.15), using only data for the years 2001 and 2003 and firms that did not switch their 4-digit industry code between these years. We find strong evidence that dropping a 'restricted' or 'prohibited' FDI policy triggered increases in foreign investment and acquisitions. The FD estimates for the effect of this policy category are negative for all measures of foreign ownership and statistically significant at the 5% level in all cases except for the *MFO dummy*. The policy liberalization increased the average foreign ownership share by 0.6 percentage points (or 5%) and the probability of being (wholly) foreign owned by around 4% (11%). Note that, despite our focus on the major policy reform, some of these estimates are smaller than in the full panel. The reason is that they identify only changes in foreign ownership *within* firms after controlling for contemporaneous changes at the 3-digit industry level through fixed effects. We cannot reject the hypothesis that encouraging FDI had no impact on foreign ownership.

[37] The estimated coefficients of the lagged control variables show the same pattern as discussed in the previous section for the unrestricted industries, and they are not reported to save space.

5.6 Effects of Ownership Restructuring on Firm Performance

5.6.1 Difference-in-differences model

We now turn to our empirical analysis of the paper's main question: How does the optimal allocation of ownership rights affect firm performance? To answer this question, we exploit the changes in foreign ownership triggered by the major FDI policy reform in 2002, which can be considered exogenous to Chinese firms. We compare the performance of firms that increased their foreign ownership share after the liberalization to a suitable control group by estimating the following difference-in-differences (DiD) model:

$$\ln \textit{performance}_{it} = \beta\, TX_i \times \textit{after}_t + \boldsymbol{\tau}\, \mathbf{T}_{jt} + \alpha_i + \alpha_t + \varepsilon_{it}. \tag{5.16}$$

This model examines the time variation in firm i's $\textit{performance}_{it}$ (in logs), most importantly the capital-labor ratio K/L and real output Y; other performance variables examined include the capital-wage bill ratio K/wL as an alternative measure of input intensity, the real capital stock K, employment L, and the real wage per employee w as measures of input quantities and quality, respectively. We restrict the main analysis to the years surrounding the major policy reform $t = \{2001, 2003\}$, and the dummy $\textit{after}_t$ takes a value of one for 2003 (and zero for 2001).[38] We also focus on the liberalized industries in which an FDI restriction (restricted or prohibited policy) is abolished in 2002. We define a firm in these industries as treated and set the alternative treatment group indicators TX_i to one if the firm's foreign ownership share increases across the respective acquisition threshold $X = \{0, 25, 50, 100\}$: (i) above 0%, reflecting the fact that FDI was entirely prohibited in some industries, (ii) to or above 25%, (iii) to or above 50%, reflecting the explicit equity requirement of *majority Chinese ownership*, or (iv) to 100%, reflecting the *Sino-Foreign joint venture* requirement (see Section 5.3.2). Note that these treatment groups are overlapping because some firms cross several ownership thresholds. In the following, we will refer to the firms for which the foreign ownership share increases across one of these thresholds as 'foreign-acquired' firms.[39] Arguably, the foreign investors increasing their ownership shares in these firms exploited the liberalization in FDI restrictions to reoptimize their ownership decisions. Provided that multinational firms are profit-maximizing, this change reveals a move to optimal ownership. The coefficient of the interaction term $TX_i \times \textit{after}_t$ is the main parameter of interest, denoted by β. Proposition 5.3 predicts $\beta > 0$ for the performance variables K/L and Y. $\mathbf{T}_{jt}$ is the same vector of input and output tariffs as in Section 5.5 (with coefficient vector $\boldsymbol{\tau}$), and ε_{it} denotes the error term. Importantly, the firm fixed effect α_i accounts for all time-invariant differences across firms (including initial conditions, industry affiliation, and treatment group status), and the year fixed effect α_t accounts for the first difference (i.e., the performance change in the control group).

What is the appropriate control group to which the treated firms should be compared? To obtain a valid estimate of β, the firms in the control group must be strictly comparable, such that their performance in 2003 provides a valid counterfactual scenario for the treated firms. More specifically, we require that the treatment is random (or strongly ignorable, in the terminology of the treatment effects literature)

[38] Bertrand et al. (2004) point out that serial correlation may invalidate inference in DiD estimations with multiple years of data. To avoid this issue, we follow their suggestion of focusing on a single pre-treatment and a single post-treatment year. The analysis first concentrates on short-term effects after 2 years, and we will subsequently consider also medium-term effects after 3-6 years.

[39] Firms switching their industry code between 2001 and 2003 are excluded from the analysis (as in the FD estimations of Section 5.5.3) because they may undergo major changes unrelated to ownership.

conditional on observables.[40] This requirement may be violated by two types of selection, at the firm level and at the industry level, respectively. First, it is a well-known fact that foreign investors engage in 'cherry picking' by acquiring the firms with the most favorable growth prospect (see e.g. Blonigen et al., 2014). Second, there may be selection of liberalized industries by the Chinese government, which had some discretion in selecting the industries to be liberalized within the scope of its commitments related to the WTO accession (see the discussion in Lu et al., 2017). We address these two margins of selection by (i) choosing a narrowly comparable control group, and (ii) applying propensity score reweighting, which will be explained in the next section.

Our preferred control group consists of firms that have already been foreign owned in 2001 and remain foreign-owned in 2003 (by the same ownership threshold as applied to the treated firms) within the liberalized industries. Such firms exist in the liberalized industries because they produce different products that were not specifically regulated by the FDI Catalogue, which classifies products more narrowly than the 4-digit industry codes at which we can map the policy to firms (see Section 5.3.2). For instance, we are considering firms in the liberalized 'cable and wire' manufacturing industry, where the treated firms producing a 'cable for high voltage' faced foreign ownership restrictions in 2001 and were acquired by foreign investors in 2003, after this restriction was abolished. Some firms producing other 'electronic cable' products were already wholly foreign-owned before 2002, since FDI in this product was not restricted, so they are included in our control group.

This choice of control group addresses both of the aforementioned selection concerns. First, it avoids the cherry-picking critique, since the firms in the control group have similarly been selected by foreign investors. Hence, they are likely to share the same (potentially unobservable) characteristics that also make the treated firms attractive for foreign ownership. Second, choosing the control group from within the liberalized industries avoids potential issues related to the government's selection of liberalized industries, which may not be fully exogenous to (expected) future performance. Furthermore, firms within the same 4-digit industries can be expected to feature similar technological characteristics and produce similar products as the treated firms, so they constitute a suitable choice as a control group. In Section 5.6.5, we also examine alternative control groups and show that our main conclusions do not hinge on this particular choice.

5.6.2 Propensity score reweighting

Note that the firm fixed effects in equation (5.16) account for all time-invariant differences across the treatment and control groups that affect future performance. However, there may further be selection of the treated firms based on dynamic factors, which might introduce bias into an estimate of the within-transformed DiD model if these dynamic factors are correlated with future firm performance. To account for this type of selection, we implement the inverse propensity score reweighting estimator that accounts for pre-treatment differences in observable characteristics across the treatment and control groups (see e.g. DiNardo et al., 1996; Imbens, 2004). The idea behind this approach is to reweight the firms in the control group such that the distribution of observable pre-treatment characteristics in this group matches

[40] Formally, strong ignorability of the treatment requires two assumptions, namely (i) ignorability in mean: $E\left[\ln Y_{i2003} - \ln \bar{Y}_i | \mathbf{M}_{i2001}, treated_i\right] = E\left[\ln Y_{i2003} - \ln \bar{Y}_i | \mathbf{M}_{i2001}\right]$, and (ii) overlap: $Pr(treated_i = 1 | \mathbf{M}_{i2001}) < 1 \, \forall \, \mathbf{M}_{i2001} \in \mathcal{M}_{2001}$, where $\mathcal{M}_{2001}$ is the support of the covariates $\mathbf{M}_{i2001}$ (Wooldridge, 2010). The first assumption requires that treatment is random conditional on the initial observable covariates $\mathbf{M}_{i2001}$, which we will define when introducing the propensity score reweighting approach in Section 5.6.2. The second assumption requires that for each treated firm and each covariate in $\mathbf{M}_{i2001}$, there must exist at least one untreated firm with a similar value. The latter condition is enforced below by excluding the few treated firms off the common support.

that of the treated firms. To the extent that these observable characteristics also account for unobservable differences across the two groups, the reweighting estimator can deliver a consistent estimate of the average treatment effect on the treated. To implement this approach, we estimate the DiD model in equation (5.16) by weighted least squares, assigning a weight of one to the treated firms and a weight of $\hat{p}/(1-\hat{p})$ to firms in the control group, where $\hat{p}$ is the propensity score (see Nichols, 2007, 2008).

The propensity score $\hat{p}$ is defined as the predicted treatment probability. It is estimated in logit regressions of the treatment indicators (one by one for each acquisition threshold) on the following set of pre-treatment firm characteristics observed in 2001, the covariates $\mathbf{M}_{i2001}$. To account for the key determinants of the ownership decision according to the PRT, we include the log of firm-level capital intensity K/L, log TFP, and the interaction of both variables (as in Sections 5.4 and 5.5). The choice of additional covariates follows similar applications in the literature (in particular, Wang and Wang, 2015, for China): We include output, employment, the average wage per employee (all in logs), the firm's age, its export share (exports over total sales), and the equity share owned by the state or collectives (*SOE share*). We further include the pre-treatment trend in log TFP (the change between 2000 and 2001) to account for the fact that foreign investors tend to acquire firms after they experienced a negative productivity shock (Blonigen et al., 2014). Finally, the logit model includes 2-digit industry fixed effects to account for differences in the attractiveness of different industries for FDI. In Section 5.6.7, we show that our results are robust to variations in this choice of covariates.

Matching and reweighting methods based on the propensity score go back to Rosenbaum and Rubin (1983) and DiNardo et al. (1996) and are well-established in the literature estimating treatment effects in non-experimental settings. Recent applications in the FDI literature include Arnold and Javorcik (2009), Guadalupe et al. (2012), and Wang and Wang (2015). Our approach, which combines the DiD regression with propensity score reweighting, features three advantages over alternative available methods. First, the estimator allows us to control for the contemporaneous changes in tariffs $\mathbf{T}_{jt}$, which might also affect performance, in the multivariate regression framework. Second, it has been shown by Busso et al. (2014) that reweighting is typically superior to matching when the overlap in terms of observable covariates between treatment and control groups is good, as in our application. Third, a very attractive feature of the combined approach is its double robustness property, which means that it can deliver consistent estimates as long as either the regression model or the propensity score model is correctly specified (Imbens, 2004).

5.6.3 Main estimation results

Table 5.3 presents the main results of our empirical analysis: The reoptimization of ownership shares after China's FDI liberalization caused large and significant increases in the capital-labor ratio and output of the acquired firms. The table shows estimates of β in the DiD model of equation (5.16) for the different dependent variables indicated in the header. The four rows implement alternative definitions of treatment based on the four different acquisition thresholds ($>$0%, $\geq$25%, $\geq$50%, and 100%), so each estimate reported in the table derives from a separate regression. As described in Section 5.6.1, the estimates compare the performance change of foreign-acquired firms between 2001 and 2003 to our preferred control group of firms that were already foreign-owned in 2001 (and remained so in 2003) within the liberalized industries. Depending on the definition of treatment, there are between 123 and 174 firms in the treatment group. To account for initial observable differences that may affect treatment, the control group observations have been reweighted by the estimated propensity score $\hat{p}$, as described

in Section 5.6.2; the estimates of $\hat{p}$ and covariate balancing tests are discussed below.

We first discuss the effects of foreign acquisitions on input ratios, summarized in columns 1 and 2 of Table 5.3. As predicted by Proposition 5.3, the capital-labor ratio K/L increases after foreign acquisitions (column 1). Since this effect is estimated relative to a reweighted control group of comparable, foreign-owned firms, it can be interpreted as the consequence of optimal ownership restructuring after the FDI liberalization. Intuitively, integration increases the firm's capital-labor ratio because the foreign investor's incentives to contribute to the firm's capital stock increase, whereas the Chinese firm's incentives to optimally hire labor diminish. In line with this argument, the increase in K/L is highest for the high acquisition thresholds of 50% and 100%. These effects are statistically significant at the 5% confidence level and they are large in economic terms. The estimated coefficient implies an increase of 27% in K/L for wholly foreign-acquired firms.[41] Note that the effects are smaller and insignificant when we define treatment by the foreign ownership threshold of 25%, which constitutes the legal threshold for receiving official FIE status under Chinese law. This finding might be explained by the fact that this threshold was never legally binding, whereas the other thresholds reflect explicit ownership restrictions that were initially in place for a subset of the liberalized industries. Furthermore, firms with official FIE status receive preferential tax treatment, so firms crossing the 25% threshold might do so for tax reasons rather than considerations about incomplete contracts. The increase in the input ratio is confirmed and the estimated magnitudes are even larger when we measure labor input by the real wage bill (K/wl, column 2).

Table 5.3: *Effects of optimal ownership restructuring on firm performance: 2001 vs. 2003*

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	K/L	K/wL	K	L	w	Y	N (1-3)/(4-6)
T0 × *after*	0.138**	0.153*	0.268***	0.130**	-0.0149	0.294***	3,512/3,518
	(0.0648)	(0.0827)	(0.0636)	(0.0585)	(0.0621)	(0.0639)	
T25 × *after*	0.0336	0.0742	0.214***	0.181***	-0.0405	0.254***	3,198/3,204
	(0.0635)	(0.0800)	(0.0585)	(0.0577)	(0.0616)	(0.0619)	
T50 × *after*	0.187**	0.296***	0.248***	0.0615	-0.109	0.146*	2,266/2,270
	(0.0742)	(0.103)	(0.0655)	(0.0627)	(0.0812)	(0.0805)	
T100 × *after*	0.239***	0.318***	0.209***	-0.0485	-0.0821	-0.0810	1,416/1,420
	(0.0715)	(0.100)	(0.0642)	(0.0573)	(0.0771)	(0.0796)	

Note: The table reports estimates of β in equation (5.16) for the years 2001 and 2003, with the dependent variables indicated in the header (all in logs). Each coefficient estimate in the table derives from a separate regression, which controls for weighted input and output tariffs as well as firm and year fixed effects. Treated firms are defined as those increasing their foreign ownership share between 2001 and 2003 to or across different thresholds $X = \{0, 25, 50, 100\}$ in the liberalized industries (where FDI was prohibited or restricted before 2002, but not after). The control group consists of firms that are foreign-owned (by the same thresholds as the treated firms) in both years within the liberalized industries, and they are reweighted by the estimated propensity score as described in Section 5.6.2. The last column lists the number of observations N for columns 1-3 and 4-6 in each row. Standard errors clustered by firm are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

Columns 3 to 5 provide direct evidence concerning the effects of ownership restructuring on the inputs in line with our PRT model. The estimated positive effect on K/L and K/wL is driven first and foremost by a substantial increase in the real capital stock for foreign-acquired firms, irrespective of how we define the acquisition threshold (column 3). At the same time, the evidence on employment is

[41] This number is computed as $[\exp(0.239) - 1] \times 100\%$ from the estimate in column 1 of Table 5.3.

mixed (column 4). Labor input increased only for firms crossing the low ownership thresholds, but not for those acquired by at least 50% by foreign investors. This ambiguous effect can be rationalized by the countervailing effects discussed in Section 5.2.3: the negative incentive effect reduces labor provision by the Chinese firm, while the positive size effect increases L. However, the head count of employees L might not adequately measure the aggravated underinvestment into labor. Instead of reducing the number of workers, the firm might reduce its effort into hiring and training of workers, as these activities are much more difficult to enforce in a contract. Such underinvestment might be visible in the wage per employee as a proxy for the average quality of the workforce. In line with this argument, we find that the wage weakly decreased in the acquired firms relative to the control group, though these effects are insignificant (column 5). Overall, our findings provide support for our PRT model, which predicts that the foreign multinational's investment incentives into relationship-specific capital should increase, while the Chinese firm's incentives for labor provision may deteriorate after an increase in foreign ownership towards the optimum. As a result of these shifts in input provision, the capital-labor ratio increased.

Note that our results highlight an important identification challenge in testing the PRT. A popular approach in the international economics literature has used the capital-labor ratio as a proxy for the headquarter intensity parameter η from the PRT in order to explain the relative prevalence of vertical integration (or intra-firm trade). However, K/L is endogenous to the ownership decision, as we have demonstrated theoretically in Section 5.2.2 and empirically in this section: Deeper integration increases K/L because it tilts the investment incentives towards the foreign owner, which typically contributes to the firm's capital stock but is less involved in hiring labor. Regressing measures of integration on the firm-specific capital-labor ratio in an attempt to test the PRT hence suffers from a reverse causality issue that follows from the same theoretical framework.[42]

We quantify the performance gains from optimal ownership restructuring in terms of real output Y in column 6 of Table 5.3. Based on our PRT model, we expect that the increase from the restricted towards the optimal ownership share after FDI liberalization leads to overall efficiency gains and raises output in the acquired firms. Indeed, we find that the optimal allocation of ownership rights can have large positive effects on firm performance. Firms increasing their foreign ownership share across the thresholds of 0%, 25%, or 50% between 2001 and 2003 experienced significant output gains over this period, as predicted by Proposition 5.3. Somewhat surprisingly, there is no evidence for short-term output gains by 2003 for firms that are fully acquired, an observation that we will return to below. Note that the effects identified for the lower acquisition thresholds are economically sizable. The move to optimal ownership allowed these firms to increase their production by 16-34% relative to the highly comparable, reweighted control group. These findings show that the optimal choice of ownership matters substantially for firm performance.

Two considerations might rationalize why the effects on firm output are insignificant for firms that moved to 100% foreign ownership. First, it is conceivable that some foreign investors choose too high ownership shares immediately after liberalization, for instance due to information frictions. Within our PRT framework, one would expect that the resulting reduction in investment incentives of the Chinese firm (in line with the estimated effects on L and w) tends to depress output, in particular for higher ownership shares. However, the model predicts that for a move to the optimal ownership share, the increased incentives by the foreign owner should typically outweigh these negative effects and lead to a

[42] To ameliorate this concern, we have hence used industry-level and lagged firm-level measures of capital intensity as proxies for η when exploring correlations in Section 5.4.

net increase in production. Hence, the fact that the output effect is non-positive for the 100% threshold might indicate that some of the foreign investors have chosen excessively high ownership shares, thereby exacerbating the production inefficiency. In line with this conjecture, the average foreign ownership share of these firms drops from 100% to 73% in 2004 and remains stable at this level over the subsequent years (among the firms that continue to be observed). This pattern might reflect a correction of the initial ownership decision. Indeed, if we restrict the treatment group for $T100$ to firms that remain wholly foreign owned over the subsequent two years, we find a positive estimate corresponding to output gains of 9.2%, which remains insignificant however. Second, beyond our static model, there may exist short-term adjustment costs from organizational restructuring and adjustment processes within the firm, which delay some of the efficiency gains. This issue may be particularly pronounced in wholly acquired firms, where the ownership change is most radical and hence the adjustment costs are largest. We provide some evidence supporting this idea when considering medium-term effects in Section 5.6.4.

How well does the propensity score reweighting estimator account for initial differences across firms in the treatment and control groups? The estimates of the logit models used to obtain the propensity scores are summarized in Table 5.B.1 (columns 1-4) in Appendix 5.B. They show that most of the selected covariates are individually significant in several logit regressions, suggesting that the treatment and control groups differ significantly in terms of these pre-treatment characteristics. They jointly explain a relevant part of the acquisition probability, reflected in pseudo R^2 values in the range of 0.082-0.156. Note that, consistent with the PRT, the treated firms are initially less capital intensive than the already foreign-owned firms, which form the control group. These initial observable differences are successfully accounted for by the propensity score reweighting approach. To illustrate this, the second half of Table 5.B.1 (columns 5-8) shows the same logit regressions reweighted by the estimated propensity score. It reveals that none of the covariates remains individually significant in these weighted regressions. The approach is also successful in finding a similar firm from the control group for each firm in the treatment group. There are very few exceptions of treated firms whose covariates lie off the common support and which have therefore been excluded from the main analysis (one to five firms, depending on the definition of treatment). The pseudo R^2 drops below 0.01 in the reweighted logit regressions, indicating that the initial characteristics have no explanatory power for the treatment after reweighting.

The success of the propensity score reweighting approach is also confirmed by the standard covariate balancing tests, which compare sample averages of the pre-treatment characteristics before and after reweighting (not reported). For all treatment definitions and all but one of the covariates, t-tests cannot reject the null hypothesis of zero differences in the reweighted sample means at any conventional significance level. The sole exception is the *SOE share*, which differs substantially across the two groups by construction, since we have imposed a minimum requirement on the initial foreign ownership share in the control group. This difference cannot be fully eliminated by reweighting, but it is no longer significant conditional on the other covariates, as shown in Table 5.B.1. Most importantly, initial differences in the main performance variables of interest are substantially reduced: The remaining absolute bias ranges between 1.1% and 6.4% for the initial capital-labor ratio, it is never greater than 10.4% for output, and generally insignificant for all outcome variables considered.

5.6.4 Medium-term effects

In the previous section, we have concentrated on the immediate impact of ownership restructuring in the first year after the FDI liberalization. However, this restructuring process may be costly in the short

term, and the subsequent efficiency gains within the firm may take some years to materialize. For these reasons, the short-term effects estimated above might underestimate the gains from the optimal allocation of ownership. To investigate the medium-term effects, we apply the DiD model to performance gains between 2001 and the subsequent years 2004-2007 (for which firm data are available).

In each estimation, we continue to focus on one pre-treatment year (2001) and one post-treatment year (each of the years 2004-2007) to allow for valid inference (see Bertrand et al., 2004). The DiD model is specified as in equation (5.16), with the $after_t$ dummy equal to one in the post-treatment year, and the treatment dummies TX_i defined by firms increasing their foreign ownership across the different thresholds between the two years. This approach allows for both ownership restructuring and firm performance to respond with some delay to the liberalization. Also, we continue to combine the DiD model with the propensity score reweighting approach to account for initial differences across firms that might affect selection into treatment.

We first examine the medium-term impact on various performance variables before studying the full evolution of the output effects over time. Table 5.4 summarizes the effects on inputs and output over the five-year period between 2001 and 2006. We find that the estimated effects on the input ratios K/L and K/wL are positive also over the medium term in almost all estimations, but they are much smaller than over the short term and typically insignificant (columns 1-2). Instead, we find statistically significant and economically large positive effects on the quantities of both inputs K and L in the foreign-acquired firms defined by the thresholds of 0%, 25%, and 50% (columns 3-4). The medium-term effect on wages is negative, as over the short term, but it remains mostly insignificant (column 5). The medium-term effects on input ratios, quantities, and wages suggest that the positive short-term effect on the factor ratio between 2001 and 2003 identified above is not due to a lasting change in the firm's technology towards more capital intensive production. Instead, the results are in line with the logic of our PRT model: Optimal ownership restructuring triggers immediate adjustments in the investment incentives of both the local firm and the foreign investor, which lead to efficiency gains, an increase in output, and hence more use of both K and L over the medium term. The weakly negative wage effects corroborate our expectation that the quality of labor input provision by the Chinese firms tends to deteriorate after acquisitions. Most importantly, the output gains increase substantially over the medium term (column 6).

Table 5.4: *Medium-term effects of optimal ownership restructuring: 2001 vs. 2006*

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	K/L	K/wL	K	L	w	Y	N (1-3)/(4-6)
T0 × *after*	0.110	0.130	0.424***	0.302***	-0.0226	0.511***	2,536/2,550
	(0.0841)	(0.0872)	(0.0920)	(0.0698)	(0.0784)	(0.104)	
T25 × *after*	-0.0231	0.161*	0.288***	0.300***	-0.184**	0.410***	2,344/2,360
	(0.106)	(0.0932)	(0.106)	(0.0735)	(0.0916)	(0.0991)	
T50 × *after*	0.113	0.131	0.429***	0.315***	-0.0179	0.355***	1,726/1,734
	(0.0945)	(0.111)	(0.103)	(0.0841)	(0.0656)	(0.121)	
T100 × *after*	0.0836	0.137	0.0492	-0.0428	-0.0571	0.158	1,112/1,120
	(0.0901)	(0.103)	(0.0843)	(0.0757)	(0.0673)	(0.111)	

Note: The table reports estimates of β in equation (5.16) for the years 2001 and 2006, with the dependent variables indicated in the header (all in logs). The definitions of the treatment and control groups, the control variables, and the reweighting approach are analogous to the main estimations (see the note to Table 5.3). The dummy variable $after_t$ is redefined to equal zero in 2001 and one in 2006. Standard errors clustered by firm are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

Figure 5.5: *Output effects over time*

Note: The figure illustrates estimates of β in equation (5.16) for the pre-treatment year 2001 and alternative post-treatment years 2003-2007, with log output as the dependent variable. The definitions of the treatment group (indicated by the alternative dummy variables $TX_i, X = \{0, 25, 50, 100\}$), the control group, the control variables, and the reweighting approach are analogous to the main estimations (see the note to Table 5.3). The dummy variable $after_t$ is redefined to equal zero in 2001 and one in the respective post-treatment year.

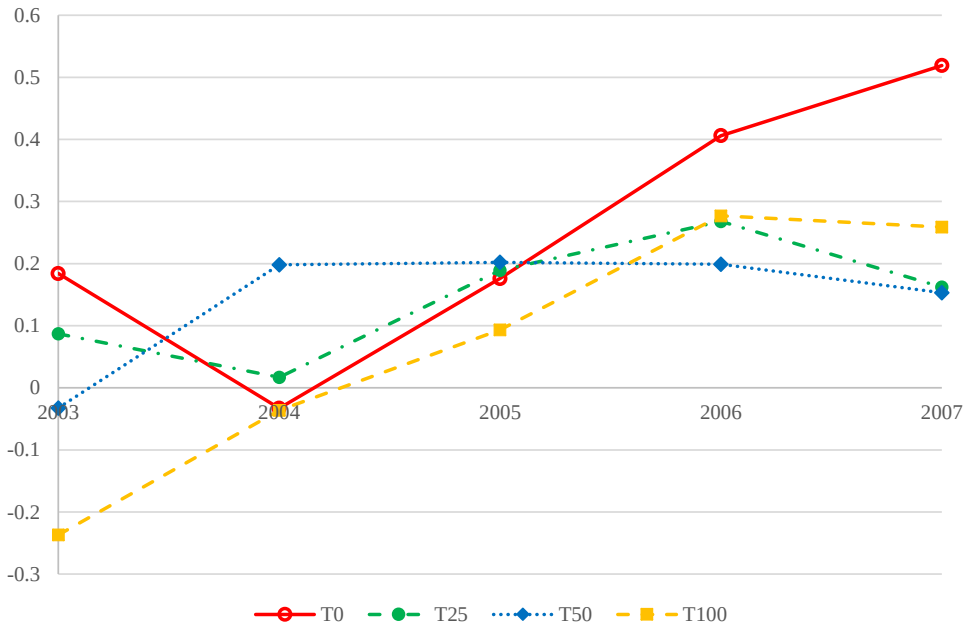
Figure 5.5 focuses on the effect of ownership on output, our main performance variable of interest, and illustrates the annual evolution of the estimated output gains over time. Note that each of the point estimates of β in the figure is based on a separate DiD estimation. The output gains show an increasing tendency over time, which is most clearly visible for the highest and the lowest acquisition thresholds. The increase over time is substantial. The average output gains by 2007 are twice as large as the short-term effect in 2003 for acquisitions defined by the 0% and 50% thresholds. Over the medium term of 4-6 years, the point estimate for the output effect is also positive for firms reaching the 100% threshold, though it remains statistically insignificant throughout. For all other acquisition thresholds, the output gains are significantly positive in each of the years 2004-2007. These findings confirm our conjecture that the output gains from ownership reallocation over the medium-term exceed the initial effect identified in the main analysis. The figure further suggests that optimal ownership restructuring might even have favorable dynamic effects on output growth rather than just a one-time level effect.

The analysis so far has focused on output as our preferred measure of firm performance. As discussed in Section 5.2.3, this choice has the two key advantages that our PRT model delivers a clear prediction for the output effect and this measure is easily obtained from the data. An alternative performance measure typically considered in the literature is TFP. Should the optimal ownership restructuring after liberalization also increase firm productivity? The answer to this question depends on the definition of TFP. Note that in our model, TFP is defined as a technology parameter determined in a stochastic innovation process at the entry stage, which is exogenous to subsequent firm decisions and does therefore *not* react to ownership changes. Instead, the efficiency gains in the model materialize purely through reoptimized input choices, conditional on TFP as we have defined it so far. Alternatively, one may adapt a broader definition of TFP encompassing also other features of the production process that allow firms to produce

more output using the same quantity of inputs. Such features may include, for instance, the suitability or in-time delivery of inputs. If the decisions on these features are costly, relationship-specific, and non-contractible, they may be subject to underinvestment and hence endogenous to ownership. However, the predicted effect of ownership changes then depends on which party makes these non-contractible decisions. In the next step, we consider the possibility that our empirical measure of TFP reflects also these features of the production process, and we examine the effects of optimal ownership restructuring on TFP. Since these possible adjustments to the production process are likely to take time, we examine the evolution of productivity effects over several years.

Figure 5.6 illustrates that there are productivity gains from optimal ownership restructuring, which materialize after 4-6 years. The effects on TFP are estimated year by year in analogy to the output effects in Figure 5.5. Over the short term of 2-3 years, the effects on TFP are small and vary around zero for the four alternative definitions of treatment. However, the point estimates tend to increase over time, as in the case of output, and they are all positive 4-6 years after the liberalization. These productivity gains are significant at the 5% level in 2006 for all treatment definitions except for the majority threshold ($T50$), and they reach up to 68% in 2007. We obtain similar results when using labor productivity as an alternative dependent variable (not reported). Within our model, these findings can be interpreted as efficiency gains from optimal ownership that improve the production process beyond the reoptimized choice of input quantities. These additional gains take some years to materialize.

Figure 5.6: *Productivity effects over time*



Note: The figure illustrates estimates of β in equation (5.16) for the pre-treatment year 2001 and alternative post-treatment years 2003-2007, with log TFP as the dependent variable. The definitions of the treatment group (indicated by the alternative dummy variables TX_i, $X = \{0, 25, 50, 100\}$), the control group, the control variables, and the reweighting approach are analogous to the main estimations (see the note to Table 5.3). The dummy variable $after_t$ is redefined to equal zero in 2001 and one in the respective post-treatment year.

5.6.5 Alternative control groups

Our baseline choice of the control group, consisting of foreign-owned firms in the liberalized industries, has a number of advantageous features discussed above, but it is only one of several possible choices. This section demonstrates that our main insights are remarkably robust to alternative choices of the control group. These variations further allow us to address possible challenges to identification, relate our findings to the existing literature, and strengthen our main conclusions. To this end, we return to the main DiD analysis for 2001 vs. 2003 and consider as alternative control groups: (i) foreign-owned firms in unregulated industries, (ii) major exporters within liberalized industries, (iii) domestic-acquired firms within liberalized industries, and (iv) foreign-acquired firms in unregulated industries. The fourth exercise constitutes our most stringent and important robustness check.

To consistently estimate the average treatment effect of foreign acquisitions on the treated, firms in the control group should not be affected by the treatment. This requirement is sometimes referred to as the stable unit treatment value assumption (SUTVA). The SUTVA may be challenged by the existence of spillovers from FDI to other domestic firms. As argued by Lu et al. (2017), domestic firms in China may be indirectly affected by the inflow of foreign capital, e.g. through increased competition in the domestic output market. Note that for our main analysis this concern is ameliorated by the fact that only a small number of firms are acquired between 2001 and 2003, and they produce different products than the control group firms, which were already foreign-owned before the liberalization. Furthermore, the firms in this preferred control group export on average 46-57% of their sales (depending on the applied ownership threshold), so they compete mostly on world markets rather than in the domestic market, which limits the scope for spillovers through competition.

To address remaining concerns related to spillovers, we conduct a cross-industry comparison by choosing firms that are foreign-owned in both years in unregulated industries (that were not regulated by the FDI Catalogue, neither in 2001 nor in 2003) as an alternative control group. These firms certainly produce different products and are hence very unlikely to compete directly with the treated firms. The approach allows us to relax the standard assumption by assuming that the SUTVA holds across industry clusters, but not necessarily within clusters (similar to Girma et al., 2015). In our cross-industry specifications, we modify the logit model for estimating the propensity score used in reweighting to account for the possibility that the Chinese government selected the liberalized industries. Following the considerations in Lu et al. (2017), we augment the logit model used to estimate the propensity scores by the following set of covariates specific to the 4-digit industry: the number of firms, average employment, and average wage per firm (all in logs), the average firm age, and the total exports to sales ratio.[43] From the logit model, we also omit the 2-digit industry fixed effects, which would otherwise fully explain the treatment in some cases. The estimation results reported in columns 1-2 of Table 5.5 confirm our main findings. The effects of foreign acquisitions on the capital-labor ratio are significantly positive and larger for treatment defined by the higher acquisition thresholds, and positive and significant output gains are identified for the low thresholds.

An alternative way of addressing concerns related to spillovers, while maintaining the focus on narrowly comparable firms within the liberalized industries, is by choosing major exporters as a control group. Since these firms are competing on world markets, spillovers through competition in the domes-

[43] This selection of industry-level matching covariates is motivated by possible considerations by the Chinese government in relaxing the FDI restrictions, such as promoting exports, protecting infant industries, and avoiding adverse effects on the labor market. Note that Lu et al. (2017) adopt a broader definition of liberalized industries, which includes for instance also those industries where FDI was newly encouraged. Hence, their preferred set of covariates differs slightly from ours.

Table 5.5: *Alternative control groups*

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Control group in header:	Foreign-owned (unregulated)		Major exporters (liberalized)		Domestic-acquired (liberalized)		Foreign-acquired (unregulated)	
	K/L	Y	K/L	Y	K/L	Y	K/L	Y
T0 × after	0.127**	0.336***	0.263***	0.234**	0.156*	0.361***	0.165**	0.312***
	(0.0575)	(0.0756)	(0.0796)	(0.104)	(0.0820)	(0.0725)	(0.0700)	(0.0984)
N	21,594	21,628	808	810	1,606	1,606	1,864	1,870
T25 × *after*	0.0842*	0.283***	0.164**	0.220**	0.109	0.286***	0.0690	0.273***
	(0.0480)	(0.0735)	(0.0743)	(0.0919)	(0.0845)	(0.0707)	(0.0590)	(0.0855)
N	19,314	19,346	898	900	1,650	1,650	2,044	2,050
T50 × *after*	0.237***	0.0919	0.290***	-0.00586	0.197*	0.179**	0.189**	0.156
	(0.0567)	(0.0927)	(0.0719)	(0.0852)	(0.101)	(0.0845)	(0.0822)	(0.104)
N	14,046	14,074	954	956	1,582	1,582	1,608	1,610
T100 × *after*	0.208***	-0.0235	0.224***	-0.0181	0.268	0.109	0.139**	0.164
	(0.0578)	(0.122)	(0.0753)	(0.0854)	(0.193)	(0.0998)	(0.0678)	(0.155)
N	9,226	9,254	1,174	1,178	1,622	1,624	1,642	1,648

Note: The table reports estimates of β in equation (5.16) for the years 2001 and 2003, with the dependent variables indicated in the header (all in logs). The definitions of the treatment group and the control variables are analogous to the main estimations (see the note to Table 5.3). The control group consists of foreign-owned firms in unregulated industries (in both years) in columns 1-2, firms exporting the majority of their sales (in both years) in liberalized industries in columns 3-4, domestic-acquired firms (identified by changes in registration types as in Wang and Wang, 2015) in liberalized industries in columns 5-6, and foreign-acquired firms (by the same thresholds as the treated firms) in unregulated industries in columns 7-8. The reweighting approach in columns 1-2 and 7-8 includes industry-level covariates and excludes industry fixed effects relative to the baseline approach, which is used in columns 3-6 and described in Section 5.6.2. Standard errors clustered by firm are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

tic product market are likely to be negligible. We define major exporters as firms that export the majority of their sales both in 2001 and 2003, and use them as the control group in DiD regressions restricted to liberalized industries. To maintain a sufficiently large pool of comparable firms in the control group, we do not maintain the additional restriction of foreign ownership, but instead keep all major exporters that are *not* foreign owned in both years. Since we have implemented a strict minimum requirement on the export share in the control group, we exclude this variable from the (otherwise unchanged) set of covariates used to estimate the propensity score, by which observations are reweighted as in the main analysis. The estimation results in columns 3-4 of Table 5.5 strongly confirm the positive effects of foreign acquisitions after liberalization on K/L and also (for low acquisition thresholds) on output.

Wang and Wang (2015) suggest that for identifying the causal effects of foreign ownership, domestic-acquired firms constitute a suitable control group for foreign-owned firms. They argue that domestic-acquired firms undergo a similar process of ownership restructuring which might involve adjustments within the firm that are unrelated to FDI. Hence, these firms share many important features with the treatment group but not the 'foreignness' of the investor. Note that our intention in this paper is not to identify the distinguishing features of foreign compared to domestic acquisitions; by contrast, our main objective lies in quantifying the gains from an optimal allocation of ownership rights. This question does not hinge on the nationality of the investor per se; our approach merely exploits the fact that foreign investors were previously restricted and hence their ownership share is initially suboptimal and can increase to the optimal level due to FDI liberalization. In light of our predictions derived from the PRT, one might expect qualitative adjustments within domestic-acquired firms that are similar to those triggered by foreign acquisitions, as long as these domestic acquisitions constitute a reallocation

of ownership rights towards the optimum. However, since policy (and other) restrictions on domestic ownership did not change between 2001 and 2003, the driving force behind contemporaneous domestic acquisitions remains opaque. By contrast, for firms acquired by foreign investors in the liberalized industries, we clearly expect that initial inefficiencies due to contractual frictions are reduced by the optimal reallocation of ownership rights. Therefore, one might expect the performance gains induced by foreign acquisitions in these industries to exceed those caused by domestic acquisitions.

We identify domestic acquisitions in the data from changes in firms' registration types following the baseline approach taken by Wang and Wang (2015). We then compare the performance of foreign-acquired firms to domestic-acquired firms within liberalized industries in DiD regressions of equation (5.16), while reweighting observations by the estimated propensity score (in analogy to the main analysis). The estimation results in columns 5-6 of Table 5.5 reveal large positive point estimates of β for K/L as the dependent variable, but these are only weakly significant in two cases and insignificant in the other two cases. This finding is not too surprising provided that domestic acquisitions are also likely to involve ownership restructuring towards a more efficient allocation of property rights, and since it seems less clear which stakeholder is involved in the labor and capital input choices. More strikingly, we find strong evidence for positive output effects from foreign acquisitions relative to domestic acquisitions. The estimated output gains are positive throughout, even larger than in the main analysis, and significant at the 5% level for all but the highest acquisition threshold. These results suggest that foreign acquisitions of previously restricted firms are associated with output gains relative to other ownership changes, in line with our expectations based on the PRT model. However, these effects might also be partially attributable to benefits of foreign ownership per se.

As a last exercise, we compare the treated firms, which are foreign-acquired in liberalized industries, to *other foreign-acquired* firms in unregulated industries. Note that this novel approach combines two key advantages of the previous exercises and constitutes an even more ambitious identification strategy. In this setup, the treatment and control group are identical both in terms of the foreignness of the investor and the dynamic adjustment processes that might be triggered by any ownership changes. Since we further account for initial observable differences across firms and industries by the propensity score reweighting approach (implemented as described for the cross-industry specifications in columns 1-2), the only salient difference between the treatment and control group firms in this setup is whether FDI was liberalized in their industry or not. Should we expect to find any performance differences between these two groups of firms? Potentially yes, because the efficiency losses caused by the misallocation of ownership rights prior to liberalization are likely to be larger in liberalized industries. Whatever motivates increased foreign ownership in the control group firms, it is unlikely to be removing a similar distortion as in the liberalized industries, since the foreign ownership share could already be chosen freely in the unregulated industries in 2001.

The results of this important robustness check strongly confirm our theoretical predictions. The estimated effects of foreign acquisitions on K/L and Y in liberalization vs. unregulated industries are summarized in columns 7-8 of Table 5.5. We find significantly positive effects on the capital-labor ratio of around 15-21% for all definitions of treatment except for the 25% acquisition threshold. The point estimates for output imply large output gains of 17-37%, though these are only statistically significant

for the lower acquisition thresholds.[44] These estimates ensure that the observed performance effects are not due to the foreignness of the investor or due to some dynamic adjustment that might take place in any firm changing ownership. They corroborate our hypothesis that it is indeed the optimal reallocation of ownership rights which caused the shift in input ratios and the ensuing output gains. Hence, these findings lend strong empirical support to Proposition 5.3 and they confirm that the efficiency gains from an optimal allocation of ownership rights are quantitatively substantial, potentially even exceeding 30% in terms of short-term output gains, as shown in the main analysis.

5.6.6 Discussion

Our main empirical analysis has identified and quantified the effects of a reallocation of ownership rights towards the optimum on firm performance. Guided by our theoretical model of Section 5.2, we have interpreted the estimated effects as consequences of a reduction in inefficiencies arising from incomplete contracts. This interpretation of our results seems warranted in light of the fact that both general patterns of foreign ownership in unregulated Chinese industries (see Section 5.4) as well as the changes in input ratios induced by the acquisitions (see Section 5.6.3) are in line with the predictions of our property-rights theory of the multinational firm. However, it should be noted that our findings are more general than this specific model setup. For instance, one may envision an alternative theory of the multinational firm based the transaction-cost approach (see e.g. Grossman and Helpman, 2003, 2004, 2005), in which ownership can entirely eliminate underinvestment problems due to contractual frictions and hence increase efficiency and firm output. However, such a model would not necessarily predict the increase in capital-labor ratios consistent with the PRT, for which we have found robust support in the data. More generally, we view our results as quantifying the performance gains from optimal ownership reallocation.

Can alternative mechanisms, unrelated to optimal ownership restructuring explain our findings? One might conjecture that the performance effects of foreign acquisitions identified above derive in part from other benefits of foreign ownership, such as technology transfers, improved financial conditions, or foreign market access. If the foreign investor simply transfers a better and more capital-intensive technology to the Chinese firm, this might explain the increase in output and the capital-labor ratio. Also, the literature has shown that MNEs improve the financial conditions in acquired firms, e.g. by providing access to internal capital markets (see Alfaro and Chen, 2012; Desai et al., 2004; Manova et al., 2015). Guadalupe et al. (2012) further argue that access to foreign markets via the parent company can foster innovation activities within foreign-acquired firms. We have three reasons to believe that the performance effects identified in this paper cannot be explained by these mechanisms alone and are indeed caused by optimal ownership restructuring.

First, our evidence from foreign-acquired firms in China is in line with the view that foreign investors contribute to the capital stock of their subsidiaries – either through direct provision of capital goods or (co-)financing of investments. However, we view the provision of finance not as an alternative explanation, but as a direct consequence of the shift in investment incentives predicted by the PRT. Note that in principle, in the absence of contractual imperfections, the foreign firm should have similar incentives

[44] An alternative approach to identifying this effect exploits a triple difference specification, i.e., a variant of equation (5.16) that includes a triple interaction term $foreign\text{-}acquired_i \times liberalized_j \times after_t$ (along with the self-explanatory indicator variables and all two-way interactions as covariates) in the sample of all firms in liberalized and unregulated industries. To account for both selection by investors (within and across industries) and selection by the government (across industries), the observations are reweighted by a propensity score estimated from a logit regression of the triple interaction term on the same pre-treatment industry and firm covariates as before. Such triple difference estimations (not reported) lead to very similar conclusions.

to provide credit to its Chinese supplier whether or not it acquires any shares in the supplier. Our PRT model provides an explicit rationale for why the foreign investor's incentives to provide relationship-specific investments into the supplier's capital stock improve after an increase in the foreign ownership share. Therefore, we view the provision of capital as one of the channels through which the optimal reallocation of ownership rights can boost firm output, as predicted by our theoretical model.

Second, we find no evidence in our data that foreign acquisitions increased access to export markets, and thereby induced innovative activity in Chinese firms. We test directly for this mechanism by using exports (in logs or normalized by sales) and the share of revenue in new products (in levels or logs, as a measure of product innovation) as alternative dependent variables in the DiD model of equation (5.16). If market access through foreign ownership and innovation were important for the performance gains, we would expect to find positive effects in these regressions. However, in these estimations (which are not reported to save space), we do not find any significant positive effects of foreign acquisitions on exports or product innovation. Therefore, export market access through foreign acquisitions, which Guadalupe et al. (2012) have found to be important for Spanish firms, cannot explain the firm-level output gains that we have identified in this paper.

Third, and most importantly, we have found similar effects on K/L and Y as in our main analysis even when comparing the treated firms to other foreign-acquired firms in unregulated industries (see columns 7-8 of Table 5.5). If benefits related to the foreignness of the investor were responsible for the observed capital and output gains, we should not find any significant performance differences across foreign-acquired firms in liberalized and other industries.[45] The reason is that better technology, finance, or market access through foreign ownership should similarly benefit firms irrespective of previous ownership restrictions. The fact that foreign-acquired firms in liberalized industries improved their performance significantly compared to other, contemporaneously foreign-acquired firms cannot be explained by these alternative mechanisms. These findings constitute strong evidence that the performance gains found in this paper are indeed caused by the organizational restructuring from a previously inefficient allocation of property rights (under the policy restrictions) towards optimal ownership.

5.6.7 Robustness checks concerning propensity scores

Table 5.B.2 in Appendix 5.B demonstrates that our main result are robust to varying the set of covariates in the logit model used to estimate the propensity score. The table reports estimates of the main DiD model for 2001 and 2003 with the baseline configuration of treatment and control groups, while varying the reweighting approach across different columns. Since Manova et al. (2015) argue that foreign investors tend to acquire more financially vulnerable firms, we account for this potential dimension of selection by including the following measures for the firm's initial financial conditions in the logit model in columns 1-2 (following Wang and Wang, 2015): the liquidity ratio $\big((\textit{current assets} - \textit{current liabilities})/\textit{total assets}\big)$ and the leverage ratio $(\textit{total liabilities}/\textit{total assets})$. To account for observable differences that may describe the varying attractiveness of different liberalized industries for FDI, we include in columns 3-4 the same 4-digit industry-level covariates as in the cross-industry analysis of Section 5.6.5. Since the evolution of performance variables other than TFP may signal an attractive target for foreign investors, we add pre-treatment trends in the main performance variables Y and K/L

[45] Note that observations in these cross-industry regressions have also been reweighted to account for initial differences in covariates, which include the firm's size, capital intensity, and industry characteristics. Hence, the effects cannot be attributed to higher initial financial vulnerability of the firms in liberalized industries.

(defined as changes between 2000 and 2001 in logs) in columns 5-6. Finally, we account for varying regional attractiveness for FDI through province fixed effects in columns 7-8. In all of the resulting regressions, we find the baseline results to be qualitatively unchanged and also the magnitudes of the estimated effects on K/L and Y are in the same ballpark as in the main analysis.

5.7 Conclusions

In this paper, we have provided a first quantification of the gains in firm performance that can be achieved by the optimal allocation of ownership rights within multinational firms. We have exploited a unique liberalization of China's policy restrictions on foreign ownership to show that a move to optimal ownership can immediately boost firm output by up to 34% and induces even larger performance gains over the medium term of five years. By comparing foreign-acquired firms to foreign-owned firms (within liberalized industries) and even to other foreign-acquired firms (in unregulated industries), and applying a propensity-score reweighting approach, we have ensured that our results are not driven by selection or other benefits of foreign ownership. We have also demonstrated that foreign ownership patterns in China in general, and the changes in input ratios and output induced by the liberalization in particular, can be rationalized by a property-rights theory of the multinational firm featuring partial ownership.

Our contribution fills an important gap in the literature on firm organization. A large body of research has sought to explain the extent of firm boundaries by the fact that firms optimally allocate ownership rights in order to minimize the inefficiencies arising from incomplete contracts. What has been entirely missing from this literature is a sense of the magnitudes of these inefficiencies and the potential gains from optimal ownership. We provide direct evidence on optimal ownership restructuring and a first quantification of the associated performance gains. These findings demonstrate that the optimal decision on firm boundaries matters substantially for firm performance in practice. The size of these gains underlines the importance of research geared towards understanding firms' organizational decisions.

What are the policy implications of our findings? Should developing countries continue to liberalize foreign ownership restrictions (which remain widespread, see OECD, 2016)? Our analysis of the firms that received FDI after the liberalization suggests a positive evaluation of China's FDI reform in 2002. Abolishing restrictions on FDI and other policies that prevent the optimal allocation of ownership rights may hence lead to substantial efficiency and output gains in the affected firms. However, in order to obtain explicit policy recommendations, a comprehensive assessment of the aggregate implications is needed. The positive effects of the policy reform may well be magnified through the entry and exit of firms and reallocation of resources across firms, though there might also be short-term adjustment costs. Quantifying these aggregate effects of allowing for optimal ownership is an important task for future research.

Our quantification exercise can be seen in a broader context of the literature aiming to explain the large differences in aggregate productivity between developing countries, such as China, and the US. Hsieh and Klenow (2009) provide evidence that a large part of these differences are due to the misallocation of resources across firms. However, misallocation may be driven by a variety of frictions in output or factor markets, information flows, or contractual relationships, and the relative contribution of these frictions remains an open question. Our findings indicate that the inefficiencies due to contractual frictions and the misallocation of ownership rights might play an important role in explaining the large cross-country productivity gaps.

Bibliography

Ackerberg, Daniel A., Kevin Caves, and Garth Frazer, "Identification Properties of Recent Production Function Estimators," *Econometrica*, 2015, *83* (6), 2411–2451.

Alfaro, Laura and Maggie Xiaoyang Chen, "Surviving the Global Financial Crisis: Foreign Ownership and Establishment Performance," *American Economic Journal: Economic Policy*, 2012, *4* (3), 30–55.

_ , Pol Antràs, Davin Chor, and Paola Conconi, "Internalizing Global Value Chains: A Firm-Level Analysis," Technical Report 21582, National Bureau of Economic Research 2015.

Antràs, P., *Global Production: Firms, Contracts, and Trade Structure*, Princeton University Press, 2015.

Antràs, Pol, "Firms, Contracts, And Trade Structure," *The Quarterly Journal of Economics*, 2003, *118* (4), 1375–1418.

_ and Davin Chor, "Organizing the Global Value Chain," *Econometrica*, 2013, *81* (6), 2127–2204.

_ and Elhanan Helpman, "Global Sourcing," *Journal of Political Economy*, 2004, *112* (3), 552–580.

_ and _ , "Contractual Frictions and Global Sourcing," in Elhanan Helpman, Dalia Marin and Thierry Verdier, ed., *The Organization of Firms in a Global Economy*, Harvard University Press, 2008, pp. 55–83.

_ and Stephen R. Yeaple, "Chapter 2 - Multinational Firms and the Structure of International Trade," in Elhanan Helpman Gita Gopinath and Kenneth Rogoff, eds., *Handbook of International Economics*, Vol. 4, Elsevier, 2014, pp. 55–130.

Arnold, Jens Matthias and Beata Smarzynska Javorcik, "Gifted Kids or Pushy Parents? Foreign Direct Investment and Plant Productivity in Indonesia," *Journal of International Economics*, 2009, *79* (1), 42–53.

Bache, Peter Arendorf and Anders Laugesen, "Trade Liberalisation and Vertical Integration," Economics Working Papers 2013-14, Department of Economics and Business Economics, Aarhus University 2013.

Berkowitz, Daniel, Hong Ma, and Shuichiro Nishioka, "Recasting the Iron Rice Bowl: The Reform of China's State-Owned Enterprises," *The Review of Economics and Statistics*, 2017, *99* (4), 735–747.

Bernard, Andrew B., J. Bradford Jensen, Stephen J. Redding, and Peter K. Schott, "Intrafirm Trade and Product Contractibility," *American Economic Review: Papers & Proceedings*, May 2010, *100* (2), 444–448.

Bertrand, Marianne, Esther Duflo, and Sendhil Mullainathan, "How Much Should We Trust Differences-In-Differences Estimates?," *The Quarterly Journal of Economics*, 2004, *119* (1), 249–275.

Bircan, Çağatay, "Optimal Degree of Foreign Ownership under Uncertainty," Technical Report 166, European Bank for Reconstruction and Development 2014.

Blonigen, Bruce A., Lionel Fontagné, Nicholas Sly, and Farid Toubal, "Cherries for Sale: The Incidence and Timing of Cross-border M&A," *Journal of International Economics*, 2014, *94* (2), 341–357.

Brandt, Loren, Johannes Van Biesebroeck, and Yifan Zhang, "Creative Accounting or Creative Destruction? Firm-level Productivity Growth in Chinese Manufacturing," *Journal of Development Economics*, 2012, *97* (2), 339–351.

_ , Johannes Van Biesebroeck, Luhang Wang, and Yifan Zhang, "WTO Accession and Performance of Chinese Manufacturing Firms," *American Economic Review*, 2017, *107* (9), 2784–2820.

Busso, Matias, John DiNardo, and Justin McCrary, "New Evidence on the Finite Sample Properties of Propensity Score Reweighting and Matching Estimators," *The Review of Economics and Statistics*, 2014, *96* (5), 885–897.

Cai, Hongbin and Qiao Liu, "Competition and Corporate Tax Avoidance: Evidence from Chinese Industrial Firms," *The Economic Journal*, 2009, *119* (537), 764–795.

Chen, Wenjie, "The Effect of Investor Origin on Firm Performance: Domestic and Foreign Direct Investment in the United States," *Journal of International Economics*, 2011, *83* (2), 219–228.

Corcos, Gregory, Delphine M. Irac, Giordano Mion, and Thierry Verdier, "The Determinants of Intrafirm Trade: Evidence from French Firms," *Review of Economics and Statistics*, 2013, *95* (3), 825–838.

Defever, Fabrice and Farid Toubal, "Productivity, Relationship-specific Inputs and the Sourcing Modes of Multinationals," *Journal of Economic Behavior & Organization*, 2013, *94*, 345–357.

Desai, Mihir A, C Fritz Foley, and James R Hines, "A Multinational Perspective on Capital Structure Choice and Internal Capital Markets," *The Journal of Finance*, 2004, *59* (6), 2451–2487.

DiNardo, John, Nicole M. Fortin, and Thomas Lemieux, "Labor Market Institutions and the Distribution of Wages, 1973-1992: A Semiparametric Approach," *Econometrica*, 1996, *64* (5), 1001–1044.

Dollar, David and Shang-Jin Wei, "Das (Wasted) Kapital: Firm Ownership and Investment Efficiency in China," Technical Report 07/9, International Monetary Fund 2007.

Eppinger, Peter and Bohdan Kukharskyy, "Contracting Institutions and Firm Boundaries," Technical Report 100, University of Tübingen 2017.

Federico, Stefano, "Outsourcing Versus Integration at Home or Abroad and Firm Heterogeneity," *Empirica*, 2010, *37* (1), 47–63.

_ , "Headquarter Intensity and the Choice Between Outsourcing versus Integration at Home or Abroad," *Industrial and Corporate Change*, 2012, *21* (6), 1337–1358.

Feenstra, Robert C. and Gordon H. Hanson, "Ownership and Control in Outsourcing to China: Estimating the Property-Rights Theory of the Firm," *The Quarterly Journal of Economics*, 2005, *120* (2), 729–761.

Fernandes, Ana P. and Heiwai Tang, "Determinants of Vertical Integration in Export Processing: Theory and Evidence from China," *Journal of Development Economics*, 2012, *99* (2), 396–414.

Gandhi, Amit, Salvador Navarro, and David A Rivers, "On the Identification of Production Functions: How Heterogeneous is Productivity?," 2016. University of Wisconsin – Madison, mimeo.

Girma, Sourafel and Holger Görg, "Evaluating the Foreign Ownership Wage Premium Using a Difference-in-differences Matching Approach," *Journal of International Economics*, 2007, *72* (1), 97–112.

_ **and** _ , "Multinationals' Productivity Advantage: Scale or Technology?," *Economic Inquiry*, 2007, *45* (2), 350–362.

_ **, Yundan Gong, Holger Görg, and Sandra Lancheros**, "Estimating Direct and Indirect Effects of Foreign Direct Investment on Firm Productivity in the Presence of Interactions Between Firms," *Journal of International Economics*, 2015, *95* (1), 157–169.

Grossman, Gene M. and Elhanan Helpman, "Integration versus Outsourcing in Industry Equilibrium," *Quarterly Journal of Economics*, 2002, *117* (1), 85–120.

_ **and** _ , "Outsourcing versus FDI in Industry Equilibrium," *Journal of the European Economic Association*, 2003, *1* (2-3), 317–327.

_ **and** _ , "Managerial incentives and the international organization of production," *Journal of International Economics*, 2004, *63* (2), 237–262.

_ **and** _ , "Outsourcing in a Global Economy," *Review of Economic Studies*, 2005, *72* (1), 135–159.

Grossman, Sanford J. and Oliver D. Hart, "The Costs and Benefits of Ownership: A Theory of Vertical and Lateral Integration," *Journal of Political Economy*, 1986, *94* (4), 691–719.

Guadalupe, Maria, Olga Kuzmina, and Catherine Thomas, "Innovation and Foreign Ownership," *American Economic Review*, 2012.

Harris, Robert S. and David Ravenscraft, "The Role of Acquisitions in Foreign Direct Investment: Evidence from the U.S. Stock Market," *The Journal of Finance*, 1991, *46* (3), 825–844.

Hart, Oliver and John Moore, "Property Rights and the Nature of the Firm," *Journal of Political Economy*, 1990, *98* (6), 1119–1158.

Hsieh, Chang-Tai and Peter J. Klenow, "Misallocation and Manufacturing TFP in China and India," *The Quarterly Journal of Economics*, 2009, *124* (4), 1403–1448.

Huang, Yasheng, *Capitalism with Chinese Characteristics: Entrepreneurship and the State*, Cambridge University Press, 2008.

Imbens, Guido W., "Nonparametric Estimation of Average Treatment Effects Under Exogeneity: A Review," *The Review of Economics and Statistics*, 2004, *86* (1), 4–29.

Kalinova, Blanka, Angel Palerm, and Stephen Thomsen, "OECD's FDI Restrictiveness Index: 2010 Update," OECD Working Papers on International Investment 2010/03, OECD Publishing, Paris 2010.

Kohler, Wilhelm and Marcel Smolka, "Global Sourcing and Firm Selection," *Economics Letters*, 2014, *124* (3), 411–415.

_ **and** _ , "Global Sourcing of Heterogeneous Firms: Theory and Evidence," CESifo Working Paper Series 5184, CESifo Group Munich 2015.

Kohler, Wilhelm K. and Marcel Smolka, "Sourcing Premia with Incomplete Contracts: Theory and Evidence," *The B.E. Journal of Economic Analysis & Policy*, 2011, *11* (1), 10.

Kukharskyy, Bohdan, "Relational Contracts and Global Sourcing," *Journal of International Economics*, 2016, *101* (C), 123–147.

_ **and Michael P. Pflüger**, "Relational Contracts and the Economic Well-being of Nations," Technical Report 5394, IZA Discussion Papers, Bonn 2010.

Levinsohn, James and Amil Petrin, "Estimating Production Functions Using Inputs to Control for Unobservables," *Review of Economic Studies*, 2003, *70* (2), 317–341.

Lu, Yi, Zhigang Tao, and Lianming Zhu, "Identifying FDI Spillovers," *Journal of International Economics*, 2017, *107*, 75 – 90.

Manova, Kalina, Shang-Jin Wei, and Zhiwei Zhang, "Firm Exports and Multinational Activity under Credit Constraints," *Review of Economics and Statistics*, 2015, *97* (3), 574–588.

Melitz, Marc J., "The Impact of Trade on Intra-Industry Reallocations and Aggregate Industry Productivity," *Econometrica*, 2003, *71* (6), 1695–1725.

Nichols, A., "Causal Inference with Observational Data," *Stata Journal*, 2007, *7* (4), 507–541(35).

_ , "Erratum and Discussion of Propensity-score Reweighting," *Stata Journal*, 2008, *8* (4), 532–539(8).

Nunn, Nathan and Daniel Trefler, "The Boundaries of the Multinational Firm: An Empirical Analysis," in Elhanan Helpman, Dalia Marin, and Thierry Verdier, ed., *The Organization of Firms in a Global Economy*, Harvard University Press, 2008, pp. 55–83.

_ **and** _ , "Incomplete Contracts and the Boundaries of the Multinational Firm," *Journal of Economic Behavior & Organization*, 2013, *94*, 330–344.

_ **and** _ , "Domestic Institutions as a Source of Comparative Advantage," in Elhanan Helpman, Kenneth Rogoff, and Gita Gopinath, eds., *Handbook of International Economics*, Vol. 4, Elsevier, 2014, chapter 5.

OECD, "OECD FDI Regulatory Restrictiveness Index (Edition 2016)," 2016.

Olley, G Steven and Ariel Pakes, "The Dynamics of Productivity in the Telecommunications Equipment Industry," *Econometrica*, 1996, *64* (6), 1263–1297.

Qin, Julia Ya, "Trade, Investment and Beyond: The Impact of WTO Accession on China's Legal System," *The China Quarterly*, 2007, *191*, 720–741.

Rosenbaum, Paul R. and Donald B. Rubin, "The Central Role of the Propensity Score in Observational Studies for Causal Effects," *Biometrika*, 1983, *70* (1), 41–55.

Schwarz, Christian and Jens Suedekum, "Global Sourcing of Complex Production Processes," *Journal of International Economics*, 2014, *93* (1), 123–139.

Sheng, Liugang and Dennis Tao Yang, "Expanding Export Variety: The Role of Institutional Reforms in Developing Countries," *Journal of Development Economics*, 2016, *118*, 45 – 58.

Swenson, Deborah L., "Foreign Mergers and Acquisitions in the United States," in "Foreign Direct Investment" NBER Chapters, National Bureau of Economic Research, 1993, pp. 255–284.

Tomiura, Eiichi, "Foreign Outsourcing, Exporting, and FDI: A Productivity Comparison at the Firm Level," *Journal of International Economics*, 2007, *72* (1), 113–127.

UNCTAD, "World Investment Report 2011 – Non-equity Modes of International Production and Development," Technical Report, United Nations Conference on Trade and Development 2011.

_ , "World Investment Report 2013 – Global Value Chains: Investment and Trade for Development," Technical Report, United Nations Conference on Trade and Development 2013.

Wang, Jian and Xiao Wang, "Benefits of Foreign Ownership: Evidence from Foreign Direct Investment in China," *Journal of International Economics*, 2015, *97* (2), 325–338.

Wooldridge, Jeffrey M., *Econometric Analysis of Cross Section and Panel Data*, Vol. 2 of *MIT Press Books*, The MIT Press, 2010.

Yeaple, Stephen Ross, "Offshoring, Foreign Direct Investment, and the Structure of U.S. Trade," *Journal of the European Economic Association*, 2006, *4* (2-3), 602–611.

Zhang, Kevin H., *China as the World Factory* Routledge Studies in the Growth Economies of Asia, Taylor & Francis, 2006.

5.A Mathematical Appendix

5.A.1 Derivation of optimal ownership share and proof of Proposition 5.1

Taking the partial derivative of $\Psi(\cdot)$ in equation (5.11) with respect to β_F and simplifying yields:

$$\frac{\partial \Psi(\beta_F)}{\partial \beta_F} \times \left[\frac{\alpha}{1-\alpha} \left(\frac{\eta}{\beta_F} - \frac{1-\eta}{1-\beta_F} \right) \Big(1 - \alpha\eta\beta_F - \alpha(1-\eta)\left[1-\beta_F\right] \Big) + \Big(\alpha - 2\alpha\eta \Big) \right] = 0.$$

The first two factors are non-negative for any $\beta_F \in (0,1)$ and $w, r > 0$, so we can concentrate on the term in large brackets, which can be simplified to obtain:

$$(2\eta - 1)\,\beta_F^2 + 2\eta\,(\alpha - \alpha\eta - 1)\,\beta_F - \eta\,(\alpha - \alpha\eta - 1)\,.$$

Setting this term equal to zero and solving the quadratic form delivers the optimal revenue share β_F^* given in equation (5.13). Finally, s^* in equation (5.12) is obtained by rearranging equation (5.4).

The optimal revenue share β_F^* increases weakly in the headquarter intensity η, as

$$\frac{\partial \beta_F^*}{\partial \eta} = \frac{\left[2\alpha\eta(1-\eta) - 1 + 2\sqrt{\eta(1-\eta)(1-\alpha\eta)(\alpha\eta+1-\alpha)} \right] \left[2\alpha\eta(\eta-1) - (1-\alpha) \right]}{2(2\eta-1)^2 \sqrt{\eta(1-\eta)(1-\alpha\eta)(\alpha\eta+1-\alpha)}}$$

is non-negative. This can be seen by noting that for α approaching 1 and $\eta = 0.5$ each of the two terms in square brackets approaches its supremum, which is negative in both cases (and this expression as well as β_F^* are not defined). Hence, both terms are negative for all other admissible parameter values, the denominator is always non-negative, and therefore the entire expression is non-negative. It follows from the strictly monotonic relationship between s^* and β_F^* described in equation (5.12) that an increase in η, raising β_F^*, implies a higher optimal ownership share s^*. This completes the proof of Proposition 5.1.

5.A.2 Proof of Proposition 5.3

In Section 5.2.2, we have derived the optimal input choices conditional on the chosen organizational form s, which result in the capital-labor ratio in equation (5.8). We have further shown in Proposition 5.2 that abolishing foreign ownership restrictions induces optimal ownership restructuring in previously constrained firm pairs, which increase their ownership share s and hence β_F. Since any increase in β_F raises the capital-labor ratio in equation (5.8):

$$\frac{\partial \left(K_j^*(i)/L_j^*(i) \right)}{\partial \beta_F(s)} = \frac{\eta}{1-\eta} \frac{w}{r} \frac{1}{\left[1-\beta_F(s)\right]^2} > 0,$$

so does the optimal ownership restructuring. This proves part (i) of Proposition 5.3.

Taking the partial derivative of the second-best output level $x_j^*(i)$ from equation (5.9) with respect to $\beta_F(s)$ yields:

$$\frac{\partial x_j^*(i)}{\partial \beta_F(s)} = x_j^*(i) \left[\frac{\eta}{1-\alpha} \beta_F(s)^{-1} - \frac{1-\eta}{1-\alpha} \left[1-\beta_F(s)\right]^{-1} \right],$$

which is positive iff

$$\frac{\eta}{\beta_F(s)} > \frac{1-\eta}{1-\beta_F(s)} \iff \eta - \eta\beta_F(s) > \beta_F(s) - \eta\beta_F(s) \iff \eta > \beta_F(s).$$

Hence, an increase in $\beta_F(s)$ (due to an increase from $\bar{s}$ to s^*) raises output towards the second-best level if $\eta > \beta_F(s)$. This condition holds for any $\eta > 0.5$, as can easily be verified using equation (5.13) and seen in Figure 5.1(a). This proves part (ii) of Proposition 5.3.

To examine input quantities, we continue to focus on the effect of an increase in $\beta_F(s)$ implied by an increase in s towards the optimal level s^*. Taking the partial derivative of $K_j^*(i)$ from equation (5.7) with respect to $\beta_F(s)$ yields:

$$\frac{\partial K_j^*(i)}{\partial \beta_F(s)} = \Theta(i)\eta A_j^{\frac{1}{1-\alpha}} \alpha^{\frac{1}{1-\alpha}} \left(\beta_F(s) r^{-1}\right)^{\frac{1-\alpha+\alpha\eta}{1-\alpha}} \left([1-\beta_F(s)]\, w^{-1}\right)^{\frac{\alpha(1-\eta)}{1-\alpha}} \frac{1-\beta_F(s)-\alpha(1-\eta)}{(1-\alpha)\beta_F(s)(1-\beta_F(s))}$$

The sign of this derivative is the same as the sign of the fraction. This expression is always positive evaluated at the optimal β_F^* from equation (5.13), so the impact of optimal ownership restructuring on capital is clearly positive.

Taking the partial derivative of $L_j^*(i)$ from equation (5.6) with respect to $\beta_F(s)$ yields:

$$\frac{\partial L_j^*(i)}{\partial \beta_F(s)} = \Theta(i)(1-\eta) A_j^{\frac{1}{1-\alpha}} \alpha^{\frac{1}{1-\alpha}} \left(\beta_F(s) r^{-1}\right)^{\frac{\alpha\eta}{1-\alpha}} \left([1-\beta_F(s)]\, w^{-1}\right)^{\frac{1-\alpha\eta}{1-\alpha}} \frac{\alpha\eta - \beta_F(s)}{(1-\alpha)\beta_F(s)(1-\beta_F(s))}$$

The sign of this derivative is the same as the sign of the fraction. This expression is always negative evaluated at the optimal β_F^* from equation (5.13). Hence, a marginal increase in s towards the optimum s^* discourages labor investments. However, this effect may be overcompensated (for high values of η and α) by increased labor demand through the output increase caused by optimal ownership restructuring (see part (ii) of Proposition 5.3).

To see how the net effect on employment can be positive, consider the following numerical example, where we specify $\delta(s) = s$ for simplicity and set $A = w = r = 1$, $\beta = 0.5$, $\alpha = 0.9$, and $\eta = 0.95$. Suppose the initial policy is $\bar{s} = 0.5$, which implies $\beta_F(\bar{s}) = 0.77$. Abolishing this policy restriction entails an increase in the ownership share and hence an increase in the revenue share to the optimal value $\beta_F^* = 0.91$. With this parametrization, both the marginal employment effect $\partial L_j(i)/\partial \beta_F(s)$ evaluated at $\bar{s}$ as well as the total change in employment from $L_j(i)|_{\bar{s}}$ to $L_j^*(i)$ is positive.

5.B Additional Tables

Table 5.B.1: *Logit estimations of the propensity score and reweighted results*

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Logit regressions				Reweighted logit regressions			
	T0	T25	T50	T100	T0	T25	T50	T100
ln TFP	0.289	0.166	-0.710**	-0.222	-0.0129	-0.0326	-0.00237	0.125
	(0.330)	(0.303)	(0.354)	(0.473)	(0.326)	(0.299)	(0.359)	(0.455)
ln K/L	-0.515***	-0.486***	-0.602***	-0.519***	0.0191	0.0382	0.0429	0.137
	(0.137)	(0.128)	(0.153)	(0.196)	(0.134)	(0.127)	(0.159)	(0.196)
ln TFP × ln K/L	-0.0252	-0.0244	0.0623	-0.117	0.00280	0.0171	0.0155	0.0389
	(0.0614)	(0.0557)	(0.0623)	(0.0769)	(0.0604)	(0.0543)	(0.0627)	(0.0805)
ln L	-0.379*	-0.272	-0.615***	-0.581*	0.0344	0.0653	0.0605	0.229
	(0.208)	(0.196)	(0.225)	(0.304)	(0.202)	(0.194)	(0.229)	(0.308)
ln w	-0.586***	-0.570***	-0.418*	0.0987	0.0506	0.0604	0.0369	-0.0606
	(0.188)	(0.178)	(0.214)	(0.277)	(0.190)	(0.182)	(0.215)	(0.290)
Age	0.0422***	0.0484***	0.0604***	0.0306	-0.00344	-0.00687	0.00340	-0.0278
	(0.0122)	(0.0129)	(0.0212)	(0.0335)	(0.0123)	(0.0134)	(0.0211)	(0.0428)
Exports/ sales	-0.780**	-0.560*	-1.100***	-1.323***	0.0355	0.0575	0.139	0.164
	(0.348)	(0.309)	(0.358)	(0.410)	(0.344)	(0.310)	(0.354)	(0.394)
SOE share	1.531***	1.987***	3.148***		0.122	0.157	0.106	
	(0.346)	(0.337)	(0.542)		(0.338)	(0.331)	(0.554)	
ln Y	0.196	0.157	0.543**	0.741**	-0.0300	-0.0787	-0.0511	-0.322
	(0.225)	(0.211)	(0.243)	(0.322)	(0.221)	(0.212)	(0.255)	(0.344)
Pre-trend *TFP*	-0.147	0.0365	0.0177	0.339*	0.00170	0.0341	0.0329	0.0218
	(0.137)	(0.134)	(0.146)	(0.205)	(0.136)	(0.134)	(0.144)	(0.202)
N	1,707	1,543	1,090	619	1,706	1,541	1,085	615
Pseudo R^2	0.141	0.156	0.176	0.082	0.002	0.005	0.007	0.008
Log-likelihood	-331.7	-349.4	-241.8	-163.9	-139.6	-159.2	-107.6	-68.46
χ^2	109.1	129.6	103.3	29.34	0.847	2.513	2.607	2.156

Note: The table reports coefficient estimates from logit regressions with the dependent variable indicated in the header (treatment dummies, indicating foreign acquisitions between 2001 and 2003 defined by different thresholds). Columns 1-4 report the regressions that predict the propensity score used for reweighting in the main estimations (in Table 5.3). Columns 1-4 report the same regressions reweighted by the estimated propensity score. All regressions include 2-digit industry fixed effects. Standard errors clustered by firm are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

Table 5.B.2: *Robustness concerning the estimation of the propensity score*

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Financial conditions		Industry covariates		Pre-trends Y and K/L		Province fixed effects	
	K/L	Y	K/L	Y	K/L	Y	K/L	Y
T0 × *after*	0.138**	0.295***	0.135**	0.280***	0.139**	0.295***	0.141**	0.294***
	(0.0650)	(0.0639)	(0.0651)	(0.0651)	(0.0648)	(0.0639)	(0.0661)	(0.0668)
N	3,512	3,518	3,512	3,518	3,512	3,518	2,924	2,928
T25 × *after*	0.0336	0.254***	0.0374	0.244***	0.0326	0.256***	0.0589	0.282***
	(0.0636)	(0.0621)	(0.0631)	(0.0624)	(0.0635)	(0.0621)	(0.0639)	(0.0653)
N	3,196	3,202	3,198	3,204	3,198	3,204	3,038	3,044
T50 × *after*	0.183**	0.150*	0.195***	0.137*	0.185**	0.148*	0.172**	0.152*
	(0.0760)	(0.0821)	(0.0746)	(0.0811)	(0.0746)	(0.0812)	(0.0770)	(0.0877)
N	2,258	2,262	2,262	2,266	2,266	2,270	2,112	2,116
T100 × *after*	0.238***	-0.0880	0.237***	-0.0953	0.244***	-0.0820	0.198***	-0.0365
	(0.0718)	(0.0800)	(0.0744)	(0.0795)	(0.0713)	(0.0794)	(0.0720)	(0.0921)
N	1,416	1,420	1,416	1,420	1,416	1,420	1,350	1,354

Note: The table reports estimates of β in equation (5.16) for the years 2001 and 2003, with the dependent variables indicated in the header (all in logs). The definitions of the treatment and control groups and the control variables are analogous to the main estimations (see the note to Table 5.3). The propensity scores used in reweighting are estimated as in the main equations (see Table 5.B.1 for the baseline set of covariates) with the following added covariates: liquidity and leverage ratios in columns 1-2; the number of firms, employment, and wage per employee (all in logs), average firm age and the exports to sales ratio at the 4-digit industry level in columns 3-4; the pre-trends in output and the capital-labor ratio (annual differences in logs between 2000 and 2001) in columns 5-6; and province fixed effects in columns 7-8. Standard errors clustered by firm are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

Chapter 6

Contracting Institutions and Firm Boundaries*

Contractual frictions are widely known to shape firm boundaries. But do better contracting institutions, which reduce these frictions, induce firms to be more or less deeply integrated? This paper provides a large-scale investigation of this question using a unique micro dataset of ownership shares across half a million firm pairs worldwide. We uncover strong evidence that better contracting institutions in subsidiaries' countries favor deeper integration, particularly in relationship-specific industries. We formally show that these findings can be explained by a generalized Property-Rights Theory of the firm featuring partial ownership, while they are at odds with the canonical Transaction-Cost Theory.

*This chapter is based on joint work with Bohdan Kukharskyy.

6.1 Introduction

In his seminal contribution, Coase (1937) raised one of the most fundamental questions in economics: What determines firm boundaries? Virtually every firm has to decide whether to cooperate with its business partners at arm's-length or integrate them to some degree into its boundaries. A profound understanding of this integration decision is required more than ever in the age of globalization, characterized by the emergence of multinational corporations that span their boundaries across several countries.[1] All theoretical explanations of firm boundaries provided to date recognize the fundamental importance of contractual imperfections (cf. Gibbons, 2005), which arguably depend on the quality of contracting institutions. Thus, the large international differences in judicial quality prevailing across the globe should play a key role in shaping firm boundaries. Yet, the direction of this effect is a priori not clear. Do better contracting institutions induce firms to be more or less deeply integrated? To answer this question, this paper develops a parsimonious theoretical model of the relationship between contracting institutions and firm boundaries and tests its predictions using a unique micro dataset of ownership shares across half a million firm pairs worldwide.

While existing theories of the firm agree on the importance of contractual imperfections for shaping firm boundaries, they make *opposite* predictions regarding the effect of contracting institutions on the optimal degree of integration. To illustrate this point, consider the two classical theories of the firm: the Transaction-Cost Theory (TCT) by Williamson (1971, 1975, 1985) and the Property-Rights Theory (PRT) by Grossman and Hart (1986) and Hart and Moore (1990).[2] Under both theories, we examine a production relationship between two parties, a firm's headquarters (HQ) and a manufacturing producer. If courts cannot fully enforce contracts between the parties, and if the producer needs to invest into relationship-specific inputs, then these investments are plagued by a hold-up problem. According to the TCT, the HQ can eliminate the resulting inefficiencies by integrating the producer into firm boundaries at the expense of an exogenous governance cost. In this theory, better contracting institutions in the producer's country mitigate the hold-up problem in arm's-length transactions, and therefore make integration less attractive. By contrast, the PRT argues that contractual imperfections cause hold-up inefficiencies even *within* firm boundaries. By integrating the producer, the HQ obtains residual control rights over non-contractible inputs, but undermines the producer's incentives to invest into these inputs. According to the PRT, better contracting institutions in the producer's country reduce the need to incentivize the producer, and therefore make integration more attractive.[3] Hence, the predictions of the two theories regarding the effect of contracting institutions on firm boundaries are diametrically opposed.

Which of the two competing hypotheses finds empirical support? As a first glance at the relationship between the degree of integration and contracting institutions, Figure 6.1(a) plots the ownership shares of more than half a million firm pairs from more than one hundred countries in 2014, recorded in the Orbis database by Bureau van Dijk (BvD), against the rule of law index of the subsidiary's country –

[1] UNCTAD (2011, 2013) estimates that multinationals accounted for one quarter of world GDP and participated in 80% of world trade in 2010, with more than one third of world exports traded within multinational firms' boundaries. See Barba Navaretti and Venables (2004) for stylized facts on the growing significance of multinationals over time.

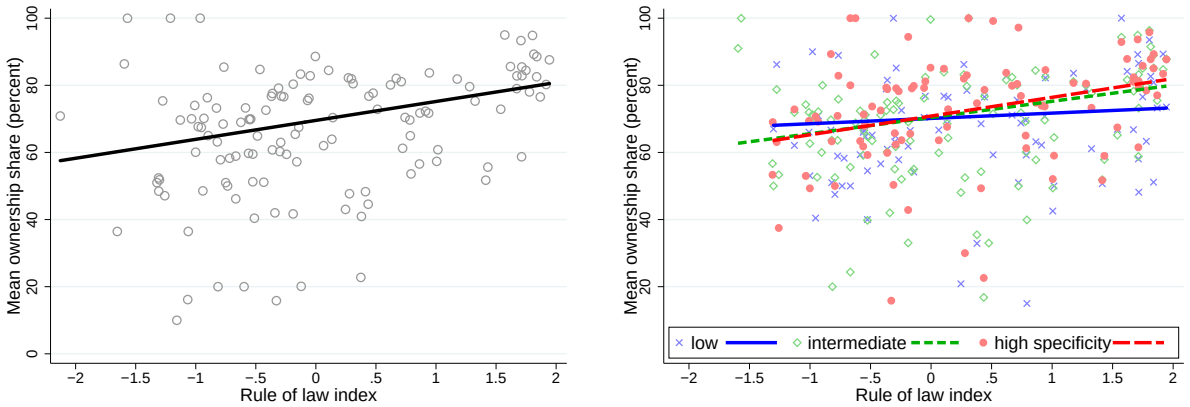
[2] These two theories are among the most acclaimed in organizational economics (see e.g. Whinston, 2003). They have also been instrumental in understanding the boundaries of multinational firms; see Grossman and Helpman (2002, 2003, 2005) for a TCT view, and Antràs (2003) as well as Antràs and Helpman (2004, 2008) for a PRT perspective.

[3] The original work by Grossman and Hart (1986) and Hart and Moore (1990) does not provide a testable prediction regarding the effect of contracting institutions on firm boundaries. Hence, we formally derive this PRT prediction in our paper.

a standard measure of the quality of contracting institutions.[4] We observe a positive and significant correlation between ownership shares and judicial quality, suggesting that subsidiaries tend to be more deeply integrated in countries with better contracting institutions. This observation runs counter to the TCT logic, but it is in line with the PRT.

Figure 6.1: *Ownership shares and contracting institutions*

(a) *By subsidiary's country* (b) *By country and relationship-specificity*



Note: The graphs plot mean values of ownership shares (exceeding 10%) of firm pairs against the rule of law index of the subsidiary's country in the cross-section of 2014. In Figure 6.1(a), the ownership shares are arithmetic means by the subsidiary's country. In Figure 6.1(b), the ownership shares are arithmetic means by the subsidiary's country and the relationship-specificity category of the subsidiary's industry, whereby 'low' relationship-specificity means that the industry contains zero differentiated or reference-priced products according to the liberal Rauch (1999) classification, for 'intermediate' specificity the share of these products lies between zero and one, and 'high' reflects a share equal to one. The lines are obtained from univariate regressions of the mean ownership shares on the rule of law index, whereby each observation is weighted by the underlying number of firm pairs. In Figure 6.1(a), the estimated slope parameter is 5.639 with a t-value of 2.65 (based on robust standard errors), the R^2 is 0.026, and the sample is based on N=605,547 firm pairs. In Figure 6.1(b), for low relationship-specificity the slope is 1.577 (t=0.90, R^2=0.002, N=25,751), for intermediate specificity it is 4.818 (t=2.15, R^2=0.023, N=104,014), and for high specificity it is 5.583 (t=3.10, R^2=0.027, N=101,172).

It is well-known that contractual imperfections *per se* do not necessarily lead to hold-up inefficiencies. It is the combination of contractual incompleteness and relationship-specificity – defined as the extent to which investments have a higher value within a given relationship than outside of it – that causes 'lock-in' and hold-up (see e.g. Joskow, 2005). Hence, one would expect a differential impact of contracting institutions depending on the degree of relationship-specificity of a subsidiary's investments. More precisely, a high degree of relationship-specificity magnifies the respective effect of contracting institutions predicted by either theory of the firm. According to the TCT, the negative impact of contracting institutions on the attractiveness of integration is particularly pronounced in industries with high degrees of relationship-specificity, since the hold-up problem in those industries is relatively more severe (see Antràs, 2015). Conversely, as we formally show in this paper, the PRT predicts a positive interaction effect between the quality of contracting institutions and the degree of relationship-specificity of producers' goods on the optimal ownership share. Figure 6.1(b) provides a first assessment of this interaction effect. It distinguishes subsidiaries' industries by their degree of relationship-specificity, measured by the share of differentiated goods according to the Rauch (1999) classification. The figure reveals that the positive correlation between ownership shares and the rule of law index is strongest for subsidiaries

[4]For the purpose of illustration, we plot mean ownership shares aggregated by country (Figure 6.1(a)) or by country and relationship-specificity of the subsidiary (Figure 6.1(b)), instead of the myriad of individual observations at the level of firm pairs. However, the regression lines are weighted by the underlying numbers of firm pairs, so they represent correlations in the raw data. The data are described in detail in Section 6.3.2.

operating in industries with a high degree of relationship-specificity; the correlation is slightly weaker for intermediate degrees of relationship-specificity, and it is small and insignificant in non-specific industries. Hence, the patterns observed in the raw data do not match the TCT view but square well with the PRT predictions.

To formalize the above arguments, we set up a theoretical model in the spirit of the PRT. Motivated by our empirical setup, this framework generalizes the conventional PRT of the multinational firm (as in Antràs, 2003) in three respects. First, the key novel feature of our model is that the HQ faces a continuous integration decision and chooses the equilibrium ownership *share* in the producer, rather than deciding only between the two extreme cases of full integration and arm's-length transactions. Second, we allow for partial contractibility, in the sense that courts can verify and enforce a fraction of the producer's investments into inputs, and this fraction may vary across countries with different quality of contracting institutions (as in Acemoglu et al., 2007; Antràs and Helpman, 2008). Third, rather than assuming that the producer's inputs are fully customized to a given relationship, we consider partial relationship-specificity and allow the degree of customization to vary across industries.[5] Given that the producer's investments are not fully contractible, the parties bargain over the distribution of the surplus from the relationship ex-post (i.e., after all inputs have been produced). This setup is characterized by a hold-up problem and ex-ante underinvestment by the producer. As a result, the HQ's ownership decision involves a simple trade-off between her *share* of the surplus generated by the relationship and the *size* of this surplus.[6] Intuitively, an increase in the ownership share shifts residual control rights between the two parties: It increases the HQ's outside option but reduces that of the producer. This improves the HQ's bargaining position and allows her to reap a larger share of the surplus in ex-post negotiations. Yet, the producer anticipates the stronger hold-up associated with a higher ownership share of the HQ, so his underinvestment becomes more severe, and the surplus size decreases.

Our generalized PRT model delivers the following two key predictions, which rationalize the patterns observed in Figure 6.1. First, it predicts a positive relationship between the optimal ownership share and the quality of contracting institutions in the producer's country. Intuitively, if courts can verify and enforce contracts on a larger share of inputs, the need for incentivizing the producer's investments decreases. Consequently, the HQ integrates the producer more deeply in order to obtain a larger fraction of the surplus. Second, the positive effect of contracting institutions on the optimal ownership share is predicted to be more pronounced in industries with a higher relationship-specificity. To develop the intuition behind this result, it is important to understand first that a high relationship-specificity mitigates the negative effect of a higher ownership share on the producer's investment incentives. In industries with a high degree of relationship-specificity, where inputs have little value on the outside market, the producer's potential outside option is small and of little importance for his ex-post payoff and ex-ante underinvestment. Hence, an increase in the ownership share, reducing the producer's outside option, has only a weak negative effect on his investment incentives. Conversely, in industries with a low degree of relationship-specificity, the producer's potential outside option is large, and any increase in the ownership share strongly aggravates the underinvestment problem. Hence, an improvement in contracting institutions allows the HQ to disproportionately increase the optimal ownership share in industries with a high degree of relationship-specificity, where increased ownership disincentivizes the producer's invest-

[5] Note that a PRT with partial contractibility and relationship-specificity has been considered before by Antràs (2015); our novel framework combines these features with the explicit modeling of non-zero outside options for both parties and continuous ownership shares, which are required to explain the above-mentioned empirical patterns.

[6] For clarity, we refer to the HQ as 'she' and the producer as 'he' throughout the paper.

ments less. To summarize, our PRT model predicts a positive direct effect of the quality of contracting institutions and a positive interaction effect of contracting institutions and relationship-specificity on the optimal ownership share.

We test the model's predictions using unique data on global ownership links from Orbis. This database provides an extensive account of multinational firms' ownership structures at an exceptional level of detail. It reports the ownership shares of roughly 300,000 headquarters in more than half a million subsidiaries worldwide. The dataset is uniquely suited for the purpose of our study because it combines three key advantages: First, it includes *firm-pair* specific ownership data along with information on the countries, industry affiliations, and other characteristics of *both* firms. Second, it covers headquarters and subsidiaries in more than one hundred countries around the globe. And third, it includes both domestic and international ownership links as well as multiple ownership links for some firms. These features allow us to provide a large-scale investigation of the relationship between firm boundaries and contracting institutions, while thoroughly controlling for heterogeneity across countries, industries, and firms on both sides of the ownership link.[7] We enrich the database with various country-level proxies for contracting institutions and industry-level measures of relationship-specificity to test the theoretical predictions.

Our empirical analysis proceeds in two steps. In the first step, we scrutinize the correlation between the quality of contracting institutions and firm boundaries illustrated in Figure 6.1(a). To this end, we regress firm-pair specific ownership shares on the rule of law index, while controlling for a large set of observable characteristics (of the subsidiary's country, the ownership structure, and bilateral investment costs) and a battery of fixed effects (for the subsidiary's industry as well as the HQ's country-industry). Conditional on all of these covariates, we find that firms own significantly higher shares of their subsidiaries in countries with better judicial quality. This finding supports our first theoretical prediction.

In the second step, we move towards a more stringent test of the theory by regressing ownership shares on an interaction term between the quality of contracting institutions in the subsidiary's country and the relationship-specificity of its industry (scrutinizing the pattern from Figure 6.1(b)). This approach allows us to effectively account for all observable or unobservable characteristics of the subsidiary's country using fixed effects, thereby addressing first-order concerns related to omitted variables (such as cultural traits or other institutions). In our preferred specification, we further control for bilateral investment costs by country-pair fixed effects. The estimates yield a positive interaction effect of country-level judicial quality and industry-level relationship-specificity on the depth of integration, which is both statistically and economically significant. This finding supports the second key prediction of our model: The positive effect of contracting institutions on the ownership share is more pronounced in industries with a higher relationship-specificity.

Our main findings are robust to addressing several challenges to identification. We find very similar results using various alternative proxies for contracting institutions and approximating relationship-specificity by a firm-pair specific measure based on the duration of a relationship. In an important set of robustness checks, we accommodate remaining concerns regarding omitted variables. To this end, we allow for the effects of economic development and other institutions on firm boundaries to differ arbitrarily across industries by including interaction terms of these country characteristics with subsidiary industry dummies (following Levchenko, 2007). We further ensure that our results are not driven by

[7] The only alternative dataset covering multinational firms in many countries is by Dun & Bradstreet. To the best of our knowledge, this database documents neither ownership intensities nor multiple ownership links per firm.

firm heterogeneity among subsidiaries or headquarters. Moreover, the effect of contracting institutions can even be identified from within-firm variation across different ownership links of the same HQ. To address the possibility that selection into production countries may be driven by factors correlated with contracting institutions, we estimate a two-stage model and correct for this type of selection following Heckman (1979). Allowing for non-linearities in ownership decisions in an ordered logistic regression framework provides additional support for our theoretical predictions. Finally, we exploit the historic origins of countries' legal systems as an exogenous source of variation in contracting institutions using instrumental variables and propensity score matching techniques (as in Nunn, 2007). The robustness of our findings to all of these checks lends strong support to the PRT.

Our paper is closely related to the work by Antràs (2015), who investigates the role of contracting institutions for firms' integration decisions both theoretically and empirically. Similar to our analysis, he contrasts the effects of contracting institutions governing the producer's investments on the relative attractiveness of vertical integration in the TCT and the PRT. Yet, our theoretical contributions differ in two respects. First, we model the integration decision as a continuous (rather than a binary) variable.[8] Second, and most importantly, we derive a clear testable prediction of the PRT regarding the *interaction* effect of contracting institutions and relationship-specificity on firm boundaries. Our empirical approach is also substantially different. While Antràs (2015) approximates the relative attractiveness of foreign integration vs. outsourcing using industry-level data on U.S. intra-firm imports, we test our predictions using a global dataset of ownership shares at the level of firm pairs.[9] Although the author considers the interaction between a country's contracting institutions and an industry's relationship-specificity in their impact on U.S. intra-firm trade, the evidence remains inconclusive. By contrast, our micro data yield strong evidence for the positive interaction effect of contracting institutions and relationship-specificity on the depth of integration.

To the best of our knowledge, Acemoglu et al. (2009) is the only existing firm-level study of the link between contracting institutions and vertical integration in a large cross-section of countries. The authors combine data on primary and secondary activities within a given firm with U.S. industry-level input-output (I-O) tables to construct a vertical integration index, designed to approximate the firm's propensity to own a vertically integrated supplier.[10] The authors do not find a significant relationship between this index and a country's contracting costs, but they document a higher degree of vertical integration in countries that have both higher contracting costs and greater financial development. A

[8] In so doing, we relate to Antràs and Helpman (2008), who show that the HQ's optimal share of ex-post surplus increases in the contractibility of a supplier's inputs according to the PRT. However, the authors treat this share as a hypothetical construct, which cannot be freely chosen by the HQ. We complement their findings by allowing the HQ to choose from a continuum of ownership shares and show how the integration decision depends on contracting institutions and relationship-specificity. Previous theoretical contributions have studied partially integrated production processes across *multiple* producers, either organized sequentially along the value chain (Antràs and Chor, 2013; Alfaro et al., 2015) or simultaneously contributing to a single production stage (Schwarz and Suedekum, 2014). However, none of these papers considers partial integration of a *single* firm.

[9] The U.S. intra-firm trade data has become a workhorse tool in empirical studies of international integration decisions (see Yeaple, 2006; Nunn and Trefler, 2008, 2013; Antràs and Chor, 2013). Corcos et al. (2013) have taken this approach to the firm level using French customs data. In line with the PRT, they find a positive correlation between contract enforcement in the foreign country and the share of intra-firm imports. Other studies on intra-firm trade using firm-level data from a single country include Carluccio and Fally (2012), Defever and Toubal (2013), Kohler and Smolka (2014, 2015), and Tomiura (2007), all of which find patterns consistent with the PRT.

[10] This approach has subsequently been used by Alfaro and Charlton (2009) and Alfaro et al. (2015) to characterize the organization of multinational production activities. Recent evidence on US firms with multiple domestic plants (Atalay et al., 2014) or with multinational affiliates (Ramondo et al., 2016) suggests that integrated firm pairs do not necessarily engage in intra-firm trade even if they are vertically linked via I-O tables. Note that we do not rely on I-O tables to identify vertical links in this paper. Also, our theoretical explanation of the integration decision is not restricted to vertical links, nor does it presuppose any intra-firm trade, as producers in the model may sell their output to final consumers.

key advantage of our data is that we observe ownership intensities across firm *pairs*, which allows us to identify the effect of subsidiary country institutions while controlling for the potentially confounding role of HQ country institutions using fixed effects. In contrast to Acemoglu et al. (2009), our approach uncovers a robust positive link between the quality of contracting institutions in a subsidiary's country and the integration intensity.

Our theoretical and empirical results further contribute to the literature that aims to discriminate between the two prominent theories of the firm: TCT vs. PRT. The fact that the predictions of the TCT differ in important ways from those of the PRT is generally known among economists (see Gibbons, 2005; Whinston, 2003). A substantial body of empirical research has assessed various predictions of the two theories using data from a single firm, industry, or country.[11] Yet, we are unaware of any empirical investigation attempting to contrast these differential predictions using micro data from multiple countries. We suggest a twofold explanation for the scarcity of empirical evidence on this fundamental question. First, theories of the firm have mostly concentrated on the two 'extreme' cases of full integration and arm's-length contracting. Thus, the fact that commercial transactions between independent firms are rarely observed in the data poses a major challenge for testing these theories. We make progress in this discourse by deriving predictions regarding the *intensity* of integration and test them using data on firm ownership *shares*. Second, theoretical predictions are often formulated in terms of abstract concepts (e.g. marginal returns on investments or quasi-rents), which are extremely difficult to capture empirically. The key explanatory variables in our analysis are contracting institutions and relationship-specificity, for which we can obtain well-established proxies from readily available data.

The remainder of the paper is organized as follows. Section 6.2 sets up a PRT model of the firm and derives our key predictions regarding optimal ownership shares. Section 6.3 describes our empirical approach and the data. Section 6.4 presents our estimation results and a multitude of robustness checks. Section 6.5 concludes.

6.2 Theoretical Model

6.2.1 Set-up

Consider a simple game between a firm's headquarters (H) and a (manufacturing) producer (M). Since the latter may eventually be owned to some degree by the former, we frequently refer to M as the subsidiary. The two parties can be located in the same or in different countries. Each firm is run by one owner-manager. The HQ possesses the idea (a blueprint) for the production of a differentiated final good, and the producer has the capacity to implement this idea. Without loss of generality, we normalize both parties' ex-ante outside options to zero.[12] Assuming constant elasticity of substitution (CES) preferences over varieties of a differentiated final good implies the following iso-elastic demand for a single variety:

$$x = Dp^{-1/(1-\alpha)}, \quad 0 < \alpha < 1,$$

[11] For instance, Masten (1984) studies procurement decisions of a large aerospace company, Baker and Hubbard (2003, 2004) examine ownership shares in the U.S. truck industry, Feenstra and Hanson (2005) consider the ownership structure of processing trade in China, and Acemoglu et al. (2010) investigate technological determinants of vertical integration in the UK. See also Klein (2015) for reviews of other empirical studies.

[12] Throughout the paper, we use 'ex-ante' to describe the point in time before the relationship-specific investments are sunk and 'ex-post' to describe the period thereafter. As will become clear below, both parties may have non-zero outside options ex-post.

whereby x and p denote quantity and price, respectively, $D > 0$ is a demand shifter, and α is a parameter related to the elasticity of substitution between any two varieties, $\sigma = 1/(1-\alpha)$. This demand function yields the following revenue:

$$R = x^{\alpha} D^{1-\alpha}. \tag{6.1}$$

Final goods are produced by M with a continuum of (*m*anufacturing) inputs $m(i)$, indexed by points on the unit interval, $i \in [0,1]$. One unit of $m(i)$ is produced with one unit of labor. Without loss of generality, we normalize the unit production costs of $m(i)$ to one. M combines these inputs into final goods according to the Cobb-Douglas production function:

$$x = \exp\left[\int_0^1 \ln m(i) \mathrm{d}i\right]. \tag{6.2}$$

Throughout the analysis, we assume that M is indispensable for the production of x, in the sense that H cannot manufacture final goods without M.[13]

Firms operate in an environment of contractual incompleteness, i.e., courts cannot fully verify and enforce all of the subsidiary's investments into intermediate inputs. To formalize this idea, we adopt the notion of partial contractibility from Acemoglu et al. (2007) and Antràs and Helpman (2008). More specifically, we assume that investments into inputs in the range $[0,\mu]$, with $0 \leq \mu \leq 1$, can be stipulated in an enforceable ex-ante contract, while investments into the remaining inputs cannot be verified by the courts and are therefore non-contractible. Following these authors, we interpret μ as the quality of contracting institutions in M's country. The idea behind this notion of contracting institutions is that a more efficient judicial system can enforce contracts over a wider range of product characteristics (see also Chapter 4 in Antràs, 2015). Clearly, there might also be technological factors that affect the degree of contractibility μ. Our modeling of μ as a country-specific variable reflects the notion that, for any given production technology, better contracting institutions are ceteris paribus more efficient at enforcing contracts. To consider an illustrative example, only well-functioning courts are able to verify whether high-tech inputs, such as computer chips, are produced according to the required standard. Hence, production of computer chips is contractible in countries with high judicial quality, but non-contractible in countries with bad contracting institutions.

We assume that M's inputs are customized to H's blueprint, and hence are (partially) relationship-specific. More precisely, by selling an input on the outside market, one can recoup only a fraction $(1-\rho)$ of the production costs, whereby $\rho \in [0,1]$ measures the degree of relationship-specificity. For $\rho = 0$, M's inputs have the same value for an outside party as within the current relationship, whereas $\rho = 1$ represents the case of fully relationship-specific inputs.[14] In what follows, we treat ρ as an industry-specific variable, i.e., subsidiaries in industries with a high ρ produce highly relationship-specific inputs (see also Antràs, 2015).

Since some of M's inputs are non-contractible ex-ante, H and M bargain over the surplus ex-post, i.e., after investments have been made. In anticipation of ex-post bargaining, H chooses ex-ante the optimal ownership share $s \in [0,1]$ in M. What are H's costs and benefits of choosing a higher s? The answer

[13] This assumption can be justified by the fact that H lacks either the production capacity or the expertise required to assemble the final goods (or both). This is the reason why the two parties need to form a relationship in the first place.

[14] Our modeling of relationship-specificity presupposes the existence of a perfectly competitive outside market. The assumption that M's inputs have a lower value for a tertiary party (as compared to the current relationship) reflects the idea that an outside buyer would have to incur additional costs to customize these inputs to her production process. This reduced-form approach can be rationalized by a richer model of the outside market along the lines of Grossman and Helpman (2001, 2002).

to this question crucially depends on the underlying theory of the firm. According to the Transaction-Cost Theory (TCT), a higher ownership share translates into a higher ability of H to 'dictate' to M the amount of non-contractible inputs, but it involves additional governance costs. The Property-Rights Theory (PRT) challenges this view by arguing that, regardless of the ownership share, H cannot enforce non-contractible inputs by fiat, and that the hold-up problem prevails even within firm boundaries. Instead, the PRT sees the role of ownership in shaping H's residual control rights, i.e., the authority to determine the use of M's inputs under circumstances that are not specified in a contract. As discussed in more detail below, the two theories provide diametrically opposed predictions regarding the effect of contracting institutions on the equilibrium ownership structure of the firm. We develop our main theoretical argument based on the PRT and provide a short discussion of the alternative predictions of the TCT in Section 6.2.4.

Following the PRT approach, we assume that ex-post negotiations take the form of generalized Nash bargaining. More precisely, each party obtains his or her outside option (i.e., the payoff in the absence of trade) plus a fraction of the ex-post gains from trade (the so-called quasi-rent), defined as revenue minus both parties' outside options. Let $\beta \in (0,1)$ denote the share of the quasi-rent accruing to H (henceforth, H's bargaining power), while the remaining share $(1-\beta)$ goes to M. If the bargaining breaks down, intermediate inputs can be sold on the outside market. Each party's outside option depends on the fraction of inputs he or she possesses. The HQ has enforceable ownership rights over contractible inputs $m(i)$, $i \in [0,\mu]$. The extent to which each party has residual control rights over non-contractible inputs depends on H's ownership share in M. More specifically, H controls the fraction s of non-contractible inputs, while M controls the remaining share $(1-s)$ of $m(i)$, $i \in [\mu,1]$. This 'zero-sum' notion of outside options builds on the idea of residual control rights by Grossman and Hart (1986), who argue that, "if one party gets rights of control, then this diminishes the rights of the other party to have control" (p. 693).[15]

The timing of events is as follows. In t_1, H chooses the ownership share s in M.[16] In t_2, H stipulates the amount of contractible inputs to be produced by M and commits to compensating him for the associated production costs. In t_3, M invests into non-contractible inputs and provides the amount of contractible inputs stipulated in the ex-ante contract. In t_4, the parties bargain over the surplus from the relationship. In t_5, final goods are produced and sold, and the revenue is distributed among the parties according to the agreements reached in t_2 and t_4. In the following section, we solve this game by backward induction.

6.2.2 Equilibrium

Before characterizing the subgame perfect equilibrium of the game outlined above, it is instructive to consider first the *hypothetical* case of complete contracts. If courts could perfectly verify and enforce investments into all intermediate inputs, the parties would stipulate the amount of $m(i), i \in [0,1]$, which maximizes the joint surplus:

$$\max_{\{m(i)\}_{i=0}^{1}} \pi = R - \int_0^1 m(i)\mathrm{d}i.$$

[15] The reader familiar with Antràs and Helpman (2004, 2008) will notice two differences in our modeling of outside options compared to their approach. First, while M's outside option in Antràs and Helpman (2004, 2008) is set to zero regardless of the ownership structure, it is equal to zero in our framework only under full integration (i.e., $s = 1$). Second, if the bargaining breaks down, H in the current framework cannot produce final goods on her own (cf. also footnote 13).

[16] Following Grossman and Hart (1986) and Hart and Moore (1990), we do *not* assume a direct cost of acquisition of (a larger share of) M. Our results remain qualitatively unchanged if we introduce a fixed cost of integration into the model.

Solving this maximization problem using equations (6.1) and (6.2) yields the first-best (FB) amount of inputs:

$$m(i) = \alpha R \equiv m^{FB} \quad \forall i \in [0,1], \tag{6.3}$$

whereby $R = D\alpha^{\alpha/(1-\alpha)}$ is obtained from plugging equations (6.2) and (6.3) into equation (6.1) and solving the resulting expression for R.

Consider now the relevant case of contractual incompleteness, introduced in Section 6.2.1. In t_4, each party obtains his or her outside option plus a fraction of the quasi-rent (Q), defined as follows:

$$Q = R - (1-\rho)(1-s)\int_{\mu}^{1} m(i)\mathrm{d}i - \left[(1-\rho)s\int_{\mu}^{1} m(i)\mathrm{d}i + (1-\rho)\int_{0}^{\mu} m(i)\mathrm{d}i\right], \tag{6.4}$$

whereby R is given by equation (6.1). The second term on the right-hand side represents M's outside option, which is equal to the outside value $(1-\rho)$ of the fraction $(1-s)$ of non-contractible inputs $m(i), i \in [\mu, 1]$. The term in the square brackets denotes H's outside option and consists of the outside value of the fraction s of non-contractible inputs, as well as the outside value of contractible inputs $m(i), i \in [0, \mu]$.

In t_3, M anticipates the outcome of Nash bargaining from period t_4 and chooses the amount of non-contractible inputs which maximizes her payoff from the ex-post negotiations net of production costs of these inputs:[17]

$$\max_{\{m(i)\}_{i=\mu}^{1}} \pi_M = (1-\rho)(1-s)\int_{\mu}^{1} m(i)\mathrm{d}i + (1-\beta)Q - \int_{\mu}^{1} m(i)\mathrm{d}i. \tag{6.5}$$

Using equations (6.1), (6.2), and (6.4), the solution to this maximization problem yields the optimal amount of non-contractible (n) inputs:

$$m(i) = \delta\alpha R \equiv m_n \qquad \forall i \in [\mu, 1], \tag{6.6}$$

as a function of revenue, obtained from plugging equations (6.2) and (6.6) into equation (6.1):

$$R = \left(\left[\exp\int_{0}^{\mu} \ln m(i)\mathrm{d}i\right]^{\alpha} (\delta\alpha)^{\alpha(1-\mu)} D^{1-\alpha}\right)^{\frac{1}{1-\alpha(1-\mu)}}, \tag{6.7}$$

whereby

$$\delta \equiv \frac{1-\beta}{1-\beta+s(1-\rho)+\rho\beta}. \tag{6.8}$$

Since $0 < \delta \leq 1$ for all $\beta \in (0,1)$ and $\rho, s \in [0,1]$, it can be seen immediately from the comparison of equations (6.3) and (6.6) that $m_n \leq m^{FB}$ for any given level of R. Intuitively, M anticipates ex-post hold-up with respect to non-contractible inputs and underinvests into these inputs compared to the first-best level.

The magnitude of M's underinvestments into non-contractible inputs (the size of m_n) depends crucially on the ownership share and the degree of relationship-specificity. Since these dependencies are key to understanding the main predictions derived in the next section, we formulate:

Lemma 6.1. *For any given level of revenue, the subsidiary's investments into non-contractible inputs (i) decrease in the ownership share, and (ii) this negative effect is mitigated by a higher relationship-*

[17] Note that contractible inputs do not enter M's maximization problem, since they are chosen by H in t_2, and M is fully compensated for the associated production costs.

specificity.

Proof. For part *(i)*, note that $\frac{\partial m_n}{\partial s}\big|_R < 0$ is implied by $\frac{\partial \delta}{\partial s} < 0$ from equation (6.8). For part *(ii)*, the cross partial-derivative of m_n with respect to s and ρ is $\frac{\partial^2 m_n}{\partial s \partial \rho}\big|_R = \frac{1-(1-\rho)(s-\beta)}{[1+(1-\rho)(s-\beta)]^3}\alpha(1-\beta)R$. Since $(s-\beta) \in (-1,1)$ for all $s \in [0,1]$ and $\beta \in (0,1)$, we immediately have $\frac{\partial^2 m_n}{\partial s \partial \rho}\big|_R > 0$ for all $\alpha \in (0,1)$, $\rho \in [0,1]$, and $R > 0$.

The intuition behind the first part of Lemma 6.1 derives from the fact that an increase in s ceteris paribus decreases M's outside option, and thereby worsens his ex-post bargaining position. If M expects to receive a smaller payoff ex-post, his ex-ante incentives to invest into m_n decrease. To understand the second part of Lemma 6.1, consider two different industries, one with a very high relationship-specificity (ρ approaching one) and one with a low relationship-specificity (ρ close to zero). In the highly relationship-specific industry, M's investments have only a small value on the outside market. Hence, a change in the ownership share s has little effect on M's outside option and on his payoff (see equation (6.5)). In other words, if the relationship-specificity is high, H can increase the ownership share without reducing M's investment incentives too much at the margin. By contrast, in an industry with a low degree of relationship-specificity, there is potentially much to gain for M on the outside market. Thus, any change in the ownership share affecting this relatively large outside option has a substantial impact on M's payoff. As a result, an increase in the ownership share strongly aggravates the underinvestment problem if the relationship-specificity is low. Generalizing this argument for all values of ρ, we conclude that a higher relationship-specificity mitigates the negative effect of an increased ownership share on the subsidiary's investment incentives.

Consider now H's optimization problem. In t_2, the HQ stipulates the amount of contractible inputs that maximizes her payoff from Nash bargaining net of the compensation for these inputs:

$$\max_{\{m(i)\}_{i=0}^{\mu}} \pi_H = (1-\rho)s(1-\mu)m_n + (1-\rho)\int_0^{\mu} m(i)\mathrm{d}i + \beta Q - \int_0^{\mu} m(i)\mathrm{d}i, \tag{6.9}$$

subject to M's participation constraint (PC), obtained from plugging equation (6.6) into equation (6.5):

$$\pi_M = (1-\beta)Q - (1-\mu)\left[1-(1-\rho)(1-s)\right]m_n \geq 0, \tag{6.10}$$

whereby Q and m_n are given by equations (6.4) and (6.6), respectively.[18] In our baseline analysis, we assume that M's PC is fulfilled and non-binding (i.e., $\pi_M > 0$), and solve the unconstrained maximization problem from equation (6.9). There are two reasons for this approach. First, it allows us to illustrate the HQ's key trade-off in the simplest possible manner. Second, we show in Appendix 6.A.1 that M's PC is slack for the vast majority of relevant parameter values. Intuitively, the need to incentivize M typically implies a more stringent upper bound on the optimal ownership share than the PC would. Nevertheless, we verify in Section 6.2.4 that our key predictions are qualitatively unchanged if the PC is binding and H solves the optimization problem from equation (6.9) with equation (6.10) as an equality constraint.

Using equations (6.4), (6.6), (6.7), and (6.8) in equation (6.9), and solving H's maximization problem for the optimal number of contractible (c) inputs, we obtain:

$$m(i) = \omega\alpha R \equiv m_c \qquad \forall i \in [0,\mu], \tag{6.11}$$

[18] The HQ also accounts for M's incentive compatibility constraint (ICC), which ensures that M utilizes non-contractible inputs $(1-s)(1-\mu)m_n$ within the current relationship rather than selling them on the outside market. Formally, the ICC is fulfilled whenever M's payoff from Nash bargaining is not smaller than his ex-post outside option, i.e., $(1-\rho)(1-s)(1-\mu)m_n + (1-\beta)Q \geq (1-\rho)(1-s)(1-\mu)m_n$. Notice that $Q \geq 0$ is a sufficient condition for M's ICC to hold. Since this condition is implied by M's PC from equation (6.10), the ICC may be ignored whenever the PC is fulfilled.

as a function of revenue, obtained from inserting equation (6.11) into equation (6.7):

$$R = \delta^{\frac{\alpha(1-\mu)}{1-\alpha}} \omega^{\frac{\alpha\mu}{1-\alpha}} \alpha^{\frac{\alpha}{1-\alpha}} D, \tag{6.12}$$

whereby

$$\omega \equiv \frac{s\alpha(1-\rho)(1-\mu) - \beta^2(1-\rho)\left[1-\alpha(1-\mu)\right] + \beta\left[1+s(1-\rho)-\alpha(1+s)(1-\mu)(1-\rho)\right]}{\left[1-\alpha(1-\mu)\right]\left[\rho+\beta(1-\rho)\right]\left[1-\beta+s(1-\rho)+\rho\beta\right]}. \tag{6.13}$$

In t_1, H chooses the optimal ownership share by solving the following maximization problem:

$$\max_s \pi_H = (1-\rho)s(1-\mu)\delta\alpha R - \rho\mu\omega\alpha R + \beta[R-(1-\rho)(1-\mu)\delta\alpha R - (1-\rho)\mu\omega\alpha R], \tag{6.14}$$

keeping in mind M's PC from equation (6.10). Plugging equations (6.8), (6.12), and (6.13) into equation (6.14), we obtain from the first-order condition the optimal ownership share:

$$s^*(\mu,\rho) = \frac{1+\beta^2(1-\rho)-2\beta-\alpha(1-\beta)(1-\mu)[1-\beta(1-\rho)]}{(1-\rho)[\beta+\alpha(1-\beta)(1-\mu)]}. \tag{6.15}$$

Plugging this ownership share as well as equations (6.8), (6.12), and (6.13) into equation (6.14), it can be shown that H's maximum profits from the relationship are positive for all admissible parameter values.

6.2.3 Comparative statics and testable predictions

In this section, we use comparative statics analysis to derive testable predictions regarding the effect of contracting institutions on the optimal ownership share. The relationship between s^* and μ is summarized in

Proposition 6.1. *The optimal ownership share increases in the quality of contracting institutions.*
Proof. $\frac{\partial s^*}{\partial \mu} = \frac{\alpha(1-\beta)^2}{(1-\rho)[\beta+\alpha(1-\beta)(1-\mu)]^2} > 0 \;\forall\; \alpha,\beta \in (0,1), \mu \in [0,1], \rho \in [0,1)$.

To understand the intuition behind this result, consider the trade-off faced by H when choosing s^*. On the one hand, a higher ownership share increases the HQ's outside option, and thereby raises H's profits specified in equation (6.9). On the other hand, a higher s^* reduces M's payoff (see equation (6.5)) and aggravates the ex-post hold-up from the viewpoint of M. This worsens M's ex-ante underinvestment in non-contractible inputs (see the first part of Lemma 6.1), and reduces the revenue from equation (6.7). Simply put, by choosing a higher ownership share in the subsidiary, the HQ trades off a larger *fraction* of surplus against a larger surplus *size*. When contracting institutions improve, the range of non-contractible inputs shrinks. This reduces the need for incentivizing M by giving him residual control rights. Hence, H optimally retains a larger share of the surplus for herself by choosing a higher ownership share s^*.

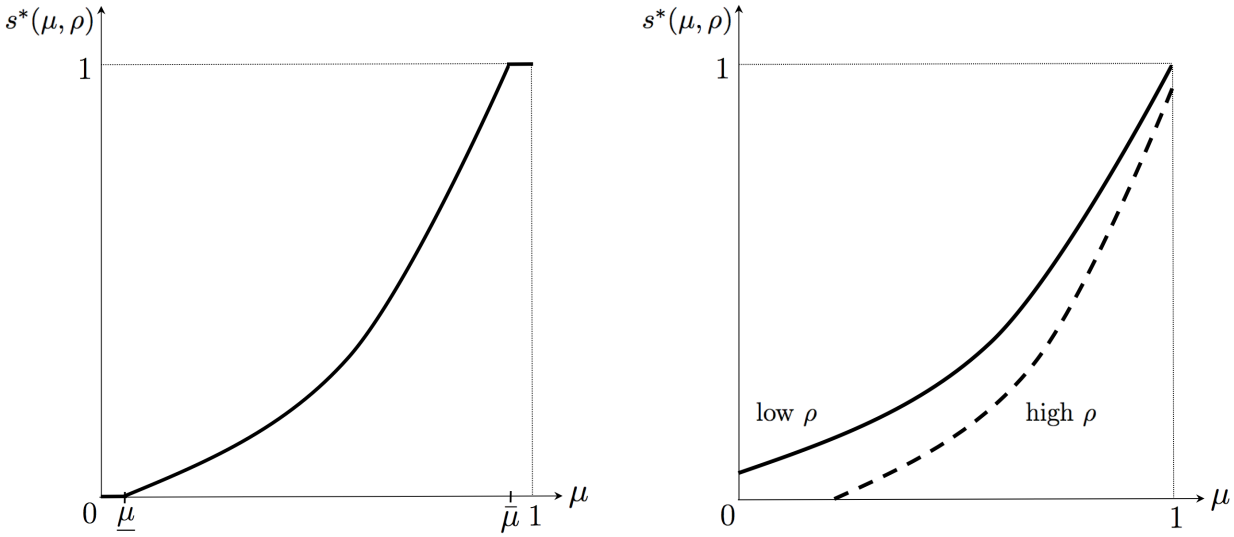
Figure 6.2(a) illustrates the positive relationship between s^* and μ derived in Proposition 6.1.[19] In an environment of poor contracting institutions, where μ is below some threshold $\underline{\mu}$, the HQ optimally chooses an ownership share of zero in order to provide maximal incentives for M. For $\mu \in (\underline{\mu}, \overline{\mu})$, the optimal ownership share increases monotonically in μ, reflecting the fact that better contracting institutions can enforce contracts on a wider range of inputs, and thereby substitute for the need to incentivize M's investment. For very high institutional quality, above the threshold $\overline{\mu}$, the HQ maximizes her share of the surplus by choosing full ownership. It should be noted that, for some parameter combinations, $\underline{\mu}$

[19] As depicted in the figure, the second-order derivative of s^* with respect to μ is positive: $\frac{\partial^2 s^*}{\partial \mu^2} = \frac{2\alpha^2(1-\beta)^3}{(1-\rho)[\beta+\alpha(1-\beta)(1-\mu)]^3} > 0$. The threshold values $\underline{\mu} = \frac{\beta[2-\alpha(2-\rho)]-\beta^2(1-\alpha)(1-\rho)+\alpha-1}{\alpha(1-\beta)[1-\beta(1-\rho)]}$ and $\overline{\mu} = \frac{\beta[3(1-\alpha)-\rho(1-2\alpha)]-\beta^2(1-\alpha)(1-\rho)+\alpha(2-\rho)-1}{\alpha(1-\beta)[2-\rho-\beta(1-\rho)]}$ can easily be derived from $s^*(\underline{\mu}) = 0$ and $s^*(\overline{\mu}) = 1$, respectively.

may lie below zero and $\overline{\mu}$ may exceed one. If $\underline{\mu} < 0$ and $\overline{\mu} > 1$, the optimal ownership share s^* lies strictly within the unit interval and it is strictly increasing in the quality of contracting institutions for all values of μ.

Figure 6.2: *Optimal ownership share* s^*

(a) *Direct effect of* μ (b) *Interaction between* μ *and* ρ



Consider next the interaction effect between μ and ρ in their impact on s^*, which is summarized in

Proposition 6.2. *The positive effect of contracting institutions on the optimal ownership share is stronger in industries with a higher degree of relationship-specificity.*

Proof. $\frac{\partial^2 s^*}{\partial\mu\partial\rho} = \frac{\alpha(1-\beta)^2}{(1-\rho)^2[\beta+\alpha(1-\beta)(1-\mu)]^2} > 0 \;\forall\; \alpha,\beta \in (0,1), \mu \in [0,1], \rho \in [0,1)$.

The intuition behind this key result builds on the insights from Proposition 6.1 and Lemma 6.1: According to Proposition 6.1, the optimal ownership share is monotonically increasing in the quality of contracting institutions. Also, Lemma 6.1 shows that the negative effect of a higher ownership share on M's investments into non-contractible inputs is mitigated in highly relationship-specific industries. Hence, if contracting institutions improve, H increases the optimal ownership share more strongly in industries with a higher degree of relationship-specificity, where the adverse effect of a higher s^* on M's investments is less severe.

Figure 6.2(b) illustrates the interaction effect between contracting institutions μ and relationship-specificity ρ. It plots the optimal ownership share s^* as a function of μ for a low value of ρ (solid line) and for a high value of ρ (dashed line). Reflecting Proposition 6.2, the line is steeper for the highly relationship-specific industry at all levels of μ. The more specific M's investments, the less does an increase in the optimal ownership share disincentivize these investments. Hence, H can exploit an improvement in institutional quality by increasing her ownership share more strongly in the highly relationship-specific industry.

Note that, while the effect of ρ on the *slope* of $s^*(\mu)$ is clear-cut, its effect on the *level* of s^* is a priori ambiguous. In the case depicted in Figure 6.2(b), the dashed line lies strictly below the continuous line. However, for alternative parameter combinations, it may lie strictly above this line or intersect it

once in the unit interval.[20] This ambiguity is explained by the interplay of two opposing effects: On the one hand, an increase in relationship-specificity ρ decreases M's outside option and reduces his ex-ante investments. On the other hand, a rise in ρ increases the surplus that M can obtain within the relationship (the quasi-rent from equation (6.4)), which improves his investment incentives. The relative magnitude of these opposing effects depends on M's bargaining power $(1-\beta)$.[21] In particular, if M's bargaining power is relatively low (the case depicted in the figure), he puts a high weight on his outside option compared to the quasi-rent. As a result, the net effect of an increase in relationship-specificity is that it provides disincentives, which must be compensated by a lower ownership share. In the alternative case, if M has a relatively high bargaining power, the positive effect of an increase in relationship-specificity on the quasi-rent dominates, which incentivizes M's ex-ante investments and allows H to retain a higher ownership share. Importantly, the positive interaction effect of μ and ρ on s^* predicted by Proposition 6.2 holds regardless of which case prevails.

6.2.4 Extensions and discussion

Before turning to an empirical test of Propositions 6.1 and 6.2, it is worth pausing to discuss their generality. In Section 6.2.4, we show that our main predictions continue to hold if M's participation constraint is binding, whereas considering ex-ante transfers would yield uninteresting results. In Section 6.2.4, we provide a generalization of the benchmark framework that incorporates joint production and allows us to relate our results to the seminal contributions by Antràs (2003) and Antràs and Helpman (2004). Finally, Section 6.2.4 provides a brief discussion of the predictions by the Transaction-Cost Theory regarding the effect of contracting institutions on firm boundaries, which are diametrically opposed to those of the PRT.

Participation constraint and ex-ante transfers

Recall that our baseline results were derived under the assumption that M's PC from equation (6.10) is not binding for any optimal ownership share given by equation (6.15). In Appendix 6.A.1, we provide a sufficient condition for M's PC to be non-binding and show that it is fulfilled for the vast majority of relevant parameter values. Nevertheless, we verify that our main theoretical results continue to hold also in those cases for which M's PC is binding. A tedious but straightforward analysis of H's maximization problems from equations (6.9) and (6.14), subject to M's PC from equation (6.10), yields the optimal ownership share:

$$s^*_{\text{PC}} = \frac{1-\beta-\alpha(1-\mu)[1-\beta(1-\rho)]}{\alpha(1-\rho)(1-\mu)}.$$

It can be verified that both the first-order derivative of this share with respect to μ as well as the cross-partial derivative with respect to μ and ρ are positive for all $\alpha, \beta \in (0,1), \mu, \rho \in [0,1)$:

$$\frac{\partial s^*_{\text{PC}}}{\partial \mu} = \frac{1-\beta}{\alpha(1-\rho)(1-\mu)^2} > 0, \quad \frac{\partial^2 s^*_{\text{PC}}}{\partial \mu \partial \rho} = \frac{1-\beta}{\alpha(1-\rho)^2(1-\mu)^2} > 0.$$

Hence, Propositions 6.1 and 6.2 continue to hold in case of a binding PC.

[20] Evaluating $\frac{\partial s^*}{\partial \rho}$ at the lower ($\underline{\mu}$) and upper ($\overline{\mu}$) threshold values of μ reported in footnote 19 yields $\frac{\partial s^*}{\partial \rho}|_{\mu=\underline{\mu}} = -\frac{\beta}{1-\rho} < 0$ and $\frac{\partial s^*}{\partial \rho}|_{\mu=\overline{\mu}} = \frac{1-\beta}{1-\rho} > 0$. Bearing in mind Proposition 6.1, there is a unique threshold $\hat{\mu}$ such that the dashed line is underneath the solid line for $\mu < \hat{\mu}$, and it lies above the solid line for $\mu \geq \hat{\mu}$. Solving $\frac{\partial s^*}{\partial \rho} = 0$ for μ yields the cutoff $\hat{\mu} = \frac{\alpha(1-\beta)+2\beta-1}{\alpha(1-\beta)}$. Since $\hat{\mu}$ is not restricted to the unit interval, the dashed line may cross the solid line at $\mu < 0$, $\mu > 1$, or $\mu \in [0,1]$.

[21] Formally, m_n from equation (6.6) increases in ρ if and only if $\beta < s$, and it decreases in ρ if this inequality if reversed.

Notice that our benchmark analysis does not allow for ex-ante lump sum transfers (side payments), which are frequently assumed in the literature to ensure that the entire surplus from the relationship accrues to one party (the HQ). As shown in Appendix 6.A.2, allowing for these transfers in the present context would result in an uninteresting case of zero optimal ownership shares, regardless of the quality of contracting institutions. To understand the intuition behind this result, recall the key trade-off faced by H in our model: By choosing a higher ownership share, H weighs a higher share of surplus against a larger surplus size. If she can extract the entire surplus from M via ex-ante transfers, this trade-off vanishes and H's only objective is to maximize the surplus size. Since both M's investments in non-contractible inputs and the overall revenue decrease in s (see equations (6.6) and (6.7)), H's optimal ownership share in the presence of ex-ante transfers is always equal to zero. To generate a non-trivial trade-off from the viewpoint of the HQ, the baseline model does not allow for ex-ante transfers.

Headquarter intensity

So far, we have assumed that all investments required for production are borne solely by M. One might wonder whether our predictions extend to the case in which both parties invest into relationship-specific and non-contractible inputs, resulting in a two-sided hold-up problem. To tackle this question, we introduce an element of joint production by assuming the Cobb-Douglas production technology from Antràs and Helpman (2004):

$$x = \left(\frac{h}{\eta}\right)^{\eta} \left(\frac{m}{1-\eta}\right)^{(1-\eta)}, \tag{6.16}$$

whereby h represents headquarter services provided by H, and $\eta \in (0,1)$ captures the relative importance of headquarter services in the production process (henceforth, headquarter intensity or HI). Each unit of h is produced with one unit of labor. Without loss of generality, we normalize H's unit production costs to one. As in the benchmark model, we assume that M produces a continuum of manufacturing inputs $m = \exp\left[\int_0^1 \ln m(i) \mathrm{d}i\right]$, whereby only the fraction $\mu \in [0,1]$ of the inputs $m(i)$ is contractible, while the remaining fraction $(1-\mu)$ cannot be verified and enforced by the courts. As before, we also assume that the parties can recoup a fraction $(1-\rho)$ of the production costs of manufacturing inputs on the outside market, whereby $\rho \in [0,1]$ captures the degree of relationship-specificity. To keep our model simple, we assume that headquarter services h are fully non-contractible and entirely relationship-specific. The timing of the game is identical to the one presented in the main text, apart from the period t_3, in which H now provides headquarter services, while M simultaneously and non-cooperatively invests into non-contractible manufacturing inputs and provides the amount of contractible manufacturing inputs stipulated in period t_2. This set-up implies a two-sided hold-up problem and ex-ante underinvestment by both parties. As shown in Appendix 6.A.3, solving the model yields the following optimal ownership share:

$$s^*_{\text{HI}} = \frac{1+\beta^2(1-\rho)\big[1-\alpha[1-\mu(1-\eta)]\big]-2\beta-\alpha\big[1-\mu(1-\eta)-\beta[2-\rho(1-\eta)-\mu(2-\rho)(1-\eta)]\big]}{(1-\rho)\Big[\beta-\alpha\big[\beta\big(1-\mu(1-\eta)\big)-(1-\eta)(1-\mu)\big]\Big]}. \tag{6.17}$$

Before discussing the effect of contracting institutions on the optimal ownership share, two remarks are in order. First, since s^*_{HI} from equation (6.17) reduces to s^* from equation (6.15) for $\eta = 0$, the equilibrium presented in this section generalizes the results of the one-sided hold-up game analyzed in Section 6.2.2. Second, the optimal ownership share increases in the headquarter intensity η for all

permissible values of $\alpha, \beta, \eta \in (0,1)$, $\mu \in [0,1]$, and $\rho \in [0,1)$:

$$\frac{\partial s^*_{\text{HI}}}{\partial \eta} = \frac{\alpha(1-\alpha)(1-\mu)(1-\beta)^2}{(1-\rho)\Big[\beta - \alpha\left[\beta\big(1-\mu(1-\eta)\big) - (1-\eta)(1-\mu)\right]\Big]^2} > 0.$$

This result squares well with the findings by Antràs and Helpman (2004) and the general logic of the PRT: As the headquarter intensity increases (i.e., manufacturing inputs become relatively less important in the production process), the need for incentivizing M decreases and the relative attractiveness of integration increases. Given that firm boundaries in the current framework constitute a continuous choice variable, our results complement the previous literature that has modeled firm boundaries as a binary decision between integration and arm's-length contracting.

Consider now the effect of contracting institutions on the optimal ownership share. Both the first-order derivative of s^*_{HI} with respect to μ, as well as the cross-partial derivative of s^*_{HI} with respect to μ and ρ are positive for all permissible parameter values:

$$\frac{\partial s^*_{\text{HI}}}{\partial \mu} = \frac{\alpha(1-\alpha\eta)(1-\eta)(1-\beta)^2}{(1-\rho)\Big[\beta - \alpha\left[\beta\big(1-\mu(1-\eta)\big) - (1-\eta)(1-\mu)\right]\Big]^2} > 0, \quad \frac{\partial^2 s^*_{\text{HI}}}{\partial \mu \partial \rho} = \frac{1}{(1-\rho)}\frac{\partial s^*_{\text{HI}}}{\partial \mu} > 0.$$

Hence, Propositions 6.1 and 6.2 continue to hold in the extended model in which both parties invest into relationship-specific and non-contractible inputs.[22]

Transaction-Cost Theory and Other Theories of the Firm

How do contracting institutions affect firm boundaries under the alternative assumptions of the TCT? And how does this effect depend on the relationship-specificity of the subsidiary's industry? Since these questions have been treated theoretically by Antràs (2015), we abstain from developing the TCT model in this paper, but rather provide a brief discussion of his results. According to the TCT, contracting institutions and relationship-specificity play no role under *integration* because the HQ can enforce the integrated producer's ex-ante investments by fiat (at the expense of exogenous governance costs). Yet, under *arm's-length contracting*, an improvement of contracting institutions increases the HQ's profits, in particular in industries with higher degrees of relationship-specificity.[23] Intuitively, a higher quality of contracting institutions mitigates the ex-post hold-up that plagues commercial transactions between independent parties and alleviates the ex-ante underinvestment into relationship-specific inputs. This effect is particularly pronounced in industries with high degrees of relationship-specificity since the hold-up problem in those industries is relatively more severe. As a result, the relative attractiveness of integration vs. arm's-length contracting under the TCT *decreases* in the quality of contracting institutions, and it decreases more strongly in highly relationship-specific industries (see equation (8.8) in Antràs, 2015). Notice that these TCT-based predictions point in the opposite direction compared to those suggested by the PRT, i.e., Propositions 6.1 and 6.2.[24] Since the TCT delivers the alternative hypotheses to the null hy-

[22] We have also explored the interaction effect $\frac{\partial^2 s^*_{\text{HI}}}{\partial \mu \partial \eta}$, but the sign of this cross-partial derivative turns out to be ambiguous.

[23] These results are reported in equation (5.14) in Antràs (2015) and are formally derived in his Appendix 2.

[24] Although the TCT predictions in Antràs (2015) are derived for a binary (rather than a continuous) integration decision, one can easily envision a simple extension of his model in which the HQ's profits are a convex combination of profits under integration (I) and arm's-length contracting (A), $\pi = s^*\pi_I + (1-s^*)\pi_A$, whereby the optimal ownership share $s^* \in [0,1]$ governs the weight of π_I vs. π_A in the HQ's profit function. Hence, when π_A increases due to an improvement in contracting institutions, the HQ has an incentive to reduce s^* in order to reap a higher π.

potheses from the PRT, testing our theoretical predictions empirically allows us to discriminate between these two alternative theories of the firm.

6.3 Empirical Implementation

6.3.1 Econometric specifications

The point of departure for our econometric specifications is the optimal ownership share predicted by our model in equation (6.15). To concentrate on the effect of contracting institutions and their interaction with relationship-specificity, we specify econometric models that are linear in parameters, while relying on large sets of fixed effects to control for the other determinants of ownership shares from the theoretical model. We implement our empirical tests of Propositions 6.1 and 6.2 in two steps.

As a first step, we investigate Proposition 6.1 by estimating the following econometric model:

$$S_{hm} = \gamma \cdot C_\ell + \boldsymbol{\varphi} \cdot \mathbf{X}_{hm} + \varepsilon_{hm}, \tag{6.18}$$

where S_{hm} denotes the share in subsidiary m (active in industry i and country ℓ) that is owned by headquarters h (which may be active either in the same or in a different industry and country, respectively). The explanatory variable of primary interest is the quality of contracting institutions C_ℓ in the subsidiary's country ℓ, and γ is the key parameter to be estimated. The vector $\mathbf{X}_{hm}$ contains sets of control variables (with associated coefficient vector $\boldsymbol{\varphi}$), which vary by specification, and ε_{hm} is an error term. In our preferred specification, $\mathbf{X}_{hm}$ includes observable characteristics of the subsidiary's country, of the ownership structure, and proxies for bilateral investment costs specific to the country pair, as well as full sets of fixed effects (FE) for the subsidiary's industry and the HQ's country-industry. These FE play an important role in absorbing other determinants of S_{hm} according to our theory: the demand elasticity (reflected in α), relationship-specificity (ρ), and headquarter intensity (η), to the extent that these variables are industry-specific; as well as the country-industry-specific component of the HQ's bargaining power (β). The specific list of all variables, their measurement, and the data sources are detailed in Section 6.3.2.

As shown in Proposition 6.1, based on the PRT we expect a positive effect of the quality of contracting institutions on the ownership share, reflected in $\gamma > 0$. That is, for subsidiaries located in countries with better contracting institutions, we should observe ceteris paribus higher degrees of integration. Intuitively, firms in such countries can write enforceable contracts on a larger share of the subsidiary's inputs, and hence, there is less need to incentivize a subsidiary's ex-ante investments by giving him residual control rights.[25] However, we caution against interpreting the estimated coefficient as a causal effect of contracting institutions. While we control for various confounding factors, there remains a concern regarding omitted variables specific to the subsidiary's country, such as cultural characteristics or other features of the institutional environment, that may be correlated with C_ℓ and also affect S_{hm}. Hence, we interpret our estimates $\hat{\gamma}$ as conditional correlations.

We make progress towards identifying the causal effect of contracting institutions on firm boundaries by estimating the differential effect of contracting institutions across industries that differ in their

[25] A negative estimate of γ would lend support to the alternative hypothesis suggested by the TCT (see Section 6.2.4).

relationship-specificity. To test this interaction effect, we set up the econometric model:

$$S_{hm} = \beta_1 \cdot (C_\ell \times R_i) + \beta_2 \cdot C_\ell + \beta_3 \cdot R_i + \boldsymbol{\phi} \cdot \mathbf{Y}_{hm} + e_{hm}, \tag{6.19}$$

where we are mainly interested in the interaction term of contracting institutions C_ℓ in the subsidiary's country ℓ and the relationship-specificity R_i of the subsidiary's industry i. This approach allows us to address the concern regarding omitted variables by including subsidiary country FE in the vector of control variables $\mathbf{Y}_{hm}$, in addition to all of the elements of $\mathbf{X}_{hm}$ from equation (6.18). In our preferred specification of equation (6.19), we further control for country-pair FE to absorb all unobserved features of bilateral investment costs. The parameters to be estimated are $\beta_j, j \in \{1, 2, 3\}$, and the vector $\boldsymbol{\phi}$, and e_{hm} denotes the error term.

As derived formally in Proposition 6.2, the PRT predicts a positive interaction effect, i.e., $\beta_1 > 0$. Intuitively, a higher relationship-specificity mitigates the negative effect of the ownership share on the subsidiary's investments, and therefore allows the HQ to increase her ownership share more strongly in response to better contracting institutions. Thus, cross-country differences in institutional quality should have a stronger positive effect on the ownership share in industries with a higher degree of relationship-specificity.[26]

By exploiting the interaction between country-level institutions and industry-level technological characteristics, equation (6.19) resembles a difference-in-differences model, where we control for the respective first differences with country and industry FE. It is reminiscent of the econometric models traditionally used to assess the effect of institutions on international trade patterns, as discussed by Nunn and Trefler (2014).[27] However, there are two crucial differences between our model and this approach. First, by looking at ownership shares, we examine the intensity of investment links instead of trade flows. Second, our micro data analysis exploits variation across different subsidiary countries and industries within a given HQ's country-industry cell (and even within HQ in some robustness checks), in contrast to the analysis of comparative advantage, which is typically conducted at the aggregate level of industries and countries.[28]

We estimate both models by Ordinary Least Squares (OLS) due to the well-known complications that arise in non-linear models when interpreting interaction terms (see Ai and Norton, 2003), as in the case of our main variable of interest.[29] We compute two-way cluster-robust standard errors following the procedure suggested by Cameron et al. (2011).[30] First, we cluster at the level of the key explanatory variable, i.e., the level of the subsidiary's country in equation (6.18) and the subsidiary's country-industry level in equation (6.19). Second, we cluster at the level of the HQ in all OLS estimations to account for interdependencies across a given HQ's ownership decisions.

Before turning to the description of the data sources, three comments on the empirical implementation of our model are in order. First, while the key mechanism of the PRT is well understood in organizational economics, it has typically been applied to discrete ownership choices between integration and arm's-length contracting. A novelty of our approach lies in considering varying intensities of

[26] Note that the TCT yields the opposite prediction for the sign of the interaction effect (see Section 6.2.4).

[27] Acemoglu et al. (2007), Costinot (2009), Levchenko (2007), and Nunn (2007) show that contracting institutions can constitute a source of comparative advantage in international trade.

[28] Ma et al. (2010) and Wang et al. (2014) are recent exceptions analyzing the role of institutions for firm-level exports.

[29] In a robustness check, we allow for non-linearities using an ordered logistic regression model (see Section 6.4.3).

[30] Estimations are implemented using the Stata routine `reghdfe` provided by Correia (2014), which efficiently absorbs our high-dimensional FE and allows for both multi-way clustering of standard errors as well as the use of instrumental variables.

integration across firm pairs. However, one may be skeptical of whether *marginal* changes in owners' equity *shares* can indeed have significant effects on the incentives of the subsidiary's manager. Direct evidence for the relevance of this mechanism is provided for instance by Baker and Gompers (1999) using data on the equity shares held by CEOs in newly public firms. The authors show that reductions in the CEO's equity share due to the participation of venture capitalists or due to an initial public offering significantly reduce the CEO's incentives.

Second, and related, our theoretical model assumes that both parties' residual control rights change in response to a marginal increase in the ownership share at any point in the unit interval. In practice, however, particular thresholds of ownership shares (e.g. 50% or 100%) may be critical for obtaining residual control rights. We exploit the full variation in ownership shares in our main analysis but address the possibility of non-linearities by considering a discrete choice across ownership categories in a robustness check in Section 6.4.3.

Third, by approximating the contractibility μ of the subsidiary's investments with the quality of contracting institutions in the subsidiary's country, we have implicitly assumed that courts in this country are responsible for enforcing the subsidiary's investment decisions. This assumption indeed seems to reflect the prevailing legal practice in many countries.[31] It seems possible, however, that μ may also be affected by the quality of contracting institutions in the HQ's country – either directly, if courts in the HQ's country rule over contracts between the two firms; or indirectly, if multinationals transfer their institutional practices to their subsidiaries (see Chari et al., 2010). Note that we control for the potentially confounding role of contracting institutions in the HQ's country via FE.

6.3.2 Data sources

All data on firms and firm pairs used in our analysis are taken from the Orbis database provided by Bureau van Dijk (BvD). Most importantly, we observe firms' direct ownership shares (in percent) in their subsidiaries in 2014, which we use as our measure of S_{hm}.[32] We further exploit information on the firms' main activities (industry affiliations in the form of four-digit NAICS 2012 codes), their founding years, employment, and key balance sheet items in 2013. The three key advantages of the Orbis database for our purpose are the availability of firm-pair specific ownership information, its large international coverage, and the fact that it includes both domestic and international ownership links. The database is unique in encompassing all three of these features.[33]

Our sample includes firms that are classified by BvD as medium, large, and very large. We consider only those HQ that are classified as 'industrial companies', thereby excluding pension funds, public authorities, and financial companies. Furthermore, we restrict our analysis to direct ownership shares

[31] For instance, the European Council Regulation (EC) No 44/2001 (see http://eur-lex.europa.eu/LexUriServ/LexUriServ.do?uri=OJ:L:2001:012:0001:0023:en:PDF) cites the default legal principle that "jurisdiction is generally based on the defendant's domicile" (in our context, the subsidiary's country). This principle typically applies to contracts between firm pairs within the EU (and potentially also to cases in which one of the two firms is an EU resident), unless specified otherwise by the contracting parties. Also, the Chinese Law on Sino-Foreign Equity Joint Ventures explicitly stipulates that "All activities of an equity joint venture shall be governed by the laws and regulations of the People's Republic of China" (see http://english.mofcom.gov.cn/article/lawsdata/chineselaw/200301/20030100062855.shtml, both accessed on May 10, 2017).

[32] These data were provided by BvD in the form of two customized data extractions in 2014 (for balance sheets and other firm variables) and 2015 (for ownership shares). The exact date of the ownership information varies by observation and depends on the latest information available, but it refers to the year 2014 in the majority of cases. We additionally have biannual information on ownership shares for the years 2004, 2006, 2008, 2010, and 2012, which we use in robustness checks.

[33] The Orbis ownership data have previously been used to study the international transmission of shocks through multinationals (Cravino and Levchenko, 2017), the hierarchical complexity of business groups (Altomonte and Rungi, 2013), and the effect of managerial culture on firm boundaries (Kukharskyy, 2016; Gorodnichenko et al., 2017).

of at least 10%, a conventional threshold for direct investment. We implement these sample restrictions because we are interested in HQ that have a (potentially long-term) economic interest in the target firm – as described by our model – and do not merely invest due to (short-term) portfolio considerations. The resulting sample includes information on direct ownership shares for 605,547 firm pairs of 288,450 headquarters from 113 countries owning 562,243 subsidiaries in 126 countries in the cross-section of 2014. The median HQ has only a single subsidiary, which is typically located in the same country; yet, one third of all HQ owns shares in at least two subsidiaries, and 11% of them are multinational firms owning foreign subsidiaries (which make up around one fifth of the observations in our data). In this sample, the mean ownership share is 75%, with a standard deviation of 30. Full ownership is most common (46% of all firm pairs) and 18% of the ownership shares range between 50 and 51%. We explicitly address these distributional features in Section 6.4.3.

We take the 'rule of law' index from the Worldwide Governance Indicators (Kaufmann et al., 2010) as our baseline measure of the quality of a country's contracting institutions C_ℓ. This measure is a weighted average of a number of variables that reflect experts' and practitioners' assessments of the effectiveness and predictability of judicial quality and the enforcement of contracts in a given country and year. We use this index as our main measure since it is available for a large number of countries and is well-established in the literature as a valid proxy for the quality of a country's contracting institutions (see e.g. Antràs, 2015; Nunn, 2007; Nunn and Trefler, 2014). However, we test the sensitivity of our results to using a wide range of alternative proxies for the quality of contracting institutions, which we describe in Section 6.4.3. Table 6.B.1 provides a full list of subsidiary countries in our data, ranked by the rule of law index. The table shows that both the number of subsidiaries and the average ownership shares tend to be lower in countries with poor contracting institutions.

Our industry-level measure of relationship-specificity R_i is taken from Antràs and Chor (2013), who compute it from the Rauch (1999) classification of products by their degree of horizontal differentiation.[34] This classification distinguishes three categories of goods: (i) homogenous (traded on an organized exchange), (ii) reference-priced (not sold on an organized exchange, but reference prices are quoted in trade publications), and (iii) differentiated (all residual goods). For each industry, our baseline measure of R_i is calculated as the share of product codes in the industry that are classified as differentiated or reference-priced.[35] The idea underlying this approach is that, unlike homogenous goods, differentiated goods are customized to the specific needs of a buyer-seller relationship. The more differentiated goods there are within a given industry, the thinner is the outside market for the typical goods produced in this industry, and hence, the higher is the relationship-specificity. The size of the sample for our main regression analysis is reduced by the availability of the Rauch relationship-specificity measure for the subsidiary's industry. Summary statistics for the main estimation sample are provided in Table 6.B.2.

The vectors of control variables $\mathbf{X}_{hm}$ and $\mathbf{Y}_{hm}$ include, in addition to various sets of FE, the following observable characteristics of the subsidiary's country and industry, the ownership structure, and

[34] These data are available on the authors' websites at the six-digit 2002 U.S Input-Output industry classification level. We map them to the four-digit NAICS 2012 level using official correspondence tables provided by the Bureau of Economic Analysis: http://www.bea.gov/industry/xls/2002DetailedItemOutput.xls and by the US Census Bureau: http://www.census.gov/eos/www/naics/concordances/concordances.html (both accessed on April 4, 2016).

[35] Due to ambiguities for some goods, there are two versions of the Rauch (1999) classification, a 'conservative' and a 'liberal' one, whereby the former maximizes and the latter minimizes the number of goods that are classified as differentiated. Following Alfaro et al. (2015) and Antràs and Chor (2013), we use the liberal classification in our baseline analysis and the conservative version in robustness checks. Also, reference-priced goods may be understood as either differentiated or homogenous. We treat reference-priced goods as differentiated in our baseline analysis and classify them as non-differentiated in robustness checks.

the country pair, which may affect the depth of integration. For the subsidiary's country, we take the log of GDP as a measure of country size; the log of GDP per capita as a proxy for the income and wage level; the log of the endowment ratio (K_ℓ/L_ℓ), defined as the real capital stock divided by employment (average hours worked by employed persons), as a measure of relative factor abundance; and the average years of schooling as a proxy for the human capital stock (Barro and Lee, 1996). These variables are taken from the Penn World Tables (version 8.1; see Feenstra et al., 2013) for the year 2010. To control for industry-level differences in capital intensity, we include the log of the capital-to-employment ratio (K_i/L_i) of the median firm by industry in the Orbis dataset for the year 2013, along with a Heckscher-Ohlin-type interaction term of the industry's capital intensity with the endowment ratio: ln $(K_i/L_i)\times$ ln (K_ℓ/L_ℓ). We further control for three variables capturing the nature of the ownership structure: (i) the number of subsidiaries of the HQ, (ii) the number of shareholders of the subsidiary, and (iii) a dummy indicating domestic (as opposed to international) ownership links.[36] The first two variables capture the complexity of the business group (cf. Altomonte and Rungi, 2013; Schwarz and Suedekum, 2014). The motivation for including the third variable is that international ownership may involve additional fixed costs that affect the ownership shares. We proxy for bilateral investment costs by using a set of standard gravity control variables from the CEPII dataset (Head et al., 2010): the distance between the most populous cities in log kilometers, the time zone difference in hours, and indicator variables for countries sharing a common border, official language, or (current or past) colonial link.

Finally, the vectors $\mathbf{X}_{hm}$ and $\mathbf{Y}_{hm}$ also include control variables for other characteristics of the institutional environment in the subsidiary's country. These characteristics, such as financial development or intellectual property rights protection, are correlated with the quality of contracting institutions, and they may also affect the integration decision. To ensure that the rule of law index does not pick up the effects of other types of institutions, we control for them using a set of proxies that have previously been used in the international economics literature (see e.g. Nunn and Trefler, 2014; Javorcik, 2004): financial development, approximated by the sum of private credit and stock market capitalization divided by GDP from the World Bank's Global Financial Development Database (GFDD) in 2012; labor market flexibility, defined as one minus the rigidity of employment index from the World Bank's Doing Business Reports (based on Botero et al., 2004), averaged over the period 2004-2009 (the years when the index was reported); the index of intellectual property rights (IPR) protection developed by Park (2008) in 2010 (the last available year); and the expropriation risk score in the first quarter of 2014, based on expert assessments by the information services company IHS Markit.[37]

6.4 Estimation Results

6.4.1 Ownership shares and contracting institutions

Table 6.1 summarizes our estimation results for different specifications of equation (6.18). It substantiates the fact illustrated in Figure 6.1(a) that subsidiaries are more deeply integrated in countries with better contracting institutions. Column 1 reports a positive unconditional correlation between the owner-

[36] Due to missing observations for the variables provided in Orbis, we define the number of subsidiaries (shareholders) as the maximum value of the number of subsidiaries (shareholders) reported by BvD (which may include non-manufacturing firms, public entities, or private persons) and the number of subsidiaries (shareholders) actually observed in the database.

[37] The score assesses the "risk that the government will expropriate or nationalise assets". A key advantage of the country risk scores by IHS Markit is that they distinguish the risk of expropriation by the government from the risk that the judicial system may not enforce contracts between private parties, which we exploit in a robustness check.

ship share and the rule of law index, which is highly significant. The coefficient estimate increases as we successively add control variables for characteristics of the subsidiary's country and industry, the ownership structure (column 2), and bilateral investment costs (column 3).[38] Ownership shares tend to be lower in countries with a more educated population, while the other country characteristics are insignificant. The industry's capital intensity enters negatively, while the Heckscher-Ohlin interaction term is weakly positive. Domestic ownership links are characterized by lower shares than international links, presumably reflecting the idea that investing abroad is associated with additional fixed costs, which are only worth paying in case of a substantial stake in a foreign company. The average ownership share also decreases (somewhat mechanically) in the number of shareholders. Most proxies for bilateral investment costs seem to have no significant effect on ownership shares, but we find a stable negative correlation with the common language dummy.

The positive correlation between the depth of integration and the quality of contracting institutions is also confirmed in columns 4-5 of Table 6.1, where we control for unobservable factors using large sets of FE. After adding FE for the subsidiary's country, the HQ's country, and the HQ's industry in column 4, the point estimate is reduced almost by half, but it remains highly significant. The estimate is very similar when we further include HQ country-industry FE in column 5. In column 6, we control for several other dimensions of institutional quality in the subsidiary's country, namely financial development, labor market flexibility, IPR protection, and expropriation risk. The results indicate that ownership shares are higher in more financially developed countries, those with better IPR protection, and (perhaps surprisingly) countries with higher expropriation risk. Most importantly, the estimated coefficient for the rule of law index remains positive and significant at the one percent level. This preferred estimate suggests that firms choose ownership shares which are higher by around 11 percentage points in countries with a rule of law index that is higher by one standard deviation (in the cross-country sample in 2014).

We interpret this result as indicative evidence for Proposition 6.1, which predicts that firms choose deeper integration of their subsidiaries located in countries with better contracting institutions. However, despite the large set of control variables included in our preferred specification from column 6 of Table 6.1, there may be unobserved country-specific factors that are correlated with contracting institutions and firm boundaries. Hence, these regressions cannot identify the causal effect of contracting institutions on ownership shares. This caveat notwithstanding, we can state that after controlling for a wide range of relevant covariates and a host of unobservables, we continue to find a positive correlation between the quality of contracting institutions in the subsidiary's country and the depth of integration, in line with the PRT.

Before turning to the analysis of the interaction effect, we briefly pause to compare our results to the existing literature. Note that the implications of our estimates differ from Acemoglu et al. (2009), who find no significant correlation between their firm-level index of vertical integration and various proxies for contracting institutions. One potential explanation is that they examine (primary and secondary) activities that are fully integrated within the *same firm*, whereas we examine partial ownership shares *across* firms. However, besides these conceptual differences, the discrepancy in results may also be explained by the confounding role of contracting institutions in the HQ's country. To see this, note that Acemoglu et al. (2009) examine the degree of vertical integration of a given firm by examining input-output relationships between its various economic activities, all of which are presumably located

[38]The number of observations declines with more demanding specifications in Table 6.1 because control variables are missing for some firm pairs and observations are dropped if they are fully explained by the FE.

Table 6.1: *Ownership shares and contracting institutions*

Dependent variable: ownership share	Plain	Control variables		Fixed effects (FE)		Other
		subsidiary	bilateral	3 FE	2 FE	institutions
	(1)	(2)	(3)	(4)	(5)	(6)
Rule of law (subsidiary country)	5.639***	13.78***	13.41***	7.657***	7.439***	11.12***
	(2.130)	(3.552)	(3.298)	(1.727)	(1.761)	(2.540)
ln GDP (subsidiary country)		1.868	3.264	0.528	0.180	-0.626
		(3.450)	(3.061)	(0.906)	(1.008)	(3.995)
ln GDP per capita (subsidiary country)		-6.764	-6.286	-2.913	-2.265	0.606
		(4.790)	(4.847)	(2.931)	(3.034)	(7.956)
ln (K_ℓ/L_ℓ) (subsidiary country)		-8.084	-9.103	-2.932*	-2.413	-2.715
		(5.954)	(5.582)	(1.556)	(1.640)	(4.978)
ln years of schooling (subsidiary country)		-48.31**	-45.47*	-25.72**	-26.66**	-57.51***
		(23.27)	(24.42)	(10.47)	(10.70)	(14.92)
ln (K_i/L_i) (subsidiary industry)		-5.132**	-4.753*			
		(2.578)	(2.625)			
ln $(K_\ell/L_\ell)\times$ ln (K_i/L_i) (subsidiary)		0.639*	0.585*	0.168	0.106	0.0570
		(0.339)	(0.347)	(0.173)	(0.170)	(0.159)
Domestic ownership link dummy		-8.920***	-15.01***	-8.432***	-7.779***	-3.942**
		(1.995)	(4.333)	(1.602)	(1.564)	(1.917)
Number of subsidiaries (headquarters)		-0.0158*	-0.0134	-0.00854	-0.0137*	-0.0114
		(0.00863)	(0.00877)	(0.00629)	(0.00705)	(0.00777)
Number of shareholders (subsidiary)		-0.787**	-0.788**	-0.770**	-0.739**	-0.665**
		(0.366)	(0.367)	(0.360)	(0.347)	(0.318)
ln distance			-1.186	0.0639	-0.114	1.552
			(2.376)	(0.681)	(0.663)	(1.022)
Time zone difference (hours)			-1.084*	-0.330	-0.271	-0.335
			(0.646)	(0.422)	(0.409)	(0.340)
Common border dummy			-1.298	0.672	0.533	1.622
			(2.909)	(1.076)	(0.990)	(1.321)
Common language dummy			-5.979***	-4.279**	-4.518**	-3.301*
			(1.801)	(1.808)	(1.714)	(1.708)
Colonial link dummy			1.981	1.759	1.882	1.005
			(1.907)	(2.351)	(2.228)	(2.328)
Financial development (subsidiary country)						0.0441***
						(0.0137)
Labor market flexibility (subsidiary country)						-6.369
						(4.508)
IPR protection (subsidiary country)						8.078*
						(4.466)
Expropriation risk (subsidiary country)						11.21***
						(2.195)
Subsidiary industry fixed effects (FE)	no	no	no	yes	yes	yes
Headquarters country and industry FE	no	no	no	yes	nested	nested
Headquarters country-industry FE	no	no	no	no	yes	yes
Observations	605,547	525,653	524,597	524,206	522,792	433,108
R^2	0.026	0.127	0.133	0.229	0.265	0.278

The table reports estimates of (variations of) equation (6.18). Standard errors clustered by subsidiary country and by HQ are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

in the same country. Yet, as shown by Antràs and Helpman (2008), contracting institutions have the opposite effect on firm boundaries depending on whether they affect the contractibility of a HQ's or a subsidiary's inputs. If both the subsidiary and the HQ are located in the same country, a single country-level proxy for contracting institutions may confound these opposing effects. Our data features two important advantages that enable us to resolve this issue: They include international ownership links as well as multiple links for some firms. This allows us to fully control for contracting institutions in the

HQ's country by FE. Provided that the HQ's investments are fully governed by contracting institutions in the HQ's country, we can identify the positive relationship between contracting institutions in the subsidiary's country and the depth of integration.

6.4.2 Ownership shares, contracting institutions, and relationship-specificity

In Table 6.2, we provide empirical evidence supporting the idea that better contracting institutions have a stronger positive effect on ownership shares for subsidiaries producing highly relationship-specific inputs, as predicted by Proposition 6.2. The table develops our preferred specification of equation (6.19) step by step. In column 1, we examine the correlation without any control variables, which reveals a positive estimate of the interaction term. It suggests that the positive correlation between the rule of law index and ownership shares is concentrated in industries with high relationship-specificity, in line with Figure 6.1(b).

Table 6.2: *Ownership shares, contracting institutions, and relationship-specificity*

Dep. var.: ownership share	Plain	Controls	3 FE	4 FE	Cty-ind. FE	Pair FE
	(1)	(2)	(3)	(4)	(5)	(6)
Rule of law × specificity (subsidiary)	6.975***	3.327	2.748**	2.464**	3.476***	3.432***
	(1.634)	(2.270)	(1.172)	(1.085)	(0.973)	(0.909)
Rule of law (subsidiary country)	-1.703	7.734***	4.924***			
	(1.502)	(2.261)	(1.200)			
Specificity (subsidiary industry)	-2.090	0.197				
	(1.324)	(2.160)				
Control variables from Table 6.1	no	yes	yes	yes	yes	yes
Subsidiary industry FE	no	no	yes	yes	yes	yes
Subsidiary country FE	no	no	no	yes	yes	nested
Headquarters country and industry FE	no	no	yes	yes	nested	nested
Headquarters country-industry FE	no	no	no	no	yes	yes
Country-pair FE	no	no	no	no	no	yes
Observations	230,937	187,716	187,587	187,587	186,092	186,110
R^2	0.024	0.128	0.228	0.242	0.288	0.295

The table reports estimates of (variations of) equation (6.19). Standard errors clustered by subsidiary country-industry and by HQ are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

In column 2, when we control for the observable variables from the third column of Table 6.1, the point estimate for the interaction term becomes smaller and insignificant.[39] However, significance is restored as we add FE for the HQ's country and industry along with subsidiary industry FE in column 3. As an important step towards identification, we add subsidiary country FE in column 4. Note that this specification is superior to simple cross-country regressions, as it identifies the effect of country-level institutions across industries with varying degrees of relationship-specificity after controlling for any unobservable country characteristics. We control for further potential confounding factors, such as international differences in financing conditions of a given industry, by adding HQ country-industry FE in column 5. Finally, to account for unobserved bilateral factors like cultural differences or ethnic ties, we include country-pair FE to arrive at our preferred specification in column 6. In all of the regressions with FE, we find a significantly positive estimate for the interaction effect. The estimated size of the effect

[39] For relationship-specificity, we find insignificant estimates with switching signs, which is not surprising given the ambiguity predicted by our model (see Section 6.2.3). The coefficient estimates for the control variables are not reported to save space. We choose not to include the additional institutional measures from column 6 of Table 6.1 because they would substantially reduce the sample and are subsequently controlled for via subsidiary country FE.

is quite stable across all specifications in columns 2 through 6. A quantitative interpretation of the preferred estimate suggests that an improvement in contracting institutions by one standard deviation would increase the ownership share by 3.4 percentage points *more* for a subsidiary in a highly relationship-specific industry (composed of differentiated goods only) compared to a subsidiary in a non-specific (homogenous) industry.

Our estimation results provide strong support for Proposition 6.2, and hence for the PRT. In line with this theoretical prediction, we find that firms choose ceteris paribus deeper integration of subsidiaries in countries with better contracting institutions, and this effect increases in the relationship-specificity of the subsidiary's industry. Intuitively, there is less need to incentivize the subsidiary's ex-ante investments if contracting institutions are better and courts are able to enforce contracts on a wider range of issues. This mechanism is more pronounced for subsidiaries producing relationship-specific inputs, since the adverse effect of a higher ownership share on the subsidiary's underinvestment is mitigated in industries with higher degrees of relationship-specificity. Therefore, contracting institutions have a disproportionately positive effect on the integration decision in relationship-specific industries.

As mentioned in the introduction, the only empirical analysis to date that has studied the interaction effect of contracting institutions and relationship-specificity on firm boundaries is by Antràs (2015, see his Table 8.9). Using the ratio of intra-firm imports to arm's-length imports by U.S. industries as a dependent variable, he finds a negative and significant interaction effect, which is however not robust to more stringent testing. In particular, the estimate is rendered insignificant after controlling for interactions of GDP per capita with industry dummies. We propose three possible explanations for the difference between these inconclusive results and our positive estimates, which are robust to the same test and several other checks, as shown in the next section. First, since our sample covers HQ in many countries, while Antràs (2015) considers only U.S. imports, the differences may be explained by particular features of the U.S. economy. Second, larger ownership shares do not necessarily translate into higher intra-firm trade volumes. In particular, the U.S. data classify all firm pairs with ownership shares exceeding 6% as related parties, and there might be large trade volumes between partially integrated firms. Third, and related, compositional effects within industries may mask the positive interaction effect that we have identified at the level of firm pairs.[40]

6.4.3 Robustness analysis

This section discusses the following challenges to identification: measurement of key explanatory variables, potentially remaining confounding factors, robustness across samples, selection into production countries, non-linearities in ownership decisions, and the possibility of reverse causality. We use several variations of our preferred specifications as well as alternative estimation techniques to find that our main results are remarkably robust to addressing each of these challenges.

Alternative measurement of key variables

As a first step, Table 6.3 experiments with alternative measures of our key explanatory variables. In Panel A, we reestimate equation (6.18) using several alternative proxies for the quality of contracting institutions C_ℓ, and in Panel B, we employ the same proxies to estimate the interaction effect in equation (6.19). All regressions are based on our preferred specifications from the final columns of Tables 6.1

[40] While Antràs (2015) controls for productivity dispersion within industries, this term approximates only one of several relevant dimensions of micro-level heterogeneity that we account for in Section 6.4.3.

and 6.2, respectively. We use the following proxies for contracting institutions, which are described in detail in Table 6.B.3 and defined such that a higher value indicates better institutional quality: the index of contract enforcement between private parties by IHS Markit (column 1), the law and order component of the International Country Risk Guide by Political Risk Services (PRS, column 2), the inverse distance to the frontier in enforcing contracts from the World Bank's Doing Business (WBDB) database (column 3), the index of legal formalism developed by Djankov et al. (2003, column 4), the index of property rights freedom by the Heritage foundation (column 5), and the enforceability of contracts measure by Business Environmental Risk Intelligence (BERI, column 6). As reported in Panel A, the conditional correlation with the ownership share is estimated to be positive for each of the six proxies for contracting institutions, and it is highly significant in four cases. An even clearer picture emerges from Panel B. The estimated interaction effects of our alternative proxies for contracting institutions with relationship-specificity are positive and significant at the five percent level in all cases (and at the one percent level in all but one case). This allows us to conclude that our main estimation results are robust to using a wide range of alternative measures of the quality of contracting institutions.

In Panel C of Table 6.3, we examine alternative measures of relationship-specificity. The first three columns use variants of the Rauch (1999) classification. In column 1, we reclassify referenced-priced goods as non-differentiated (instead of differentiated), but adhere to the liberal classification. Next, we use the conservative (rather than liberal) variant of the classification, alternatively denominating referenced-priced goods as differentiated (column 2) or non-differentiated (column 3). Again, all three regressions confirm the significantly positive interaction effect. In columns 4 to 6, we use a proxy for relationship-specificity that is specific to the firm pair. We compute the number of years in which a given firm pair is consecutively observed in our biannual data for 2004-2014, normalized by the maximum number of observed years, and call this measure the 'relationship duration'. This approach builds on the recent theoretical work by Martin et al. (2016), which suggests that the duration of firm relationships may serve as a revealed measure of their specificity. Intuitively, long-term relationships are more likely to involve relationship-specific investments that have a lower value on the outside market. We compute the relationship duration for all firm pairs in our data and can therefore estimate our preferred specification with the alternative measure in a much larger sample (see column 4). The estimated interaction effect is positive and significant, as predicted by Proposition 6.2. Since the relationship duration is specific to the firm pair, our setup further allows us to control for firm-specific effects. To this end, we gradually add firm FE for the HQ (column 5) and the subsidiary (column 6) to our specification. While this narrows down our sample to firms with multiple ownership links, it allows us to control for unobservable firm characteristics that may affect the integration decision, such as the managers' time preferences (see Kukharskyy, 2016). We find a highly significant interaction effect within HQ, which remains significant at the ten percent level even when identified from variation within headquarters *and* within subsidiaries. These estimates strongly corroborate our main findings.

Table 6.3: *Alternative measurement of key variables*

Dep. var.: ownership share	IHS Markit	PRS	WBDB	Djankov	Heritage	BERI
	(1)	(2)	(3)	(4)	(5)	(6)
A. First-order effect for alternative measures of contracting institutions						
Contracting institutions	11.56***	4.836	1.270	3.528**	10.45***	8.080***
	(3.597)	(3.119)	(1.169)	(1.707)	(2.063)	(2.056)
Observations	433,108	433,108	433,108	430,271	433,108	426,996
R^2	0.276	0.275	0.275	0.279	0.278	0.282
B. Interaction effect for alternative measures of contracting institutions						
Contracting institutions × specificity	4.314***	3.816***	2.344***	2.449***	3.427***	1.765**
	(1.523)	(0.896)	(0.862)	(0.759)	(0.925)	(0.866)
Observations	185,650	186,110	186,110	184,829	186,110	173,621
R^2	0.295	0.295	0.295	0.296	0.295	0.297

C. Interaction effect for alternative measures of relationship-specificity

	Rauch specificity measures			Relationship duration		
	diff.	conserv.	diff. & cons.	baseline	HQ firm FE	2 firm FE
Rule of law × specificity (alternative)	1.135*	1.667***	3.401***			
	(0.598)	(0.620)	(1.292)			
Rule of law × relationship duration				1.388***	1.778***	2.887*
				(0.481)	(0.456)	(1.692)
Relationship duration				0.909	1.789***	2.552
				(0.571)	(0.516)	(2.100)
HQ firm FE	no	no	no	no	yes	yes
Subsidiary firm FE	no	no	no	no	no	yes
Observations	186,110	186,110	186,110	523,322	362,339	19,679
R^2	0.295	0.295	0.295	0.281	0.615	0.846

Panel A reports estimates of equation (6.18), including all control variables and FE from column 6 of Table 6.1. Panels B and C report estimates of equation (6.19), including all control variables and FE from column 6 of Table 6.2. Panels A and B use the alternative measures of the quality of contracting institutions listed in the header, which are described in Table 6.B.3. Panel C uses alternative measures of relationship-specificity described in the text. Standard errors clustered by subsidiary country and by HQ (Panel A) or by subsidiary country-industry and by HQ (Panels B and C) are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

Additional control variables

In this section, we return to our baseline measures and account for potentially confounding factors by including additional control variables in our preferred specifications. We start by controlling for a variety of differential effects of subsidiary country and industry characteristics, then explicitly account for firm heterogeneity, and finally control for remaining unobservables, including HQ firm FE.

Table 6.4 first addresses the possibility that country-specific variables may have differential effects across industries. Even after controlling for country FE, the interaction effect in equation (6.19) might be confounded by other characteristics of the subsidiary's country, such as economic development or other institutions, which are positively correlated with the quality of contracting institutions. If these country characteristics affect the firms' integration decisions and if they have a different effect in more specific industries, this may bias our estimates. Moreover, subsidiary country characteristics may affect the ownership decisions through channels other than relationship-specificity. To account for all of these

channels, we adopt a very general approach that controls for *arbitrary* effects of country-specific factors across industries. We begin by controlling for the differential effects of economic size and economic development by adding two full sets of interaction terms of subsidiary industry dummies with GDP and GDP per capita in the subsidiary's country to our main specification of equation (6.19). This approach was first developed by Levchenko (2007) for studying exports and adopted by Antràs (2015) in a context similar to our paper. Column 1 of Table 6.4 shows that our main interaction effect is fully robust to this important robustness check. In columns 2-3, we proceed analogously by controlling for interaction terms of subsidiary industry dummies with proxies for endowments (capital-labor ratio and human capital) and the quality of other types of institutions (financial development, labor market flexibility, IPR protection, and expropriation risk) in the subsidiary's country, respectively.[41] We find that these tests do not alter our previous conclusions, as the estimated main interaction effect maintains a similar magnitude as in the baseline regression and is always significant at conventional levels.

Table 6.4: *Controlling for differential effects of subsidiary country and industry characteristics*

Dep. var.: ownership share	GDP	Endowments	Institutions	Industry K/L	Combination 1	Combination 2
	(1)	(2)	(3)	(4)	(5)	(6)
Rule of law × specificity	3.318**	3.907***	4.499**	3.402***	3.073*	4.641***
	(1.472)	(1.251)	(1.867)	(0.909)	(1.815)	(1.664)
Observations	186,110	186,110	152,642	186,110	157,166	156,707
R^2	0.298	0.298	0.309	0.296	0.307	0.307

The table reports estimates of equation (6.19). All regressions include the control variables and fixed effects from column 6 of Table 6.2. In addition, we control for interactions of a full set of subsidiary industry dummies with the following characteristics of the subsidiary country: GDP and GDP per capita in column 1, endowments (capital-labor ratio and human capital) in column 2, and other institutions (financial development, labor market flexibility, IPR protection, and expropriation risk) in column 3. Column 4 includes interactions of a full set of subsidiary country dummies with the subsidiary industry's capital intensity. In column 5, we control for interactions of industry dummies with GDP per capita, the capital-labor ratio, and financial development, as well as interactions of industry capital intensity with country dummies. In column 6, we control for interactions of industry dummies with human capital, financial development, and expropriation risk, as well as interactions of industry capital intensity with country dummies. Standard errors clustered by subsidiary country-industry and by HQ are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

It is also conceivable that technological features of the subsidiary's industry have varying effects on ownership shares across country characteristics other than the ones considered in columns 1-3. To the extent that these technological features are reflected in the capital intensity of the subsidiary's industry, we can control for them very generally through interaction terms of capital intensity with subsidiary country dummies. Our main finding is confirmed after adding these control variables in column 4 of Table 6.4. We can apply an even more stringent tests by combining several of the aforementioned sets of interaction terms in a single regression. Note that including all of the above interaction effects makes inference impossible due to the large number of covariates. We therefore report the results of regressions including two different combinations of interaction terms (though other combinations yield similar results). The specification reported in column 5 includes four sets of variables that seem a priori most important: interaction terms of industry dummies with GDP per capita, the capital-labor ratio, and financial development, as well as interaction terms of country dummies with industry-level capital intensity. In the last column, we continue to control for arbitrary effects of industry capital intensity across subsidiary countries, but add interactions of industry dummies with the set of country characteristics that are individually

[41] These variables are defined in Section 6.3.2.

significant in Table 6.1: human capital, financial development, and expropriation risk. We continue to find a significantly positive interaction effect between contracting institutions and relationship-specificity in these highly demanding specifications and conclude that differential effects of other relevant country and industry characteristics cannot explain our main findings.

Table 6.5 addresses potentially remaining concerns about omitted variables related to the characteristics of the individual firms. While we have abstracted from firm heterogeneity in our theoretical model, differences across firms – both headquarters and subsidiaries – may potentially play a role for ownership decisions. For instance, one might suspect that particularly large and productive subsidiary firms are more lucrative investment targets, therefore attracting higher ownership shares; alternatively, one might argue that large and productive firms are more likely to be listed on the stock exchange and thus characterized by widespread shareholdings. In either case, if firms producing relationship-specific goods can grow larger on average (e.g. due to market power), and if these firms tend to locate in countries with better contracting institutions (e.g. due to better infrastructure), then neglecting firm heterogeneity might bias the estimate of our main interaction effect. One could construct similar narratives for other dimensions of firm heterogeneity. For this reason, we control for the following observable characteristics of the subsidiary firm, which may be relevant for ownership shares: firm size (measured by $\ln$ *employment*), labor productivity (defined as $\ln$(*value added*/*employment*)), the firm's age, its capital intensity (defined as $\ln$(*capital*/*employment*)), and a shareholder dummy, indicating whether the subsidiary itself holds any shares in other firms. These variables are lagged by one year (based on unconsolidated financial accounts in Orbis in 2013), which ameliorates potential concerns regarding reverse causality. The estimation results reported in columns 1 and 2 of Table 6.5 strengthen our main findings, as the interaction effect is estimated to be even larger than in our main regression. They further reveal that ownership shares are in fact higher for larger, more productive, and older subsidiaries, while they are lower for subsidiaries that are more capital intensive or are shareholders themselves.

In the next step, we control for firm heterogeneity among HQ. We include the same lagged firm characteristics as in the case of subsidiaries, except for the shareholder dummy, which is replaced by a subsidiary dummy, indicating whether the HQ itself is owned (to some degree) by other firms in our data. Columns 3 and 4 of Table 6.5 confirm the positive estimates for the rule of law index and its interaction effect with relationship-specificity, both of which are highly significant in this further reduced sample. Larger, more productive, and capital intensive HQ tend to own higher ownership shares; age enters negatively; and evidence on the subsidiary dummy is inconclusive. In the last two columns of Table 6.5, we include the control variables for both firms. We find that both the direct and the interaction effect have the predicted signs, but only the former is statistically significant. The insignificance of the interaction effect is likely to be explained by the reduced sample, which is smaller by an order of magnitude compared to our main estimation. The remaining sample covers only 21 subsidiary countries, predominantly EU members, which are characterized by similar contracting institutions, and hence little variation in the rule of law index. Thus, the lack of significance in this sample is not surprising. Given that controlling for firm variables drastically reduces our sample, we omit them in our main analysis. Most importantly, the positive point estimates for our interaction effect in these robustness checks suggest that firm heterogeneity does not cause an upward bias in our main regression.

In our main analysis, we have abstracted from political barriers to FDI. However, many countries maintain policy restrictions on foreign equity ownership. Such restrictions on foreign equity are measured by the OECD's FDI Regulatory Restrictiveness Index across 62 countries and 22 industries in

Table 6.5: *Firm heterogeneity*

Dep. var.: ownership share	Subsidiary firm controls		Headquarters firm controls		Both firm controls	
	(1)	(2)	(3)	(4)	(5)	(6)
Rule of law	5.348**		6.406***		4.307**	
	(1.861)		(1.439)		(1.437)	
Rule of law × specificity		4.309**		7.446***		2.650
		(1.774)		(2.538)		(3.455)
Employment (subsidiary)	1.972***	2.650***			-0.0281	0.697***
	(0.505)	(0.165)			(0.516)	(0.263)
Labor productivity (subsidiary)	0.257	0.922***			-1.376*	-0.546
	(0.451)	(0.235)			(0.710)	(0.416)
Age (subsidiary)	0.0291**	0.0184**			0.0540***	0.0353**
	(0.0119)	(0.00887)			(0.0117)	(0.0179)
Capital intensity (subsidiary)	-0.631***	-0.355***			-1.292***	-0.856***
	(0.191)	(0.101)			(0.261)	(0.162)
Shareholder dummy (subsidiary)	-3.157***	-3.411***			-4.068***	-4.365***
	(0.383)	(0.352)			(0.456)	(0.603)
Employment (headquarters)			2.320***	2.031***	2.759***	2.373***
			(0.401)	(0.210)	(0.377)	(0.221)
Labor productivity (headquarters)			1.672***	1.649***	2.253***	2.273***
			(0.237)	(0.242)	(0.224)	(0.304)
Age (headquarters)			-0.0753***	-0.0778***	-0.0694***	-0.0682***
			(0.00791)	(0.0165)	(0.00699)	(0.0203)
Capital intensity (headquarters)			0.838***	0.677***	1.226***	0.987***
			(0.220)	(0.119)	(0.141)	(0.147)
Subsidiary dummy (headquarters)			1.505	2.156***	0.880	1.852*
			(1.389)	(0.810)	(1.509)	(1.085)
Observations	106,829	53,353	84,520	36,153	42,757	19,325
R^2	0.229	0.263	0.241	0.269	0.259	0.295

Columns 1, 3, and 5 report estimates of equation (6.18), including all the control variables and FE from column 6 of Table 6.1. Columns 2, 4, and 6 report estimates of equation (6.19), including all the control variables and FE from column 6 of Table 6.2. In addition, we control for one-year lags of the listed firm-level control variables for the subsidiary firm (columns 1-2), for the HQ (columns 3-4), and for both firms (columns 5-6). Standard errors clustered by subsidiary country and by HQ (columns 1, 3, and 5) or by subsidiary country-industry and by HQ (columns 2, 4, and 6) are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

2014.[42] We modify our main estimation equations to include this index for the subsidiary's country plus an interaction term of the index with a dummy variable indicating foreign ownership relationships, since FDI restrictions are expected to have a differential effect on international investments. Unreported estimation results reveal that this robustness check leaves our main estimates for the effect of rule of law and its interaction with relationship-specificity unaffected in terms of both economic size and statistical significance.

Table 6.6 controls for remaining sources of unobserved heterogeneity in the data. While we have already addressed primary concerns regarding omitted variables, one can still envision more intricate narratives of potentially confounding factors. For instance, institutions in the HQ country may have a differential effect on ownership shares across industries (in addition to the direct effects that we have controlled for throughout). Similarly, the effects of other characteristics of the HQ country might vary across subsidiary industries. The most general way to address these issues is by augmenting our baseline specifications with an additional set of FE for each combination of HQ countries and subsidiary

[42]The data are obtained from http://www.oecd.org/investment/fdiindex.htm. We focus on foreign equity restrictions, since we find that other types of restrictions captured by the index (e.g. on screening, approval, and foreign personnel) do not affect ownership shares.

industries. The results of estimating these augmented specifications, reported in the first two columns of Table 6.6, support our main predictions.

So far, we have controlled for the (technological) fundamentals of both parties' industries via HQ and subsidiary FE. However, it is conceivable that industry-pair specific factors may also affect ownership shares. For instance, Antràs and Chor (2013) show that the integration decision is affected by the interaction of 'downstreamness' of the subsidiary's industry with the demand elasticity for final goods, which may be interpreted as an industry characteristic of the HQ in case the subsidiary is a vertically integrated supplier. To control for these (and other) unobservables, columns 3 and 4 of Table 6.6 add industry-pair FE to the previous specifications. Again, the size and significance of both key estimates hardly change, which is why these computationally intensive FE are omitted in the main analysis.

Table 6.6: *Additional unobservable effects and within-firm estimates*

Dep. var.: ownership share	HQ cty-subsid. ind. FE		+ Industry-pair FE		+ HQ firm FE	
	(1)	(2)	(3)	(4)	(5)	(6)
Rule of law	10.82***		10.39***		7.227***	
	(2.674)		(2.651)		(2.073)	
Rule of law × specificity		3.980***		3.619***		3.603*
		(1.273)		(1.319)		(2.072)
HQ country-subsidiary industry FE	yes	yes	yes	yes	yes	yes
Industry-pair FE	no	no	yes	yes	yes	yes
HQ firm FE	no	no	no	no	yes	yes
Observations	431,348	185,125	420,630	180,361	283,334	102,055
R^2	0.299	0.321	0.337	0.356	0.656	0.679

Columns 1, 3, and 5 report estimates of equation (6.18), including all the control variables and FE from column 6 of Table 6.1. Columns 2, 4, and 6 report estimates of equation (6.19), including all the control variables and FE from column 6 of Table 6.2 plus the indicated FE. Standard errors clustered by subsidiary country and by HQ (columns 1, 3, and 5) or by subsidiary country-industry and by HQ (columns 2, 4, and 6) are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

A significant advantage of our data over those used in previous studies is that we can separately identify both firms that form an ownership link – the HQ and the subsidiary. To fully exploit this advantage, we proceed by identifying the effect of contracting institutions from variation across different subsidiary countries and industries *within* the same HQ. For this purpose, we add HQ firm FE to the previous specifications (including also all previously introduced FE). This approach implicitly restricts the sample to HQ that hold ownership shares in at least two subsidiaries in different countries (or industries for the interaction effect). Column 5 of Table 6.6 shows that, within a given firm, the HQ chooses higher ownership shares in those subsidiaries that are located in countries with better contracting institutions. The within-firm estimate for the effect of rule of law is highly significant and of the same order of magnitude as in the baseline estimation. Also, as reported in column 6, the estimated interaction effect of rule of law with relationship-specificity has a similar magnitude as before, though it is estimated with less precision within HQ and only significant at the 10% level. Overall, the evidence from these highly demanding robustness checks lends further support to our main predictions.

Subsamples and panel

In Table 6.7, we explore the robustness of our preferred specifications in alternative sample configurations. Panel A shows estimates of equation (6.18), including all control variables and FE from column 6

Table 6.7: *Subsamples and panel*

Dep. var.: ownership share	Subsamples				Panel	
	S≥15%	FDI	Vertical	High-income	Baseline	+ FE
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. First-order effect						
Rule of law	11.16***	7.089**	11.01***	6.829***	7.585***	7.130***
	(2.574)	(2.996)	(2.470)	(1.333)	(0.577)	(0.541)
Subsidiary industry-year FE					yes	yes
HQ industry-year & country-year FE					yes	nested
HQ industry-country-year FE					no	yes
Observations	419,766	65,220	319,236	367,874	1,917,452	1,910,195
R^2	0.294	0.192	0.269	0.307	0.224	0.259
Panel B. Interaction effect						
Rule of law × specificity	2.402***	4.505**	3.608***	3.590***	2.727**	2.297**
	(0.823)	(1.796)	(1.065)	(1.248)	(1.081)	(0.912)
Subsidiary industry-year & country-year FE					yes	yes
HQ industry-year & country-year FE					yes	nested
Country-pair FE					yes	nested
HQ industry-country-year FE					no	yes
Country-pair-year FE					no	yes
Observations	180,753	34,752	136,234	145,401	788,813	778,994
R^2	0.307	0.318	0.297	0.320	0.236	0.289

Panel A reports estimates of equation (6.18), including all the control variables and FE from column 6 of Table 6.1. Panel B reports estimates of equation (6.19), including all the control variables and FE from column 6 of Table 6.2. Column 1 includes only ownership shares of at least 15%. Column 2 restricts the sample to international ownership links (FDI). Column 3 restricts the sample to HQ and subsidiaries in different industries (vertical). Column 4 restricts the sample to high-income subsidiary countries with above-median GDP per capita. Columns 5 and 6 use the biannual panel from 2004 to 2014, controlling for the indicated time-varying FE in addition to observables. Standard errors clustered by subsidiary country and by HQ (Panel A) or by subsidiary country-industry and by HQ (Panel B) are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

of Table 6.1, while Panel B considers the preferred specification of equation (6.19). In our main estimation sample, we have restricted ownership shares to a minimum of 10% to exclude small investments, which may be driven by portfolio considerations rather than lasting business interests. When increasing this threshold to 15% in column 1, we continue to find positive estimates for rule of law and the interaction effect. As noted in the introduction, a substantial share of the literature concerning the effects of contracting institutions on firm boundaries has thus far concentrated on international investments and on vertical buyer-seller relationships (see e.g. Antràs and Chor, 2013; Antràs, 2015). While our theory and empirical analysis are more general, we now show that our results are relevant to this literature, as they continue to hold even if we exclude ownership links in the domestic country or in the same industry. In the sample of foreign direct investments (FDI) in column 2, we find a slightly lower estimate for the rule of law index in the first specification, but a larger interaction effect. Both estimates are significant at the five percent level, although the sample size shrinks by a factor of five to seven. Our baseline estimates hardly change when we restrict the sample to subsidiaries active in a different four-digit NAICS industry

from their owner, which we call 'vertical' relationships (column 3).[43] One might suspect that the quality of contracting institutions varies mainly between developed and developing countries, but less among OECD countries, which make up the bulk of observations in the Orbis database. To verify that our estimates are not driven by particular features of developing countries, we restrict the sample to subsidiaries located in high-income countries in column 4.[44] This approach has the additional advantage of excluding subsidiaries located in several small island states, including so-called 'offshore financial centers', where firms invest mainly for tax reasons. In this subsample, there remains substantial variation in the rule of law index, which has a standard deviation of 0.60 (compared to 0.86 in the main estimation sample). The estimated coefficients of the rule of law index and the interaction effect are positive and significant also in the high-income sample.

In the last two columns of Table 6.7, we turn to the full panel of ownership shares that we observe biannually for 2004-2014. By pooling these data, we substantially increase the sample and can additionally exploit the time variation in ownership shares for identification. Note that we control for large sets of time-varying FE in addition to all the observable control variables from Table 6.1 in column 5, and we further extend the set of FE in the last column. These panel regressions confirm our positive and significant estimates for the rule of law index and the interaction effect with relationship-specificity.

Selection into countries

In our main analysis, we have taken the location of the subsidiary as given and focused on the HQ's choice of the integration intensity. As predicted by our theoretical model, we find that differences in contracting institutions across countries shape the intensity of integration. However, a HQ's location choice, i.e., the selection of the production country, is also likely to be driven by contracting institutions and other country characteristics. Under certain conditions, this location choice can affect our analysis of the intensive margin of integration. In particular, one may envision that in practice, the HQ solves a two-stage decision problem, choosing first whether or not to produce in a given country, and in the second stage, deciding on the depth of integration (the optimal ownership share). Depending on what explains the location choice, such a decision structure might introduce selection bias to our estimations. Note that the direction of this bias is a priori unclear, as it depends on how the variables that drive selection in the first stage are correlated with the ownership shares and our key explanatory variables.

To address this issue, we estimate two-stage models that apply the selection correction proposed by Heckman (1979). The first-stage selection equation explains a dummy $O_{h\ell}$, which indicates whether or not we observe ownership shares (of at least 10%) of HQ h in any subsidiary in country ℓ, by the following probit regression:

$$\Pr(O_{h\ell} = 1|\mathbf{V}_{h\ell}) = \Phi\left(\boldsymbol{\nu} \cdot \mathbf{V}_{h\ell}\right), \tag{6.20}$$

whereby $\Pr$ denotes probability and $\Phi(\cdot)$ is the standard normal distribution function. The vector $\mathbf{V}_{h\ell}$ (with associated coefficient vector $\boldsymbol{\nu}$) includes all the country-specific and country-pair specific variables contained in $\mathbf{X}_{hm}$ from equation (6.18) (see column 6 of Table 6.1) as well as HQ country and industry FE. In addition, following the approach by Helpman et al. (2008), we include in $\mathbf{V}_{h\ell}$ a 'religious distance'

[43] This definition reflects the notion that subsidiaries active in a different industry from their parent are less likely to replicate the activity of the HQ, but instead the two firms find themselves at different positions along the (vertical) value chain. The same definition has been used for instance by Alfaro and Charlton (2009) and Fajgelbaum et al. (2015). As noted in footnote 10, our theoretical argument does not presuppose the existence of supply-use relationships between the two firms.

[44] As a threshold, we choose the median GDP per capita across all subsidiary countries from the main estimation sample. This is a very high threshold, as the country just below the median is Spain.

variable, which captures the dissimilarity in the religious beliefs across country pairs.[45] This approach reflects the idea that similar religious beliefs may induce people to engage in economic activity and invest in the other country, while we have no reason to believe that they also affect the intensity of integration. Since the religious distance variable is excluded in the second-stage models, it contributes to identification. Given that we do not observe the HQ's business partners in countries for which the dummy $O_{h\ell}$ is equal to zero, we add one observation with $O_{h\ell} = 0$ for each country not selected by a given HQ. Since this procedure inflates the dataset with zeros, estimating equation (6.20) for the entire sample reaches computational limits, so we estimate it instead by HQ country or group of countries.[46] From the probabilities predicted by equation (6.20), we compute the inverse Mills ratio (IMR), the so-called non-selection hazard. The IMR is then included in the second-stage models, given by equations (6.18) and (6.19), to correct for potential selection bias.

Table 6.8: *Selection into countries*

Dep. var.: ownership share	Baseline		Heckman correction	
	(1)	(2)	(3)	(4)
Rule of law	11.12***		10.80***	
	(2.540)		(2.528)	
Rule of law × specificity		3.432***		3.769***
		(0.909)		(1.062)
IMR			-1.965*	1.654**
			(1.150)	(0.805)
Observations	433,108	186,110	432,954	152,567
R^2	0.278	0.295	0.278	0.303

Columns 1 and 3 report estimates of equation (6.18), including all the control variables and FE from column 6 of Table 6.1. Columns 2 and 4 report estimates of equation (6.19), including all the control variables and FE from column 6 of Table 6.2. Columns 1 and 2 repeat the baseline estimates. In columns 3 and 4, we correct for selection by including the inverse Mills ratio (IMR), predicted by probit regressions of equation (6.20), estimated separately for 29 HQ countries or country groups. Standard errors clustered by subsidiary country and by HQ (columns 1 and 3) or by subsidiary country-industry and by HQ (columns 2 and 4) are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

The estimation results from the two-stage Heckman models provide some evidence for the relevance of selection and strengthen our previous findings. The unreported first-stage probit regressions reveal for most of the HQ countries (or groups of countries) that religious distance tends to decrease the probability of an ownership link, in line with expectations. Furthermore, we find in most cases that subsidiaries are more likely to be observed in countries with a higher rule of law index. The second-stage regressions are reported in Table 6.8, which repeats for convenience the baseline results (from the last columns of Tables 6.1 and 6.2) in the first two columns. The added IMR in columns 3 and 4 turns out to be significant, indicating a selection bias in the absence of a correction. More precisely, the coefficient of the IMR is weakly negative in column 3, and it is positive and significant in the preferred model of column 4, which includes also country-pair FE. Most importantly, the estimates for the effects of rule of law and its interaction with relationship-specificity are positive and highly significant in these Heckman regressions. Moreover, the interaction effect is larger than the baseline estimate, indicating a downward bias due to selection. Hence, our main estimate without the selection correction is conservative in the sense

[45] Our 'religious distance' variable is taken from Spolaore and Wacziarg (2016) and represents a population-weighted measure of the similarity of religions based on a categorization by the World Christian Database (chosen to maximize country coverage in the Orbis data). We set religious distance to zero for domestic pairs.

[46] Depending on the number of observations per country in Orbis, we estimate equation (6.20) separately for 29 individual HQ countries or groups of countries, defined by the world regions indicated in Table 6.B.1.

that it tends to underestimate the positive interaction effect of contracting institutions and relationship-specificity on ownership shares.

Ordered logistic regression for ownership categories

There are two potential concerns related to the use of linear estimation methods in our baseline analysis. First, the linear model might not be ideally suited to explain the ownership share S_{hm}, which lies between zero and one, because it does not respect the variable's natural upper and lower bounds. As a symptom of this issue, OLS estimation may yield predicted values outside the unit interval, which is the case for a small share of observations in our preferred specification of equation (6.19). Second, OLS estimation does not allow for non-linear effects of the explanatory variables on ownership shares, although such non-linearities may be relevant in practice. For instance, there may be critical threshold values of the ownership share above which the owner obtains additional control rights. In terms of our model, one might envisage that full residual control rights over non-contractible inputs lie with the party that owns more than 50% of the firm. Similarly, the HQ may face additional obstacles when trying to exercise her residual control rights unless she owns 100% of the subsidiary. It is therefore conceivable that a firm's choice *between* ownership categories, such as majority and full ownership, is more sensitive to judicial quality compared to the choice of ownership shares *within* an ownership category (e.g. between 63% and 64%), implying non-linearities.

To address both of these concerns, we estimate an ordered logistic (henceforth, ordered logit) regression model using maximum likelihood methods. The dependent variable in this model is the categorical variable $\tilde{S}$, which takes on three distinct values for categories of ownership shares: 1 for minority (10-49.99%), 2 for majority (50-99.99%), and 3 for full ownership (100%). The ordered logit model, which describes the probabilities that the ownership share S_{hm} lies in each category, is specified as follows:

$$\Pr(\tilde{S} = 1|\mathbf{Z}_{hm}) = \Lambda(\chi_1 - \boldsymbol{\xi}\mathbf{Z}_{hm}) \tag{6.21}$$

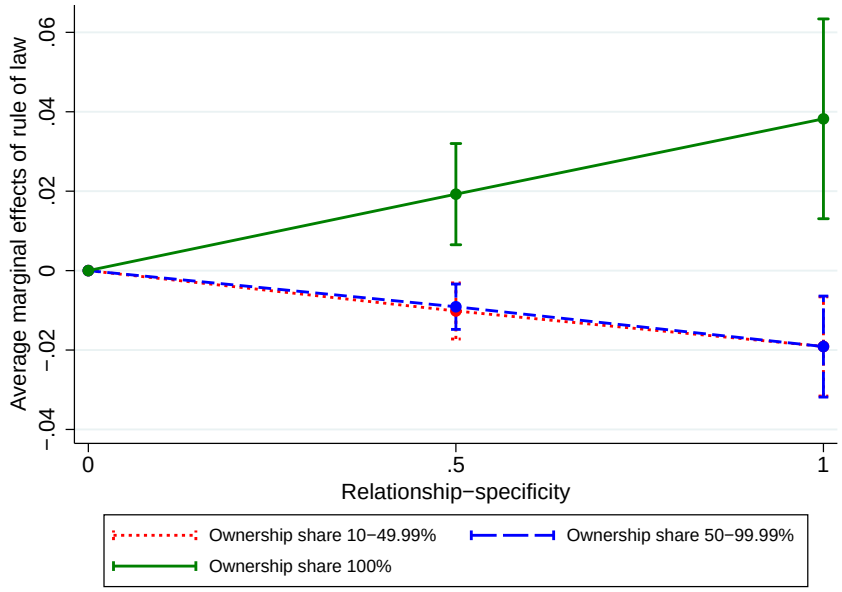
$$\Pr(\tilde{S} = 2|\mathbf{Z}_{hm}) = \Lambda(\chi_2 - \boldsymbol{\xi}\mathbf{Z}_{hm}) - \Lambda(\chi_1 - \boldsymbol{\xi}\mathbf{Z}_{hm}) \tag{6.22}$$

$$\Pr(\tilde{S} = 3|\mathbf{Z}_{hm}) = 1 - \Lambda(\chi_2 - \boldsymbol{\xi}\mathbf{Z}_{hm}) \tag{6.23}$$

where $\Lambda(\cdot)$ is the logistic function, $\mathbf{Z}_{hm}$ is the vector of all explanatory variables from the preferred specification of equation (6.19), and χ_1, χ_2, as well as $\boldsymbol{\xi}$ are the parameters to be estimated. This modeling approach allows us to take into account the natural upper and lower bounds of ownership shares and to address potential non-linearities. On the downside, the approach ignores the variation in ownership shares within each category, which amounts to assuming that institutions play a negligible role for explaining the choice of ownership within categories. Therefore, the ordered logit model constitutes an important robustness check, but it is not strictly preferable to our baseline approach. When estimating the model, we further need to apply two changes compared to our baseline estimation due to technical limitations. First, to keep the model computationally feasible, we cannot account for country-industry FE for the HQ, but instead include FE by HQ industry and by country pair. Second, standard errors are clustered only by the subsidiary's country-industry combination.

Figure 6.3 illustrates the ordered logit estimation results in terms of so-called 'average marginal effects' (AMEs) of rule of law on the probability that the ownership share lies in different categories, which are evaluated at three values of relationship-specificity (0, 0.5, and 1). These AMEs are obtained by computing the predicted effect of a marginal improvement in rule of law on the ownership share for each firm

Figure 6.3: *Average marginal effects from ordered logit model for ownership categories*



Note: The figure depicts estimated average marginal effects (AMEs) of rule of law by relationship-specificity from the ordered logit regression model specified in equations (6.21) to (6.23), alongside 95% confidence intervals. The solid line represents the AMEs of rule of law on the probability of full ownership ($S_{hm} = 1$), the dashed line represents the AMEs of rule of law on the probability of majority ownership ($S_{hm} \in [0.5, 0.9999]$), and the dotted line represents the AMEs of rule of law on the probability of minority ownership ($S_{hm} \in [0.1, 0.4999]$). Standard errors are clustered at the subsidiary country-industry level. The number of observations is 187,605 and the pseudo R^2 is 0.2215.

pair, at the observed levels of all other covariates, and averaging these marginal effects across all observations for a given category of relationship-specificity (using Stata's `margins` command). Consider first the solid line, representing the AMEs of rule of law on full ownership. The point estimates are significantly positive for intermediate and high relationship-specificity and increasing in the latter. By contrast, the AMEs of rule of law on minority and majority ownership (the dashed and dotted lines, respectively) are estimated to be negative and decreasing in relationship-specificity. Somewhat surprisingly, there is no noticeable difference in the AMEs between minority and majority ownership. Note, however, that these AMEs cannot be considered in isolation, but need to be interpreted relative to the other ownership categories. Taken together, the estimates suggest that better contracting institutions increase the propensity for HQ to choose full ownership *relative to* lower ownership shares, and this effect is stronger in industries with a higher relationship-specificity. Hence, the ordered logit estimates support our main hypotheses. They further add the insight that, empirically, better institutions particularly favor full ownership over joint ventures.

Instrumental variables

Since we regress micro-level ownership shares on aggregate variables, measured at the levels of industries and countries, reverse causality does not appear to be a relevant issue when estimating equations (6.18) and (6.19). We might, however, imagine that the government of a country which has attracted many large foreign investments (in relationship-specific industries) would have particularly strong incentives to improve the quality of domestic contracting institutions. While a large bulk of foreign investment need not be reflected in high average ownership shares at the firm level, we nevertheless address the pos-

sibility of reverse causality by using instrumental variables (IV). We adopt the standard approach of using the historic origin of a country's legal system as an IV for the rule of law index (see Nunn, 2007). For this purpose, we rely on the classification of legal origins developed by La Porta et al. (1998) and revised by La Porta et al. (2008) into British common law or civil law of French, German, or Scandinavian origin.[47] We choose British common law as the base category and use three indicator variables for the other categories. Since legal origins are pre-determined, they are exogenous to ownership structures and can therefore resolve a possible reverse causality issue. In addition, the IV approach also tackles other potential biases due to omitted variables, discussed Section 6.4.3, or due to measurement error in our proxy for contracting institutions.

Table 6.9: *Instrumental variables*

	First-order effect		Interaction effect	
	First stage	Second stage	First stage	Second stage
Dep. var. in the header:	Rule of law	Ownership share	Rule of law × specificity	Ownership share
	(1)	(2)	(3)	(4)
Rule of law		13.29**		
		(5.320)		
Rule of law × specificity				3.524**
				(1.437)
French legal origin dummy	-0.587***			
	(0.155)			
German legal origin dummy	-0.0123			
	(0.0895)			
Scandinavian legal origin dummy	-0.480*			
	(0.253)			
French legal origin × specificity			-1.126***	
			(0.0774)	
German legal origin × specificity			-0.547***	
			(0.0784)	
Scandinavian legal origin × specificity			0.125*	
			(0.0730)	
Observations	433,108	433,108	186,110	186,110
Partial R^2 (excluded IV)	0.331		0.429	
F-statistic (excluded IV, Kleibergen-Paap)	11.15		110.7	
P-value of F-test	0.0000		0.0000	

The table reports estimation results of 2SLS regressions. Column 1 reports the first-stage estimates and column 2 reports the second-stage estimates of equation (6.18), in which we instrument rule of law by a set of legal origin dummies, including all the control variables and FE from column 6 of Table 6.1. Column 3 reports the first-stage estimates and column 4 reports the second-stage estimates of equation (6.19), where we instrument the interaction of rule of law × specificity by interactions of legal origin dummies with specificity, including all the control variables and FE from column 6 of Table 6.2. Standard errors clustered by subsidiary country and by HQ (columns 1-2) or by subsidiary country-industry and by HQ (columns 3-4) are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

Table 6.9 reports the results of two-stages least squares (2SLS) estimations of our preferred specifications of equations (6.18) and (6.19). The first column reports the first-stage estimation results of regressing the rule of law index on the legal origin dummies for the subsidiary's country. It shows that countries with a legal system of French and Scandinavian origin have a significantly lower rule of law index (conditional on the covariates), and that these differences explain a substantial share of the variation

[47] The original classification includes the Socialist tradition as a fifth category. La Porta et al. (2008) reclassify the Socialist countries by French or German civil law, from which their legal systems originated and to which many of them reverted after the break-up of the Soviet Union. We follow this revised approach.

in judicial quality, as evidenced by a high partial R^2 of 0.33. The F-test for significance of the excluded IV yields a Kleibergen-Paap F-statistic of 11.15, exceeding the Stock and Yogo (2002) critical value for a 10% maximal IV bias relative to OLS, so legal origin serves as a strong IV. The second-stage regression in column 2 yields a positive and significant estimate for the instrumented rule of law index, confirming our previous findings. In columns 3 and 4, we instrument for the interaction term between the rule of law index and relationship-specificity in equation (6.19) using interactions of legal origin dummies with relationship-specificity. The first-stage regression (column 3) reveals that these interaction terms are both individually and jointly significant, with a very high F-statistic of the excluded IV and a partial R^2 of 0.43. In the second-stage regression, summarized in column 4, we find a positive and significant interaction effect of a similar magnitude as in the OLS estimations, which supports Proposition 6.2.

Propensity score matching

The critical assumption for the validity of the IV approach to estimating equation (6.19) is that the historical origins of countries' legal systems have no *differential* effect (by relationship-specificity) on firm boundaries in 2014 other than through contracting institutions, conditional on all control variables. This exclusion restriction may be violated if legal origins are correlated with other cultural or institutional characteristics that also shape firm boundaries differentially across industries. Such a violation of the exclusion restriction might be the reason for the slightly larger estimate of the interaction effect in our 2SLS regression compared to OLS, which is in conflict with the upward bias of OLS that we would expect due to reverse causality.[48] To address a potential violation of the exclusion restriction, we continue to follow Nunn (2007) and implement Propensity Score Matching (PSM). The idea of PSM, which goes back to Rosenbaum and Rubin (1983, 1984), is to select observations from treatment and control groups that are similar based on observable characteristics, assuming that they are also similar in terms of unobservables.

In our application, we seek to compare similar firm pairs involving subsidiaries in countries with favorable and unfavorable contracting institutions. Therefore, we select all observations of subsidiaries located in countries whose legal system is of British origin ($L_{hm} = 1$), which has been shown to be most favorable for investors, and match them to the most comparable observation of a subsidiary located in a country with French legal origin ($L_{hm} = 0$) *in the same industry*. Comparability is determined by the propensity score, i.e., the predicted value of the indicator P_{hm}, as explained by the following probit regression:

$$P_{hm} = \Pr(L_{hm} = 1|\mathbf{W}_{hm}) = \Phi(\boldsymbol{\zeta} \cdot \mathbf{W}_{hm} + \upsilon_{hm}), \tag{6.24}$$

where we match observations on the variables summarized in the vector $\mathbf{W}_{hm}$ (with associated coefficients $\boldsymbol{\zeta}$), and υ_{hm} is an error term. In the baseline PSM approach, $\mathbf{W}_{hm}$ includes the following variables: GDP per capita of the subsidiary's country, capital intensity and intangible assets intensity of the HQ's industry, a dummy variable for domestic (vs. international) ownership links, and $\ln$ *employment* of the subsidiary firm.[49] Capital intensity and intangible assets intensity serve as proxies for the relative importance of the HQ's inputs in the production process, an important determinant of the severity of hold-up problems identified in the incomplete-contracts literature (see Antràs, 2015, and our model extension in

[48] An alternative, less problematic explanation is that the IV approach corrects for a downward bias in the OLS estimates due to measurement error or omitted variables.

[49] Capital intensity is defined as the logarithm of total capital over total employment, and intangible assets intensity is defined as the logarithm of total intangible assets over total fixed assets, both measured in the HQ's industry in 2013. The firm-level variables are defined in Section 6.4.3.

Section 6.2.4). To better control for firm heterogeneity, we then vary the set of matching variables $\mathbf{W}_{hm}$ by successively adding the following characteristics of the subsidiary firm: age, capital intensity, and the shareholder dummy. Based on the predicted propensity score $\widehat{P}_{hm}$ from equation (6.24), we match observations within a given subsidiary industry with their so-called 'nearest neighbor' (with replacement), i.e., the single observation with the most similar propensity score, while restricting observations to the common support.[50]

For the matched observations, we construct the ratio of ownership shares for the subsidiary in the British legal origin country (B) over the one located in the French legal origin country (F). The logarithm of this ratio is then regressed on our preferred measure of relationship-specificity R_i:

$$\ln\left(S_{hmB}/S_{hmF}\right) = \psi_1 + \psi_2 \cdot R_i + \xi_{hmBF}, \tag{6.25}$$

with coefficients ψ_1 and ψ_2, and an error term ξ_{hmBF}. Standard errors are clustered at the level of the industry i in which subsidiary m is active. Since the contracting institutions in British legal origin countries are more favorable for investors, Proposition 6.2 would predict higher ownership shares for subsidiaries in these countries producing more relationship-specific goods, which translates into an estimate $\widehat{\psi}_2 > 0$.

Table 6.10 reports our results from estimating equation (6.25). We start in column 1 by combining all possible observations in the same industry involving subsidiaries from a British and a French legal origin country, which results in almost 100 million pairs of observations. The regression reveals a positive estimate for the coefficient of relationship-specificity, confirming our intuition based on the PRT. However, ownership shares may differ between these pairs for a variety of reasons other than legal origins. Therefore, we restrict the analysis to matched firm-pair observations, which are similar in terms of the variables contained in $\mathbf{W}_{hm}$. For all variants of $\mathbf{W}_{hm}$, we find estimates $\widehat{\psi}_2$ which are positive and significant (at least at the ten percent level) in columns 2-5. These estimates, which lie in the range of 0.44-0.62, are smaller than in the unmatched sample. This finding is in line with the expected direction of a bias that would arise from reverse causality or from omitted variables positively correlated with contracting institutions. Overall, the PSM results lend further support to our main hypothesis that better contracting institutions increase the depth of integration between firms more strongly in relationship-specific industries.

[50] Formally, we choose for each observation involving a subsidiary with British legal origin the observation involving a subsidiary with French legal origin in the same industry i for which the absolute difference in propensity scores is smallest. This procedure is implemented by the Stata module `psmatch2` provided by Leuven and Sianesi (2015). A similar approach has been adopted by Ma et al. (2010) using firm-level data.

Table 6.10: *Propensity score matching*

Dep. var.: ln (S_{hmB}/S_{hmF})	Unmatched	Baseline PSM	+ Age	+ Capital int.	+ Shareholder
	(1)	(2)	(3)	(4)	(5)
Specificity (subsidiary industry)	0.907***	0.442*	0.619***	0.504***	0.451**
	(0.310)	(0.232)	(0.228)	(0.189)	(0.195)
Observations	94,959,461	9,578	9,571	7,078	7,070
R^2	0.002	0.001	0.001	0.001	0.001

The table reports estimates of equation (6.25). The first column considers the unmatched sample of all possible combinations of observations involving one subsidiary in a British and one in a French legal origin country. Columns 2-5 are restricted to the sample of (nearest neighbor) pairs matched based on the propensity score predicted by variants of equation (6.24). In column 2, observations are matched via the following variables: GDP per capita of the subsidiary's country, capital intensity and intangible assets intensity of the HQ's industry, a dummy variable for domestic ownership links, and ln *employment* of the subsidiary firm. Columns 3-5 successively add to equation (6.24) the subsidiary firm variables listed in the header. Standard errors clustered by subsidiary industry are reported in parentheses. Asterisks indicate significance levels: * p<0.10, ** p<0.05, *** p<0.01.

6.5 Concluding Remarks

The fundamental role of contractual imperfections in shaping firm boundaries is widely accepted in the economic discipline. However, there is no consensus on whether reducing these imperfections eventually leads to more deeply integrated firms or a stronger reliance on markets. We contribute to this debate by developing a generalized Property-Rights Theory of the firm, which suggests that better contracting institutions increase the willingness of headquarters to obtain a larger ownership share in their subsidiaries, and that this effect is particularly pronounced in industries with a high degree of relationship-specificity. Using unique micro data on global ownership links across firm pairs, we find strong empirical support for these predictions. Our findings are confirmed for a variety of different measures of contracting institutions and relationship-specificity. They are also robust to controlling for a host of unobservable factors and industry-specific effects of economic development or other institutions. Finally, we corroborate our results by using legal origins as an exogenous source of institutional quality in instrumental variables and propensity score matching techniques.

What are the policy implications of our findings? Policymakers in developing countries may hope to attract foreign direct investment by improving the quality of domestic contracting institutions. Perhaps surprisingly, the Transaction-Cost Theory would suggest that such improvements discourage (rather than encourage) foreign ownership, since they facilitate market-based transactions and thus undermine the incentive for FDI. This paper has demonstrated that the Property-Rights Theory confirms the policymakers' intuition: Better contracting institutions should induce investors to choose higher degrees of integration. This intuition is strongly supported by our extensive empirical analysis of global firm pairs. Furthermore, we show that an improvement of contracting institutions has a particularly strong effect on the integration intensity in industries with a high degree of relationship-specificity. Since relationship-specific industries are typically also characterized by high technology and information content, improving judicial quality may entail further favorable outcomes through spillovers from FDI.

Bibliography

Acemoglu, Daron, Pol Antràs, and Elhanan Helpman, "Contracts and Technology Adoption," *American Economic Review*, 2007, *97* (3), 916–943.

_ , Rachel Griffith, Philippe Aghion, and Fabrizio Zilibotti, "Vertical Integration and Technology: Theory and Evidence," *Journal of the European Economic Association*, 2010, *8* (5), 989–1033.

_ , Simon Johnson, and Todd Mitton, "Determinants of Vertical Integration: Financial Development and Contracting Costs," *Journal of Finance*, 06 2009, *64* (3), 1251–1290.

Ai, Chunrong and Edward C. Norton, "Interaction Terms in Logit and Probit Models," *Economics Letters*, 2003, *80* (1), 123–129.

Alfaro, Laura and Andrew Charlton, "Intra-industry Foreign Direct Investment," *American Economic Review*, December 2009, *99* (5), 2096–2119.

_ , Pol Antràs, Davin Chor, and Paola Conconi, "Internalizing Global Value Chains: A Firm-Level Analysis," Technical Report 21582, National Bureau of Economic Research 2015.

Altomonte, Carlo and Armando Rungi, "Business Groups as Hierarchies of Firms: Determinants of Vertical Integration and Performance," Working Paper Series 1554, European Central Bank 2013.

Antràs, P., *Global Production: Firms, Contracts, and Trade Structure*, Princeton University Press, 2015.

Antràs, Pol, "Firms, Contracts, And Trade Structure," *The Quarterly Journal of Economics*, 2003, *118* (4), 1375–1418.

_ and Davin Chor, "Organizing the Global Value Chain," *Econometrica*, 2013, *81* (6), 2127–2204.

_ and Elhanan Helpman, "Global Sourcing," *Journal of Political Economy*, 2004, *112* (3), 552–580.

_ and _ , "Contractual Frictions and Global Sourcing," in Elhanan Helpman, Dalia Marin and Thierry Verdier, ed., *The Organization of Firms in a Global Economy*, Harvard University Press, 2008, pp. 55–83.

Atalay, Enghin, Ali Hortaçsu, and Chad Syverson, "Vertical Integration and Input Flows," *American Economic Review*, 2014, *104* (4), 1120–48.

Baker, George P. and Thomas N. Hubbard, "Make versus Buy in Trucking: Asset Ownership, Job Design, and Information," *American Economic Review*, 2003, *93* (3), 551–572.

_ and _ , "Contractibility and Asset Ownership: On-board Computers and Governance in U.S. Trucking," *Quarterly Journal of Economics*, 2004, *119*, 1443–1479.

Baker, Malcolm P. and Paul A. Gompers, "Executive Ownership and Control in Newly Public Firms: The Role of Venture Capitalists," November 1999. Harard Business School, mimeo.

Barba Navaretti, Giorgio and Anthony Venables, *Multinational Firms in the World Economy*, Princeton University Press, 2004.

Barro, Robert J and Jong Wha Lee, "International Measures of Schooling Years and Schooling Quality," *American Economic Review*, 1996, *86* (2), 218–23.

Botero, Juan C., Simeon Djankov, Rafael La Porta, Florencio Lopez de Silanes, and Andrei Shleifer, "The Regulation of Labor," *The Quarterly Journal of Economics*, 2004, *119* (4), 1339.

Broda, Christian and David E. Weinstein, "Globalization and the Gains from Variety," *The Quarterly Journal of Economics*, 2006, *121* (2), 541.

Cameron, A. Colin, Jonah B. Gelbach, and Douglas L. Miller, "Robust Inference With Multiway Clustering," *Journal of Business & Economic Statistics*, 2011, *29* (2), 238–249.

Carluccio, Juan and Thibault Fally, "Global Sourcing under Imperfect Capital Markets," *The Review of Economics and Statistics*, 2012, *94* (3), 740–763.

Chari, Anusha, Paige P. Ouimet, and Linda L. Tesar, "The Value of Control in Emerging Markets," *Review of Financial Studies*, 2010, *23* (4), 1741–1770.

Coase, Ronald H., "The Nature of the Firm," *Economica – New Series*, 1937, *4* (4), 386–405.

Corcos, Gregory, Delphine M. Irac, Giordano Mion, and Thierry Verdier, "The Determinants of Intrafirm Trade: Evidence from French Firms," *Review of Economics and Statistics*, 2013, *95* (3), 825–838.

Correia, Sergio, "REGHDFE: Stata Module to Perform Linear or Instrumental-variable Regression Absorbing any Number of High-dimensional Fixed Effects," Statistical Software Components, Boston College Department of Economics 2014.

Costinot, Arnaud, "On the Origins of Comparative Advantage," *Journal of International Economics*, 2009, *77* (2), 255–264.

Cravino, Javier and Andrei A. Levchenko, "Multinational Firms and International Business Cycle Transmission," *Quarterly Journal of Economics*, 2017.

Crozet, Matthieu and Pamina Koenig, "Structural Gravity Equations with Intensive and Extensive Margins," *Canadian Journal of Economics*, 2010, *43* (1), 41–62.

Defever, Fabrice and Farid Toubal, "Productivity, Relationship-specific Inputs and the Sourcing Modes of Multinationals," *Journal of Economic Behavior & Organization*, 2013, *94*, 345–357.

Djankov, Simeon, Rafael La Porta, Florencio Lopez de Silanes, and Andrei Shleifer, "Courts," *The Quarterly Journal of Economics*, 2003, *118* (2), 453–517.

Fajgelbaum, Pablo, Gene M Grossman, and Elhanan Helpman, "A Linder Hypothesis for Foreign Direct Investment," *The Review of Economic Studies*, 2015, *82* (1), 18–121.

Feenstra, Robert C. and Gordon H. Hanson, "Ownership and Control in Outsourcing to China: Estimating the Property-Rights Theory of the Firm," *The Quarterly Journal of Economics*, 2005, *120* (2), 729–761.

_ , **Robert Inklaar, and Marcel Timmer**, "The Next Generation of the Penn World Table," Technical Report 19255, National Bureau of Economic Research 2013.

Gibbons, Robert, "Four Formal(izable) Theories of the Firm?," *Journal of Economic Behavior & Organization*, 2005, *58*, 200–245.

Gorodnichenko, Yuriy, Bohdan Kukharskyy, and Gerard Roland, "Culture and Global Sourcing," 2017. University of Tübingen, mimeo.

Grossman, Gene M. and Elhanan Helpman, "Integration vs. Outsourcing in Industry Equilibrium," CESifo Working Paper Series 460, CESifo Group Munich 2001.

_ **and** _ , "Integration versus Outsourcing in Industry Equilibrium," *Quarterly Journal of Economics*, 2002, *117* (1), 85–120.

_ **and** _ , "Outsourcing versus FDI in Industry Equilibrium," *Journal of the European Economic Association*, 2003, *1* (2-3), 317–327.

_ **and** _ , "Outsourcing in a Global Economy," *Review of Economic Studies*, 2005, *72* (1), 135–159.

Grossman, Sanford J. and Oliver D. Hart, "The Costs and Benefits of Ownership: A Theory of Vertical and Lateral Integration," *Journal of Political Economy*, 1986, *94* (4), 691–719.

Hart, Oliver and John Moore, "Property Rights and the Nature of the Firm," *Journal of Political Economy*, 1990, *98* (6), 1119–1158.

Head, Keith, Thierry Mayer, and John Ries, "The Erosion of Colonial Trade Linkages after Independence," *Journal of International Economics*, 2010, *81* (1), 1–14.

Heckman, James J., "Sample Selection Bias as a Specification Error," *Econometrica*, 1979, *47* (1), 153–61.

Helpman, Elhanan, Marc Melitz, and Yona Rubinstein, "Estimating Trade Flows: Trading Partners and Trading Volumes," *The Quarterly Journal of Economics*, 2008, *123* (2), 441–487.

Javorcik, Beata, "The Composition of Foreign Direct Investment and Protection of Intellectual Property Rights: Evidence from Transition Economies," *European Economic Review*, 2004, *48* (1), 39–62.

Joskow, Paul L., "Vertical Integration," in Claude Menard and Mary M. Shirley, eds., *Handbook of New Institutional Economics*, Boston, MA: Springer US, 2005, pp. 319–348.

Kaufmann, Daniel, Aart Kraay, and Massimo Mastruzzi, "The Worldwide Governance Indicators: Methodology and Analytical Issues," 2010. Policy Research Working Paper 5430, World Bank.

Klein, Peter G., "The Make-or-Buy Decision: Lessons from Empirical Studies," in C. Ménard and M. Shirley, eds., *Handbook of New Institutional Economics*, Springer, 2015.

Knack, Stephen and Philip Keefer, "Institutions And Economic Performance: Cross-Country Tests Using Alternative Institutional Measures," *Economics and Politics*, 1995, *7* (3), 207–227.

Kohler, Wilhelm and Marcel Smolka, "Global Sourcing and Firm Selection," *Economics Letters*, 2014, *124* (3), 411–415.

_ **and** _ , "Global Sourcing of Heterogeneous Firms: Theory and Evidence," CESifo Working Paper Series 5184, CESifo Group Munich 2015.

Kukharskyy, Bohdan, "Relational Contracts and Global Sourcing," *Journal of International Economics*, 2016, *101* (C), 123–147.

La Porta, Rafael, Florencio Lopez de Silanes, and Andrei Shleifer, "The Economic Consequences of Legal Origins," *Journal of Economic Literature*, June 2008, *46* (2), 285–332.

_ **, Florencio Lopez de Silanes, Andrei Shleifer, and Robert W. Vishny**, "Law and Finance," *Journal of Political Economy*, 1998, *106* (6), 1113–1155.

Leuven, Edwin and Barbara Sianesi, "PSMATCH2: Stata Module to Perform Full Mahalanobis and Propensity Score Matching, Common Support Graphing, and Covariate Imbalance Testing," 2015.

Levchenko, Andrei A., "Institutional Quality and International Trade," *Review of Economic Studies*, 2007, *74* (3), 791–819.

Ma, Yue, Baozhi Qu, and Yifan Zhang, "Judicial quality, Contract Intensity and Trade: Firm-level Evidence from Developing and Transition Countries," *Journal of Comparative Economics*, 2010, *38* (2), 146–159.

Martin, Julien, Isabelle Méjean, and Mathieu Parenti, "Revealed Input Specificity: Insights from Firm-to-firm Export Data," 2016. Université Libre de Bruxelles, mimeo.

Masten, Scott E., "The Organization of Production: Evidence from the Aerospace Industry," *Journal of Law and Economics*, 1984, *27*, 403–417.

Nunn, Nathan, "Relationship-specificity, Incomplete Contracts, and the Pattern of Trade," *Quarterly Journal of Economics*, 2007, *122* (2), 569–600.

_ **and Daniel Trefler**, "The Boundaries of the Multinational Firm: An Empirical Analysis," in Elhanan Helpman, Dalia Marin, and Thierry Verdier, ed., *The Organization of Firms in a Global Economy*, Harvard University Press, 2008, pp. 55–83.

_ **and** _ , "Incomplete Contracts and the Boundaries of the Multinational Firm," *Journal of Economic Behavior & Organization*, 2013, *94*, 330–344.

_ **and** _ , "Domestic Institutions as a Source of Comparative Advantage," in Elhanan Helpman, Kenneth Rogoff, and Gita Gopinath, eds., *Handbook of International Economics*, Vol. 4, Elsevier, 2014, chapter 5.

Park, Walter G, "International Patent Protection: 1960–2005," *Research policy*, 2008, *37* (4), 761–766.

Ramondo, Natalia, Veronica Rappoport, and Kim J. Ruhl, "Intrafirm Trade and Vertical Fragmentation in U.S. Multinational Corporations," *Journal of International Economics*, 2016, *98* (C), 51–59.

Rauch, James E., "Networks versus Markets in International Trade," *Journal of International Economics*, June 1999, *48* (1), 7–35.

Rosenbaum, Paul R. and Donald B. Rubin, "The Central Role of the Propensity Score in Observational Studies for Causal Effects," *Biometrika*, 1983, *70* (1), 41–55.

_ **and** _ , "Reducing Bias in Observational Studies Using Subclassification on the Propensity Score," *Journal of the American Statistical Association*, 1984, *79* (387), 516–524.

Schwarz, Christian and Jens Suedekum, "Global Sourcing of Complex Production Processes," *Journal of International Economics*, 2014, *93* (1), 123–139.

Spolaore, Enrico and Romain Wacziarg, "Ancestry, Language and Culture," in Victor Ginsburgh and Shlomo Weber, eds., *The Palgrave Handbook of Economics and Language*, London: Palgrave Macmillan UK, 2016, pp. 174–211.

Stock, James and Motohiro Yogo, "Testing for Weak Instruments in Linear IV Regression," NBER Technical Working Papers 0284, National Bureau of Economic Research 2002.

Tomiura, Eiichi, "Foreign Outsourcing, Exporting, and FDI: A Productivity Comparison at the Firm Level," *Journal of International Economics*, 2007, *72* (1), 113–127.

UNCTAD, "World Investment Report 2011 – Non-equity Modes of International Production and Development," Technical Report, United Nations Conference on Trade and Development 2011.

_ , "World Investment Report 2013 – Global Value Chains: Investment and Trade for Development," Technical Report, United Nations Conference on Trade and Development 2013.

Wang, Yongjin, Yanling Wang, and Kunwang Li, "Judicial Quality, Contract Intensity and Exports: Firm-level Evidence," *China Economic Review*, 2014, *31*, 32 – 42.

Whinston, Michael D., "On the Transaction Cost Determinants of Vertical Integration," *Journal of Law, Economics, & Organization*, 2003, *19* (1), 1–23.

Williamson, Oliver, "The Vertical Integration of Production: Market Failure Considerations," *American Economic Review*, 1971, *61* (2), 112–123.

_ , *Markets and Hierarchies: Analysis and Antitrust Implications*, The Free Press, New York, 1975.

_ , *The Economic Institutions of Capitalism*, The Free Press, New York, 1985.

Yeaple, Stephen Ross, "Offshoring, Foreign Direct Investment, and the Structure of U.S. Trade," *Journal of the European Economic Association*, 2006, *4* (2-3), 602–611.

6.A Mathematical Appendix

6.A.1 Participation constraint

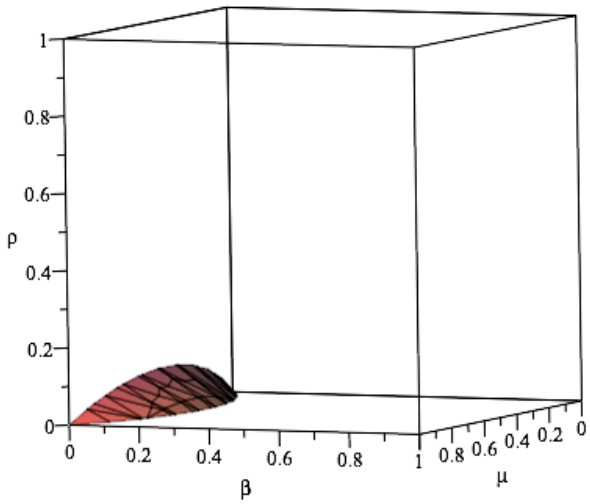
Utilizing equations (6.4), (6.6), (6.8), (6.11), (6.12), (6.13), and (6.15), as well as the definition of $\alpha = \frac{\sigma-1}{\sigma}$ in equation (6.10), we obtain after simplification the sufficient condition for which the optimal ownership share s^* from the viewpoint of H does *not* violate M's participation constraint:

$$\sigma[\rho+\beta(1-\rho)]+\mu^2(\sigma-1)^2(1-\beta)(1-\rho)-\mu(\sigma-1)\Big[\sigma\left[1-2\rho-\beta(1-\rho)\right]-(1-\beta)(1-\rho)\Big]\geq 0.$$

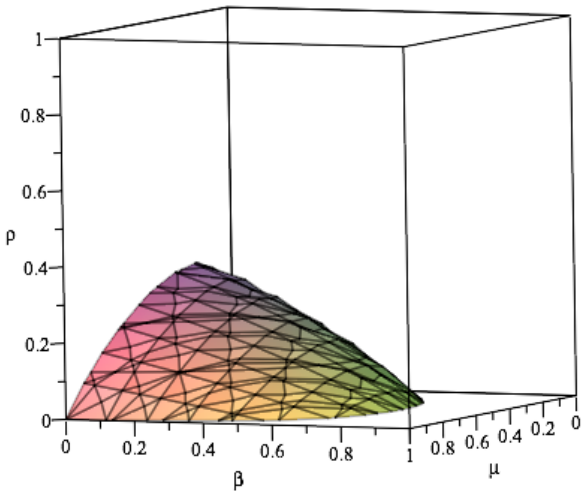
A tedious but straightforward analysis shows that this inequality is more likely to hold the higher ρ and β, less likely to hold the higher σ, and is ambiguously affected by a change in μ. To assess the overall likelihood of this inequality to hold for various *combinations* of parameter values, we fix the value of σ and depict all possible combinations of $\beta \in (0,1)$ and $\mu, \rho \in [0,1]$ which fulfill the above-mentioned condition with equality. The value of $\sigma = 2.25$ assumed in Figure 6.A.1(a) is the mean value in Crozet and Koenig (2010), obtained from estimating a structural model of international trade using French firm-level data. The plane depicted in this figure illustrates the parameter combinations for which M's PC is fulfilled with equality, while it is slack (i.e., $\pi_M > 0$) for any combination of β, μ, and ρ above this plane, and it would be violated (i.e., $\pi_M < 0$) below this plane. As can be seen from Figure 6.A.1(a), M's PC is fulfilled (and can hence be ignored) for the vast majority of parameter values. In Figure 6.A.1(b), we choose an alternative value of $\sigma = 13$, reflecting the mean value estimated by Broda and Weinstein (2006) for five-digit industries, which may be considered a rather high value for the average elasticity of substitution. Compared to Figure 6.A.1(a), M's PC is binding for a larger subset of the parameter space. Nevertheless, it is still non-binding for the vast majority of permissible parameter values.

Figure 6.A.1: *Combinations of β, μ, and ρ which satisfy M's PC with equality*

(a) $\sigma = 2.25$



(b) $\sigma = 13$



6.A.2 Ex-ante transfers

Assume that, after the optimal ownership share is chosen (i.e., in period t_1), H charges from M a transfer (participation fee) T. This transfer can be positive or negative, and it ensures that M is just indifferent between participating in the current relationship and obtaining his ex-ante outside option (normalized to zero).[51] Formally, the equilibrium transfer satisfies the following condition:

$$\pi_M - T = 0 \tag{6.A.1}$$

whereby π_M is given by equation (6.5). Since the transfer is conducted in t_1, it does not affect M's maximization problem in period t_3. Hence, the optimal amount of non-contractible inputs m_n continues to be given by equation (6.6).

Under consideration of the ex-ante transfer, H's pure profit reads $\pi_{HT} = \pi_H + T$, whereby π_H is given by equation (6.9) and T is determined by equation (6.A.1). H's objective function in period t_2 reads:

$$\max_{\{m(i)\}_{i=0}^{\mu}} \pi_{HT} = R - (1-\mu)m_n - \int_0^{\mu} m(i)\mathrm{d}i, \tag{6.A.2}$$

whereby m_n is given by equation (6.6). Notice that, in the presence of ex-ante transfers, H reaps the entire surplus from the relationship. Using equations (6.6), (6.7), and (6.8), the maximization problem from equation (6.A.2) yields the optimal amount of contractible inputs:

$$m(i) = \theta\alpha R \equiv m_c \qquad \forall i \in [0,\mu],$$

as a function of equilibrium revenue (obtained from plugging equation (6.A.2) into equation (6.7)):

$$R = \delta^{\frac{\alpha(1-\mu)}{1-\alpha}} \theta^{\frac{\alpha\mu}{1-\alpha}} \alpha^{\frac{\alpha}{1-\alpha}} D, \tag{6.A.3}$$

whereby

$$\theta \equiv \frac{1 + s(1-\rho) - \beta(1-\rho) - \alpha(1-\beta)(1-\mu)}{[1-\alpha(1-\mu)]\,[1 + s(1-\rho) - \beta(1-\rho)]}. \tag{6.A.4}$$

In period t_1, H maximizes $\pi_{HT} = R - (1-\mu)\delta\alpha R - \mu\theta\alpha R$ via the choice of s, whereby δ, R, and θ are given by equations (6.8), (6.A.3), and (6.A.4), respectively. The first-order condition of this maximization problem yields the following optimal ownership share:

$$s^* = -\frac{\rho\beta}{1-\rho},$$

which is negative. To understand the intuition behind this result, notice from equation (6.8) that $s^* = -\frac{\rho\beta}{1-\rho}$ would fully eliminate M's underinvestment (since $\delta|_{s=s^*} = 1$). With ex-ante transfers, H obtains the entire surplus from the relationship and maximizes the overall surplus by choosing the lowest possible ownership share, which is equal to zero regardless of contracting institutions.

[51] This assumption can be justified by assuming an infinitely elastic supply of M agents competing for a given relationship.

6.A.3 Headquarter intensity

M's maximization problem in period t_3 continues to be given by equation (6.5). Bearing in mind the new production function from equation (6.16), this maximization problem delivers M's reaction function:

$$m(i) = (1-\eta)\delta\alpha R \equiv m_n \qquad \forall i \in [\mu, 1], \tag{6.A.5}$$

whereby δ is given by equation (6.8). In t_3, H chooses the amount of h which maximizes her share of the quasi-rent from equation (6.4) minus production costs of headquarter services: $\max \pi_H = \beta Q - h$.[52] This maximization problem yields the optimal amount of non-contractible headquarter services:

$$h_n = \eta\beta\alpha R, \tag{6.A.6}$$

as a function of revenue (obtained from plugging equations (6.16), (6.A.5), and (6.A.6) into equation (6.1)):

$$R = \left(\left[\exp \int_0^{\mu} \ln m(i) \mathrm{d}i \right]^{\alpha(1-\eta)} \beta^{\alpha\eta}\, \delta^{\alpha(1-\eta)(1-\mu)}\, \alpha^{\alpha[1-\mu(1-\eta)]} (1-\eta)^{-\alpha\mu(1-\eta)} D^{1-\alpha} \right)^{\frac{1}{1-\alpha[1-\mu(1-\eta)]}}. \tag{6.A.7}$$

In t_2, H chooses the amount of contractible inputs that maximizes her profit:

$$\max_{\{m(i)\}_{i=0}^{\mu}} \pi_H = (1-\rho)s(1-\mu)m_n + (1-\rho)\int_0^{\mu} m(i)\mathrm{d}i + \beta Q - \int_0^{\mu} m(i)\mathrm{d}i - h_n, \tag{6.A.8}$$

subject to M's participation constraint ($\pi_M \geq 0$), whereby m_n, h_n, and R are given by equations (6.A.5), (6.A.6), and (6.A.7), respectively. To keep the exposition as simple as possible, we assume in what follows that M's PC is fulfilled. It should be noted, however, that our results continue to hold in case of a binding PC. Utilizing equations (6.4), (6.A.5), (6.A.6), and (6.A.7) in equation (6.A.8), and solving H's maximization problem yields the optimal amount of contractible manufacturing inputs and the associated revenue:

$$m(i) = (1-\eta)\kappa\alpha R \equiv m_c \qquad \forall i \in [0, \mu], \qquad R = \delta^{\frac{\alpha(1-\mu)}{1-\alpha}} \kappa^{\frac{\alpha\mu}{1-\alpha}} \alpha^{\frac{\alpha}{1-\alpha}} D, \tag{6.A.9}$$

whereby

$$\kappa \equiv \frac{\beta - \alpha[\beta\eta - \delta(1-\rho)(1-\mu)(s-\beta)]}{[(1-\rho)\beta + \rho]\Big[1 - \alpha\left[1-\mu(1-\eta)\right]\Big]}, \tag{6.A.10}$$

and δ is given by equation (6.8).

In t_1, H chooses the optimal ownership share by solving the following maximization problem:

$$\max_s \pi_H = (1-\rho)s(1-\mu)\delta\alpha R - \rho\mu\kappa\alpha R + \beta[R - (1-\rho)(1-\mu)\delta\alpha R - (1-\rho)\mu\kappa\alpha R] - \eta\beta\alpha R.$$

Utilizing equations (6.8), (6.A.9), and (6.A.10) therein, we obtain from the first-order condition the optimal ownership share presented in equation (6.17).

[52] Recall that h is assumed to be fully relationship-specific, and hence, it does not affect H's outside option.

6.B Data Appendix

Table 6.B.1: *List of countries by rule of law index and average ownership shares*

Rank	ISO	Rule of law	Subsidiaries	Average share	Rank	ISO	Rule of law	Subsidiaries	Average share
1	FIN[m]	1.94	6,040	87.59	64	GHA[a]	-0.19	7	80.33
2	DNK[m]	1.92	10,161	80.24	65	BMU[c]	-0.19	176	41.68
3	NOR[m]	1.87	14,571	76.55	66	MKD[h]	-0.24	321	59.43
4	NZL[b]	1.84	1,946	88.55	67	MAR[a]	-0.27	310	76.59
5	CHE[j]	1.84	9,814	82.54	68	BGR[h]	-0.29	3,277	76.78
6	SWE[m]	1.81	19,203	89.27	69	BRA[c]	-0.29	6,043	62.98
7	NLD	1.80	35,248	94.86	70	IND[e]	-0.30	9,731	60.26
8	AUT	1.78	11,998	78.03	71	SEN[a]	-0.31	60	79.16
9	AUS[b]	1.75	8,016	84.45	72	TUN[a]	-0.33	1	15.83
10	LUX	1.72	2,505	82.86	73	TTO[c]	-0.34	5	42.00
11	SGP[f]	1.71	2,115	85.47	74	LKA[e]	-0.36	612	77.58
12	CAN	1.71	2,231	58.71	75	THA[f]	-0.36	3,045	60.76
13	GBR[l]	1.71	60,985	93.33	76	SRB[h]	-0.37	1,858	76.73
14	HKG[d]	1.67	4,147	82.80	77	BIH[h]	-0.42	688	72.59
15	DEU	1.67	62,260	79.02	78	MHL[b]	-0.43	13	51.13
16	IRL	1.62	3,582	85.58	79	ZMB[a]	-0.47	9	84.73
17	LIE[j]	1.57	44	95.00	80	MDA[i]	-0.49	156	64.90
18	ISL[m]	1.54	615	72.87	81	MWI[a]	-0.51	3	40.40
19	USA	1.43	27,863	55.62	82	VNM[f]	-0.52	626	59.50
20	JPN	1.41	41,368	51.73	83	JAM[c]	-0.53	13	51.26
21	BEL	1.33	14,519	75.30	84	PHL[f]	-0.54	1,016	69.96
22	FRA	1.28	59,175	79.59	85	CHN[d]	-0.55	16,691	69.78
23	EST[i]	1.18	1,644	81.87	86	COL[c]	-0.56	1,059	69.88
24	MLT[n]	1.02	567	60.81	87	IDN[f]	-0.56	213	59.73
25	TWN[d]	1.01	2,016	57.34	88	ALB[h]	-0.59	125	73.26
26	CZE	0.94	7,800	83.70	89	DOM[c]	-0.60	1	20.00
27	PRT[n]	0.94	10,168	71.71	90	TZA[a]	-0.62	30	68.97
28	ISR[g]	0.91	2,405	72.28	91	MEX[c]	-0.67	79	58.90
29	CYP[h]	0.87	335	71.74	92	KEN[a]	-0.67	12	46.17
30	BRB[c]	0.86	4	56.56	93	GAB[a]	-0.72	1	58.28
31	QAT[g]	0.79	30	53.60	94	SLV[c]	-0.73	3	50.01
32	KOR	0.79	5,764	64.97	95	BFA[a]	-0.75	1	51.00
33	SVN[h]	0.78	1,463	69.67	96	KAZ[g]	-0.77	4,447	85.42
34	ESP	0.74	38,436	70.43	97	PER[c]	-0.77	684	68.46
35	MUS[a]	0.72	86	61.23	98	CUB[c]	-0.79	32	57.82
36	LTU[i]	0.71	1,266	81.08	99	GUY[c]	-0.81	2	20.00
37	LVA[i]	0.67	1,377	82.06	100	EGY[a]	-0.82	250	63.14
38	POL[i]	0.62	13,165	80.12	101	CIV[a]	-0.83	16	73.21
39	ARE[g]	0.51	201	72.85	102	NPL[e]	-0.90	2	65.00
40	URY[c]	0.48	3	77.67	103	PRY[c]	-0.90	5	76.31
41	MYS[f]	0.44	6,091	76.76	104	RUS	-0.93	36,431	70.15
42	BWA[a]	0.44	5	44.57	105	BGD[e]	-0.94	12	48.55
43	CYM[c]	0.41	171	48.31	106	DZA[a]	-0.95	91	67.44
44	OMN[g]	0.38	57	40.92	107	MDG[a]	-0.96	1	100.00
45	CPV[a]	0.37	2	22.76	108	LBN[g]	-0.98	18	67.68
46	CRI[c]	0.31	5	80.49	109	PAK[e]	-1.00	141	60.07
47	HUN[i]	0.30	5,560	81.93	110	UKR[i]	-1.01	9,969	74.00
48	JOR[g]	0.28	51	47.70	111	BLR[i]	-1.03	31	70.02
49	SVK	0.27	3,204	82.22	112	MOZ[a]	-1.06	2	36.45
50	BHR[g]	0.25	32	43.02	113	LBR[a]	-1.07	2	16.16
51	GRC[h]	0.14	2,299	70.38	114	ARG[c]	-1.13	842	69.66
52	ITA	0.13	45,664	63.92	115	BDI[a]	-1.16	1	10.00
53	HRV[h]	0.11	1,708	84.46	116	GTM[c]	-1.21	1	100.00
54	SAU[g]	0.06	312	62.02	117	IRN[g]	-1.26	14	47.10
55	GEO[g]	0.00	51	88.59	118	ECU[c]	-1.28	74	75.36
56	ZAF[a]	-0.05	1,825	82.83	119	BOL[c]	-1.30	6	51.84
57	ROU[h]	-0.06	5,227	76.15	120	UZB[g]	-1.31	9	48.46
58	NAM[a]	-0.07	4	75.50	121	NGA[a]	-1.31	21	52.40
59	KNA[a]	-0.12	2	20.10	122	AGO[a]	-1.32	1	51.00
60	RWA[a]	-0.13	5	83.34	123	SYR[g]	-1.57	2	99.98
61	MNE[h]	-0.14	100	67.49	124	IRQ[g]	-1.59	34	86.37
62	KWT[g]	-0.16	77	57.21	125	ZWE[a]	-1.65	3	36.46
63	TUR[g]	-0.17	5,818	65.41	126	VEN[c]	-2.13	4	70.85

The table lists ISO country codes, sorted in descending order by the rule of law index, the number of subsidiaries in our data, and the average direct ownership share by country. Lower average ownership shares are highlighted in darker shades of gray. Countries grouped by world regions for the selection model in Section 6.4.3 are indicated by: [a] Africa, [b] Oceania, [c] Latin America and Caribbean, [d] China, Taiwan, and Hong Kong, [e] South Asia, [f] South East Asia, [g] Central and South West Asia, [h] South East Europe, [i] Eastern Europe, [j] Switzerland and Liechtenstein, [k] France and Monaco, [l] UK incl. Gibraltar, [m] Northern Europe, [n] South West Europe.

Table 6.B.2: *Summary statistics for main estimation sample*

Variable	Obs.	Mean	Std. Dev.	Min.	Max.
Direct ownership share (percent, Orbis)	230,937	74.805	29.813	10	100
Rule of law index (Worldwide Governance Indicators, World Bank)	230,937	0.87	0.914	-1.595	1.944
Specificity (baseline, subsidiary industry, based on Rauch, 1999)	230,937	0.949	0.162	0	1
ln GDP (subsidiary country, 2010, PWT)	230,772	14.053	1.29	7.636	16.38
ln GDP per capita (subsidiary country, 2010, PWT)	230,772	10.086	0.559	6.013	11.655
ln capital-labor ratio (K_ℓ/L_ℓ) (subsidiary country, 2010, PWT)	188,120	7.86	0.974	2.846	9.756
ln years of schooling (subsidiary country, 2010, PWT)	230,153	1.103	0.112	0.538	1.286
ln capital intensity ln (K_i/L_i) (subsidiary country, Orbis)	230,937	3.277	0.988	-3.061	8.616
ln distance (km, CEPII)	229,770	5.978	1.203	1.007	9.883
Contiguity dummy (CEPII)	229,770	0.048	0.214	0	1
Common language dummy (CEPII)	229,770	0.042	0.2	0	1
Time zone difference (hours, CEPII)	229,101	0.517	1.81	0	12
Colonial link dummy (CEPII)	229,770	0.02	0.14	0	1
Domestic ownership link dummy (Orbis)	230,937	0.798	0.402	0	1
Number of subsidiaries (headquarter, Orbis)	230,937	15.191	41.898	1	888
Number of shareholders (subsidiary, Orbis)	230,937	1.914	3.83	1	421
Financial development (subsidiary country, GFDD)	200,438	182.055	70.789	13.804	582.734
Labor market flexibility (subsidiary country, World Bank)	230,734	0.633	0.168	0.352	1
IPR protection index (subsidiary country, Park, 2008)	223,467	1.112	0.459	-2.253	1.696
Expropriation risk (IHS Markit)	230,349	-.502	0.538	-1.358	3.868
Contract enforcement (IHS Markit)	230,349	0.771	0.626	-3.179	1.481
Law and order (PRS)	230,287	0.771	0.687	-1.61	1.824
Enforcing contracts (WBDB)	230,817	0.7	0.839	-2.837	2.245
Legal formalism index (Djankov et al., 2003)	227,727	0.178	0.682	-2.174	2.667
Property rights freedom (Heritage foundation)	230,809	0.798	0.943	-1.559	1.88
Enforceability of contracts (BERI)	197,694	0.558	0.912	-1.61	1.782
Specificity (conservative, subsidiary industry, based on Rauch, 1999)	230,937	0.966	0.114	0.114	1
Specificity (differentiated, subsidiary industry, based on Rauch, 1999)	230,937	0.722	0.379	0	1
Relationship duration (years/10, Orbis)	230,937	0.255	0.306	0	1
ln employment (subsidiary, Orbis)	109,421	3.727	1.573	0	11.658
ln (value added/ employment) (subsidiary, Orbis)	57,877	4.429	1.039	-6.136	13.333
ln (capital/ employment) (subsidiary, Orbis)	100,649	2.188	2.478	-7.767	14.708
ln (value added/ employment) (subsidiary, Orbis)	134,290	0.693	0.364	0	1
Shareholder dummy (subsidiary, Orbis)	144,545	0.287	0.452	0	1
Firm age (subsidiary, Orbis)	143,679	19.495	18.241	0	813

The table reports summary statistics of all variables used in the empirical analysis for the full estimation sample of Table 6.2.

Table 6.B.3: *Alternative proxies for the quality of contracting institutions*

Measure	Source	Description
Contract enforcement	IHS Markit	Inverse measure of the "risk that the judicial system will not enforce contractual agreements between private-sector entities" (2014, first quarter).
Law and order	Political Risk Services (PRS)	This component of the International Country Risk Guide is designed to measure "the strength and impartiality of the legal system" and "popular observance of the law" (2014).
Enforcing contracts	World Bank Doing Business (WBDB)	The distance to the frontier in enforcing contracts reflects the "time, cost and procedural complexity to resolve a standardized commercial dispute between two domestic businesses" involving "the breach of a sales contract" (2014).
Legal formalism	Djankov et al. (2003)	The index "measures substantive and procedural statutory intervention in judicial cases at lower-level civil trial courts".
Property rights freedom	Heritage foundation	The index reflects a "qualitative assessment of the extent to which a country's legal framework allows individuals to freely accumulate private property, secured by clear laws that are enforced effectively by the government" (2014).
Enforceability of contracts	Business Environmental Risk Intelligence (BERI)	Measures the "relative degree to which contractual agreements are honored and complications presented by language and mentality differences" (Knack and Keefer, 1995).

Chapter 7

Service Offshoring and Firm Employment

Major technological advances have recently spurred a new wave of offshoring in services, which used to be non-tradable. Should service workers in developed countries worry about their jobs? Trade theory has given a nuanced answer to this question, suggesting that efficiency gains from offshoring may counteract direct job losses, which leaves the predicted net effect ambiguous. This paper investigates the employment effects of service offshoring in a newly combined and exceptionally detailed panel dataset, covering almost the entire universe of German firms' service imports over the years 2001-2013. It exploits firm-specific export supply shocks by partner countries and service types as an instrumental variable to find that service offshoring has *increased* firm employment. In line with the canonical trade in tasks model, the employment gains are greater in firms with higher initial levels of service offshoring.

7.1 Introduction

There is a widespread fear in the developed world that domestic jobs are endangered by offshoring to low-wage countries. A popular narrative suggests that these anxieties played a decisive role in the US presidential elections of 2016.[1] More generally, the increasing public concern about offshoring is evidenced by the rising media attention to this topic over time, as documented for the UK and the US by Amiti and Wei (2005) and Mankiw and Swagel (2006), and more recently discussed by Hummels et al. (2016). In Germany, offshoring has been found to increase job loss fears, particularly among high-skilled workers (Geishecker et al., 2012). These workers may indeed be threatened by a 'new wave' of offshoring in services (Bardhan and Kroll, 2003) that has been triggered by groundbreaking innovations in information and communication technologies (ICT) over the past decades. The spread of portable computers, broadband internet, and the smartphone has made services tradable globally that could only be provided locally before. As a consequence, the share of service trade in world GDP has doubled from around 3% to 6% between 1985 and 2016.[2] Since jobs in the newly tradable commercial services are typically skill-intensive and high-paying (Jensen, 2008), the question arises: Are these 'good' domestic jobs lost due to service offshoring?

In theory, the employment effects of offshoring are not as clear-cut as one might expect. While the relocation itself obviously reduces domestic employment, the associated cost savings entail a productivity effect that can help offshoring firms to expand their output, which ameliorates the job losses and might even turn them into job creation. The seminal 'trade in tasks' model by Grossman and Rossi-Hansberg (2008), formalizing these ideas, has typically been applied to offshoring of manufactured inputs, but similar effects can be expected for services.[3] The ambiguous theoretical predictions call for an empirical investigation into the employment effects of service offshoring.

The lack of appropriate micro data on service offshoring firms has severely restricted empirical investigations in the past, as establishing causality is difficult at the industry level. Only recently, firm-level data on service trade have become available and opened the door to a more rigorous causal analysis. Such a firm-level analysis can build on recent extensions of the trade in tasks model that show how the relocation and productivity effects of offshoring vary across heterogeneous firms.[4] However, the same models also predict that larger and more productive firms are more active in offshoring, raising an important endogeneity issue that needs to be addressed empirically.

This paper's goal is to estimate the causal effects of service offshoring on firm employment in Germany, the world's second largest importer of services. Service offshoring is measured by imports of tradable commercial services, reflecting the fact that the relocation of a required service to another country necessarily entails subsequent imports of this service. The paper builds on an exceptionally detailed

[1] For instance, the political analysis website *FiveThirtyEight* reports that Donald Trump received substantially more votes in counties with a higher share of routine jobs, which are potentially threatened by automation or offshoring (see https://fivethirtyeight.com/features/trump-was-stronger-where-the-economy-is-weaker/). Autor et al. (2017) stress the impact of imports from China on the election outcome.

[2] For comparison, goods trade increased by around 40% relative to GDP over the same period. These numbers are based on service trade data from the World Trade Organization (WTO) as well as data on GDP and goods trade from the World Bank's World Development Indicators.

[3] The early contribution by Jones and Kierzkowski (1990) focused on the offshoring of services, not manufactured inputs. The beneficial effects of cost savings from offshoring for domestic labor have previously been investigated by Arndt (1997, 1998a,b), Egger and Falkinger (2003), Jones and Kierzkowski (2001), and Kohler (2004a,b). In addition to the negative relocation effect and the positive productivity effect, the trade in tasks model further predicts an ambiguous relative-price effect (see Deardorff, 2001a,b, for an earlier account).

[4] See Egger et al. (2015, 2016), Groizard et al. (2014), and Sethupathy (2013).

panel dataset, covering almost the entire universe of German firms' service imports by partner countries and service types over the period from 2001 to 2013. These data are newly combined with firm-level employment and balance sheet information for the purpose of this analysis. The resulting firm panel provides an almost comprehensive picture of German service offshoring.

I use these micro data to investigate the firm-level employment effects of both changes in the volume of offshoring (the intensive margin) and new offshoring (the extensive margin). The main analysis at the intensive margin instruments for service offshoring using firm-specific export supply shocks by partner countries and service types in fixed effects regressions. This instrumental variable (IV) strategy follows Hummels et al. (2014), who use it to analyze offshoring of manufactured inputs by Danish firms. The IV exploits the fact that firms' importing behavior is highly firm-specific and stable over time. Holding the initial import mix of a firm constant, the time variation in its partner countries' exports to the rest of the world by service types is used as a revealed measure of changes in these countries' comparative advantage. The rationale behind this approach is that a German firm, which initially imports a given service from a given country, benefits disproportionately from an improvement in the country's comparative advantage in this particular service, and can thus expand its offshoring activities. The crucial assumption of the IV approach is that foreign exports to the rest of the world are uncorrelated with the German firm's employment growth, except through offshoring (conditional on the control variables, which include industry-year fixed effects). Under this exclusion restriction, it serves to identify the causal effect of service offshoring on firm employment.

The main finding of this paper is that service offshoring has *increased* domestic employment in German firms. The estimated elasticity of 6.0-7.6% suggests that this effect is economically sizable, and the IV approach ensures that the positive effect is not driven by simultaneity or omitted variables. Instead, the two-stage least squares estimates can be interpreted as evidence for a cost savings effect from offshoring, which allows firms to expand and hire more workers. In line with this interpretation, service offshoring is found to boost firm output, labor productivity, and the use of other inputs. Moreover, the employment gains are greater in firms with higher initial levels of service offshoring. These patterns are fully in line with the functioning of the productivity effect in Grossman and Rossi-Hansberg (2008) and can be rationalized by recent offshoring models featuring firm heterogeneity (see Antràs et al., 2017; Egger et al., 2016). Intuitively, if foreign services become cheaper, the firm experiences cost savings that tend to raise its output and can increase the use of domestic inputs, including employment. However, these cost savings are approximately zero for the first offshored worker, as they work only through the infra-marginal tasks, which have previously been offshored. Consequently, a firm that has previously offshored more services experiences greater cost savings, and hence more favorable employment effects. These predictions are borne out strongly in the data, suggesting that service offshoring entails substantial productivity effects that benefit domestic employment in the firm.

The analysis proceeds by digging deeper into the rich micro data to shed some light on three interesting features of the employment effects at the intensive margin of service offshoring. First, splitting up service imports into high-income and low-income source countries reveals that the latter entails more favorable employment and wage effects. This pattern is in line with the expectation that services provided by low-income countries are more complementary to domestic employment in German firms in terms of task and skill requirements and are therefore more likely to benefit domestic workers (see Groizard et al., 2014). Second, one might suspect that service imports reflect not only activities that were previously conducted by the firm itself, but also those sourced from other German suppliers prior to offshoring.

However, sample splits confirm that service offshoring increased employment irrespective of contemporaneous changes in domestic sourcing. Furthermore, the employment gains turn out to be even stronger for narrow offshoring (of the service type corresponding to the firm's own industry) and for firms in the service sector, where the substitution of domestic suppliers might be less relevant because firms engage in service activities themselves. Third, since the combined dataset includes information on all sizeable foreign direct investment (FDI) links of German firms, it can be used to shed some first light on possible differences in the employment effects between intra-firm and arm's length service offshoring. The data show that the bulk of German service offshoring takes place via outsourcing, which strongly boosts domestic employment.

The positive employment effects at the intensive margin of service offshoring prove robust to several modifications of the IV strategy, alternative treatment of the data, and addressing various challenges to identification. They are also confirmed in different subperiods, in dynamic panel data models, and after accounting for possible selection into the dataset or into the service offshoring activity.

The analysis is completed by considering the extensive margin of service offshoring in a difference-in-differences propensity score matching approach. It compares employment changes in firms that start service offshoring for the first time to changes in a matched control group with similar initial conditions. This analysis reveals that newly offshoring firms experience non-negative employment effects over the subsequent one to five years. While the average treatment effect on the offshoring firms is even positive in the full sample pooled over all years, detailed annual analysis reveals that the estimated effects of starting service offshoring are insignificant in most years, but never negative.

This paper contributes to the literature estimating the labor market effects of offshoring, pioneered by Feenstra and Hanson (1996a,b, 1999). Focusing predominantly on the offshoring of manufactured inputs, numerous studies have applied their proxy for offshoring, which is based on industry-level imports and input-output tables, to the analysis of industry-level wages and employment.[5] The same approach has further been applied to the analysis of worker-level wages.[6] Recently, Hummels et al. (2014) have brought the analysis to firm-worker data for Denmark, which allows them to measure (and instrument for) offshoring using firm-level import data. Their approach has been adapted by Antràs et al. (2017) and Bernard et al. (2018) to investigate the impact of offshoring on firms' other sourcing and innovation activities. An alternative firm-level approach has exploited the activities of multinational enterprises' foreign affiliates to measure offshoring.[7] Matching methods have been used to analyze the employment effects at the extensive margin of offshoring by Monarch et al. (2017), using firm offshoring events in the US, and by Moser et al. (2015), exploiting qualitative information from an establishment-level survey in Germany. The majority of these studies find that offshoring of manufactured inputs has small adverse effects on low-skilled workers' domestic employment or wages.

Amiti and Wei (2005, 2009a) were the first to apply the approach developed by Feenstra and Hanson to services. They analyze the relationship between employment and service offshoring at the industry level in the UK and the US, where they find mixed evidence and rather small correlations. Crinò (2010b,

[5]This literature is reviewed by Feenstra and Hanson (2003). A recent contribution in this vein is by Wright (2014), who finds evidence for negative effects of offshoring on employment and positive effects on output in US industries.

[6]See Ebenstein et al. (2014, 2015) for the US as well as Geishecker and Görg (2008) and Baumgarten et al. (2013) for Germany.

[7]Examples include Head and Ries (2002) for Japan, Muendler and Becker (2010) for Germany, as well as Harrison and McMillan (2011) and Sethupathy (2013) for the US. Note that this approach restricts attention to intra-firm offshoring and is less suitable to study the offshoring of services, since my data reveal that the majority of service offshoring by German firms is to unrelated parties, not to foreign affiliates or investors.

2012) further investigates the relative employment effects across skill groups in the US and Europe. He finds that service offshoring favors high-skilled employment, similar to offshoring of manufactured inputs.[8] These findings are confirmed in worker-level wage data for the UK by Geishecker and Görg (2013), who also rely on the industry-level service offshoring measure. The two studies most closely related to this paper are Crinò (2010a) and Hijzen et al. (2011), who examine the link between service offshoring and firm-level employment. Crinò (2010a) applies matching methods to a cross-section of Italian firms to investigate the employment effects at the extensive margin of service offshoring, which turn out to be insignificant as in this paper. While he compares firms that import services to those that do not, the German firm panel data allow me to extend this approach to a difference-in-differences setup, which examines *changes* in employment in firms that *start* importing services, thereby controlling for time-invariant confounding factors. Hijzen et al. (2011) use detailed import data for the UK to show that increasing service offshoring coincides with higher employment growth. However, their approach does not address the likely endogeneity of offshoring, so they conclude that the positive correlation may be explained either by efficiency gains from offshoring or by simultaneity. This paper applies an IV strategy to eliminate the second possibility, and thus provides first causal evidence for the firm-level employment gains from increased service offshoring.

More broadly, the paper contributes to the literature analyzing service trade (surveyed by Francois and Hoekman, 2010), which is small compared to the abundance of research on goods trade. A distinguishing feature is the intangible and non-storable nature of services, which traditionally could be provided only face to face (Hill, 1977). By relaxing this requirement, advancements in ICT have contributed to a steep decline of international trade costs for many services over the past decades (see Hoekman and Braga, 1997; Freund and Weinhold, 2002). Despite the differences to goods trade, previous studies found that aggregate service trade is well explained by the traditional gravity equation (Kimura and Lee, 2006; Head et al., 2009). It was not until recent years that firm-level service trade data have become available, which have uncovered that service-trading firms are a small group, with a similar and even more pronounced heterogeneity in terms of performance than found for goods traders (see most prominently Breinlich and Criscuolo, 2011, for the UK).[9] The key advantage of the dataset used in this paper relative to previous studies is that it combines (close to) full coverage of German firms' service trade with information on firm employment and other key characteristics, which allows for a comprehensive analysis of causal effects.

The paper is organized as follows: Section 7.2 describes the rich micro dataset of German firms' service trade compiled for this study. Section 7.3 describes the econometric model and the within-firm correlations between employment and service offshoring found in simple OLS regressions. Section 7.4 develops the IV strategy and uses it to estimate the employment effects at the intensive margin of service offshoring, followed by several deeper investigations. Section 7.5 analyzes the role of selection and the extensive margin of offshoring. The final section concludes with a brief discussion of the findings and their implications.

[8] Criscuolo and Garicano (2010), Jensen and Kletzer (2005, 2010), and Liu and Trefler (2011) develop alternative approaches for identifying tradable services from US occupational data to investigate the labor market implications of service offshoring. A related strand of the literature has studied the effect of service offshoring on productivity; see Amiti and Wei (2009b) and Winkler (2010) for industry-level studies and Crinò (2008) for a firm-level analysis.

[9] Similar patterns have been documented by Ariu (2016) for Belgium, Federico and Tosti (2016) for Italy, Gaulier et al. (2010) for France, Kelle and Kleinert (2010) for Germany (using the same service trade database as this paper), Morikawa (2015) for Japan, Wolfmayr et al. (2013) for Austria, and Damijan et al. (2015) for four other European countries.

7.2 Data

7.2.1 Data sources

The panel dataset of German firms used in this paper combines information from three sources: the Statistics on International Trade in Services (SITS), the Corporate Balance Sheet Statistics (USTAN), and the Microdatabase Direct Investment (MiDi). These confidential micro datasets are provided by the Research Data and Service Centre (RDSC) of the Deutsche Bundesbank (the German central bank) on site for research purposes. Only recently these datasets have been linked at the RDSC, and this paper is among the first to exploit information combined from all three sources.[10] This link is essential for the paper's objective because SITS does not contain information on firm employment, which is hence taken from USTAN and MiDi (see Section 7.2.3 on how the data are combined).

The Statistics on International Trade in Services (SITS, Biewen et al., 2013) provides highly detailed panel data on imports and exports of services by firm, month, partner country, and service category for around 22,000-26,000 firms over the years 2001-2013 (a full list of all countries and service categories is provided in the documentation). These data are collected by the Deutsche Bundesbank for the purpose of compiling the German balance of payments. Service trade flows are defined as transactions between German residents and non-residents, which correspond to modes 1 (cross-border trade), 2 (consumption abroad), and 4 (presence of natural persons) according to the WTO's General Agreement on Trade in Services (GATS). Transactions via commercial presence (GATS mode 3), such as purchases by foreign affiliates, are not included in this definition, which is ideal given the paper's focus on domestic employment effects. The data contain the universe of German firms' service trade for all transactions exceeding the reporting threshold of € 12,500. The high level of detail and the comprehensive coverage of these data make them uniquely suited to address the research question posed in this paper. These features represent key advantages over the vast majority of firm-level service trade data previously used in the literature, which are based on firm surveys.[11] In particular, the SITS data allow me to construct a firm-specific instrument for service offshoring (see Section 7.4.1). The analysis in this paper focuses on tradable commercial services, which are typically the subject of the offshoring debate. Therefore, it excludes all service trade classified as government services, incidental payments, private transfers, royalties and license fees, as well as travel and transport services.[12] Any references to total service imports throughout the paper refer to these tradable commercial services. In anticipation of the IV strategy, the remaining service categories are grouped into ten service types – such as communications, engineering, or research and development (R&D) services – according to Table 7.A.1, which broadly follows Biewen et al. (2013). The SITS are also aggregated over months to combine them with the other annual datasets.

Information on firm employment comes from two data sources. The first is the German Corporate Balance Sheet Statistics (USTAN, see Deutsche Bundesbank, 1998; Stöss, 2001). This dataset contains detailed balance sheets and income statements of a large number of non-financial German firms. Crucially for this paper, it includes information on the number of employees. The USTAN data are collected by the Deutsche Bundesbank for the purpose of credit assessments. The balance sheet data are used and

[10] MiDi has previously been used in combination with USTAN (e.g. by Jäckle and Wamser, 2010; Muendler and Becker, 2010), or in combination with SITS (Biewen et al., 2012). Eppinger (2014) provides first explorations of a dataset linking all three sources.

[11] Of the datasets used in all previous studies known to me, only the Belgian data used by Ariu (2016) are comparable to the SITS in terms of coverage.

[12] These payments cannot be thought of as offshoring, as is also argued by Head et al. (2009). Liu and Trefler (2011) make an analogous selection. The main findings are robust to a narrower selection, as discussed in Section 7.4.7.

carefully validated by central bank staff to assess the value of securitized, non-marketable claims or bills of exchange, which are presented as collateral to the central bank by commercial banks. The analysis excludes all consolidated balance sheets pertaining to corporations, as well as balance sheets referring to a short fiscal year. The full USTAN dataset used in this paper covers the years 1999-2013 and includes around 22,000-29,000 firms per year over the period 2001-2013.

The second source of employment data is the Microdatabase Direct Investment (MiDi, see Lipponer, 2011; Schild and Walter, 2017), which contains information on all German firms with inward or outward stocks of FDI above a reporting threshold. This threshold has been unchanged since 2002 at a minimum of 10% shares or voting rights in an affiliate with a balance sheet total exceeding € 3 million. Most importantly in the context of this paper, MiDi contains information on the number of employees and turnover for all firms involved in FDI, i.e., each multinational enterprise (MNE) and each foreign-invested enterprise (FIE) in Germany. The bulk of information contained in MiDi serves to provide a detailed picture of German firms' FDI links, including information on the country and industry of their foreign affiliates or investors. These data are exploited in this paper to identify German firms which have a related party in the country and industry to which they offshore services.

Bilateral data on international trade in services across country pairs, required to construct the IV (in Section 7.4.1), is taken from the UN Comtrade database.[13] Since information is frequently missing at more disaggregate levels, I use service trade data for a set of service codes corresponding largely to the first level of the Extended Balance of Payments Services (EBOPS) 2002 classification. The EBOPS codes are matched to the service types in the SITS data according to the correspondence reported in Table 7.A.1.

7.2.2 Service offshoring by service types, countries, and industries

Which service types do German firms offshore most frequently and to which main countries? The service trading activities of German firms have previously been described by Kelle and Kleinert (2010) and Biewen et al. (2013) based on earlier vintages of the SITS data. Hence, my description of general patterns in the data is brief and focuses on service offshoring, defined as the imports of other commercial services. The interested reader is referred to the Online Appendix for further details.

By far the most popular type of service for imports is other business services, a diverse category including administrative services, advertising, and marketing. These types of services are offshored by 50,840 firms in the full SITS sample and account for 24% of the aggregate import value. Only insurance services account for a larger share (33%), but these imports are concentrated among a small number of firms (mainly in the insurance and finance sectors), and they are excluded in a robustness check. The next largest import shares are in R&D (10.5%), Computer & IT (8.6%), and engineering services (6.7%).

The bulk of German service offshoring is to high-income countries. The top source countries are the US and the UK, followed by several neighboring countries (Switzerland, the Netherlands, and France), reflecting the fact that gravity is also relevant for trade in services (see Kimura and Lee, 2006; Head et al., 2009). India, known as a popular source country for service offshoring, ranks only 20th. The US are also the main source country for each individual service class with three exceptions: financial services (UK), engineering (UK), and construction services (Austria). It seems plausible that gravity is strongest in construction services, which are likely to be provided mainly via the presence of natural persons.

[13] These data are obtained from https://comtrade.un.org/data/.

How do firms' service offshoring activities vary across industries? As pointed out by Kelle and Kleinert (2010), German firms from all sectors engage in importing services. The largest value of service imports goes to the manufacturing sector, where other business services account for the largest share of imports. In the service sector, narrow offshoring is rather prevalent: Firms in the communication, construction, and insurance industries all import predominantly the same type of services as their own industry provides. Note, however, that most of the variation in service offshoring activities is across firms within narrowly defined industries, as discussed in more detail Section 7.2.5.

7.2.3 Data preparation

The three micro-level datasets introduced in Section 7.2.1 are combined via firm identification numbers, which are identical in SITS and MiDi. They are matched to USTAN via a correspondence table provided by the Deutsche Bundesbank (see Schild et al., 2017), resulting in the new combined dataset that covers around 57,000 firms annually over the years 2001-2013. Since USTAN does not cover the full population of German firms, and since by far not all firms are involved in service trade or FDI, the three data sources overlap only imperfectly. Nevertheless, the combined sample of firms for which we know both their employment and whether they import services (around 31,000 firms per year) includes around 83% of the total import value of other commercial services reported in SITS. This impressive coverage is due to the fact that the (almost fully covered) MNEs and FIEs in MiDi as well as the firms reporting to USTAN tend to be large compared to the average firm and hence account for a disproportionate share of total service imports. These firms make up on average 25% of aggregate employment and 53% of aggregate turnover in the underlying non-financial private business sector in every year.[14] Overall, the newly combined dataset provides an almost comprehensive picture of German firms' service offshoring activities and the firms in this dataset employ a substantial share of the entire German labor force.

Since SITS covers the entire universe of German firms' service trade (above a low reporting threshold), firms from USTAN and MiDi that do not show up in SITS have negligible service imports and exports, which can hence be set to zero. Similarly, one can be sure that a firm in USTAN or SITS is not involved in any economically significant inward or outward FDI if it does not report to MiDi.

Information on firm employment, the key outcome variable analyzed in this paper, is taken from either USTAN or MiDi, with preference given to the former data source, as it also contains the balance sheet information used to construct additional control variables in the empirical analysis. Due to the importance of firm employment, I implement a number of consistency checks to validate this variable and eliminate potential outliers, as described in the Online Appendix.

An important variable required for the empirical analysis that is not readily observable in the data is the physical capital stock. It is constructed by the perpetual inventory method, closely following the procedure applied by Bachmann and Bayer (2014) to the USTAN data, which is outlined in the Online Appendix.

Firm productivity has been established in both theoretical and empirical work as an important determinant of firms' offshoring activities (see Antràs and Helpman, 2004; Kohler and Smolka, 2014) and may be affected by service offshoring (Amiti and Wei, 2009b). It is approximated in this paper by labor

[14]These figures are based on data from the German national accounts from the German Statistical Office (see https://www.destatis.de/EN/). The non-financial private business sector is defined by excluding the sectors not covered in the micro data: private households, public administration, and the financial sector.

productivity (LP), defined as real output per worker.[15]

7.2.4 Summary statistics of firm variables

The subsequent analysis of correlations (in Section 7.3.2) draws on the entire combined firm dataset, subject to the availability of control variables and other required information. For the main analysis of employment effects at the intensive margin of service offshoring (in Section 7.4), the sample needs to be restricted to firms with positive service imports and employment information observed in several years over the period 2002-2013. The reason is that the identification strategy, which will be described in detail in Section 7.4.1, relies on time variation in imports. It further depends on the availability of aggregate service trade data from Comtrade, which reduces the estimation sample slightly. Since the main focus of the empirical analysis is on changes at the intensive margin, the full estimation sample used for this analysis (corresponding to the estimates reported in column 1 of Table 7.3) is described in more detail in this section.

The full estimation sample is an unbalanced panel of 7,018 firms (around 3,400 per year), which account for 71% of total service imports in SITS. Even though the number of firms in this sample is small compared to the original data sources, they account for the bulk of all service offshoring and around half of aggregate employment and turnover in the combined USTAN and MiDi datasets. These service importers represent around 11% of aggregate employment and 25% of aggregate turnover in the German non-financial private business sector. The shares are even higher in the manufacturing sector, where the full estimation sample accounts for 30% of aggregate employment and 41% of aggregate turnover (due to a higher coverage rate of USTAN in this sector).

Table 7.1 provides summary statistics of several key variables in the USTAN and SITS samples as well as the full estimation sample. The combined dataset created for this paper provides the first opportunity to examine the characteristics of German service importers compared to other firms. Service importers are a small group, among which import volumes are heavily concentrated in even fewer firms, as the numbers provided in the previous paragraph suggest. This heterogeneity is also visible in several dimensions of firm size and performance. By comparing the USTAN sample with the subset of SITS for which additional firm information is available from the other sources, it can be seen that service importers are larger in terms of employment, turnover, and capital stock; they pay higher wages and are more productive (in terms of LP) than other firms in USTAN. Furthermore, the table suggests that firms included in the estimation sample are even larger on average in terms of employment and turnover, which seems plausible since these firms are successfully offshoring services over several years. This fact can rationalize why the firms in the estimation sample account for the bulk of total German service imports.

7.2.5 Stylized facts about the micro-structure of service offshoring

The data reveal that the service importing structure is highly firm-specific and stable over time within firms, two stylized facts that are important for the IV approach. Out of the possible 1,790 country-service type combinations, positive imports are observed for 1,003 combinations in the full estimation sample.[16]

[15] Throughout the paper, wages are deflated by the consumer price index (CPI), while turnover and output are deflated by industry-level producer price indices (complemented by the CPI whenever missing), both obtained from the German Statistical Office.

[16] In the SITS, I can distinguish between ten aggregate service types and more than 200 source regions (countries and territories). Excluding small islands and overseas territories, and keeping only those countries on which import data is available both in SITS and at the aggregate level from Comtrade, I am left with 179 source countries.

Table 7.1: *Summary statistics*

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	USTAN sample			SITS sample			Full estimation sample		
Variable	Mean	SD	N	Mean	SD	N	Mean	SD	N
employment	216	2,204	276,357	622	4,267	98,723	1,026	6,170	40,350
ln *employment*	3.682	1.737	276,357	4.912	1.671	98,723	5.475	1.556	40,350
importer	0.123	0.329	320,398	0.871	0.336	318,932	1.000	0.000	40,350
imports	887	23,507	320,398	3,750	92,920	318,932	19,758	220,607	40,350
ln *imports*	5.772	2.143	39,559	5.296	2.007	277,680	7.171	2.030	40,350
# *countries*	0.767	3.994	320,398	3.450	6.565	318,932	9.778	12.594	40,350
# *service types*	0.271	0.922	320,398	1.543	1.338	318,932	2.924	1.898	40,350
# *country-service types*	8.509	18.989	41,310	4.568	11.418	318,932	14.650	24.842	40,350
ln *narrow offshoring*	5.829	2.300	4,206	5.343	2.084	61,160	7.338	2.475	6,366
ln *intra-firm offshoring*	6.358	2.677	2,187	6.031	2.417	11,693	6.628	2.586	5,625
exporter	0.037	0.188	320,398	0.318	0.466	318,932	0.436	0.496	40,350
ln *exports*	6.926	2.437	11,802	6.235	2.258	101,360	7.856	2.412	17,588
MNE	0.052	0.222	320,398	0.111	0.314	318,932	0.350	0.477	40,350
FIE	0.052	0.222	320,398	0.194	0.395	318,932	0.582	0.493	40,350
ln *turnover*	8.947	2.009	307,217	10.727	1.655	97,220	11.298	1.590	40,199
goods exporter	0.287	0.452	320,398	0.075	0.264	318,932	0.285	0.452	40,350
ln *capital*	7.701	2.241	252,059	9.265	2.216	35,310	9.803	2.140	15,240
ln *wage*	3.734	0.568	267,276	4.063	0.458	39,503	4.078	0.373	16,576
ln *output*	4.346	2.013	307,049	6.324	1.647	39,536	6.876	1.564	16,536
ln *LP*	0.891	1.051	269,716	1.107	1.066	38,306	1.162	0.991	16,536
ln *material inputs*	10.745	2.164	256,419	12.406	1.965	36,940	12.925	1.876	15,900
ln *domestic services*	6.727	2.090	318,029	8.934	1.588	40,328	9.463	1.537	16,178

Note: The table lists the mean, standard deviation (SD), and number of observations (N) of important variables used in the analysis for three different samples: The USTAN sample (51,926 firms, columns 1 to 3), the SITS sample (83,482 firms, columns 4 to 6), and the preferred estimation sample for the analysis at the intensive margin of service offshoring (7,018 firms, columns 7 to 9). Source: RDSC of the Deutsche Bundesbank, SITS, MiDi, and USTAN, 2001-2013, own calculations.

The median firm-year observation in this sample has positive imports in five different service types and 29 different country-service type combinations. These service offshoring strategies vary considerably across firms, such that only a small number of country-service type combinations feature positive imports from many firms.[17] For instance, only six firms report positive imports for the median country-service type combination in 2007, the middle of the sample period. These numbers illustrate that typically very few firms share the same service offshoring strategy. Instead, the importing structure is highly firm-specific, presumably driven by the specific needs of individual firms.

Even within narrowly defined industries, service importing choices vary substantially across firms, similar to the patterns reported by Hummels et al. (2014) for goods trade. To substantiate this point, I regress the import share of a firm by country and service class, observed in the first year that the firm is importing services, on different sets of dummies.[18] While 3-digit industry dummies explain only 10% of the variation in these shares, a regression on firm dummies yields an R^2 of 82% (or 42% when firms with a single observation are dropped). This finding casts doubt on the standard approach in the offshoring literature, which applies the same input coefficients to all firms in an industry based on aggregate input-output tables. The German data reveal that this concern applies also to service trade, calling for a firm-level approach to measuring (and instrumenting) service offshoring, as the one pursued in this paper. I

[17] In 2007, the most popular service imports are 'other business services' (in particular the category 'advertising, commercial, and administrative services') from Switzerland, followed by imports of the same service type from the UK and the US.

[18] The data from this first year for each firm constitute the pre-sample and will be omitted in the main empirical analysis for the reasons described in Section 7.4.1.

return to this issue in a robustness check in Section 7.4.7.

To illustrate the stability of firms' importing strategies over time, I compute the share of the total service import value in the full estimation sample that is accounted for by the firm-country-service type combinations that are observed in the first importing year for each firm. The data show that these firm-country-service type combinations account for 57% of the total service import value in the full estimation sample of all subsequent years. For the average firm in the estimation sample, these initial country-service type combinations even amount to 73% of their total import value. This observation reflects a high degree of persistence in firms' service importing strategies over time, which is comparable to the persistence observed in Danish goods trade (Hummels et al., 2014). These features of the data allow me to hold the initial import shares by country and service type constant when instrumenting for the intensive margin of service offshoring in the empirical analysis.

Despite the stable importing structure, there is also considerable within-firm variation in the value of service imports over time, reflected in a coefficient of variation of 0.429 (the standard deviation of service imports based on the within-firm variation relative to the mean) in the full estimation sample. The empirical analysis in Section 7.4 investigates to what extent this variation in service offshoring can explain changes in firm employment.

7.3 Econometric Model and Within-firm Correlations

7.3.1 Econometric model

To investigate the relationship between service offshoring activities and employment of firm i, active in industry j and year t, the following econometric model is specified:

$$\ln \textit{employment}_{i,t} = \beta \cdot \textit{service offshoring}_{i,t} + \boldsymbol{\varphi} \cdot \mathbf{X}_{i,t-1} + \alpha_{j,t} + \alpha_i + \varepsilon_{i,t}, \tag{7.1}$$

where the key explanatory variable $\textit{service offshoring}_{i,t}$ represents either the logarithm of the value of service imports $\ln \textit{service imports}_{i,t}$ (for the main analysis at the intensive margin of offshoring), or a dummy variable indicating positive service imports (for the analysis at the extensive margin of offshoring), or alternative measures of service offshoring that will be introduced in the analysis below. Firm employment is the main dependent variable of interest in this paper. Yet, other firm performance measures will also be considered as alternative dependent variables in equation (7.1). The main parameter of interest is β, the partial effect of service offshoring on firm employment.

Importantly, equation (7.1) includes firm fixed effects α_i, which absorb any time-invariant components of the firm's market environment and geographic location, its productivity, size, ownership structure, and other firm characteristics. In addition, the industry-year fixed effects $\alpha_{j,t}$ absorb any shocks to demand, factor markets, or technology that are common to all firms in an industry (defined at the 2-digit level of the NACE Rev. 1.1 classification). The equation further includes the following firm-level control variables $\mathbf{X}_{i,t-1}$ lagged by one year (with associated coefficients $\boldsymbol{\varphi}$): dummy variables indicating whether the firm is a service exporter, an MNE, or an FIE (both defined by the MiDi thresholds), as well as $\ln turnover$.

Equation (7.1) is the firm-level analogue to the employment equations estimated by Amiti and Wei (2005, 2009a).[19] It can be derived as a conditional labor demand function from a standard cost minimization problem of the firm (see e.g. Hamermesh, 1993). For this purpose, it is assumed that wages are exogenous to firms and absorbed by industry-year fixed effects. This assumption is innocuous to the extent that wages are set at the industry level by unions, which play an important role in the German labor market. Nevertheless, it is relaxed in a robustness check in Section 7.4.7.

OLS regressions would yield consistent estimates of β only under the strong assumption that the error term $\varepsilon_{i,t}$ in equation (7.1) is uncorrelated with service offshoring (and the other explanatory variables), e.g. because it is due to random measurement error. While Section 7.3.2 does examine OLS estimates of β to illustrate patterns in the full available dataset, I abstain from making this assumption, as it seems unlikely to hold in practice. In particular, the alternative measures of *service offshoring*$_{i,t}$ may be correlated with the error term due to the simultaneity of employment and offshoring decisions, both of which are affected by output demand, technology, and labor supply shocks. Therefore, the OLS estimates presented below are interpreted as conditional correlations, whereas causal analysis is relegated to Sections 7.4 and 7.5.

7.3.2 Within-firm correlations

Table 7.2 demonstrates that various measures of service offshoring activity are positively correlated with employment and other dimensions of performance in German firms. It reports the estimated conditional correlations between various firm variables (in logs), indicated in the header, and several alternative measures for service offshoring, indicated in each row. Each cell reports an estimate of β based on a different variant of the within-transformed equation (7.1).[20] The estimate shown in the top cell in column 1 examines the correlation of employment with the firm extensive margin of service offshoring, as measured by the service importer dummy. The FE estimate suggests that firms do not experience a significant increase or decrease in terms of employment as they start (or stop) service offshoring. There is also no significant correlation of the wage per employee with the importer dummy (see column 2). However, the estimates reported in columns 3-4 in the first row show that service offshoring firms are larger in terms of output and more productive as measured by LP compared to non-offshoring firms. These findings extend the results of Breinlich and Criscuolo (2011), who have estimated performance 'premia' of service importers in similar regressions using cross-sectional variation across UK firms. Table 7.2 reveals that these performance differences in terms of output and productivity are significant (though much smaller) even when identified from time variation within firms in Germany. Finally, the service offshoring dummy is positively correlated with sourcing of material inputs but uncorrelated with sourcing of domestic services (columns 5-6).[21]

The second, third, and fourth rows of Table 7.2 extend the analysis by considering several within-firm extensive margins of service offshoring as explanatory variables in equation (7.1). They examine

[19] In contrast to their approach of first-differencing, I prefer to estimate the equation using the within transformation to consider also medium-term employment effects and use the largest possible sample. A first-differenced version of equation (7.1) is considered in Section 7.4.4.

[20] The remaining coefficient estimates are not reported to save space. The last column lists the range of the number of firms per row, which varies slightly depending on data availability. These fixed effects (FE) estimations, as well as most others in this paper, are implemented using the Stata command `reghdfe` (Correia, 2014).

[21] Expenditure on domestic services is approximated by 'other operating charges' from the firm's profit and loss account in USTAN minus its service imports reported in SITS. The item 'other operating expenditures' presumably includes a large part of the expenses on both domestic and foreign purchases of services, though this can unfortunately not be verified in the available data.

Table 7.2: *Within-firm correlations of firm performance and service offshoring*

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Employment	Wage	Output	LP	Material inputs	Domestic services	# firms (min-max)
service importer	0.00348	-0.00336	0.0368***	0.00926*	0.0377***	-0.00293	30,355-41,740
	(0.00500)	(0.00293)	(0.00465)	(0.00513)	(0.00740)	(0.00604)	
# *service types*	0.0167***	0.000149	0.0191***	0.00533**	0.0150***	-0.00524*	30,355-41,740
	(0.00205)	(0.00134)	(0.00229)	(0.00242)	(0.00353)	(0.00295)	
# *countries*	0.00859***	-0.000267	0.00801***	0.00109	0.00863***	0.000369	30,355-41,740
	(0.000847)	(0.000467)	(0.000912)	(0.000946)	(0.00155)	(0.00120)	
# *country-service types*	0.00396***	0.0000779	0.00353***	0.000658	0.00358***	-0.000171	5,265-12,154
	(0.000484)	(0.000215)	(0.000545)	(0.000459)	(0.000851)	(0.000644)	
ln *service imports*	0.0254***	0.00118	0.0330***	0.0128***	0.0283***	-0.0225***	5,079-11,766
	(0.00247)	(0.00160)	(0.00266)	(0.00282)	(0.00437)	(0.00372)	
ln *narrow offshoring*	0.0246***	-0.00496	0.0362***	0.0293***	0.0632	-0.0461***	363-1,981
	(0.00648)	(0.00753)	(0.0111)	(0.0105)	(0.0422)	(0.0147)	

Note: The table reports FE estimates of equation (7.1). Each cell corresponds to one single regression, where the dependent variable is the log of the variable indicated in the header and the explanatory variable measuring service offshoring is indicated in each row. The last column reports the range of the number of firms per row. All regressions control for lagged dummy variables indicating service exporter, MNE, and FIE status, lagged ln $turnover$, as well as fixed effects by firm and by industry-year. Standard errors reported in parentheses are robust to heteroskedasticity and autocorrelation within firms. Asterisks indicate significance levels: * p<0.1, ** p<0.05, *** p<0.01. Source: RDSC of the Deutsche Bundesbank, SITS, MiDi, and USTAN, 2001-2013, own calculations.

one by one the correlations of the firm performance variables with the numbers of source countries, service types, and country-service type combinations in which positive service imports are observed. The estimates reveal that firms which start offshoring to another country, or those adding another service type or another country-service type combination to their offshoring activities, become larger in terms of employment and output and source more materials. Plausibly, these correlations are stronger at the levels of service types or countries compared to the more disaggregate country-service type combinations. The estimates suggest that, for example, a firm experiences larger performance changes if it starts offshoring IT services for the first time (to any country), compared to the situation in which it is already importing IT services from one country (e.g. India) and starts to import them additionally from another country (e.g. China). The correlations with wages and domestic service purchases are small and insignificant throughout, but there is some evidence that offshoring another service type coincides with increased labor productivity.

The intensive margin of service offshoring is examined in in the fifth row of Table 7.2, which uses ln *service imports* as an explanatory variable. Note that the requirement of positive imports in at least two years reduces the sample substantially compared to the full sample of USTAN and MiDi firms that was used to examine the firm extensive margin. The estimate reported in column 1 suggests that an increase in service imports by 10% coincides on average with a small increase in employment by around 2.5%. The estimated coefficient is insignificant in the wage regression, while output, productivity, and materials are positively correlated with the intensive margin of offshoring. Column 6 reports a negative correlation between service offshoring and expenditure on domestic services, which seems to indicate a substitute relationship.

The sixth row applies an alternative, narrow definition of offshoring by focusing on imports in the service type that corresponds to the firm's own NACE industry code (similar to the narrow measure by Feenstra and Hanson, 1999). The correspondence between service types and industries is provided in Table 7.A.1. Note that, since only firms in the service sector can have positive values of narrow

offshoring, the sample is drastically reduced. The estimated correlations with narrow offshoring are positive for employment, output, and LP and again negative for domestic services.

Due to the obvious endogeneity issues mentioned in Section 7.3.1, the OLS estimates in Table 7.2 cannot be interpreted as causal effects of offshoring. In particular, the positive estimates for employment and output may reflect a positive efficiency gain from offshoring, but they could also be driven by demand shocks or productivity shocks inducing simultaneity bias. These issues are addressed in detail in the next sections. First, Section 7.4 investigates the employment effect of increasing service offshoring (the intensive margin) through an IV. Second, Section 7.5 scrutinizes the role of effect of starting service offshoring (the firm-level extensive margin) on employment through matching methods.

7.4 Intensive Margin of Service Offshoring

7.4.1 Identification through instrumental variable

Estimates of equation (7.1) by OLS to investigate the intensive margin of service offshoring (as shown in Table 7.2) may be subject to bias from (at least) three sources. First, firm-specific productivity or output demand shocks that are not fully captured by the industry-year fixed effects can be expected to bias the estimate of β due to simultaneity. This bias can be positive if firms that get more productive or face increasing demand grow larger in terms of employment *and* find it easier to overcome the costs associated with offshoring (see e.g. Egger et al., 2016; Sethupathy, 2013). Or simultaneity bias might be negative if firms offshore services in response to adverse shocks in their output market, such as increased import competition (as argued by Monarch et al., 2017). Second, it can be expected that any firm-specific (or regional) labor supply shock biases the estimate of β downwards. Examples of labor supply shocks that may vary across firms within industries include high wage demands by firm-level unions, skills shortages, or regional outmigration. All of these shocks would simultaneously lower firm-level employment and encourage offshoring, resulting in a downward bias in β. Third, the value of imports is an imperfect proxy for the amount of service offshoring, and this measurement error leads to attenuation bias towards zero as long as it is not systematic. While the net direction and magnitude of these combined biases is a priori not clear, they all seem to be of first-order relevance.

In order to address the apparent endogeneity issues, I exploit world export supply (WES) by the firms' trading partners as an IV for the intensive margin of service offshoring. More precisely, the IV is the firm-specific mix of countries' export supply to the rest of the world by service type. Intuitively, a German firm importing a given service from a given partner country benefits disproportionately from an improvement in this country's comparative advantage, reflected in its growing exports of this service. Thus, the firm can respond to the positive supply shock by increasing its amount of service offshoring. Arguably, the service exports of foreign countries to the rest of the world have no other direct or indirect effect on firm-level employment, nor are they affected by the German firm itself or by other determinants of the firm's employment decision (conditional on the control variables, which include industry-year fixed effects). Based on this exclusion restriction, WES can be used to instrument for service imports in equation (7.1). This idea was first proposed by Hummels et al. (2014) for Danish firms' goods trade, and

I closely follow their approach in constructing the IV for German firms' service trade.[22] The availability of international data on bilateral service trade imposes a constraint on the level of disaggregation at which the IV can be constructed (see Section 7.2.1). As a result, the IV exploits variation in service exports from 179 partner countries across ten service types over twelve years.

The firm-specific and time-varying instrument $IV_{i,t}$ is computed as the following weighted sum:

$$IV_{i,t} = \sum_{s} \sum_{c} ISH_{i,c,s,0} WES_{c,s,t}, \tag{7.2}$$

where world export supply $WES_{c,s,t}$ is the export supply by country c of service type s in year t to all countries in the world except Germany.[23] The weights are firm-level import shares $ISH_{i,c,s,0}$, defined as the firm's share of imports by country and service type in its total service imports in year $t = 0$, which indicates the first year in which positive service imports by the firm are observed. The first year is 2001 for the majority of all firms in the estimation sample, but it is a later year for firms that started to import during the sample period. All of the observations from $t = 0$ form the pre-sample, which is subsequently omitted in all estimations. As demonstrated in Section 7.2.5, German firms' service importing structure is highly firm-specific and stable over time. The fact that very few firms share the same service import mix ensures that there is substantial firm-level variation in the IV, even though $WES_{c,s,t}$ itself is not firm-specific. The stability of the importing structure further allows me to hold the import shares fixed at their pre-sample values, thereby avoiding a potential endogeneity issue due to adjustments in these shares over time. Time-invariant correlates of $ISH_{i,c,s,0}$ are fully controlled for by firm fixed effects. Following Hummels et al. (2014), the first-stage estimation uses a log-log specification, regressing $\ln$ *service imports*$_{i,t}$ on $\ln IV_{i,t}$.

7.4.2 Employment effects of increasing service offshoring

Table 7.3 presents the main finding of this paper: Increased service offshoring has significantly boosted firm-level employment in Germany over the period 2002-2013. Panel A of the table reports estimates from two-stage least squares (2SLS) regressions of equation (7.1), instrumenting for service offshoring by world export supply as described in Section 7.4.1. The first column considers the full available estimation sample of firms importing services in multiple years. In this sample, the instrumented service offshoring is estimated to increase firm-level employment by an elasticity of 7.6%, which is statistically significant at the 1% level. This positive estimate provides evidence for efficiency gains from service offshoring that lead to firm-level employment gains. These gains are also economically significant. The estimate suggests that an increase in service offshoring by 10% (or approximately € 2 million at the mean) would create around eight new jobs at the average firm in this sample, which has 1,000 employees.

The positive employment effect of service offshoring is confirmed in each of two subsamples defined by the source of the employment data – MiDi (column 2, comprising MNEs and FIEs) and USTAN (column 3). The effect is similar in terms of economic magnitude and statistical significance in both

[22] The scope, structure, and level of detail in the German service trade data is comparable to that of the Danish goods trade data used by Hummels et al. (2014). Autor et al. (2013) use a similar approach at a more aggregate level, which concentrates on the impact of Chinese imports on local labor markets in the US and has been applied to German goods trade by Dauth et al. (2014). Bernard et al. (2018) adapt the IV strategy by Hummels et al. (2014) and use export shares (instead of WES) to construct their IV, arguing that this is a better measure of comparative advantage. I conduct a robustness check to show that my main findings are unchanged when the IV is constructed using a measure of revealed comparative advantage.

[23] I use disaggregate service trade data from Comtrade by country pair and service type to compute $WES_{c,s,t}$, but this variable itself is not bilateral.

subsamples. Restricting the sample to USTAN further allows me to include the following firm-level control variables that may affect employment (in column 4): a dummy variable indicating the firm's goods exporter status, the log of the physical capital stock, and the log of the real wage per employee, all lagged by one year. The estimated effect of service offshoring changes only slightly through the inclusion of these additional control variables and implies an elasticity of 6.0%. The subsequent analysis concentrates on the preferred specifications from columns 1 and 4 of Table 7.3, with the former maximizing the sample size and the latter maximizing the set of relevant control variables.

In all of the 2SLS regressions, $\ln IV$ proves to be a strong instrument for $\ln$ *service imports*, as evidenced by the first-stage regression diagnostics in panel B of Table 7.3. The instrument is a significant predictor of service offshoring in all samples, characterized by very high F statistics and a partial R^2 of 2.7-3.7%, based on the variation within firms and within industry-years. Given the persistence of the firm-specific sourcing pattern, it is not surprising that the firm-specific export supply shocks reflected in the IV explain a substantial part of the time variation in service offshoring.[24] Hence, under the assumption that partner countries' world export supply is exogenous to employment in German firms, the 2SLS estimates can be interpreted as causal effects of increasing service offshoring.

It is instructive to compare the 2SLS estimate of β in column 4 of Table 7.3 to the corresponding OLS estimate in the same sample, reported in column 5. The 2SLS estimate is substantially greater. A similar picture emerges in the other samples (not shown), suggesting a negative bias in the OLS estimate of β of around one half to two thirds. This finding may seem surprising to readers who expect a positive simultaneity bias arising from productivity or demand shocks, which would increase both offshoring capabilities and firm employment. Three considerations may reconcile the estimation results with these expectations. First, it seems plausible that firm fixed effects account well for the positive selection of large and productive firms into offshoring.

Second, and most importantly, the apparent downward bias can be rationalized by the idea that firms mainly increase their offshoring in reaction to adverse shocks, as argued by Antràs et al. (2017) and Bernard et al. (2018).[25] Such adverse shocks include increased competition and negative cost shocks (Monarch et al., 2017), in particular to labor supply. In the following, I discuss three examples of such labor supply shocks, which can be expected to have opposite effects on firm employment and the attractiveness of service offshoring: the negotiation strategies of unions, skills shortages, and regional migration. These shocks have likely been relevant in Germany and vary within industries over time. In Germany, wages (and other working conditions) are negotiated autonomously between firms or employers' associations on the one hand and unions and works councils (worker representatives within the firms) on the other hand. Dustmann et al. (2014) stress the high degree of flexibility in the German labor market after a process of decentralization in wage bargaining over the 1990s and 2000s. Many unions and works councils exploited this flexibility and agreed to restraining wage growth in exchange for explicit job guarantees in times of economic distress.[26] These strong efforts of worker representatives

[24] The strong first-stage results reported in Panel B of Table 7.3 mirror the significant power of the IV in Hummels et al. (2014) even when they consider multiple interaction terms at the firm-worker level. The high F statistics in my application accommodate any potential concerns related to weak instruments.

[25] Bernard et al. (2018) use a variant of the same IV strategy to estimate the employment effects of offshoring of manufactured inputs in Danish firms and similarly find a negative bias in the OLS estimates.

[26] This flexibility is reflected in the possibilities for firms to opt out of collective bargaining agreements or to use so-called 'opening clauses' to pay lower wages if the works councils agree (Brändle and Heinbach, 2013). Prominent examples of job guarantees include the agreements of the trade union IG Metall with Siemens (see the Financial Times article 'Siemens promises workers jobs for life' of September 22, 2010) and more recently with Opel (see the Reuters report 'Opel agrees job guarantees, investments at German sites' of May 29, 2018).

Table 7.3: *Employment effects of increasing service offshoring in 2SLS estimations*

	(1)	(2)	(3)	(4)	(5)
A. Second-stage estimates. Dependent variable: ln *employment*					
	Full sample	MiDi	USTAN		
	2SLS	2SLS	2SLS	2SLS	OLS
ln *service imports*	0.0728***	0.0659***	0.0682***	0.0578***	0.0203***
	(0.0161)	(0.0182)	(0.0184)	(0.0164)	(0.00362)
lag *service exporter*	-0.00588	-0.00439	-0.00540	-0.00652	0.00305
	(0.0102)	(0.0118)	(0.0130)	(0.0117)	(0.0113)
lag *MNE*	0.0283**	0.0343**	0.0162	0.0240*	0.0241**
	(0.0125)	(0.0160)	(0.0128)	(0.0122)	(0.0121)
lag *FIE*	-0.0425**	-0.0356*	-0.0201	-0.0176	-0.0175
	(0.0182)	(0.0210)	(0.0194)	(0.0174)	(0.0172)
lag ln *turnover*	0.268***	0.275***	0.202***	0.256***	0.273***
	(0.0160)	(0.0173)	(0.0195)	(0.0252)	(0.0253)
lag *goods exporter*				-0.00289	-0.000540
				(0.00803)	(0.00792)
lag ln *capital*				0.152***	0.153***
				(0.0155)	(0.0157)
lag ln *wage*				-0.307***	-0.310***
				(0.0395)	(0.0402)
Observations	40,350	32,587	16,288	14,557	14,557
Firms	7,018	5,653	2,865	2,644	2,644
B. First-stage estimates. Dependent variable: ln *service imports*					
ln *IV*	0.149***	0.147***	0.158***	0.152***	
	(0.00769)	(0.00864)	(0.0119)	(0.0126)	
F statistic (excl. IV)	377.3	291.3	176.5	146.0	
p-value of F test	0.0000	0.0000	0.0000	0.0000	
Partial R^2 (excl. IV)	0.0291	0.0274	0.0366	0.0352	

Note: The table reports estimates of equation (7.1) by 2SLS (columns 1-4) and OLS (column 5). The top panel reports second-stage estimates with ln *employment* as the dependent variable. All regressions control for fixed effects by firm and by industry-year. The bottom panel reports results from the first-stage regressions with ln *imports* as the dependent variable. Standard errors reported in parentheses are robust to heteroskedasticity and autocorrelation within firms. Asterisks indicate significance levels: * p<0.1, ** p<0.05, *** p<0.01. Source: RDSC of the Deutsche Bundesbank, SITS, MiDi, and USTAN, 2001-2013, own calculations.

to secure domestic jobs have likely slowed down offshoring activities in the affected firms, compared to other firms, in which unions negotiated higher wages. The consequence is a negative correlation between employment and service offshoring. Another differential labor supply shock faced by German firms are skills shortages. Since the 2000s, firms have been struggling to fill vacancies in occupations such as engineering, IT, or health services.[27] This scarcity of qualified employees is particularly pronounced in certain industries and regions, where firms cannot increase their employment to the desired level and are therefore under pressure to find suitable service providers abroad. Similarly, a regional migration outflow, as observed in many East German regions during this period, simultaneously reduces employment

[27] The German government has set up a website listing these skills in short supply and seeking to attract foreign talent (see https://www.make-it-in-germany.com/en/for-qualified-professionals/working/demanded-professions).

and increases the need for offshoring. Immigration has the opposite effects. The negative correlation between service offshoring and employment implied by these and other labor supply shocks, which vary across firms within industries, biases the OLS estimate of β downwards.

Third, given that the true employment effect is positive, any non-systematic measurement error in service offshoring induces attenuation bias that contributes to the downward bias. Note that firm-level service import values are subject to a variety of influences, including bilateral exchange rates, firms' relative bargaining power, and other idiosyncrasies. By contrast, the aggregate exports of countries by service types, used to construct the IV, are more likely to capture fundamental changes in countries' comparative advantage in these service types, and hence reflect real changes in offshoring opportunities. The IV strategy exploits only the variation in service imports that are driven by these fundamental changes, and may hence mitigate the downward bias due to measurement error. However, it is likely that the IV cannot fully eliminate the third issue and there might be attenuation bias remaining in the 2SLS estimates. Overall the downward bias seems to dominate in the OLS regression, resulting in a larger 2SLS estimate for β.

The positive employment effects reported in Table 7.3 extend and strengthen previous findings by Amiti and Wei (2005) and Hijzen et al. (2011) for the UK. In particular, Amiti and Wei (2005) uncover mostly positive correlations between service offshoring and employment at the industry level, and Hijzen et al. (2011) show that this positive correlation is remarkably robust at the firm level. The new results presented in Table 7.3 show that the employment effect of service offshoring is positive and significant in German firms when IV techniques are applied to resolve the obvious endogeneity issues. My findings indicate that the positive correlation of service offshoring and firm employment that is also found in Germany (see Table 7.2) is not driven by confounding factors, but instead, it even seems to understate the positive causal effect.

As discussed in the introduction, the theoretical literature has shown how the cost savings from offshoring can increase domestic employment through the mechanism described as a productivity effect by Grossman and Rossi-Hansberg (2008). But when can the productivity effect of offshoring be expected to dominate the direct negative effect of relocation and generate net employment gains, as shown in Table 7.3? In their multi-country sourcing model, Antràs et al. (2017) derive explicit conditions under which a positive shock to offshoring conditions in one country can increase the firm's use of inputs from other countries, including the domestic economy (see their Proposition 3). This case emerges if the firm faces a high enough demand elasticity and international production cost differences are pronounced, such that there is a great potential for cost savings. These differences are likely to be substantial for the tradable commercial services considered in this paper. Note that the model by Antràs et al. (2017) does not determine to what extent the increased factor use arises within the firm, through increased employment, or outside the firm, through increased domestic sourcing; I will return to this question in Section 7.4.3. In Egger et al. (2016), the key variable that determines the net employment effect is the share of previously offshored tasks, since the productivity effect works through these intra-marginal tasks. To investigate this further, the employment effects are differentiated by lagged service offshoring in Section 7.4.4. In Groizard et al. (2014), the net employment effect of offshoring crucially depends on the complementarity between foreign and domestic activities: A positive effect is more likely if this complementarity is strong. Section 7.4.5 approaches this question by splitting offshoring into services sourced from high-income and low-income countries, of which the latter are presumably less similar, and hence more complementary to activities done within German firms.

7.4.3 Effects on other firm outcomes

To better understand the changes that service offshoring triggers within the firm, this section investigates its impact on other firm outcomes. It considers the wage per employee, output, and LP, as well as the firm's expenditures on materials and on domestic services (all in real terms and in logs) as alternative dependent variables in equation (7.1), estimated with and without the additional firm-level control variables from column 4 of Table 7.3. Table 7.4 summarizes the results of 2SLS estimations, in which $\ln$ *service imports* is instrumented by $\ln IV$ throughout.

How does service offshoring affect wages? In fact, a substantial part of the offshoring literature has concentrated on wage effects. On the one hand, one would expect the direct relocation effect to depress wages because it frees up domestic labor (Grossman and Rossi-Hansberg, 2008). On the other hand, the average wage within the firm may be positively affected by the productivity effect, rent-sharing mechanisms (Sethupathy, 2013), and compositional effects in the labor force.[28] The 2SLS wage regressions in the first two columns of Table 7.4 find none of these effects dominating, as the estimated coefficient is close to zero and insignificant in both specifications. The data do not allow for rejecting the hypothesis of a zero effect of increased service offshoring on wages in German firms. This ambiguous result is not surprising in light of the different countervailing forces just discussed. In combination with the employment gains from Table 7.3, these findings suggest that service offshoring indeed benefits the overall domestic workforce within the firm, which experiences increased employment without wage cuts on average.

The analysis proceeds by investigating the output effect of service offshoring. A positive output effect is an important precondition for positive employment effects. The firm can be expected to hire more domestic labor only if the efficiency gains from service offshoring induce it to increase its optimal scale of production. Hence, if the productivity effect is responsible for the positive employment effect identified above, there should also be direct evidence for a positive impact of service offshoring on firm-level output. The estimates reported in columns 3 and 4 of Table 7.4 confirm this expectation. Service offshoring significantly increases firm-level output, and the estimated output gains are substantial, implying an elasticity of 9.5-12.2%.[29] To account for the fact that employment and output are jointly driven by service offshoring, I have also experimented with estimating the equations for employment and output jointly in three-stage least squares regressions, similar to the approach developed by Wright (2014) in the context of US industry-level offshoring. These (unreported) estimates confirm the positive effects on both employment and output.

Table 7.4: *Effects on wages, output, productivity, and sourcing of other inputs*

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Wage		Output		LP		Material inputs		Domestic services	
$\ln$ *service imports*	-0.00855	0.00318	0.115***	0.0907***	0.0440	0.0416*	0.0726**	0.0851***	0.0110	0.0106
	(0.00997)	(0.00866)	(0.0248)	(0.0201)	(0.0274)	(0.0220)	(0.0328)	(0.0300)	(0.0277)	(0.0274)
Additional controls	no	yes	no	yes	no	yes	no	yes	no	yes
Observations	16,365	13,998	16,607	14,016	16,323	13,940	15,860	13,638	16,198	13,738
F statistic (excl. IV)	179.1	141.8	181.7	141.6	180.8	140.7	166.8	138.6	176.3	140.7

Note: The table reports 2SLS estimates of equation (7.1), where the dependent variable is the logarithm of the variable indicated in the header. All regressions control for the basic set of lagged firm control variables and fixed effects by firm and by industry-year. The even columns further account for the additional firm-level control variables from column 4 of Table 7.3. Standard errors reported in parentheses are robust to heteroskedasticity and autocorrelation within firms. Asterisks indicate significance levels: * p<0.1, ** p<0.05, *** p<0.01. Source: RDSC of the Deutsche Bundesbank, SITS, MiDi, and USTAN, 2001-2013, own calculations.

[28] Service offshoring has been found to increase the relative demand for high-skilled labor (see Crinò, 2010a,b, 2012).

[29] The possibility cannot be excluded that a part of this effect may work through an increase in goods prices, since output is deflated by an industry-level price index and data on firm-level prices are unavailable.

Columns 5-6 of Table 7.4 examine the effects of offshoring on labor productivity $\ln LP$. I find positive point estimates of around 4.2-4.5%, which are however only weakly significant in the specification with additional controls. These estimates provide suggestive evidence that service offshoring increases output by more than employment, in line with the magnitudes of the estimated employment and output effects, and in line with the intuition of the productivity effect. These novel, firm-level results extend the existing evidence on the positive association between service offshoring and productivity that has been found in industry-level data by Amiti and Wei (2009b) for the US and by Crinò (2008) for Western Europe.

Service offshoring also increases the use of other inputs, in line with the theory and empirical evidence in Antràs et al. (2017). The last four columns of Table 7.4 specifically examine the effects of service offshoring on the firm's sourcing of material inputs (columns 7 and 8) and domestic services (columns 9 and 10). The regressions reveal strong positive effects on material inputs purchases and a small positive but insignificant effect on domestically procured services. Comparing these results to the OLS estimates in Table 7.2, it can be inferred that the negative correlation between service offshoring and expenditure on domestic services identified above, which seems to suggest a substitutive relationship, does not reflect a causal negative effect. Instead, service offshoring may even weakly increase the sourcing of domestic services. These findings have important implications for the economy-wide effects of service offshoring. Note that in principle, the main analysis in this paper is restricted to employment effects within the offshoring firms. The non-negative effects on sourcing of material inputs and domestic services suggest that employment of domestic input factors outside of the offshoring firms have also benefited from service offshoring.

7.4.4 Heterogeneous employment effects by lagged offshoring

Table 7.5 shows interesting patterns of heterogeneity in the employment effects across firms, supporting the view that the cost savings from offshoring are responsible for the firm-level employment growth. The employment effects are expected to increase in the initial scale of offshoring because the productivity effect works through the infra-marginal units that have previously been offshored. To see this, note that there will be a negligible productivity effect (but only a negative relocation effect) on employment in a firm that starts to offshore the first job. By contrast, a firm that has previously offshored some service activities experiences the beneficial productivity effect from further reductions in offshoring costs in all infra-marginal units that have already been offshored. Hence, a positive shock to offshoring should have more favorable employment effects in firms that have already offshored more service. This prediction follows intuitively from the original trade in tasks model by Grossman and Rossi-Hansberg (2008), and it is formally derived in recent extensions of their model with heterogeneous firms by Egger et al. (2015, 2016).[30]

To assess how the employment effects vary with the level of previous service offshoring, the lagged level of service imports needs to be taken into account. Since lagged service imports and their unobserved determinants are likely to be correlated with contemporaneous offshoring and employment, introducing this variable into the fixed effects model may give rise to bias. To avoid this issue, the model is estimated in changes over time. To be precise, I specify the following first-differenced (FD) version of

[30]It should be noted that in the existing theory, the share of previously offshored tasks is the same for all offshoring firms (Egger et al., 2016), while in the data lagged offshoring (and its composition) varies considerably across similar firms within an industry.

equation (7.1):

$$\Delta \ln \textit{employment}_{i,t} = \gamma \cdot \Delta \ln \textit{service imports}_{i,t} + \boldsymbol{\psi} \cdot \boldsymbol{\Delta} \mathbf{X}_{i,t-1} + \delta_{j,t} + \nu_{i,t}. \tag{7.3}$$

In this equation, Δ denotes the FD operator (so $\Delta Y \equiv Y_{i,t} - Y_{i,t-1}$ is the annual change in variable Y), which implies that firm-specific effects are differenced out. The parameters γ and $\boldsymbol{\psi}$ need to be estimated, $\delta_{j,t}$ are industry-year fixed effects, and $\nu_{i,t}$ is the error term. The model is estimated by 2SLS, instrumenting for $\Delta \ln \textit{service imports}_{i,t}$ by $\Delta \ln IV_{i,t}$.

Columns 1-2 of Table 7.5 report the results of estimating the FD model in equation (7.3) for the two specifications with and without the additional firm-level control variables from column 4 of Table 7.3. The estimated effect of service offshoring on employment is positive and highly significant, and the strength of the IV is confirmed in the FD model. The point estimates of γ are smaller compared to the main FE estimates of β, especially in the larger sample. This seems plausible, provided that some of the employment gains from service offshoring take more than one year to materialize, and are hence not captured by the FD model.[31]

Table 7.5: *Heterogeneous employment effects by lagged offshoring*

	(1)	(2)	(3)	(4)	(5)	(6)
	First differencing		By log of lagged service imports		By high vs. low lagged service imports	
$\Delta \ln$ *service imports*	0.0479***	0.0557***	-0.118**	-0.0793	0.0145	0.0353***
	(0.0156)	(0.0157)	(0.0594)	(0.0568)	(0.0152)	(0.0124)
lagged service imports			-0.00118	-0.000398	0.00730	0.00564
			(0.00173)	(0.00131)	(0.00587)	(0.00653)
$\Delta \ln$ *service imports*			0.0254***	0.0228**	0.106***	0.105*
$\times$ lagged service imports			(0.00966)	(0.0112)	(0.0408)	(0.0571)
Additional controls	no	yes	no	yes	no	yes
Observations	31,873	11,296	31,873	11,296	31,873	11,296
F statistic (Kleibergen-Paap)	274.7	95.15	39.96	11.65	49.90	11.77

Note: The table reports 2SLS estimates of equation (7.3) with $\Delta \ln$ *employment* as the dependent variable, including additional interaction terms as regressors. Columns 1-2 report estimates of the simple FD model, columns 3-4 add the log of lagged of service imports ($\ln$ *service imports*$_{i,t-1}$) and its interaction with $\Delta \ln$ *service imports*$_{i,t}$, and columns 5-6 add a dummy variable indicating a level of lagged service imports above the median and its interaction with $\Delta \ln$ *service imports*$_{i,t}$. All regressions control for the basic set of lagged firm control variables (in first differences) and fixed effects by industry-year. The even columns further account for the additional firm-level control variables from column 4 of Table 7.3 (also in first differences). Standard errors reported in parentheses are robust to heteroskedasticity and autocorrelation within firms. Asterisks indicate significance levels: * p<0.1, ** p<0.05, *** p<0.01. Source: RDSC of the Deutsche Bundesbank, SITS, MiDi, and USTAN, 2001-2013, own calculations.

The table then provides a test of the aforementioned prediction by augmenting equation (7.3) with interaction terms of the change in service offshoring with its lagged level. In columns 3-4, the continuous measure of lagged offshoring $\ln$ *service imports*$_{i,t-1}$ and the interaction term $\Delta \ln$ *service imports*$_{i,t} \times \ln$ *service imports*$_{i,t-1}$ are added to the FD model. Excluded instruments are $\Delta \ln IV$ and the interaction term $\Delta \ln IV_{i,t} \times \ln$ *service imports*$_{i,t-1}$. In both specifications, the coefficient of the interaction term is estimated to be positive and statistically significant at the 5% level. Furthermore, the first-order effect of the change in service offshoring is estimated to be negative (and significant in the

[31] Note that the within-transformed (FE) model is preferred precisely because the IV has been designed to capture medium-to-long-term changes in service offshoring (see the discussion in Hummels et al., 2014).

specification without additional controls). These estimates suggest that the employment effects of service offshoring are negative at low initial levels of imports, and they increase in the value of services that have already been offshored. This is exactly the pattern expected based on the productivity effect of offshoring from the trade in tasks model. A higher initial level of offshoring implies greater cost savings and hence more beneficial employment effects from a positive shock to foreign service exports, as captured by the IV. The estimates in columns 3-4 of Table 7.5 provide strong support for this theoretical prediction.

Evaluating the marginal effects at different percentiles of the distribution of lagged service offshoring allows for a more detailed assessment of the heterogeneous employment effects. In the larger estimation sample (column 3), the estimates imply a negative employment effect of -1.4% for a firm at the fifth percentile of this distribution. The marginal effect is strongly increasing in lagged offshoring: It turns positive around the tenth percentile, equals 3.5% at the 25th percentile, and reaches 10.4% at the 75th percentile (three times the effect at the 25th percentile). The estimates in the sample with additional controls (column 4) suggest even more favorable employment effects, which are positive already for firms at the fifth percentile of the distribution of lagged service imports and exceed 11.4% at the 75th percentile.[32]

In the final set of regressions in Table 7.5, reported in columns 5-6, I employ as an alternative measure of previous offshoring a dummy variable indicating a level of lagged service imports above the median. This more flexible specification confirms the main insight from above: Firms with high levels of previous service imports experienced significantly more favorable employment effects, as demonstrated by the positive interaction effect. While the impact of service offshoring on firms with lagged import levels below the median are small and only significant in the second specification, firms with a high level of previous offshoring experienced employment gains of 12.8-15.1%, more than ten percentage points more.

Note that the 2SLS estimates in columns 3-6 each require two first-stage regressions. To assess the strength of the IV approach, the table reports the Wald rk F statistics for the cluster-robust rank test proposed by Kleibergen and Paap (2006). In the larger sample (columns 3 and 5), the F statistics seem reassuringly high compared to conventional thresholds for weak IV tests (see Stock and Yogo, 2005). For the smaller sample, the F statistics lie in the critical range, so the estimates should be interpreted with caution.

7.4.5 Offshoring to high-income vs. low-income countries

This section investigates differences in the employment and wage effects of service offshoring to different countries. Since the existing literature has found the employment and wage effects of offshoring in the US to differ between high-income and low-income source countries (Harrison and McMillan, 2011; Ebenstein et al., 2014), imports are broadly split up into two categories. Countries are defined to be 'high-income' countries if they have a GDP per capita above $ 38,500 in values of 2010 in all of the years 2001-2013 and 'low-income' otherwise. This income threshold corresponds approximately to the German real GDP in 2001 and to the median in the full SITS sample. The data reveal that high-income countries are the most important offshoring destinations for German firms, as they account for approximately two thirds of aggregate imports of other commercial services over the sample period (see

[32] These marginal effects are computed as $1 - \exp(\hat{\gamma} + \hat{\rho} \cdot \ln \textit{service imports}_{i,t-1})$, where $\hat{\rho}$ is the estimated coefficient of the interaction term, and the effect is evaluated at different percentiles of $\ln \textit{service imports}_{i,t-1}$ in the respective estimation sample.

also Section 7.2.2). This observation is not surprising given the high skill abundance of the high-income countries surrounding Germany and the fact that the gravity equation provides a good description of trade in services (Kimura and Lee, 2006). Since the other commercial services examined in this paper often involve high-skill intensive tasks, this analysis may reveal interesting differences compared to offshoring of manufactured inputs, which are typically thought of as low-skill intensive.

Table 7.6 reports the results of estimating equation (7.1) separately for service imports from high-income countries (columns 1-2) and from low-income countries (columns 3-4), and for both types of offshoring combined (columns 5-6). The endogenous service import variables are instrumented by two separate IVs for high-income and low-income imports, respectively, each computed according to equation (7.2), but including only the exporting countries to which the respective definition of income groups applies.

The 2SLS estimates reported in panel A identify positive employment effects of service offshoring to both country groups in columns 1-4. A comparison of the point estimates suggests that offshoring to low-income countries benefits domestic employment in German firms more strongly. This difference is confirmed in the specifications including both high-income and low-income country imports simultaneously, though the estimates are insignificant and the IVs are weak in the regression reported in column 6.

Table 7.6: *Offshoring to high-income vs. low-income countries*

	(1)	(2)	(3)	(4)	(5)	(6)
A. Dependent variable: ln *employment*						
	High-income countries		Low-income countries		High vs. low income	
ln *service imports* (high-income)	0.0753***	0.0421***			0.0191	-0.0345
	(0.0179)	(0.0152)			(0.0371)	(0.0437)
ln *service imports* (low-income)			0.0984***	0.0748**	0.131***	0.106
			(0.0262)	(0.0326)	(0.0506)	(0.0657)
Additional controls	no	yes	no	yes	no	yes
Observations	31,629	10,934	24,077	9,204	15,577	5,647
F statistic (Kleibergen-Paap)	279.7	136.1	142.5	56.39	23.22	6.660
B. Dependent variable: ln *wage*						
	High-income countries		Low-income countries		High vs. low income	
ln *service imports* (high-income)	-0.0184*	-0.0120			-0.135	-0.161
	(0.0112)	(0.0107)			(0.122)	(0.186)
ln *service imports* (low-income)			0.0712*	0.0593	0.206	0.245
			(0.0373)	(0.0372)	(0.167)	(0.263)
Observations	13,741	11,850	6,160	5,263	4,285	3,700
F statistic (Kleibergen-Paap)	110.0	84.84	21.69	14.60	1.137	0.501

Note: The table reports 2SLS estimates of equation (7.1) with ln *employment* as the dependent variable in panel A and ln *wage* as the dependent variable in panel B. All regressions control for the basic set of lagged firm control variables and fixed effects by firm and by industry-year. The even columns further account for the additional firm-level control variables from column 4 of Table 7.3 (but excluding the lagged wage in panel B). High-income countries are defined by a real GDP per capita above $ 38,500 in values of 2010 in all of the years 2001-2013; all others are low-income countries. Standard errors reported in parentheses are robust to heteroskedasticity and autocorrelation within firms. Asterisks indicate significance levels: * p<0.1, ** p<0.05, *** p<0.01. Source: RDSC of the Deutsche Bundesbank, SITS, MiDi, and USTAN, 2001-2013, own calculations.

Panel B investigates the impact on wages, again differentiating service offshoring by source country. The signs of the estimated wage effects tell a clear story, though these estimates are only weakly significant (and the IVs are weak in the combined estimations in columns 5-6): Service offshoring to

high-income countries tends to depress the wages paid to domestic workers, while offshoring to low-income countries has a positive effect on domestic wages.

Taken together, the results shown in Table 7.6 suggest that service offshoring to low-income countries, though less prevalent among German firms, entails more favorable effects for domestic workers than offshoring to high-income countries. The latter may even lead to wage cuts for domestic workers. These findings can be rationalized by a higher complementarity between German firms' domestic workforce and their service imports from low-income countries as opposed to high-income countries. They are in line with the expectation that relatively low skill-intensive commercial services imported from low-income countries complement high-skilled employment in German firms (Groizard et al., 2014). Service imports from high-income countries are more likely to be substitutes for domestic employment. These conclusions contrast with those obtained for offshoring by US MNEs in Harrison and McMillan (2011) and Ebenstein et al. (2014), who find that offshoring to high-income countries complements, while offshoring to low-income countries substitutes for domestic employment in US MNEs. Presumably, their analysis involves mainly the offshoring of manufacturing activities. The evidence presented here suggests the opposite substitution pattern and generally more favorable employment effects for service offshoring by German firms.

7.4.6 Deeper investigations of employment effects

This section provides deeper investigations into three interesting aspects of the employment effects at the intensive margin of service offshoring by further exploiting the richness of the combined micro data. First, to shed some light on the role of supplier substitution in the domestic economy, service imports are considered relative to expenditure on both domestic and foreign services. Second, the employment effects are distinguished between firms in the services and non-services sectors. Third, data on FDI links are exploited to distinguish the effects of potential intra-firm service offshoring and offshore outsourcing.

In principle, an increase in service imports may reflect one of three cases: It may either (i) substitute for activities previously done within the firm, or (ii) substitute for activities previously purchased from other domestic suppliers, or (iii) constitute the purchase of a service that the firm did not use at all before. Since most firms' sourcing requirements are rather stable over time, as indicated by the persistent international sourcing patterns discussed in Section 7.2.5, the following argument abstracts from the third possibility. Among the other two cases, the first corresponds to a widespread understanding of service offshoring both in the general public and among economists, and may be called 'genuine offshoring'. The second case does not accord well with this common definition of offshoring, but it may instead be labeled 'supplier substitution'. Since supplier substitution does not involve any relocation of tasks previously conducted in-house, there is no reason to believe that it should entail any negative employment effects within the firm, but instead jobs may be lost in other domestic firms. Supplier substitution may however have beneficial productivity effects within the firm similar to genuine offshoring, provided that the same services are provided from abroad at a lower cost. Hence, the possibility that service imports reflect supplier substitution rather than genuine offshoring implies a risk of overstating the employment effects of genuine service offshoring, which is the main subject of this paper. While supplier substitution may be of interest in itself, the available data do not allow for a comprehensive assessment of its employment effects, since domestic service providers cannot be identified. However, data on firms' expenditure on other services can be exploited to gauge the relevance of supplier substitution for the employment effects of service offshoring identified above.

Table 7.7 provides indicative evidence that supplier substitution is not the driving force behind the positive employment effects. The regressions reported in panel A split the sample into firms that experienced a contemporaneous decrease (columns 1-2) or an increase (columns 3-4) in their *service import share*, defined as the ratio of service imports over 'other operating charges', which reflects expenditure on both domestic and foreign services (see footnote 21). Note that genuine offshoring should increase the import share, while pure supplier substitution should increase this share by even more, as it changes only the composition but not the total amount of service purchases (approximated by the denominator of the import share). Hence, while the import share may change for a variety of other reasons, firms that experience a decrease in this share should be less likely to import services that they purchased domestically before, but more likely to engage in genuine offshoring. The estimates reported in columns 1 and 2 of panel A confirm that service offshoring had a significantly positive effect on employment in these firms. Columns 3 and 4 further reveal that higher service imports also increased employment in firms with a contemporaneous increase in their import shares, and as might be expected, the effect is slightly stronger in these firms (though not significantly different). These regressions demonstrate that service offshoring has boosted firm employment regardless of contemporaneous changes in domestic sourcing activities. The last two columns in panel A employ $\ln$ *service import share* as an explanatory variable (instrumented by $\ln IV$) in equation (7.1) and confirm the positive employment effects also for this relative measure of service offshoring.

In panel B of Table 7.7, firms' sectoral affiliations are used to shed further light on the employment effects. Columns 1-2 start by applying a narrow definition of offshoring along the lines of Feenstra and Hanson (1999). As in the OLS regressions of Section 7.3.2, narrow offshoring is defined as imports in the service type corresponding to the firm's own industry. For instance, imports of IT services by a software firm qualify as narrow offshoring, but imports of IT services by a consultancy or a manufacturer do not. The idea behind this definition is that imports within the firm's own industry are more likely to represent activities that could be done by the firm itself, and hence reflect more narrowly the common understanding of genuine offshoring (rather than supplier substitution). This approach reduces the estimation sample to service sector firms which import in several years the same type of service as they provide. Interestingly, the estimated elasticity is even much higher for narrow offshoring (instrumented by $\ln IV$) compared to the broad definition in the largest available sample (column 1). The estimate is rendered insignificant and the IV is weak in the specification with additional firm control variables (column 2); however, the point estimate remains positive and large also in this small sample of only 212 firms. The larger estimates may not be too surprising given that services are highly differentiated within the ten service types and firms are likely to offshore non-core tasks, such as back-office activities, which may well be complementary to the core tasks of their domestic workforce. The fact that the point estimates are substantially greater than for broad offshoring can have various explanations, but it certainly does not suggest that the employment effects in the main analysis are overstated due to supplier substitution.

In the remaining columns of panel B, the estimation samples are split into firms in the service sector (NACE codes $\geq$45, columns 3 and 4) and those in the primary and industrial sectors (defined by NACE codes $<$45, columns 1 and 2). As described in Section 7.2.4, firms across all sectors import other commercial services, but the offshoring activities of service sector firms may differ in important ways from those of other firms. Should we expect the employment effects of service offshoring to be more or less favorable in the service sector compared to other sectors? The answer to this question is a priori not clear. On the one hand, firms in the service sector are more service intensive in their sourcing

activities. Input-output tables from the German Statistical Office show that other commercial services inputs accounted for 44.2% of the total value of inputs purchased by the service sector in 2013, compared to only 11.7% for the primary and industrial sectors. Since services account for a larger share of the total costs of service sector firms, there is a greater cost savings potential from service offshoring, which can boost domestic employment. On the other hand, it seems more likely that the service imports by service sector firms substitute for activities that could in principle be done within these firms, which by definition produce services as their own output. As noted before, this type of offshoring may result in less favorable employment effects. The estimation results in the split samples reveal greater employment effects for service sector firms compared to the primary and industrial sectors in both specifications.[33] These differences across sectors may suggest that the benefits of greater cost savings dominate the threat of greater substitutability of employment in the service sector.

To what extent is service offshoring conducted via intra-firm service trade within multinational firms' boundaries or through offshore outsourcing? The MiDi data, covering all of German inward and outward FDI (above a low reporting threshold), can be exploited to provide a tentative answer to this question. These data also serve to relate the findings in this paper to the literature that has exclusively focused on the foreign activities of MNEs for measuring offshoring. To approach this question, I apply two definitions of intra-firm service offshoring. First, I classify firms' service imports from all countries and industries in which they also have a foreign affiliate or parent (according to MiDi) as intra-firm service offshoring. The industries corresponding to the imported service types are assigned according to Table 7.A.1. 25.6% of the value of German firms' service imports come from such countries and industries and may therefore be intra-firm. Second, I drop the industry requirement and include in the definition all service imports from a country where the firm reports any foreign investment link, to allow for the possibility that a related party active in a different industry also provides services. Both definitions are likely to overstate the amount of actual intra-firm offshoring. The second definition furthermore provides an upper bound for the relevance of this phenomenon, since the existence of a related party in the source country of service imports is clearly a necessary condition for intra-firm trade. 1,095 firms in the full estimation sample satisfy this condition, and the implied upper bound for intra-firm service offshoring is approximately 34.5% of aggregate imports.

These numbers show that the vast majority of service offshoring by German firms takes place via outsourcing, while intra-firm trade most likely makes up a small share in German service imports (certainly less than one third). Hence, datasets restricted only to MNEs seem useful only to a limited extent for analyzing service offshoring. Three comments on these numbers are in order. First, service imports via commercial presence (GATS mode 3) are not captured in the SITS, as they involve domestic transactions between the foreign affiliate and a service provider within the foreign country. Second, there may be measurement error in intra-firm trade, as the values of these transactions are based on transfer prices. Third, a large share of German service offshoring is to EU countries, where regulation is similar and services can presumably be outsourced easily within the single market without setting up a foreign affiliate, which might explain the relatively low share of intra-firm offshoring.

Panel C of Table 7.7 reports the estimated employment effects of (potential) intra-firm service offshoring and outsourcing. It reports 2SLS estimates of regressions including the two alternative definitions of intra-firm service offshoring (instrumented by $\ln IV$) in equation (7.1). Whether restricting the defini-

[33] This observation is in line with the differences across sectors found by Hijzen et al. (2011) in UK firm-level data. The positive employment effects in manufacturing firms complement existing evidence that service trade liberalization can boost the productivity of manufacturing firms (Arnold et al., 2011) and industries (Beverelli et al., 2017) using these services.

tion to countries and service types in which related parties are present (columns 1 and 2) or more broadly to FDI partner countries (columns 3 and 4), the evidence is mixed. In the larger sample, the employment effects are estimated to be positive and large, but at most weakly significant. In the smaller sample, the estimates are close to zero and insignificant. Since the IV turns out to be weak, these estimates do not allow for final conclusions on the impact of intra-firm service offshoring. The last two columns in panel C conclude this discussion by considering offshore outsourcing. These regressions include only service

Table 7.7: *Deeper investigations of employment effects*

	(1)	(2)	(3)	(4)	(5)	(6)
A. Service import share						
	Decreasing import share		Increasing import share		Continuous import share	
ln *service imports*	0.0656***	0.0482***	0.0775***	0.0684***		
	(0.0201)	(0.0176)	(0.0298)	(0.0256)		
ln *service import share*					0.0674***	0.0497***
					(0.0194)	(0.0168)
Additional controls	no	yes	no	yes	no	yes
Observations	6,821	5,932	7,650	6,668	16,404	13,991
F statistic (excl. IV)	123.7	97.28	69.20	58.34	162.3	130.0
B. Narrow offshoring and sector split						
	Narrow offshoring		Service sector		Non-service sectors	
ln *narrow offshoring*	0.177***	0.0769				
	(0.0618)	(0.120)				
ln *service imports*			0.0761**	0.0676*	0.0518***	0.0361**
			(0.0339)	(0.0410)	(0.0133)	(0.0141)
Additional controls	no	yes	no	yes	no	yes
Observations	6,074	944	16,602	3,778	23,501	10,697
F statistic (excl. IV)	25.67	4.857	107.1	35.65	266.1	108.7
C. Intra-firm imports and outsourcing						
	FDI by country-industry		FDI by country		Offshore outsourcing	
ln *intra-firm offshoring*	0.132	-0.00945	0.129*	0.0157		
	(0.0839)	(0.0663)	(0.0675)	(0.107)		
ln *offshore outsourcing*					0.0766***	0.0577***
					(0.0168)	(0.0166)
Additional controls	no	yes	no	yes	no	yes
Observations	5,250	1,009	5,518	1,063	39,901	14,532
F statistic (excl. IV)	10.35	12.40	18.27	3.509	312.0	144.3

Note: The table reports 2SLS estimates of equation (7.1) with ln *employment* as the dependent variable. All regressions control for the basic set of lagged firm control variables and fixed effects by firm and by industry-year. The even columns further account for the additional firm-level control variables from column 4 of Table 7.3. In panel A, columns 1-2 restrict the sample to firms with a decreasing import share, columns 3 and 4 restrict it to firms with an increasing import share, and columns 5-6 consider ln *service import share* as an explanatory variable. In panel B, columns 1-2 consider narrow offshoring of services corresponding to the firm's own industry, columns 3-4 restrict the sample to the service sector (NACE codes $\geq$45), and columns 5-6 restrict it to non-service sectors (NACE codes $<$45). In panel C, columns 1-2 consider intra-firm service offshoring defined as imports from a country-service type combination where the firm has an FDI link, columns 3-4 consider intra-firm service offshoring more broadly defined as imports from any country where the firm has an FDI link, and columns 5-6 examine offshore outsourcing (imports from countries without an FDI link). Standard errors reported in parentheses are robust to heteroskedasticity and autocorrelation within firms. Asterisks indicate significance levels: * p$<$0.1, ** p$<$0.05, *** p$<$0.01. Source: RDSC of the Deutsche Bundesbank, SITS, MiDi, and USTAN, 2001-2013, own calculations.

imports from countries where the firm does not have a related party (again instrumented by $\ln IV$). Offshore outsourcing has strong positive effects on firm-level employment of the same magnitude as total service offshoring.

7.4.7 Robustness checks

This section discusses the robustness of the main findings to various modifications in the analysis. To economize on space, only the key insights are summarized here, while the estimates and a more detailed discussion are relegated to the Online Appendix.

The positive employment effects prove robust to modifications in the construction of the IV that address several potential challenges to identification. The first modified IV uses the revealed comparative advantage (RCA) as a relative measure of export performance instead of WES in equation (7.2), since the levels of WES may be correlated across countries. The second modified IV restricts the set of destination countries for partner countries' export supply in equation (7.2) to eight non-neighboring, non-Eurozone countries that are comparable to Germany in terms of income per capita but not subject to the same demand shocks (following Dauth et al., 2014). The estimates provide further support for the employment gains from service offshoring.

I also explore the issue of measurement error in the IV due to patchy aggregate data on service trade (see e.g. Francois and Hoekman, 2010) by using only data on the country-service type combinations for which data from Comtrade are available in all of the years 2001-2013. The estimates using this modified IV yield even larger point estimates for the employment effect of service offshoring compared to the main analysis.

Interestingly, an industry-level IV, constructed using initial import shares at the industry level instead of the firm level, turns out to be weak. This finding adds to the evidence from Section 7.2.5 suggesting that the firm-specificity of the sourcing structure is a crucial feature of service offshoring.

To eliminate the possibility that very large German firms affect service trade between other countries, all firms are excluded from the estimation sample if their service imports account for more than 0.1% of the partner country's exports to the rest of the world in a given service type and year (according to Comtrade). The main findings are fully robust to excluding these few large importers.

Unobservable determinants of employment growth have been controlled for throughout the main analysis by industry-year fixed effects and within-transforming (or first-differencing) the panel data. It is conceivable, though, that larger firms, or those located in particularly dynamic regions, might have inherently higher employment growth rates and might incidentally offshore services to countries that subsequently experience particularly positive export supply shocks. To address these and related concerns, the empirical model is augmented by time-varying fixed effects by regions and by firm size groups. In an even more demanding robustness check, both of these fixed effects are combined and further allowed to vary flexibly across industries and over time. The resulting estimates strengthen my main findings.

The positive employment effects are confirmed both during the pre-crisis period (2002-2007) and during the years of the financial crisis and subsequent recovery (2008-2013). While a fully balanced panel (resulting in only 242-674 firms) yields insignificant estimates, the main results are restored when applying a milder balancing requirement, focusing on firms that offshore services in least five consecutive years.

Similar estimates are obtained when excluding specific service types that do not reflect the core of service offshoring (such as financial and insurance, or construction, communication, and personnel

services) or when focusing exclusively on other business services, the most frequently imported service type.

Including also the log of service exports, instrumented in analogy to service imports by world import demand, results in positive and large estimates of the employment effect of service offshoring. However, these estimates are insignificant and suffer from a weak IV problem in the reduced sample. Service exports by themselves also tend to increase firm employment.

The effect of adding another service type (from the same foreign country) to the offshoring portfolio can be estimated in 2SLS regressions using the firm-specific WES by partner country and year as an IV for the number of country-service types. The estimates reveal positive employment effects also for this discrete measure of service offshoring.

Robustness checks with respect to the measurement of employment provide further supporting evidence. Similar to Bachmann and Bayer (2014), I approximate employment by the ratio of the firm-level wage bill over the average wage computed at the level of eight size classes, twelve regions, and 14 broad sectors. This approach exploits the fact that the USTAN data on the wage bill is subject to consistency checks, and it builds on the rationale that average wages are rather homogenous among firms in the same size class, region, and sector. Also for approximated employment the 2SLS regressions predict strong positive effects of service offshoring.

In a final set of robustness checks, I estimate dynamic panel data models that account for dynamic implications of lagged employment. The difference GMM estimates à la Holtz-Eakin et al. (1988) and Arellano and Bond (1991) use the second lag and all available higher-order lags of the dependent variable, all lags of the exogenous variables, as well as $\ln IV$ as instruments in the FD model augmented by the lagged dependent variable. While the regression diagnostics point to specification issues in the estimation sample with additional control variables, they support the validity of the GMM approach in the larger sample (see the Online Appendix for details). These difference GMM estimates confirm the significant employment gains from service offshoring.

7.5 Selection and Extensive Margin of Service Offshoring

The main analysis in Section 7.4 has focused on the causal effects of service offshoring at the intensive margin, i.e., the variation in the extent of service imports among firms that do (and previously did) import services. This section considers the role of the extensive margin of service offshoring to answer two questions that have thus far remained open: How does selection into and out of the sample of service importing firms influence the employment effects estimated above? And what happens to employment in firms that start service offshoring for the first time?

7.5.1 Selection correction at the intensive margin

Two margins of selection are relevant for the main analysis of the employment effects at the intensive margin of service offshoring. First, since information on employment (and other covariates) is only available for the firms reporting to MiDi or USTAN, there may be sample selection into and out of these databases. Second, firms select into and out of the service offshoring activity. To the extent that these margins of selection are non-random and their determinants are correlated with the error term $\varepsilon_{i,t}$ in equation (7.1), they may bias the estimates of β in the main analysis at the intensive margin. Note that the firm fixed effects in this model account for many of the determinants of selection that one can think of,

such as the constant components of firm productivity, international experience, or superior management. Still, some of these determinants vary over time within firms and may be correlated with employment growth. Therefore, remaining concerns related to both types of selection are addressed in turn in this section.

To correct for possible selection bias in the panel, I follow the method proposed by Wooldridge (2010), building on Heckman (1979).[34] It involves estimating probit selection equations year by year and including the inverse Mills ratios $IMR_{i,t}$ predicted by these equations, interacted with year dummies δ_t, into the FD model of equation (7.3). The model is estimated by 2SLS and standard errors are bootstrapped (based on 200 replications) to account for the generated regressors $IMR_{i,t}$, of which there are two sets.

First, to correct for selection into the sample, the annual selection equations explain a dummy variable for whether the observation is included in the full estimation sample from column 1 of Table 7.3. They include as explanatory variables a set of dummy variables indicating whether the firm was a service exporter, MNE, or FIE, and whether it reported to USTAN in the previous year, as well as industry fixed effects. These variables are pre-determined, well defined for all observations in the full panel, and they are related to selection into and out of the three underlying databases. The motivation for including these explanatory variables is that firms typically show high persistence in their trading status, FDI links, and reporting behavior, so these lagged variables have predictive power for future appearance in the sample. The estimates are used to compute the first set of IMR terms.

Second, to correct for selection into service offshoring, a service importer dummy is explained in annual cross-sections by the following firm characteristics $\mathbf{M}_{i,t-1}$: dummy variables indicating whether the firm was a service importer, service exporter, MNE, or FIE in the previous year, industry fixed effects, as well as the lagged logarithm the firm's turnover, employment, and the wage per employee. These variables are identified as relevant predictors of the extensive margin of offshoring in the analysis presented below. The predicted values of the service importer dummy are used to compute the second set of IMR terms.

This twofold approach to selection correction exploits the fact that the three underlying data sources are overlapping in terms of their firm coverage. Due to this feature of the data, different sets of firms are used to estimate the different first-stage equations for selection into the sample and for selection into the service offshoring activity. In the absence of useful exclusion restrictions, the selection models are identified by assuming joint normality of the error terms in the selection equations and $\nu_{i,t}$.

Table 7.8 reports the results of estimating the selection models for the two specifications with and without the additional firm-level control variables. First, to correct for selection into the sample, the interaction terms $\delta_t \times IMR_{i,t}$ for $t = \{2002, \ldots, 2013\}$ are added to the FD model in columns 1-2. Second, columns 3-4 proceed analogously for selection into service offshoring. Third, columns 5-6 include all 24 interaction terms to correct for both selection margins. The χ^2 tests of joint significance of the $\delta_t \times IMR_{i,t}$ terms suggest that the two types of selection are only marginally relevant for explaining changes in firm employment, and only few of the interaction terms are individually significant (not reported). The estimated employment effect of service offshoring is positive and significant in all of the regressions correcting for selection. The estimates are similar in magnitude compared to the baseline FD estimates from Table 7.5, and even higher in columns 3 and 5. These results demonstrate that the main estimates are not upward biased due to selection.

[34] See also Harrison and McMillan (2011) for an application in the offshoring context.

Table 7.8: *Selection correction*

	(1)	(2)	(3)	(4)	(5)	(6)
	Sample selection		Offshoring selection		Both selection margins	
$\Delta \ln$ *service imports*	0.0477***	0.0558***	0.0635***	0.0557***	0.0630***	0.0551***
	(0.0147)	(0.0154)	(0.0187)	(0.0167)	(0.0172)	(0.0163)
Additional controls	no	yes	no	yes	no	yes
Observations	31,869	11,292	13,429	11,294	13,425	11,290
F statistic (excl. IV)	568	207	258	207	257	206
χ^2 (IMR)	21.47	14.16	34.07	17.53	40.76	30.99
p-value of χ^2-test	0.0288	0.224	0.0004	0.0931	0.0088	0.0964

Note: The table reports 2SLS estimates of equation (7.3) with Δ ln *employment* as the dependent variable, including additional interaction terms as regressors. All regressions control for the basic set of lagged firm control variables (in first differences) and fixed effects by industry-year. The even columns further account for the additional firm-level control variables from column 4 of Table 7.3 (also in first differences). Columns 1-2 correct for selection into the estimation sample by adding interaction terms $\delta_t \times IMR_{i,t}$, columns 3-4 analogously correct for selection into service offshoring status, and columns 5-6 correct for both selection margins. The χ^2 (IMR) statistic and the associated p-value refers to a test for joint significance of all the interaction terms $\delta_t \times IMR_{i,t}$. Standard errors reported in parentheses are bootstrapped based on 200 replications. Asterisks indicate significance levels: * p<0.1, ** p<0.05, *** p<0.01. Source: RDSC of the Deutsche Bundesbank, SITS, MiDi, and USTAN, 2001-2013, own calculations.

7.5.2 Analyzing the extensive margin of service offshoring by propensity score matching

In order to estimate the employment effect at the extensive margin of service offshoring, a difference in differences (DiD) model is combined with propensity score matching (PSM). This approach compares the employment changes in newly offshoring firms to those in a comparable control group of non-offshoring firms. Formally, I am interested in the average effect of starting service offshoring on the employment of offshoring firms, i.e., the average treatment effect on the treated (ATT):

$$ATT_{\tau,t} = (\Delta_{\tau,t} L_{1,i} - \Delta_{\tau,t} L_{0,i} | O_{i,t} = 1), \qquad \tau = \{1, 2, 5\}. \tag{7.4}$$

In this equation, $\Delta_{\tau,t} L_{o,i} \equiv \left(\ln \textit{employment}_{o,i,t-1+\tau} - \ln \textit{employment}_{o,i,t-1}\right)$ denotes the relative change in employment of firm i with offshoring status $o = \{0, 1\}$ between the pre-treatment year $t-1$ and year $t-1+\tau$. The treatment variable $O_{i,t}$ is an indicator for offshoring firms that start to import services in year t for the first time (within the sample period 2001-2013). The ATT in equation (7.4) is estimated for three different time horizons. In the first specification, I set $\tau = 1$ to examine the immediate employment effect in the first offshoring year. Since domestic employment may need time to adjust, I examine also the medium-term effects after two and five years ($\tau = \{2, 5\}$, respectively). The latter two exercises adopt the convention that a firm is classified as offshoring only if it continues to import services in two (respectively, in five) consecutive years.

The obvious challenge in estimating the ATT in equation (7.4) is that the employment changes of an offshoring firm i are observed only for the situation in which it is offshoring $(\Delta_{\tau,t} L_{1,i})$, but not for the *counterfactual* situation in which it is *not* offshoring $(\Delta_{\tau,t} L_{0,i})$. One way to address this challenge is by using the employment changes of a suitable control group to impute the counterfactual employment changes of the offshoring firms. As a potential control group, I consider all non-offshoring firms, i.e., those which have not previously imported any services within the sample period and do not start offshoring over the years considered to compute the ATT. However, some of these non-offshoring firms differ in important ways from the treated firms and may hence not be suitable to estimate $\Delta_{\tau,t} L_{0,i}$. The

summary statistics in Table 7.1 illustrate that offshoring firms, included in the full estimation sample, are larger than firms in the full USTAN sample in terms of average employment and turnover, and they also differ substantially in terms of other firm characteristics. Similar differences are observed when comparing non-offshoring firms to firms that start service offshoring for the first time, also in the years before they start offshoring. This pattern can be rationalized by the fact that larger firms self-select into offshoring, for instance, because they are more productive and hence better able to cover the associated fixed costs (see e.g. Antràs and Helpman, 2004).

To provide a rigorous assessment of the ATT, I implement propensity score matching (PSM). The idea behind PSM is to determine a control group of firms that are as comparable as possible to the treated firms based on a set of observed covariates $\mathbf{M}_{i,t-1}$. In their seminal work, Rosenbaum and Rubin (1983) show that this objective can be achieved by matching observations based on the treatment probability predicted by $\mathbf{M}_{i,t-1}$, i.e., the propensity score $p_{i,t} \equiv Pr(O_{i,t} = 1|\mathbf{M}_{i,t-1})$. This approach allows for consistently estimating the ATT under three assumptions.

First, the 'conditional mean independence' (or ignorability in means) assumption requires that the average outcome (the subsequent change in employment) is independent of the treatment conditional on the covariates. Loosely speaking, this assumption is fulfilled if on average the observed pre-treatment characteristics $\mathbf{M}_{i,t-1}$ account also for unobserved factors that influence selection into treatment. While standard in the matching context, the conditional mean independence assumption is crucial for identification and fundamentally untestable. However, covariate balancing tests discussed below provide reassuring evidence that the PSM approach successfully accounts for differences between treated and untreated firms.

Second, the 'stable unit treatment value' assumption requires that the treatment affects all treated firms similarly and does not affect the non-treated firms. Arguably, this assumption may be questioned in any setting in which treated firms are competing with untreated firms in imperfect product or factor markets and may hence affect their performance through spillovers. However, since service imports make up a small share of firms' purchases on average, and given that firms are not matched within narrowly defined industries or regions, any violations of this assumption should have negligible effects in my application.

Third, the 'common support' assumption requires overlap between treated and untreated firms, i.e., $Pr(O_{i,t} = 1|\mathbf{M}_{i,t-1}) < 1 \,\forall\, \mathbf{M}_{i,t-1} \in \mathcal{M}_{t-1}$, where $\mathcal{M}_{t-1}$ is the support of the covariates. Intuitively, for each treated firm and each covariate, there must exist at least one untreated firm with a similar value for the covariate. This condition is enforced by excluding the few treated firms off the common support.

The propensity score $p_{i,t}$ is estimated separately for each offshoring year t using logit regressions of the treatment dummy $O_{i,t}$ on the matching covariates $\mathbf{M}_{i,t-1}$.[35] The variables included in $\mathbf{M}_{i,t-1}$ are the firm characteristics that were used to predict selection into offshoring in Section 7.5.1, measured in the pre-treatment year $t-1$. They are chosen to make firms comparable in terms of their initial service trade and foreign investment linkages, size, employment, and average wage, measured in the pre-treatment year $t-1$. This choice is the result of solving a trade-off between data availability and match quality, and it resembles the sets of covariates used in the previous matching literature on offshoring (Crinò, 2010a; Monarch et al., 2017; Moser et al., 2015). Also, the main conclusions obtained below are insensitive to small variations in this set. The logit regressions reveal that the preferred covariates

[35] Since the sets of treatment and control group firms vary by the selected time horizon $\tau = \{1, 2, 5\}$, three different logit regressions are implemented for each year t (and strictly speaking, the propensity score should have an index τ, omitted for brevity).

are individually significant predictors of treatment in most years and jointly yield a pseudo R^2 of around 9.9-14.6%.

To obtain a single ATT_τ for each time horizon $\tau = \{1, 2, 5\}$, I implement the PSM approach pooled over all years $t = \{2002, \ldots, 2013\}$. In my preferred specification, I employ the radius matching estimator with a caliper of 0.01. This DiD PSM estimator effectively compares employment changes in offshoring firms to the average of employment changes of multiple firms in the control group, where the narrow caliper ensures that only very similar firms observed in the same year are included in the control group. The matching procedure is implemented using the Stata command `psmatch2` (Leuven and Sianesi, 2003) and standard errors are bootstrapped based on 200 replications.[36]

Note that there is a crucial difference between the DiD PSM estimator employed in this paper and cross-sectional matching estimators, previously used for instance by Crinò (2010a) in the context of service offshoring. Crinò (2010a) uses PSM to compare employment *levels* across firms that import services and those that do not in a single cross-section. By contrast, this paper defines offshorers as firms that start importing services and exploits *time variation* within firms to identify the employment effects. Thereby, the DiD PSM estimator accounts for time-invariant unobservable firm characteristics and is hence superior to simple cross-sectional matching approaches (see Heckman et al., 1997; Smith and Todd, 2005). Furthermore, the panel data at hand allow for distinguishing the immediate impact of service offshoring from medium-term changes to shed some light on the persistence of the employment effects.

7.5.3 Employment effects of starting service offshoring

The results of implementing the DiD PSM procedure pooled over all years are summarized in Table 7.9. Column 1 reports the immediate ATT of service offshoring on firm employment in the first offshoring year. Columns 2-3 consider the medium-term employment effects in firms that start offshoring services and continue to do so over two and five consecutive years, respectively. The estimated ATTs suggest that service offshoring increases firm employment by around 2% in the first offshoring year, and by a similar annualized rate over the subsequent years.

The covariate balancing tests reported at the bottom of the table suggest that the PSM procedure is effective at making the treatment and control groups comparable. The medium absolute bias reduction (MABR) in terms of the pre-treatment matching variables is 75-84%. Also, logit regressions of the estimated propensity score on the matching variables yield pseudo R^2 values in the range of 0.003-0.019, which implies that these observables have very little power to predict treatment after matching. In general, the match quality is lower for the medium-term effects. Overall, the PSM seems to account for observable pre-treatment differences in the matching covariates to a satisfactory degree. To the extent that these covariates also account for unobserved heterogeneity, the DiD PSM estimator can serve to identify the ATTs.

It is important to note that these positive matching estimates are driven by firms starting to offshore in particular years. When implementing the PSM procedure year by year, the estimated ATTs fluctuate around zero. While they are never significantly negative, only a few years show positive and significant effects. Detailed results from these annual PSM estimations are shown in the Online Appendix.

[36] While Abadie and Imbens (2008) argue that the bootstrap is not generally valid in the case of nearest neighbor matching, they expect that it is valid in the case of matching algorithms for which the number of matches increases in the sample size.

Table 7.9: *Employment effects of starting service offshoring in DiD PSM estimations*

	(1)	(2)	(3)
	Immediate effect	Two-year effect	Five-year effect
ATT	0.0218***	0.0483***	0.106***
	(0.00829)	(0.0160)	(0.0352)
Observations	167,849	135,045	62,921
Treated firms	2,507	1,136	315
MABR	0.84	0.77	0.75
Pseudo R^2	0.00347	0.0113	0.0190

Note: The table reports the estimated average treatment effects on the treated ATT for firms that start offshoring services on changes in employment over a time period of 1, 2, or 5 years, in columns 1, 2, and 3, respectively. The estimates are pooled over the years 2002-2013 and based on the preferred DiD PSM estimator (radius matching with a caliper of 0.01). Each column reports the median absolute bias reduction (MABR) in terms of the pre-treatment matching covariates as well as the pseudo R^2 from regressing these covariates on the estimated propensity score. Standard errors reported in parentheses are bootstrapped based on 200 replications. Asterisks indicate significance levels: * p<0.1, ** p<0.05, *** p<0.01. Source: RDSC of the Deutsche Bundesbank, SITS, MiDi, and USTAN, 2001-2013, own calculations.

Also, the estimated ATTs are not fully robust to the use of alternative matching algorithms (results not shown). On the one hand, a kernel-based matching algorithm (using the Epanechnikov kernel function with a bandwidth of 0.01) yields very similar estimates to those shown in Table 7.9. On the other hand, single nearest neighbor matching (with a caliper of 0.01) suggests that the immediate impact of service offshoring in the first year is small and insignificant, but it confirms the positive and significant effects over the medium term. I have also experimented with varying the set of matching covariates by excluding lagged employment or wages, or by adding the lagged change in employment, which leaves the main conclusions unchanged.

To conclude, the PSM analysis at the extensive margin of service offshoring provides no evidence that starting service offshoring has destroyed jobs within German firms. Instead, the estimated employment effects are non-negative and may even be small and positive in some years.

7.6 Concluding Discussion

Fear of job losses in developed countries has spurred a public debate about service offshoring. The existing literature studying this phenomenon has found small positive correlations between employment and service offshoring, but establishing causality has remained a challenge due to endogeneity issues and the lack of detailed micro data. To address this challenge, the paper introduces a newly combined dataset covering almost the entire universe of German firms' service trade over the period from 2001 to 2013. Using this unique dataset, I conduct a comprehensive analysis of the causal effects of service offshoring (both at the intensive and at the extensive margin) on firm employment.

The main analysis uses a firm-specific IV based on foreign export supply shocks to show that increases in service offshoring have contributed to increasing employment within German firms. Also, there is no evidence of employment losses in newly offshoring firms. These results provide strong support for the productivity effect of offshoring predicted by the canonical trade in tasks model. In line with this theory, service offshoring has also boosted firm output and the use of other inputs, and the employment gains are stronger in firms that have previously offshored more services.

These findings indicate several promising avenues for future research. Since the analysis in this paper has focused explicitly on firm-level employment, it does not fully capture possible repercussions outside the firm in the domestic economy, e.g. through supplier substitution and spillover effects. Investigating the economy-wide impact of offshoring thus remains an important task. The robust positive employment effects of service offshoring found in this paper obviously contrast with the less favorable effects typically found for the offshoring of manufactured inputs in other studies (notably Hummels et al., 2014, who use a very similar methodology). Deeper investigations into the fundamental differences between these two types of offshoring and into differences across countries seem warranted. Finally, the effects of service offshoring on the composition of employment and the wage distribution across different skill groups and occupations remain a topic of major importance. Contemporaneous work by Ariu et al. (2018) and Ornaghi et al. (2017) using rich micro data from Finland and Belgium, respectively, may shed some light on these questions.

What can we expect from service offshoring in the future? Estimates from the mid-2000s suggest that around 25% of all US jobs are potentially offshorable, including many service activities that used to be non-tradable (see Blinder, 2006, 2009; Blinder and Krueger, 2013). It also seems reasonable to expect that the range of tradable services is going to expand further with technological progress. Therefore, the bulk of service offshoring may yet lie ahead. In light of the results found for the past decade in Germany, the average worker employed in a firm that considers expanding its service offshoring activities should have little reason to worry about these future developments.

Bibliography

Abadie, Alberto and Guido W. Imbens, "On the Failure of the Bootstrap for Matching Estimators," *Econometrica*, 2008, *76* (6), 1537–1557.

Amiti, Mary and Shang-Jin Wei, "Fear of Service Outsourcing: Is It Justified?," *Economic Policy*, 2005, *20*, 308–348.

_ **and** _ , "Does Service Offshoring Lead to Job Losses? Evidence from the United States," in "International Trade in Services and Intangibles in the Era of Globalization" NBER Chapters, National Bureau of Economic Research, 2009, pp. 227–243.

_ **and** _ , "Service Offshoring and Productivity: Evidence from the US," *The World Economy*, 2009, *32* (2), 203–220.

Antràs, Pol and Elhanan Helpman, "Global Sourcing," *Journal of Political Economy*, 2004, *112* (3), 552–580.

_ **, Teresa C. Fort, and Felix Tintelnot**, "The Margins of Global Sourcing: Theory and Evidence from US Firms," *American Economic Review*, 2017, *107* (9), 2514–64.

Arellano, Manuel and Stephen Bond, "Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations," *Review of Economic Studies*, 1991, *58* (2), 277–297.

Ariu, Andrea, "Services versus Goods Trade: A Firm-level Comparison," *Review of World Economics (Weltwirtschaftliches Archiv)*, 2016, *152* (1), 19–41.

_ **, J. Bradford Jensen, Katariina Nilsson Hakkala, and Saara Tamminen**, "Winners and Losers of Service Offshoring: Firm and Worker Level Evidence," 2018. University of Munich, mimeo.

Arndt, Sven W., "Globalization and the Open Economy," *The North American Journal of Economics and Finance*, 1997, *8* (1), 71–79.

_ , "Globalization and the Gains from Trade," in Klaus Jaeger and Karl-Josef Koch, eds., *Trade, Growth, and Economic Policy in Open Economies*, New York: Springer, 1998.

_ , "Super-specialization and the Gains from Trade," *Contemporary Economic Policy*, 1998, *16* (4), 480–485.

Arnold, Jens M., Beata S. Javorcik, and Aaditya Mattoo, "Does Services Liberalization Benefit Manufacturing Firms?: Evidence from the Czech Republic," *Journal of International Economics*, 2011, *85* (1), 136–146.

Autor, David H., David Dorn, and Gordon H. Hanson, "The China Syndrome: Local Labor Market Effects of Import Competition in the United States," *American Economic Review*, 2013, *103* (6), 2121–2168.

_ , _ , _ , and Kaveh Majlesi, "A Note on the Effect of Rising Trade Exposure on the 2016 Presidential Election," Technical Report, MIT Working Paper 2017.

Bachmann, Rüdiger and Christian Bayer, "Investment Dispersion and the Business Cycle," *American Economic Review*, 2014, *104* (4), 1392–1416.

Bardhan, Ashok D. and Cynthia A. Kroll, "The New Wave of Outsourcing," Fisher Center Research Reports 1103, Fisher Center for Real Estate and Urban Economics, University of California, Berkeley, Berkeley 2003.

Baumgarten, Daniel, Ingo Geishecker, and Holger Görg, "Offshoring, Tasks, and the Skill-Wage Pattern," *European Economic Review*, 2013, *61* (C), 132–152.

Bernard, Andrew, Teresa Fort, Valerie Smeets, and Frederic Warzynski, "Offshoring and Reorganization," 2018. Aarhus University, mimeo.

Beverelli, Cosimo, Matteo Fiorini, and Bernard Hoekman, "Services Trade Policy and Manufacturing Productivity: The Role of Institutions," *Journal of International Economics*, 2017, *104* (C), 166–182.

Biewen, Elena, Daniela Harsch, and Julia Spies, "The Determinants of Service Imports: The Role of Cost Pressure and Financial Constraints," Technical Report 90 2012.

_ , Sven Blank, and Simon Lohner, "Microdatabase: Statistics on International Trade in Services," Technical Documentation, Deutsche Bundesbank 2013.

Blinder, Alan S., "Offshoring: The Next Industrial Revolution?," *Foreign Affairs*, 2006, *85*.

_ , "How Many US Jobs Might Be Offshorable?," *World Economics*, 2009, *10* (2), 41–78.

_ **and Alan B. Krueger**, "Alternative Measures of Offshorability: A Survey Approach," *Journal of Labor Economics*, 2013, *31* (2), S97–S128.

Brändle, Tobias and Wolf-Dieter Heinbach, "Opening Clauses in Collective Bargaining Agreements: More Flexibility to Save Jobs?," *Review of Economics*, 2013, *64* (2), 159–192.

Breinlich, Holger and Chiara Criscuolo, "International Trade in Services: A Portrait of Importers and Exporters," *Journal of International Economics*, 2011, *84* (2), 188–206.

Correia, Sergio, "REGHDFE: Stata Module to Perform Linear or Instrumental-variable Regression Absorbing any Number of High-dimensional Fixed Effects," Statistical Software Components, Boston College Department of Economics 2014.

Crinò, Rosario, "Service Offshoring and Productivity in Western Europe," Technical Report 220, KITeS, Centre for Knowledge, Internationalization and Technology Studies, Università Bocconi, Milano, Italy 2008.

_ , "Employment Effects of Service Offshoring: Evidence from Matched Firms," *Economics Letters*, 2010, *107* (2), 253–256.

_ , "Service Offshoring and White-Collar Employment," *Review of Economic Studies*, 2010, *77* (2), 595–632.

_ , "Service Offshoring and the Skill Composition of Labour Demand," *Oxford Bulletin of Economics and Statistics*, 2012, *74* (1), 20–57.

Criscuolo, Chiara and Luis Garicano, "Offshoring and Wage Inequality: Using Occupational Licensing as a Shifter of Offshoring Costs," *American Economic Review*, 2010, *100* (2), 439–43.

Damijan, Jože, Stefanie A. Haller, Ville Kaitila, Črt Kostevc, Mika Maliranta, Emmanuel Milet, Daniel Mirza, and Matija Rojec, "The Performance of Trading Firms in the Services Sectors – Comparable Evidence from Four EU Countries," *The World Economy*, 2015, *38* (12), 1809–1849.

Dauth, Wolfgang, Sebastian Findeisen, and Jens Suedekum, "The Rise Of The East And The Far East: German Labor Markets And Trade Integration," *Journal of the European Economic Association*, 2014, *12* (6), 1643–1675.

Deardorff, Alan V., "Fragmentation Across Cones," in Sven W. Arndt and H. Kierzkowski, eds., *Fragmentation: New Production Patterns in the World Economy*, Oxford: Oxford University Press, 2001, pp. 35–51.

_ , "Fragmentation in Simple Trade Models," *The North American Journal of Economics and Finance*, 2001, *12* (2), 121–137.

Deutsche Bundesbank, "The Methodological Basis of the Deutsche Bundesbank's Corporate Balance Sheet Statistics," Monthly Report 10 1998. 49–64.

Dustmann, Christian, Bernd Fitzenberger, Uta Schönberg, and Alexandra Spitz-Oener, "From Sick Man of Europe to Economic Superstar: Germany's Resurgent Economy," *Journal of Economic Perspectives*, 2014, *28* (1), 167–188.

Ebenstein, Avraham, Ann Harrison, and Margaret McMillan, "Why Are American Workers Getting Poorer? China, Trade and Offshoring," NBER Working Papers 21027, National Bureau of Economic Research 2015.

_ , _ , _ , **and Shannon Phillips**, "Estimating the Impact of Trade and Offshoring on American Workers Using the Current Population Surveys," *The Review of Economics and Statistics*, 2014, *96* (4), 581–595.

Egger, Hartmut and Josef Falkinger, "The Distributional Effects of International Outsourcing in a 2 x 2 Production Model," *The North American Journal of Economics and Finance*, 2003, *14* (2), 189–206.

_ **, Udo Kreickemeier, and Jens Wrona**, "Offshoring Domestic Jobs," *Journal of International Economics*, 2015, *97* (1), 112–125.

_ , _ **, Christoph Moser, and Jens Wrona**, "Offshoring and Job Polarisation between Firms," CESifo Working Paper Series 6142, CESifo Group Munich 2016.

Eppinger, Peter, "Exploiting the Potential for Services Offshoring: Evidence from German Firms," IAW Discussion Paper 108, Institute for Applied Economic Research, Tübingen 2014.

Federico, Stefano and Enrico Tosti, "Exporters and Importers of Services: Firm-Level Evidence on Italy," *The World Economy*, 2016.

Feenstra, Robert C. and Gordon H. Hanson, "Foreign Investment, Outsourcing and Relative Wages," in R. C. Feenstra, G. M. Grossman, and D. A. Irwin, eds., *Political Economy of Trade Policy: Essays in Honor of Jagdish Bhagwati*, MIT Press, 1996, pp. 89–127.

_ **and** _ , "Globalization, Outsourcing, and Wage Inequality," *American Economic Review*, 1996, *86* (2), 240–245.

_ **and** _ , "The Impact Of Outsourcing And High-Technology Capital On Wages: Estimates For The United States, 1979-1990," *The Quarterly Journal of Economics*, 1999, *114* (3), 907–940.

_ **and** _ , "Global Production Sharing and Rising Inequality: A Survey of Trade and Wages," in E. Kwan Choi and J. Harrigan, eds., *Handbook of International Trade*, Vol. 1, Blackwell, 2003, chapter 6, pp. 146–185.

Francois, Joseph and Bernard Hoekman, "Services Trade and Policy," *Journal of Economic Literature*, 2010, *48* (3), 642–692.

Freund, Caroline and Diana Weinhold, "The Internet and International Trade in Services," *American Economic Review*, 2002, *92* (2), 236–240.

Gaulier, Guillaume, Emmanuel Milet, and Daniel Mirza, "Les firmes françaises dans le commerce international de services," *Économie et Statistique*, 2010, *435* (1), 125–147.

Geishecker, Ingo and Holger Görg, "Winners and Losers: A Micro-level Analysis of International Outsourcing and Wages," *Canadian Journal of Economics*, 2008, *41* (1), 243–270.

_ **and** _ , "Services Offshoring and Wages: Evidence from Micro Data," *Oxford Economic Papers*, January 2013, *65* (1), 124–146.

_ **, Maximilian Riedl, and Paul Frijters**, "Offshoring and Job Loss Fears: An Econometric Analysis of Individual Perceptions," *Labour Economics*, 2012, *19* (5), 738–747.

Groizard, Jose L., Priya Ranjan, and Antonio Rodriguez-Lopez, "Offshoring and Jobs: The Myriad Channels of Influence," *European Economic Review*, 2014, *72* (C), 221–239.

Grossman, Gene M. and Esteban Rossi-Hansberg, "Trading Tasks: A Simple Theory of Offshoring," *American Economic Review*, 2008, *98* (5), 1978–1997.

Hamermesh, Daniel S., *Labor Demand*, Princeton University Press, 1993.

Harrison, Ann and Margaret McMillan, "Offshoring Jobs? Multinationals and U.S. Manufacturing Employment," *The Review of Economics and Statistics*, 2011, *93* (3), 857–875.

Head, Keith and John Ries, "Offshore Production and Skill Upgrading by Japanese Manufacturing Firms," *Journal of International Economics*, 2002, *58* (1), 81–105.

_ **, Thierry Mayer, and John Ries**, "How Remote Is the Offshoring Threat?," *European Economic Review, Elsevier*, 2009, *53* (4), 429–444.

Heckman, James J., "Sample Selection Bias as a Specification Error," *Econometrica*, 1979, *47* (1), 153–61.

_ **, Hidehiko Ichimura, and Petra E. Todd**, "Matching as an Econometric Evaluation Estimator: Evidence from Evaluating a Job Training Programme," *Review of Economic Studies*, 1997, *64* (4), 605–654.

Hijzen, Alexander, Mauro Pisu, Richard Upward, and Peter W. Wright, "Employment, Job Turnover, and Trade in Producer Services: UK Firm-level Evidence," *Canadian Journal of Economics*, 2011, *44* (3), 1020–1043.

Hill, T. Peter, "On Goods and Services," *Review of Income and Wealth*, December 1977, *23* (4), 315–38.

Hoekman, Bernard and Carlos Braga, "Protection and Trade in Services: A Survey," *Open Economies Review*, July 1997, *8* (3), 285–308.

Holtz-Eakin, Douglas, Whitney Newey, and Harvey S Rosen, "Estimating Vector Autoregressions with Panel Data," *Econometrica*, 1988, *56* (6), 1371–1395.

Hummels, David, Jakob Munch, and Chong Xiang, "Offshoring and Labor Markets," NBER Working Papers 22041, National Bureau of Economic Research 2016.

_ **, Rasmus Jørgensen, Jakob Munch, and Chong Xiang**, "The Wage Effects of Offshoring: Evidence from Danish Matched Worker-Firm Data," *American Economic Review*, 2014, *104* (6), 1597–1629.

Jäckle, Robert and Georg Wamser, "Going Multinational: What are the Effects on Home-Market Performance?," *German Economic Review*, 05 2010, *11*, 188–207.

Jensen, J. Bradford, "Trade in High-Tech Services," *Journal of Industry, Competition and Trade*, 2008, *8* (3), 181–197.

_ **and Lori G. Kletzer**, "Tradable Services: Understanding the Scope and Impact of Services Offshoring," in S.M. Collins and L. Brainard, eds., *Brookings Trade Forum 2005, Offshoring White-Collar Work*, Brookings Institution Press, 2005, pp. 75–133.

_ **and** _ , "Measuring Tradable Services and the Task Content of Offshorable Services Jobs," in Katharine G. Abraham, James R. Spletzer, and Michael Harper, eds., *Labor in the New Economy*, University of Chicago Press, 2010. Mimeo.

Jones, Ronald W. and Henryk Kierzkowski, "A Framework for Fragmentation," in Sven W. Arndt and Henryk Kierzkowski, eds., *Fragmentation: New Production Patterns in the World Economy*, Oxford University Press, 2001, pp. 17–34.

Jones, R.W. and H. Kierzkowski, "The Role Of Services In Production And International Trade: A Theoretical Framework," in Ronald W. Jones and Anne O. Krueger, eds., *The Political Economy of International Trade*, Basil Blackwell, Oxford, 1990.

Kelle, Markus and Jörn Kleinert, "German Firms in Service Trade," *Applied Economics Quarterly (formerly: Konjunkturpolitik)*, 2010, *56* (1), 51–72.

Kimura, Fukunari and Hyun-Hoon Lee, "The Gravity Equation in International Trade in Services," *Review of World Economics (Weltwirtschaftliches Archiv)*, 2006, *142* (1), 92–121.

Kleibergen, Frank and Richard Paap, "Generalized Reduced Rank Tests Using the Singular Value Decomposition," *Journal of Econometrics*, July 2006, *133* (1), 97–126.

Kohler, Wilhelm, "Aspects of International Fragmentation," *Review of International Economics*, 2004, *12* (5), 793–816.

_ , "International Outsourcing and Factor Prices with Multistage Production," *Economic Journal*, 2004, *114* (494), C166–C185.

_ **and Marcel Smolka**, "Global Sourcing and Firm Selection," *Economics Letters*, 2014, *124* (3), 411–415.

Leuven, Edwin and Barbara Sianesi, "PSMATCH2: Stata Module to Perform Full Mahalanobis and Propensity Score Matching, Common Support Graphing, and Covariate Balance Testing," Statistical Software Components, Boston College Department of Economics 2003.

Lipponer, Alexander, "Microdatabase Direct Investment – MiDi: A Brief Guide," Technical Documentation, Deutsche Bundesbank 2011.

Liu, Runjuan and Daniel Trefler, "A Sorted Tale of Globalization: White Collar Jobs and the Rise of Service Offshoring," NBER Working Papers 17559, National Bureau of Economic Research 2011.

Mankiw, N. Gregory and Phillip Swagel, "The Politics and Economics of Offshore Outsourcing," *Journal of Monetary Economics*, 2006, *53*, 1027–1056.

Monarch, Ryan, Jooyoun Park, and Jagadeesh Sivadasan, "Domestic Gains from Offshoring? Evidence from TAA-linked U.S. Microdata," *Journal of International Economics*, 2017, *105*, 150–173.

Morikawa, Masayuki, "Service Trade and Productivity: Firm-level Evidence from Japan," Discussion papers 15030, Research Institute of Economy, Trade and Industry (RIETI) 2015.

Moser, Christoph, Dieter Urban, and Beatrice Weder Di Mauro, "On the Heterogeneous Employment Effects of Offshoring: Identifying Productivity and Downsizing Channels," *Economic Inquiry*, 2015, *53* (1), 220–239.

Muendler, Marc-Andreas and Sascha O. Becker, "Margins of Multinational Labor Substitution," *American Economic Review*, December 2010, *100* (5), 1999–2030.

Ornaghi, Carmine, Ilke Van Beveren, and Stijn Vanormelingen, "The Impact of Service and Goods Offshoring on Employment: Firm-level Evidence," Working Paper Research 319, National Bank of Belgium 2017.

Rosenbaum, Paul R. and Donald B. Rubin, "The Central Role of the Propensity Score in Observational Studies for Causal Effects," *Biometrika*, 1983, *70* (1), 41–55.

Schild, Christopher-Johannes and Frank Walter, "Microdatabase Direct Investment 1999-2015," Data Report 2017-01 – Metadata Version 4, Deutsche Bundesbank Research Data and Service Centre 2017.

_ , Simone Schultz, and Franco Wieser, "Linking Deutsche Bundesbank Company Data using Machine-Learning-Based Classification," Technical Report 2017-01, Deutsche Bundesbank Research Data and Service Centre 2017.

Sethupathy, Guru, "Offshoring, Wages, and Employment: Theory and Evidence," *European Economic Review*, 2013, *62* (C), 73–97.

Smith, Jeffrey A. and Petra E. Todd, "Does Matching Overcome LaLonde's Critique of Nonexperimental Estimators?," *Journal of Econometrics*, 2005, *125* (1-2), 305–353.

Stock, James H. and Motohiro Yogo, "Testing for Weak Instruments in Linear IV Regression," in D. W. K. Andrews and J. H. Stock, eds., *Identification and Inference for Econometric Models: Essays in Honor of Thomas Rothenberg*, Cambridge, UK: Cambridge University Press, 2005, chapter 5, pp. 80–108.

Stöss, Elmar, "Deutsche Bundesbank's Corporate Balance Sheet Statistics and Areas of Application," *Schmoller's Jahrbuch*, 2001, *121*, 131–137.

Winkler, Deborah, "Services Offshoring and its Impact on Productivity and Employment: Evidence from Germany, 1995–2006," *The World Economy*, December 2010, *33* (12), 1672–1701.

Wolfmayr, Yvonne, Elisabeth Christen, and Michael Pfaffermayr, "Pattern, Determinants and Dynamics of Austrian Service Exports – A Firm-level Analysis," FIW Research Reports series IV-005, FIW June 2013.

Wooldridge, Jeffrey M., *Econometric Analysis of Cross Section and Panel Data*, Vol. 2 of *MIT Press Books*, The MIT Press, 2010.

Wright, Greg C., "Revisiting the Employment Impact of Offshoring," *European Economic Review*, 2014, *66* (C), 63–83.

7.A Appendix

Table 7.A.1: *List of Services*

Service type	SITS service categories	Comtrade EBOPS 2002	NACE industries
Computer & IT	513	262	72
Communications	518, 591	245	64
Insurance	400, 401, 410, 420, 440-445, 450, 451, 460	253	66
Financial	533	260	65
Construction	561, 570	249	45
Engineering	512	280, 283	742, 743
R&D	511	279	73
Personnel	514, 517, 521	310	745
Other business services	516, 519, 523, 530, 531, 540	274, 284, 269, 278, 272	740, 741, 748, 749
Other services	510, 534, 562, 594, 595	282, 287	746, 747, 80, 85, 90, 92

The table lists the correspondence between the broad service types used in the analysis, the detailed service categories recorded in SITS (see Biewen et al., 2013), the Extended Balance of Payments Services (EBOPS) 2002 classification of the UN Comtrade data, and the NACE Rev. 1.1 industry classification (for the definitions of narrow and intra-firm offshoring).