Essays in International Trade: Firm Heterogeneity and Gravity Applications

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For my loving families in India and Germany

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INTRODUCTION

Economic globalization is not a self-generating and self-sustaining process. This vulnerability of globalization is evident when examining recent events such as the UK's disruptive departure from the European Union (EU), the protracted tariff war between the United States and China, the COVID-19 trade collapse and the escalating sanctions against Russia in 2022. Despite these numerous challenges, countries are still pursuing international economic integration by building new mega-regional trade blocs like the African Continental Free Trade Area (AfCFTA) and the Regional Comprehensive Economic Partnership (RCEP). These efforts suggest that globalization, whilst fragile, can be actively shaped by conscious policy choices. Moreover, alternative forms of globalization are still possible, which uphold social objectives such as climate change mitigation, public health, regional peace and human rights (Rodrik, 2020). The optimal design of globalization and regulating institutions that can best achieve such varied goals is therefore open to scientific research and policy deliberation.

This thesis contributes to this core agenda through four self-contained chapters that address various open questions concerning international trade and its distributional consequences. The first two chapters examine the heterogeneous adjustment of firms to large-scale policy shocks and shed light on the unique behaviours of so-called 'superstar' firms. These leading firms play an integral role in the global transmission of shocks (Di Giovanni, Levchenko, and Mejean, 2020), in shaping the comparative advantage of nations (Gaubert and Itskhoki, 2021) and lobbying for trade agreements (Blanga-Gubbay, Conconi, and Parenti, 2020). Their operational decisions have profound implications for labour markets (Autor et al., 2020) and the configuration of global supply chains (Alfaro-Urena, Manelici, and Vasquez, 2020). Given their overwhelming importance to trade outcomes, these first two chapters examine how large firms differ from their smaller counterparts in two additional respects: their choice of invoicing currency and their response to reductions in non-tariff barriers (NTBs).

Both these chapters rely upon highly detailed customs databases drawn from France. These databases constitute a rich source of information on the export and import activities of individual firms and are therefore ideal for contrasting the behaviour of superstars with that of other firms. Moreover, these databases permit us to examine the effects of policy changes over relatively long time periods. Beyond the study of superstars, a second important thread connecting these first two chapters is their joint examination of new developments in the EU-Asia trade partnership. This partnership has come increasingly into focus following China's rise as a manufacturing powerhouse and the rapid industrialization of other Asian economies such as India, Indonesia and Thailand.

In particular, the EU's multi-faceted relationship with China merits deeper examination. Looking at goods trade, China stood as the EU's third-largest export destination and largest source for imports in 2020. In the case of investment, Chinese companies have allocated close to €117 billion in the EU over 2000-2020, with EU companies investing €148 billion in China (European Commission, 2021). In financial markets, China's currency (Renminbi or RMB) is now viewed as a potential rival to the long-running dominance of the US dollar (USD) and the Euro. For instance, in January 2022, the RMB became the fourth most used currency in global payments by value. This rise of the RMB is studied extensively in the first chapter, from the lens of individual French exporters and their invoicing currency decisions.

CHAPTER 1, therefore, investigates the impressive growth in RMB-denominated exports from France over 2011-2017. It documents several novel stylized facts on RMB invoicing, such as the high degree of concentration in RMB usage amongst exporters. The findings here reveal that a few large multi-product firms dominate RMB exports from France, which are almost entirely directed towards China. Besides indicating strong selection effects into invoicing in RMB, this observation also suggests that RMB has not displaced either the EUR or the USD as the preferred invoicing currency in French exports outside of China.

Based on these stylized facts, the paper then rigorously examines various selection mechanisms underpinning RMB invoicing. In a novel contribution to the literature, it shows that local currency pricing (LCP) in other foreign (extra-EU) markets is a strong predictor for the probability of a given variety being invoiced in RMB in China. This finding is indicative of firms' synchronization of invoicing strategies across export destinations. Since LCP in other markets is observed to be tightly correlated with RMB adoption and is plausibly orthogonal to shocks faced by a variety in China, it is used as an instrument to determine the causal impact of RMB invoicing on exports. Such causal estimates of invoicing currency choice on trade flows are rare in the existing literature, given the high degree of stability in invoicing patterns. This paper addresses the gap by exploiting the time variation in RMB use generated by China's currency reforms, the novel instrument on LCP and a range of fixed effects to reduce concerns of omitted variable bias. The resulting coefficients from two-stage least squares (TSLS) regressions reveal that invoicing in RMB increased exports of a given variety to China by 51.4% and exported quantity by 73.2%.

In response to the growing economic influence of Asia, the EU has also negotiated a series of free trade agreements (FTAs) with several countries in the region. The first of these agreements was the deep and comprehensive EU-South Korea FTA that entered into force in 2011, which subsequently became a template for other 'new-generation' FTAs of the EU. The heterogeneous impact of this FTA across exporters of varying sizes and along multiple margins is taken up in the second chapter, written in collaboration with Gabriel Felbermayr.

CHAPTER 2, therefore, applies a triple-difference framework with high-dimensional fixed effects to examine the impact of NTB and tariff reductions achieved by this FTA. It reports findings that contrast leading theories in international trade which suggest that, at the intensive margin, the export growth rate of smaller firms should match or even exceed that of larger firms following trade cost reductions (Arkolakis, 2010; Melitz and Ottaviano, 2008). However, the paper provides robust evidence to the contrary. Across specifications, the EU-South Korea FTA is seen to generate a substantial size premium in export growth along the intensive margin which is driven by a stronger effective reduction in broadly defined NTBs for larger exporters. On the entry margins, results align with the literature with the FTA selecting intermediate-sized firms into exporting to South Korea. Besides the usual self-selection of firms into exporting, this differential impact of the FTA (and specifically NTB cuts) along the intensive margin highlights an under-explored channel through which globalization can magnify inter-firm inequality.

CHAPTER 3 moves beyond partial equilibrium analysis of policy shocks and their heterogeneous impact on individual trading firms. It does so by shifting the focus towards the structural gravity model and computing the general equilibrium impact of a different but frequently used trade policy instrument, i.e. sanctions. In joint work with Julian Hinz and Katrin Kamin, this paper highlights the welfare impact of countries coordinating their sanctions packages. To assess the role of such sanctions coalitions, the paper concentrates on the 2012 wave of sanctions against Iran and the 2014 sanctions against Russia. These sanctions episodes are examined with a Caliendo and Parro (2015)-type model of the world economy, which is based on Hinz and Monastyrenko (2022) and similar to Chowdhry et al. (2020). Crucially, the model provides changes in welfare that would be experienced by Russia, Iran, the sanctioning states and third parties under different hypothetical setups of sanctions coalitions.

The simulation results reveal that welfare losses incurred by Iran and Russia are higher when sanctions are imposed collectively rather than unilaterally by coalition members. Moreover, China's potential cooperation in the sanctions package against Russia and Iran is seen to greatly increase the punitive force of these sanctions without imposing welfare costs on existing sanctioning states or China itself. Finally, we evaluate which countries would need to be brought on board as prospective allies to amplify the deterring effect of sanctions. In the case of Iran, the most important third party countries which would increase the punitive impact of sanctions are China, India, UAE, Brazil, Azerbaijan and Russia. For Russian sanctions, the leading potential coalition partners are seen to be China, South Korea, Belarus, Turkey, Brazil and India. Overall, these results shed light on the economic cost of conducting 'war by other means' and the vital importance of building coalitions when imposing sanctions.

CHAPTER 4 examines the role of distance, a central component of the gravity relationship, in disease transmission. Co-authored with Gabriel Felbermayr and Julian Hinz, it adapts methods and hypotheses that have been developed in the gravity literature to study the urgent question of how COVID-19 infections spread rapidly across Germany in 2020. Drawing on micro-data documenting COVID-19 case loads and the socio-economic characteristics of 401 German counties, this paper provides novel evidence that geographical proximity to the super-spreader location of Ischgl accelerated the spread of the virus in the first wave.

Based on data from Google searches, the road distance to Ischgl is seen to closely match the probability of tourists in a given county visiting this Austrian ski hotspot. Using this as a proxy for travel, we find that road distance to Ischgl is indeed an important predictor of infection rates, but — in line with expectations — not of fatality rates. The results also confirm that distance to Ischgl did not become irrelevant over time for observed cases, suggesting that early lockdown measures were effective in reducing mobility and avoiding further diffusion of the virus across German counties. Interestingly, the findings rule out other hypotheses that a county's trade with China (where the outbreak was first reported) or the presence of foreign-born residents played a role in the diffusion of the virus within Germany during the first wave. This paper has been published in the German Economic Review as well.

In summary, these four chapters bring together multiple datasets and different econometric techniques to study a broad range of issues in international economics. Overall, the findings contribute to the vibrant literature on firm heterogeneity in international trade, the distributional implications of trade policy instruments such as FTAs and sanctions as well as the role of international tourism in the cross-border diffusion of the SARS-CoV-2 virus.

Chapter 1

Local Currency Pricing by Firms: Evidence from China's Renminbi Reforms

1.1 Introduction

In January 2012, the Chinese Renminbi (RMB) ranked twentieth amongst the world's currencies used most actively for global payments. By January 2022, it stood fourth in the list after overtaking the Japanese Yen.[1] This rise in RMB use within a decade follows a series of reforms undertaken by the People's Bank of China (PBC) designed to "steadily and carefully promote the internationalization of the RMB".[2] This paper exploits detailed customs data from France, the leading RMB clearing hub within the Eurozone, in order to examine the effect of these internationalization reforms on the invoicing currency choice and export performance of firms selling to China.[3] This is the first paper that analyzes the drivers of RMB invoicing using such disaggregated firm-level data.

The study of firms' invoicing currency behaviour is important in numerous respects. First, the choice of invoicing currency is not exogenous but rather the outcome of firms' profit optimization. It therefore constitutes an additional margin of adjustment that is available to firms. Second, the choice of invoicing currency has significant distributional consequences. When goods are priced in the destination's local currency, exporters bear the risk of exchange rate fluctuations. Alternatively, if exporters invoice in the home (producer) currency, that exchange rate risk is transferred to consumers. At the macroeconomic level, the choice of invoicing currency affects exchange rate pass-through into prices (Boz, Gopinath, and Plagborg-Møller, 2019) and the international transmission of shocks (Corsetti et al., 2007). Finally, from a policy perspective,

[1] For further details, see the RMB Tracker from SWIFT. Link: https://bit.ly/3tdYNIX

[2] Extract from China's 14th Five-Year Plan (2021-2025). For the complete (translated) document see https://bit.ly/3ppEAz0

[3] For further details on France's share of global RMB payments, see "China's Currency Push", MERICS (2020). Link: https://bit.ly/34vnKan

there is growing interest from central banks such as the PBC to promote their currencies for cross-border settlements. Studying firm-level responses reveals the extent to which such efforts are successful in encouraging businesses to adapt their invoicing behaviour.

With detailed information from France on firms' invoicing currency choices over 2011-2017, the paper first reports several novel stylized facts regarding RMB adoption. The key stylized fact concerns the striking growth in France's RMB-denominated (monthly) exports to China, from €2.3 million in January 2011 to €145 million by December 2017. These RMB-denominated exports accounted for more than a third (€1.77 billion) of the total increase in French exports to China between 2011 and 2017. Despite this sharp increase, French exporters' global use of RMB remains narrow as nearly 99% of all RMB-denominated exports are directed only towards China. Hence, RMB has not displaced either the EUR or the USD as the preferred invoicing currency in French exports worldwide.

Second, RMB usage shows a high degree of concentration with old varieties (those exported to China in previous years) capturing more than 75% of RMB exports in any given year. In contrast, first-time exporters to China rarely choose the RMB. Further disaggregation of RMB invoicing reveals considerable granularity. RMB invoicing firms are few and large, and even within this set, multi-product firms that invoice several products in RMB account for more than 90% of total RMB exports to China.

Based on these stylized facts, the paper examines several drivers of the selection into RMB adoption. This includes channels such as operational hedging, strategic complementarity and fixed costs of currency management which have been suggested by existing literature to analyze firm invoicing in established currencies. In addition to the above mechanisms, the paper proposes a fourth and novel channel driving RMB use. This novel channel relates the selection of a given variety into invoicing in the (newly available) local currency in China to the use of local currencies when its exported to other extra-EU destinations.

The data clearly bear out this harmonization in local currency pricing across destinations as more than 75% of RMB-invoiced varieties are also sold in other markets in their respective local currencies. In the baseline specification, such local currency pricing of a given variety elsewhere corresponds to a 0.9 percentage point increase in the probability of RMB adoption in China. Furthermore, the likelihood of RMB invoicing is magnified for varieties that are exported in local currencies to more than two destinations and which have more than two years of experience in being exported in local currencies in other destinations. This suggests that firms may prefer to invoice certain products in local currencies across their markets but were constrained from doing so in China until the RMB internationalization reforms.

Since local currency pricing in other foreign markets is observed to be a strong predictor for RMB use, this novel channel is further exploited as an instrument to determine the causal impact of RMB invoicing on firm exports to China. Such causal estimates of invoicing choice on trade flows are scarce in the existing literature, given the high degree of stability in invoicing patterns. This paper addresses the crucial gap by exploiting the time variation in firm invoicing generated by China's currency reforms, the novel instrument on local currency pricing and a rich set of firm-specific fixed effects. The validity of the novel instrument rests on the plausible assumption that the use of local currencies when exporting to other markets is independent of export sales in China or orthogonal to shocks faced by the given variety in China. Based on this assumption, the two-stage least squares (TSLS) regressions show that invoicing in RMB increases exports of a given variety in China by 51.4% and exported quantity by 73.2%. In contrast, there is no statistically significant change in export prices. This positive effect of RMB use on export sales is indicative of the competitive advantage of invoicing in local currency in destination markets.

Overall, this paper provides i) new stylized facts concerning firm heterogeneity in the adoption of a newly available currency (RMB); ii) proposes a novel mechanism driving the selection of varieties into RMB invoicing and; iii) reports causal estimates on the impact of RMB invoicing on export revenues in China. The remainder of this paper is structured as follows. Section 1.2 briefly describes prior research on firms' invoicing currency behaviour and highlights the paper's contribution to this literature. Section 1.3 outlines the key milestones in China's gradual liberalization of the RMB and discusses the motivations underlying this recent shift in China's currency policy. Section 1.4 describes the structure of French customs data and sets out three stylized facts concerning RMB invoicing. Following this, Section 1.5 analyzes selection into RMB invoicing. The TSLS regression results on the causal impact of RMB invoicing on firm exports are reported in Section 1.6. Finally, Section 1.7 concludes with policy implications and suggested avenues for future research.

1.2 Related Literature

This paper is related to several strands of research on invoicing currency. First, it contributes to the growing literature on firm-level heterogeneity in invoicing currency choice. Papers in this literature have exploited detailed customs datasets in order to analyze the relationship between firm characteristics and various invoicing strategies such as producer currency pricing (PCP), local currency pricing (LCP) or vehicle currency pricing (VCP). For instance, Amiti, Itskhoki, and Konings (2020) use Belgian customs data to demonstrate that invoicing currency choice is an active firm-level decision, but more so for exporting than importing. They also show that firms invoice in USD when exporting commodities or homogeneous goods, whereas third currency invoicing (i.e. currencies other than the EUR or USD) is more prevalent for exports of differentiated goods. Using customs data from France, Barbiero (2020) reaffirms that firm-level characteristics matter for invoicing. French firms are observed to differ in their invoicing currency choice even within the same country-industry pair with larger firms invoicing in multiple currencies as they trade with more countries. In this paper, I shift the analysis away from established currencies and provide stylized facts that reveal substantial heterogeneity across exporters in the adoption of a newly available currency (RMB). The focus on exporters also aligns with benchmark models wherein exporting firms choose the optimal currency and importers respond by adjusting the quantity purchased(Amiti, Itskhoki, and Konings, 2020).

This literature also suggests numerous mechanisms that underpin invoicing currency choices of the firm. Amiti, Itskhoki, and Konings (2020) provide evidence that firms are more likely to invoice their exports in a given currency if they also import inputs in that currency (operational hedging channel) and if competing firms within the same industry-destination pair invoice in that currency (strategic complementarity channel). Using transaction-level customs data from the UK, Crowley, Exton, and Han (2020) document an additional channel that determines firms' invoicing currency choices. Focusing on the dominance of USD, they analyze the role played by prior currency experience of the firm. In their data, firms are more likely to invoice in USD when exporting to a new destination if they have persistently used USD in existing markets. This channel emerges from their theoretical model which features increasing returns to scale associated with the fixed costs of currency management. In this paper, I depart from Crowley, Exton, and Han (2020) by examining firm behaviour along a different extensive margin which is defined in terms of additions to the set of available invoicing currencies rather than firm entry into new markets. This allows me to uncover a fourth channel that has not been examined previously.

The underlying idea for this fourth channel is that firms are more likely to adopt the RMB when exporting to China for products that they already invoice in local currencies to other destinations. This 'local currency use channel' can be rationalised if firms face high levels of competition from local sellers or if demand is highly sensitive to exchange rate risk or if it provides convenience in settling transactions with buyers. Under these various conditions, firms may prefer to invoice certain products in local currencies in all their markets but were constrained from doing so in China until the RMB internationalization reforms. In suggesting this novel channel, the paper contributes to earlier research on LCP by firms. The findings from this literature show that firms are more likely to invoice in the currency of their export destination if they are multi-national corporations (Ito, Koibushi, et al., 2021), have transactions that are large in terms of absolute size (Goldberg and Tille, 2016), can hedge at low costs against exchange rate risks using forward contracts (Ito, Koibuchi, et al., 2010) and have increased access to hedging instruments such as derivatives (Lyonnet, Martin, and Mejean, 2016).

Finally, this paper is closely related to recent empirical research on the aggregate effects of China's RMB promotion policies. For instance, Bahaj and Reis (2020) investigate the impact of establishing swap lines with the PBC on RMB usage. By exploiting bilateral country-level data from the SWIFT Institute on cross-border payments, they estimate that swap lines led to a 13 percentage point increase in the probability of a country transacting in RMB. Boz, Casas, et al. (2020) assemble an alternative dataset that provides information on the share of different invoicing currencies in aggregate exports and imports of 102 countries over 1990-2019. Unlike SWIFT, this country-level panel data focuses on invoicing rather than payments and is not altered by the presence of international payment hubs.

With this data, the authors find little variation in invoicing currency at the country-level with a high degree of stability in the share of trade invoiced in USD or EUR. However, they do observe increasing use of RMB in African economies given their growing trade with China. Georgiadis et al. (2021) widen the scope of invoicing data constructed by Boz, Casas, et al. (2020) and find that PBC swap lines do not have a clear effect on RMB invoicing. Only when interacted with countries' trade exposure to China do these swap lines show a positive impact on RMB usage. However, their data still does not cover RMB invoicing by EU members.

This paper contributes to the above literature on the rise in RMB usage in three different respects. First, it shifts the analysis of RMB invoicing from the country-level to the firm-level and does so for a major EU economy. It documents the rich heterogeneity across firms in RMB adoption that is masked by previously used aggregate data. Second, the use of firm-level data reduces potential threats to identification resulting from endogeneity at the country-level between trade flows and the signing of swap lines with China. Third, firm-level data permits an investigation of whether RMB invoicing had a causal impact on the revenues and prices of exporters to China. In doing so, this paper also adds to recent firm-level evidence provided by Messer (2020) on the impact of foreign currency risk on exports. While Messer (2020) focuses on the impact of invoicing in the home currency on exports of Brazilian firms, I focus on the impact of invoicing in the destination currency, i.e. RMB, on French firm exports to China.

1.3 Policy Context

Since joining the World Trade Organization (WTO) in 2001, China has become a major trading economy. By 2009, it was the world's largest goods exporter accounting for 9.6% of global merchandise exports and surpassing other leading exporters such as Germany (8.9%), USA (8.4%) and Japan (4.7%). Unlike these nations however, China's domestic currency remained peripheral to international capital markets due to the numerous restrictions around its use.

Several factors however motivated a shift in China's policy towards RMB promotion and internationalization. As suggested by Eichengreen and Xia (2019), the first factor was the need to lower transaction costs for Chinese firms and to promote their international competitiveness. The second was to enhance China's financial stability by decreasing the dependence of domestic exporters and importers on the USD. Third, China hoped to reduce asymmetries in global financial markets by creating a multipolar system of international currencies that would include the RMB. Finally, symbolic reasons relating to prestige and reputation also played a significant role in the push towards RMB internationalization. In this section, I provide a brief overview of the key milestones in the RMB reform process over 2011-2017.

The first important step towards reform was undertaken in July 2009, when the PBC launched a pilot programme to promote the use of RMB for trade settlements. The scope of this pilot scheme was quite limited however, with RMB settlements restricted to five cities from mainland China in addition to Hong Kong, Macau and members of the ASEAN bloc.[4] In July 2010, the scheme was expanded to cover 20 pilot areas. In these areas, only firms that were recognised as Mainland Designated Enterprises (MDEs) could settle in RMB but now with any country in the world. The purpose of creating such a list of MDEs was to control and monitor the number of Chinese firms dealing in RMB. Obtaining this status also imposed administrative costs on firms. Despite these limitations, the pilot scheme helped foster the use of RMB in Asia and established Hong Kong as the main offshore RMB center.

The most significant expansion occurred later in February 2012 when Chinese authorities issued the *Circular Yinfa [2012] No. 23*. This new Circular permitted all firms in China to settle in RMB unless they were blacklisted due to serious violations of tax or trade laws. With this new regulation, China's RMB liberalization therefore moved from a positive to a negative list approach. By eliminating the earlier requirement that firms be designated as MDEs, the Circular expanded the number of firms eligible to trade in RMB.

[4]The five mainland cities included were Shanghai, Guangzhou, Shenzhen, Dongguan and Zhuhai.

More recently, China has taken additional measures to promote the RMB. PBC signed several swap agreements with central banks, including the European Central Bank (ECB) in 2013, to increase the liquidity of RMB and provide RMB lending of last resort to foreign firms (Bahaj and Reis, 2020). This swap arrangement had a maximum size of RMB 350 billion and €45 billion and was the second-largest swap line signed by the PBC at the time, following South Korea (RMB 360 billion). In 2016, the RMB was also included within the International Monetary Fund (IMF) basket of Special Drawing Rights (SDR), providing a further boost to the currency's international reputation as a suitable vehicle for trade transactions. The addition of RMB to the SDR is notable considering that the last change had only been in 1999 when the Euro replaced the German mark and French franc in the IMF's currency basket.

Looking at the currency composition of official foreign exchange reserves held worldwide, the share of RMB-denominated assets also increased from virtually zero in 2010 to approximately 2% in 2020.[5] While the above developments have eased several constraints on RMB use, the reform process is still ongoing. This is clearly reflected in China's 14th Five Year Plan (2021-2025), one of whose stated objectives is to "steadily and carefully promote the internationalization of the RMB" and "strengthen the construction of the RMB cross-border payment system".

Since these reforms were motivated primarily by China's domestic concerns, they can be considered as exogenous shocks to individual firms. In particular, the removal of strict eligibility criteria surrounding RMB use in 2012 and the ECB's swap line agreement with PBC in 2013 enlarged the set of invoicing currencies available to EU firms trading with China. Moreover, such an expansion in currency choices is a relatively rare occurrence that has not been investigated by prior literature on invoicing currency. By permitting exporters to also invoice their products in RMB, these reforms potentially altered firms' transaction costs and their competitiveness in China's local markets.

In the following sections, I turn towards firm-level customs data from France in order to investigate the extent to which these currency promotion policies have succeeded in fostering the use of RMB by exporting firms and the impact of invoicing in RMB on their export outcomes. Besides the availability of rich customs data, France makes for an ideal choice for examining RMB adoption given that it is the leading country within the Eurozone for RMB clearing by value.

[5]Latest data from 'Currency Composition of Official Foreign Exchange Reserves (COFER) Database', IMF. Link:https://bit.ly/3Cs6O1o

1.4 Data and Stylized Facts

1.4.1 Customs Data

To examine heterogeneity in RMB invoicing across firms, I exploit detailed customs data from France (dataset DGDDI, 2018).[6] This data provides information on exports and imports of a firm (f), disaggregated by destination or source country (d), product (p), invoicing currency (c) and time (t). In addition to the merchandise value in Euros, the dataset contains information on the quantity (weight) of traded goods. Each firm is also assigned a unique identifier ('SIREN') that can be tracked over time. The 8-digit Combined Nomenclature ('CN8') classification for products is concorded over 2011-2017 following the algorithm proposed by Bergounhon, Lenoir, and Mejean (2018). The primary variable of interest, i.e. invoicing currency, is reported only from 2011 onwards and is restricted to trade with extra-EU economies.[7] Therefore, the majority of empirical analysis in this paper will rely on customs data spanning 2011 to 2017.[8]

In all, the customs database provides information on the universe of extra-EU trade of French firms over 2011-2017. It covers more than 350,000 firms that invoice in more than 140 unique currencies, trade in more than 6900 CN8 products and with 160 extra-EU partner countries. Table 1.1 provides summary statistics for exporters in this database. It reveals that the top 100 exporters alone account for 48% of total exports over 2011-2017 and exhibit substantial diversification across destinations and CN8 products. There is substantial variation in the choice of invoicing currencies as well, even within firm-CN8 product combinations. In comparison, the smallest exporters typically invoice in just one currency.

1.4.2 Stylized Facts on RMB Adoption

By exploiting the customs data described in the previous section, I document three key stylized facts relating to heterogeneity of RMB invoicing by French exporting firms. I first explore the growth of RMB invoicing over time in comparison to the USD and

[6] The transaction-level customs data that support the findings of this study is covered by statistical secrecy and can be accessed only through a previous authorization of the French Custom Administration. The customs data is from the DGDDI (Direction Générale des Douanes et Droits Indirects – a directorate of the French Ministry of Finance). The authorization is granted by the "Comité du secret" of the CNIS (Conseil National de l'Information Statistique). The link to procedures for getting access to the data is: https ://www.comite-du-secret.fr/.

[7] France's customs agency does not collect invoicing currency data for intra-EU transactions.

[8] Since invoicing behaviour is not known prior to 2011, this prevents a proper investigation of currency switching by traded varieties. In limited cases, I do exploit customs data from the preceding period (2000-2010). Even though this data lacks information on firms' currency choices, it can be helpful in determining the export histories of firms and varieties.

Table 1.1: Summary statistics: Top exporters from France to extra-EU economies

	Top 100	100-1000	Remaining
Share in total exports	48%	29%	23%
Mean # destinations per firm	80.04	51.01	4.52
Mean # CN8 per firm	283.59	128.82	8.34
Mean # currency per firm	10.9	5.12	1.18
Mean # currency per firm-CN8	2.42	1.77	1.05
Mean # currency per firm-CN8-destination	1.08	1.03	1.01

EUR before turning to the decomposition of this RMB growth along various margins. I then examine how RMB exports are distributed across firms and provide preliminary evidence of firm selection into RMB invoicing. Finally, I show that RMB invoicing even varies within firms that exported multiple products to China over 2011-2017.

Stylized fact 1. *Strong growth in RMB-denominated exports to China*

RMB-denominated exports to China have grown significantly as shown in Figure 1.1. While the value of (monthly) RMB exports to China stood at €2.3 million in January 2011, this increased to approximately €145 million in December 2017. This increase is notable in terms of shares as well. While the EUR and USD remain dominant, the share of RMB in monthly French exports to China grew from less than 1% to 10% by the end of 2017. Moreover, these RMB-denominated exports accounted for more than a third (€1.77 billion) of the total increase (€4.97 billion) in French exports to China when comparing 2011 to 2017.

In Figure 1.2, I decompose this growth in RMB exports to China into i) growth in the number of RMB invoicing firms; ii) growth in RMB invoiced CN8 products; iii) growth in RMB exports per variety (firm-CN8 pair) and; iv) growth in density i.e. share of all possible firm-CN8 combinations which have positive RMB exports. For each year, I depict the growth rates in these margins relative to 2011. Looking at 2017, the data shows nearly 700% growth in the number of RMB-invoicing exporters relative to 2011 with a 560% growth in the number of CN8 products invoiced in RMB. In comparison, growth rates are moderate in the number of EUR (22.15%) or USD (2.18%) invoicing firms and EUR (10.21%) or USD (3.83%) invoiced products.[9] This descriptive evidence clearly indicates that the growth in RMB exports is driven by the extensive rather than the intensive margin. Thus, an increasing number of firms adopted RMB for a widening range of products exported to China.[10]

Despite this growth in RMB usage, I find that nearly 99% of all RMB-denominated exports from France were directed towards China over 2011-2017. The remaining 1%

[9] See Figure A.1 in Appendix A.2. Note that comparisons across currencies in the growth of varieties is feasible as more than 95% of varieties are exported only in one currency to China in a given year.

[10] For the sectoral distribution of these RMB exports, see Figure A.2 in Appendix A.2.

Figure 1.1: Monthly French exports to China

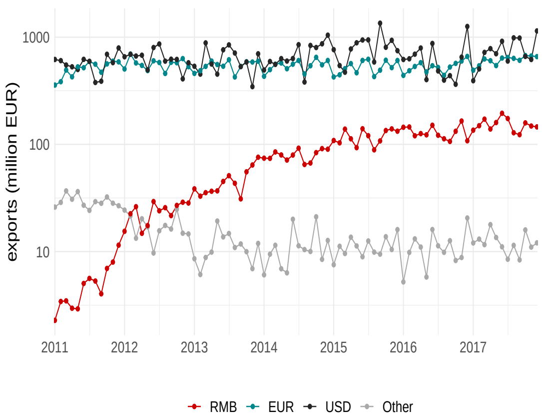
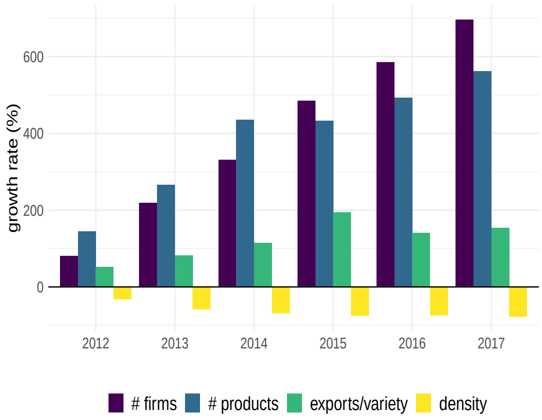


Figure 1.2: Margin decomposition



Note: Figure 1.1 plots monthly exports from France to China over 2011-2017 on a logarithmic scale. Exports are then disaggregated by invoicing choice – RMB, EUR, USD and any other currency. Figure 1.2 displays the growth in RMB exports to China (relative to 2011) along multiple margins following the decomposition strategy proposed by Bernard et al. (2009).

of exports in RMB went towards Hong Kong and Japan. Therefore, RMB was rarely used as a vehicle currency by French exporters serving other destinations during this time period. Its availability therefore did not displace the EUR or USD as the preferred invoicing currency in French exports globally.[11]

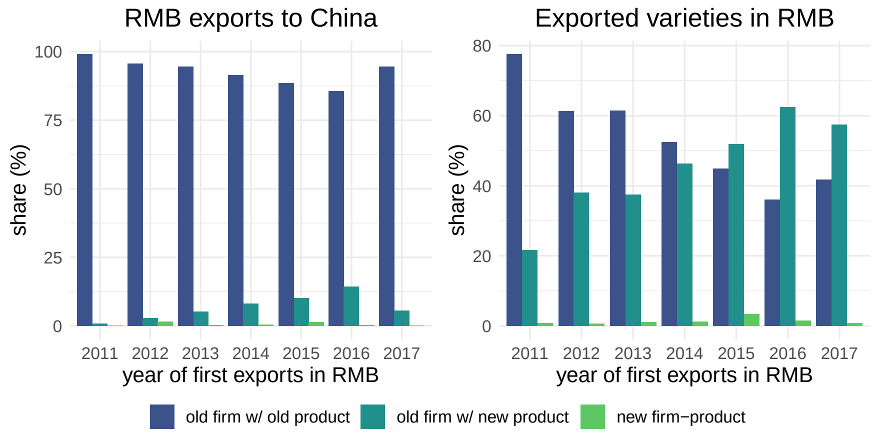
Stylized fact 2. *RMB invoicing is highly concentrated*

As discussed above, the growth in RMB use by French exporters over 2011-2017 stems predominantly from the extensive margin. This leads to the next question of whether the extensive margin comprises of firms that are new exporters to China or just new to invoicing in RMB. To examine this, I make use of the long panel of customs data spanning 2000-2017 which allows me to determine the export history of a given firm-product combination up until the time of its first year of invoicing in RMB. Since the use of RMB for settling transactions was limited to ASEAN member nations before 2011, we can reasonably assume that the first year of a French firm invoicing in RMB in the data corresponds to the actual first year of RMB adoption by the firm.

With information on export histories, I further decompose the extensive margin into the following: i) varieties that were exported to China at least once prior to the first year of exporting in RMB ('old firm w/ old product'); ii) firms that had exported to China previously but now introduce a new product to the market which is invoiced in RMB ('old firm w/ new product') and; iii) new firms that invoiced in RMB in their first

[11]On average across 2011-2017, EUR and USD-denominated exports account for 51.2% and 40.15% of total extra-EU French exports, respectively.

Figure 1.3: Decomposition of the extensive margin



Note: The horizontal axes display the year in which a given variety is first exported to China in RMB. These varieties are further differentiated into three categories by their prior exporting experience in China. The vertical axes display the share of each category in the value of France's RMB exports (left) and the number of exported varieties (right).

year of exporting to China ('new firm-product'). Figure 1.3 shows this decomposition for RMB exports in terms of value (left) and the number of varieties (right). It reveals a high degree of skewness in RMB exports, with old varieties capturing more than 75% of the total value in any given year. They also account for the majority of all varieties exported in RMB, although this share declined gradually from 77.5% in 2011 to 41.8% in 2017. This coincides with the growth observed in old firms invoicing new products in RMB. While such varieties accounted for 21% of all RMB varieties in 2011, their share surpassed 50% by 2017. In sharp contrast, first-time exporters to China rarely choose the RMB, as can be seen by their negligible shares in total RMB exports (0.5%) and RMB varieties (1.4%).

To further examine this skewness in RMB exports, Table 1.2 provides an informative snapshot from 2017 on how RMB exports are distributed across firms selling to China. The first three columns of the table indicate the number of products exported by the firm globally, to China and to China in RMB. By splicing the data in this manner, the table reveals three key facts. First, RMB invoicing is seen to be a relatively rare occurrence, with only 207 firms invoicing in this currency (rows iv - vii). Together, they account for less than 2% of all firms exporting to China in 2017. However, these RMB adopters are large, with their collective exports to China standing at approximately 17% of total French exports to China. This second fact suggests that there is considerable granularity in RMB invoicing. When looking at the set of RMB adopters, we note that firms that invoice multiple products in RMB account for more than 90% of total RMB exports (row vii). This leads to our third observation that RMB exports are

Table 1.2: RMB exports by firm type: Snapshot for 2017

	Global	China	RMB	# Firms	Share (%)	China exports	Share (%)	RMB Exports	Share (%)	RMB intensity
i)	1	1	0	1658	13.34	164.60	0.90	0.00	0.00	0.00
ii)	>1	1	0	4634	37.29	696.78	3.80	0.00	0.00	0.00
iii)	>1	>1	0	5928	47.70	14400.80	78.62	0.00	0.00	0.00
iv)	1	1	1	5	0.04	6.56	0.04	3.65	0.20	90.61
v)	>1	1	1	26	0.21	18.26	0.10	4.63	0.25	81.59
vi)	>1	>1	1	48	0.39	576.57	3.15	159.18	8.70	22.91
vii)	>1	>1	>1	128	1.03	2453.22	13.39	1661.63	90.84	64.35

Note: The table above disaggregates RMB exports to China in 2017 across firms. These firms are grouped into seven different categories, shown in each row, based on the number of products they export globally (column 1), to China (column 2) and to China in RMB (column 3). Subsequent columns document the number of firms under each category, their overall exports to China (across currencies) and exports to China in RMB (in millions). Alongside these values, the corresponding shares are also reported. The final column shows the share of RMB exports in total firm exports to China, averaged across firms within each category.

highly skewed even within the set of RMB-invoicing firms. The intensity of RMB usage of these multi-product firms is also high, with the average exporter amongst them invoicing 64.35% of its total exports to China in RMB.

Taken together, Figure 1.3 and Table 1.2 indicate that RMB usage is highly concentrated. Therefore, firms that invoice in this newly available currency are few in number, have prior experience of exporting to China, account for a substantial share of total French exports to China and tend to export multiple products in RMB. Figure A.3 in Appendix A.2 also shows that RMB-invoicing firms are larger in terms of global exports when compared to firms that never adopt the RMB. They serve more destinations, have a wider product scope, and invoice in multiple currencies globally. These superior characteristics of RMB firms provide preliminary evidence of self-selection mechanisms that may be underpinning the adoption of RMB.

Stylized fact 3. *RMB invoicing varies across products within firms*

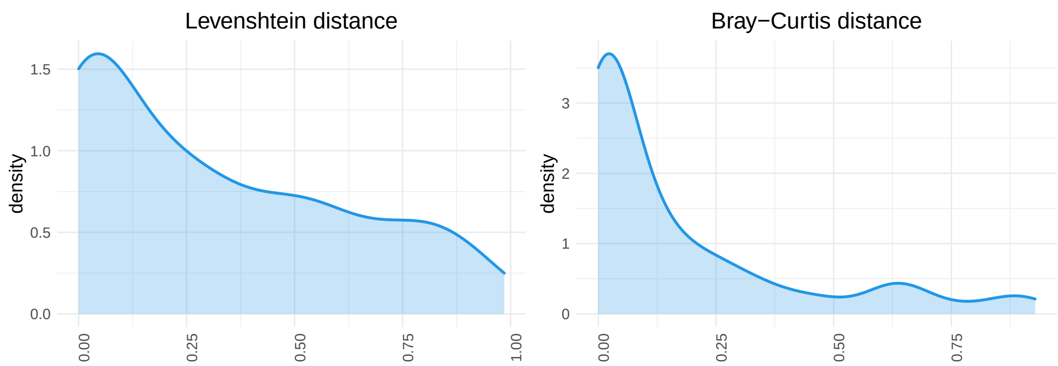
We learnt from Table 1.2 that firms invoicing multiple products in RMB dominate overall RMB exports to China. This prompts an investigation of whether RMB usage potentially varies across products within a RMB-invoicing firm. In order to examine this internal heterogeneity of the firm, I exploit a modified version of the product vector approach introduced by Fontagné, Secchi, and Tomasi (2018). In their paper, the authors construct ordered vectors of products which are exported by the firm globally (global product vector) and the subset of products exported by the firm to any given destination (local product vectors). Measures of string distances between these global and local product vectors enable them to evaluate the sparsity, fickleness and stability of export decisions made by multi-product firms.

Analogous to this approach, I construct ordered vectors of CN8 products which are exported by the firm to China (China product vector) and which are exported by the firm to China in RMB (RMB product vector). The string distance between them reveals the heterogeneity in RMB usage within the firm. For this exercise, I focus on firms that export multiple products to China, at least one of which is in RMB. For these firms, I construct two types of standard string distance measures, namely, the Levenshtein distance and the Bray-Curtis distance. Both these measures capture the extent of dissimilarity between product vectors.

The Levenshtein measure ranges from zero (identical vectors) to one (disjoint vectors) and reflects similarities in the size of product baskets exported to China and those exported to China in RMB. The Bray-Curtis measure also ranges from zero (disjoint vectors) to one (identical vectors) and reflects differences in the *shape* of these vectors.[12] Plotting these measures in Figure 1.4 reveals that there are many firms for which RMB invoicing is only partial as reflected by the range of intermediate values lying between zero and one. Since firms export 33 products to China on average, the mean Levenshtein distance indicates that 11 changes are typically required to transform their RMB product vector into the China product vector. Therefore, there is selection within the firm regarding which products are exported to China in RMB.

[12]When product vectors are coded as a sequence of zeroes and ones, the Levenshtein distance can be calculated as the number of steps (eg. replacement, addition, deletion) required to transform the RMB product vector into the China product vector. This measure is then normalized by the total number of products exported by the firm to China minus one. When the vectors are defined using product export shares instead, the Bray-Curtis measure can be computed to reflect differences in the shape (distribution of elements). For further details on the construction of these measures, refer to Appendix A.1 and Fontagné, Secchi, and Tomasi (2018).

Figure 1.4: Comparison of China and RMB product vectors



Note: The figures above are density plots of string distances measured between two ordered vectors: the vector of products exported to China and the vector of products exported to China in RMB. These string distances reflect how the RMB product vector can be transformed into the China product vector based on i) number of additions, deletions or replacements (Levenshtein) and; ii) the distribution of export shares (Bray-Curtis).

1.5 Selection into RMB Invoicing

Building on descriptive evidence provided in Section 1.4, I now investigate various channels driving selection into RMB invoicing. The first channel corresponds to the firm's hedging behaviour. Here, I examine whether a firm importing in RMB in a given year also invoices its exports in RMB. By reducing currency mismatch in its cash flows in this manner, the firm can synchronise changes in input costs with export prices and lower the exchange rate risk.[13]

The second channel relates to the fixed costs of currency risk management incurred by the firm. These fixed costs can arise from gathering experience in currency monitoring to setting up internal departments or purchasing services that can track the company's currency exposure. Here, I test whether firms that invoice in multiple currencies globally (and therefore may have already incurred such fixed costs) are more likely to adopt the RMB for settling transactions.[14]

The third channel is associated with the currency choice of competing French exporters to China. This strategic complementarity mechanism follows Crowley, Exton, and Han (2020), where competitors' marginal cost is used as an instrument for their

[13]Other forms of such operational hedging by the firm include geographical diversification of export markets and suppliers. These hedging practices are distinct from financial hedging which entails the use of derivative instruments. For more details, see Döhring et al. (2008)

[14]This is a more direct proxy for currency fixed costs than those used in previous papers such as the overall export intensity of the firm and share of Eurozone countries in firms' total exports (Amiti, Itskhoki, and Konings, 2020).

price. Specifically, I examine whether competitors' import intensity in USD affects the likelihood of an exporter switching to RMB. The rationale is that if competing firms import in USD, they are likely to export to China in USD as well due to operational hedging and costs of currency management. This in turn may reduce the probability that a firm deviates from other market players by invoicing in RMB instead.

In addition to the above three channels, I propose a novel fourth mechanism that can explain the rise in RMB invoicing by French exporters. This channel exploits information available in the customs dataset on the exporter's choice of invoicing currency to all other extra-EU destinations. If LCP emerges as the optimal strategy of the firm in these other markets, it can increase the likelihood of the firm invoicing in RMB in China as well. This local currency mechanism is partially analogous to the prior experience channel proposed by Crowley, Exton, and Han (2020) who show that firms use USD for invoicing exports to new markets when they have used it persistently in their exports to existing markets. The data confirm this possible harmonization across destinations in firms' invoicing strategy. I observe that in any given year, more than 75% of RMB-invoiced varieties were also exported to other destinations in their local currency (see Figure A.4 in Appendix A.2).

In the following section, I jointly examine these four mechanisms and their impact on RMB invoicing by French exporters to China over 2011-2017. As noted previously, firms invoicing multiple products in RMB account for the vast majority of total RMB-denominated exports to China during this period (Table 1.2), and these also exhibit ample within-firm heterogeneity in RMB adoption (Figure 1.4). In light of these stylized facts, the baseline regression for selection into RMB is specified at the level of a variety or firm-product combination.

1.5.1 Baseline Result

The baseline specification for selection into RMB invoicing is shown in equation (1.1) where the subscripts f, p and t denote the exporting firm, CN8 product and year, respectively. Since RMB invoicing is used only when firms trade with China, the specification drops the destination dimension and focuses on exports to China. The dependent variable is RMB_{fpt}, which takes the value of one for varieties that are invoiced in RMB when exported to China and zero otherwise.

The four mechanisms described above are included as $Channels^{k}_{fpt}$ and constitute the main variables of interest. Additional controls are included in the vector $\mathbf{Z}_{fpt}$ such as dummies for whether a variety is traded with other Asian economies and whether it is also imported from China. I also control for the absolute size of a variety as proxied by exports of that variety to all extra-EU destinations except China. Finally, $\mathbf{Z}_{fpt}$ includes the relative size of the variety as measured by its share in total French exports to China in the given CN8 product category.

$$RMB_{fpt} = \sum_{k=1}^{\mathcal{K}} \beta^k Channels^k_{fpt} + \mathbf{Z}_{fpt}\Gamma + \theta_{ft} + \theta_{fp} + \theta_{pt} + \epsilon_{fpt} \tag{1.1}$$

Alongside these control variables, equation (1.1) includes all possible fixed effects such that the β^k coefficients can still be identified. Supply-side shocks such as changes to firm employment, input costs, TFP or investment activity are controlled for with firm-time fixed effects. Second, firm-product fixed effects control for time-invariant factors such as the exporter's product-specific capability or market knowledge. Third, product-time fixed effects control for possible demand-side shocks common to all exporters. With this rich set of fixed effects, time-varying aggregate macroeconomic variables such as GDP growth, inflation or exchange rate fluctuations can be additionally controlled for. Therefore, the specification exploits only the variation in currency choice within firm-product combinations. This is feasible since the availability of RMB introduced additional variation in invoicing behaviour which otherwise exhibits a high degree of stability in established currencies such as the USD or EUR. Finally, the error term is clustered by firm and product.

Having outlined the baseline specification, I now turn to the construction of the four $Channels^k_{fpt}$. The first channel on firms' operational hedging is captured by a dummy indicating whether the firm imports any product in RMB in that year. However, the presence of firm-time fixed effects in equation (1.1) prevents the coefficient on this term from being identified. Therefore, I interact this with variety size, constructing a dummy for varieties that lie above the median value within a given HS-4 digit industry. The second channel on currency management fixed costs is captured by including the log of the total number of currencies used by the firm in its extra-EU exports (other than China) and excluding the use of USD given its unique status as the international reserve currency. The third channel of strategic complementarity is captured by the share of USD in competing firms' total imports (from all source countries), multiplied by their export share in the given CN8 product in China.[15] Finally, a novel channel is included in the form of a dummy that takes the value of one if the variety is exported to other destinations (excluding the US) in their respective local currencies.

Results are reported in Table 1.3 where the stringency of fixed effects employed increases as we move from columns (1) to (3). Amongst these, the most rigorous specification is that of column (3), where fixed effects are defined at the highest level of product disaggregation i.e. the CN8 product category rather than industry levels (HS-

[15] This is calculated as $\sum_{k \neq f} \frac{S_{kpt}}{1-S_{fpt}} \times importshare^{USD}_{kt}$ where $S_{fpt} = \frac{Export_{fpt}}{\sum_f Export_{fpt}}$. For further details on construction see Crowley, Exton, and Han (2020).

4 or HS-6 digit). Examining the first channel, results across the columns show that importing in RMB from China increases the probability of RMB invoicing for exported varieties above the median size. This effect is approximately 1.5 percentage points under column (3) and increases slightly to 1.7 percentage points when fixed effects are defined at the HS 4-digit industry level in column (1). This is a first confirmation of the evidence provided by previous literature on the role of operational hedging in invoicing decisions of the firm.

The results also clearly demonstrate that the second channel, relating to currency management, is pertinent for the adoption of RMB. Looking at column (1), we find that doubling the number of invoicing currencies used in exporting a variety globally generates a 3.5 percentage point increase in the probability of it being invoiced in RMB in China. The magnitude of this effect declines to 2.3 percentage points in column (3) when firm-product and product-time fixed effects are defined at the CN8 level. However, the coefficient remains statistically significant at the 1% level. Therefore, the adoption of RMB increases with the overall size of the firm's currency basket and implicitly, the firm's underlying currency management experience.

In contrast to the first two channels, the evidence on the third channel concerning strategic complementarity is mixed. While competitors' use of USD in their imports decreases RMB invoicing by 0.7 percentage points under column (1), this effect disappears in subsequent columns where firm-product and product-time fixed effects are defined at the HS 6-digit or CN8 level rather than the HS 4-digit level. Overall, the significance of this channel for RMB adoption is not as clear.

Moving to the final row of coefficients reported in Table 1.3, we observe that varieties which are invoiced in local currencies in other destinations are more likely to be priced in RMB when exported to China. This effect is statistically significant at the 5% level across the three columns and therefore withstands the inclusion of high-dimensional fixed effects. In column (3), the effect translates to an increase of 0.9 percentage points in the probability of RMB adoption for the variety. This result offers a novel insight into the synchronization of invoicing strategies across export destinations and does so by establishing a clear empirical link between LCP in other markets and the adoption of a newly available local currency. In the following section, I investigate how each of these four channels themselves exhibits heterogeneity depending on a range of variety-specific characteristics.

Table 1.3: Selection of varieties into RMB: Varying fixed effects

Dependent Variable:	RMB_{fpt}		
Model:	(1)	(2)	(3)
importing in RMB x median	0.017***	0.013**	0.015***
	(0.005)	(0.006)	(0.006)
log(#currencies)	0.035***	0.026***	0.023***
	(0.010)	(0.008)	(0.007)
competitor USD	-0.007**	-0.004	-0.002
	(0.003)	(0.004)	(0.010)
local currency	0.016**	0.012**	0.009**
	(0.006)	(0.005)	(0.004)
Firm x Time	✓	✓	✓
Firm x HS4	✓		
HS4 x Time	✓		
Firm x HS6		✓	
HS6 x Time		✓	
Firm x CN8			✓
CN8 x Time			✓
Observations	244,731	244,731	244,731
R^2	0.886	0.927	0.937

Note: The table reports estimation results following the specification outlined in equation (1.1). The dependent variable takes the value of one for varieties (firm-CN8 combinations) that are invoiced in RMB when exported to China and zero otherwise. Columns vary in the stringency of fixed effects, with the most rigorous (baseline) specification reported in column (3). Standard errors are clustered by firm and product. Significance codes: ***: 0.01, **: 0.05, *: 0.1

1.5.2 Heterogeneity in Selection

In this section, I modify the baseline specification in two ways. First, I interact each of the four channels with various product characteristics (Table 1.4). These include the following: i) a dummy if a product is differentiated as per the Rauch (1999) classification; ii) product-level trade elasticity at the HS-6 digit from the ProTEE database (Fontagné, Guimbard, and Orefice, 2020) and; iii) an industry-level measure of upstreamness at the HS-4 digit for France (Antràs and Chor, 2018). By interacting the four channels with such measures, I can examine whether the nature of the product amplifies or dampens each of the mechanisms under study.

Looking at the first channel, Table 1.4 shows that the hedging channel is not significantly different for products with different characteristics. As seen in the baseline, firm size is more crucial for driving the hedging motive than the characteristics of specific products. For the next channel, column (1) shows that currency management experience is relevant for RMB adoption only for differentiated products. This is reasonable given that homogeneous goods which are sold in organized exchanges and have an international price are likely to be exported only in USD. Column (3) also shows that this channel is less important for products that are in upstream industries. Coming to the effects of strategic complementarity, the results indicate that competitors' use of USD is less likely to induce RMB adoption in the case of differentiated varieties. Finally, Table 1.4 shows that the effect of local currency invoicing in other destinations on RMB use is marginally higher for products with higher trade elasticities.

The second modification of the baseline specification sheds light on potential non-linearities in the role of currency management fixed costs (Table 1.5). Here, I include additional dummies to the baseline that take the value of one if a variety is exported in two, three or more than three currencies to the rest of the world (column 2). Besides the total number of export invoicing currencies, the experience of invoicing in currencies from the same region may be more relevant for RMB adoption than currencies from other regions. This distinction is made in columns (3)-(5), in which I separately examine the effects of exporting in multiple currencies from Asia, Europe as well as North America (excluding USA) and South America on RMB adoption. Note that the effect of a firm exporting to any of these regions in a given year is already controlled for with the presence of firm-time fixed effects. Therefore, coefficients on these currency dummies do not simply reflect the firm's geographical diversification.[16]

[16]Due to multi-collinearity, the log of the total number of export invoicing currencies for the variety is dropped in column (2) following the inclusion of various currency dummies. However, the coefficient on this log term can still be identified in columns (3)-(5), where the currency dummies correspond only to the number of regional currencies used for export invoicing by the firm.

Table 1.4: Interactions with product characteristics

Dependent Variable:	RMB_{fpt}		
	rauch	σ_{CEPII}	upstream
Model:	(1)	(2)	(3)
import in RMB x characteristic	-0.009	0.000	0.004
	(0.008)	(0.000)	(0.007)
log(#currencies)	-0.003	0.024**	0.069***
	(0.006)	(0.010)	(0.021)
log(#currencies) x characteristic	0.032***	0.000	-0.021***
	(0.008)	(0.001)	(0.008)
competitor USD	0.032***	-0.004	-0.030
	(0.012)	(0.011)	(0.024)
competitor USD x characteristic	-0.042***	-0.000	0.013
	(0.014)	(0.000)	(0.009)
local currency	0.014***	0.018***	0.009
	(0.005)	(0.007)	(0.022)
local currency x characteristic	-0.004	0.001*	-0.000
	(0.008)	(0.001)	(0.009)
Observations	228,603	245,487	242,526
R^2	0.939	0.937	0.937

Note: The dependent variable takes the value of one for varieties (firm-CN8 combinations) that are invoiced in RMB when exported to China and zero otherwise. Each column features interactions between a given product characteristic and the four selection channels. All regressions include firm × time, firm × CN8 and CN8 × time fixed effects. Standard errors are clustered by firm and product. Significance codes: ***: 0.01, **: 0.05, *: 0.1

Looking at the results in Table 1.5, column (2) reveals that the positive effect of currency management on RMB adoption only kicks in when the variety is invoiced in four or more currencies in exports to the rest of the world. Having such a wide currency basket therefore increases the likelihood of RMB invoicing by 4.9 percentage points. Investigating regional effects, column (3) shows that the experience of invoicing in Asian currency matters for RMB adoption and is increasing in the number of Asian currencies used. For instance, invoicing the variety in four or more Asian currencies increases the likelihood of RMB adoption by nearly 12 percentage points.

This suggests that regional knowledge of currencies may be an important factor in driving selection into RMB invoicing. The effect is in addition to having large currency baskets in general, as the coefficient on log number of currencies continues to be positive and statistically significant. In comparison, there is no discernible effect of exporting in multiple currencies from Europe, North or South America on RMB use. This is apparent from columns (4)-(5), where none of the region-specific currency dummies have statistically significant coefficients.

Table 1.5: Role of currency experience

Dependent Variable:			RMB_{fpt}		
	Baseline	Global	Asia	Europe	Americas
Model:	(1)	(2)	(3)	(4)	(5)
importing in RMB x median	0.015***	0.015***	0.015**	0.015***	0.015***
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
local currency	0.009**	0.009**	0.013***	0.009**	0.010**
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
competitor USD	-0.002	-0.002	-0.003	-0.002	-0.002
	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)
log(#currencies)	0.023***		0.011**	0.021***	0.022***
	(0.007)		(0.005)	(0.007)	(0.007)
#currencies=2		0.001	0.050***	0.017	0.017
		(0.002)	(0.011)	(0.015)	(0.017)
#currencies=3		0.010*	0.069***	0.012	-0.003
		(0.006)	(0.018)	(0.013)	(0.023)
#currencies≥4		0.049***	0.120***	0.017	0.057
		(0.012)	(0.031)	(0.018)	(0.039)
Observations	244,731	244,731	244,731	244,731	244,731
R^2	0.937	0.937	0.938	0.937	0.937

Note: The dependent variable takes the value of one for varieties (firm-CN8 combinations) that are invoiced in RMB when exported to China and zero otherwise. In the columns, new dummies are introduced to equation (1.1) which correspond to the number of invoicing currencies from different regions that are used in exporting the variety. All regressions include firm × time, firm × CN8 and CN8 × time fixed effects. Standard errors are clustered by firm and product. Significance codes: ***: 0.01, **: 0.05, *: 0.1

Having examined the role of product-level heterogeneity and size of currency baskets in driving RMB adoption, I now proceed towards a closer examination of the fourth and novel mechanism – that of local currency use in other destinations. To the baseline specification, I now include an additional term where the local currency dummy is interacted with firm-product-time varying characteristics.

These characteristics include i) a dummy for varieties above the median size; ii) share of global exports of a given variety that is invoiced in local currencies abroad in a given year; iii) a dummy if the variety is invoiced in the local currency in more than two destinations in a given year and; iv) a dummy if the variety has been invoiced in a local currency for more than two years. These interaction terms will serve to illustrate how the local currency channel operates.[17] The results are reported in Table 1.6, where column (1) reports the baseline while columns (2)-(5) additionally show coefficients on the new terms.

From column (2), we observe that the local currency channel operates only for relatively large varieties that exceed the median size. For these varieties, local currency invoicing in other destinations increases the probability of RMB invoicing in China by 3.3 percentage points. From column (3), we find that RMB invoicing is also increasing in the intensity with which the variety is invoiced in local currency abroad. Finally, columns (4)-(5) reveal that the likelihood of RMB invoicing is magnified for varieties that are exported in local currencies to more than two destinations and which have more than two years of experience in being exported in local currencies. To conclude, the results here provide novel evidence that the intensity and experience of local currency use in exports to the rest of the world play an important role in driving RMB adoption in French exports to China.

[17] Data for China and the US are excluded in the construction of all these variables.

Table 1.6: Heterogeneity in the local currency channel

Dependent Variable:	RMB_{fpt}				
	Baseline	Size	Share	Destinations	Experience
Model:	(1)	(2)	(3)	(4)	(5)
local currency	0.009**	-0.021*	0.008*	0.009**	0.005
	(0.004)	(0.012)	(0.004)	(0.004)	(0.004)
local currency x characteristic		0.033***	0.017*	0.037***	0.020***
		(0.012)	(0.010)	(0.008)	(0.006)
R^2	0.937	0.937	0.937	0.938	0.937

Note: $N = 244,731$. The dependent variable takes the value of one for varieties (firm-CN8 combinations) that are invoiced in RMB when exported to China and zero otherwise. After the baseline result in column (1), subsequent columns feature interactions between the local currency dummy and a variety-specific characteristic reflecting size or local currency use. These characteristics correspond to variety size being above or below the median (column (2)), share of global exports of the variety that is invoiced in local currencies in other extra-EU destinations (column (3)), dummy for the variety being invoiced in the local currency in more than two destinations (column (4)) or for more than two years (column (5)). Standard errors are clustered by firm and product. All regressions include firm $\times$ time, firm $\times$ CN8 and CN8 $\times$ time fixed effects. Significance codes: ***: 0.01, **: 0.05, *: 0.1

1.6 Impact of RMB Invoicing on Exports

Section 1.5 provided evidence of the various mechanisms driving selection of varieties into RMB invoicing. Controlling for this selection, what impact might RMB invoicing have on export sales of French firms in China? To examine this second issue, I exploit the novel channel of local currency use as an instrumental variable (IV). The proposed IV strategy therefore exploits the fact that RMB invoiced varieties tend to be priced in local currencies even when exported to other destinations.

This result was established in the baseline estimation reported in Table 1.3, wherein local currency pricing in other markets increased the likelihood of RMB adoption in China by 0.9 percentage points. Therefore, local currency adoption in other destinations is a strong predictor for RMB invoicing and meets the 'relevance' criteria for an IV. It is also arguably independent of demand shocks experienced by the variety in China. Ex-ante, there is also no clear rationale why buyers in China would have a strong preference for varieties that are sold elsewhere in the local currencies of these other markets.[18]

[18] Note that the over-identifying restrictions test (or the J-test) cannot be conducted here as this is the case of a single endogenous variable and a single instrument. However, I do attempt to provide further confidence in this instrument through two additional checks wherein I focus on firms that already exported to China prior to the RMB reforms or limit the sample to varieties that were invoiced in RMB at least once over 2011-2017.

With these features, local currency use in other destinations can be considered a reasonable candidate to instrument for RMB adoption in China. This novel IV has not been exploited in prior research and can address the lack of causal estimates of invoicing choice on trade flows in the literature. Based on this identification strategy, equation (1.2) is estimated. Here, the dependent variable $\log(Y_{fpt})$ is either the log of firm export sales or exported quantity (in kilos) or export prices (ratio of export revenue to quantity) in China.

$$\log(Y_{fpt}) = \beta \text{RMB}^{IV}_{fpt} + \mathbf{Z}_{fpt}\Gamma + \theta_{ft} + \theta_{fp} + \theta_{pt} + \epsilon_{fpt} \tag{1.2}$$

The main variable of interest, RMB invoicing, is instrumented with a dummy that takes the value of one if the given variety is invoiced in local currencies in other destinations (excluding the US) in that year. As before, all possible fixed effects are included such that the coefficient on predicted RMB invoicing derived from the first-stage (RMB^{IV}_{fpt}) can still be identified. These fixed effects include firm-time, firm-product and product-time fixed effects to control for unobservables relating to demand or supply-side shocks as well as macroeconomic conditions such as exchange rate movements.

In addition to these fixed effects, I control for the absolute size of a variety which is measured as the log sales of the given variety to all destinations excluding China and its relative size as measured by its share in France's total exports to all destinations except China. These are captured by the vector $\mathbf{Z}_{fpt}$ in equation (1.2). Table 1.7 reports the results for both OLS and two-stage least-squares (TSLS) estimations alongside the F-test statistics. Causal interpretation of the coefficients relies on the assumption that local currency invoicing in foreign destinations is orthogonal to unobserved shocks faced by the variety in China.

Looking at the OLS coefficients in columns (1), (3) and (5), RMB invoicing is seen to have a positive and statistically significant impact on firms' export performance in China. Invoicing in RMB raises firm exports by 108.9%, exported quantity by 90.2% and export prices by 9.8%. However, these OLS estimates are likely to be biased due to the selection of older and larger varieties into RMB adoption as described in Section 1.5. Therefore, columns (2), (4), (6) report results from TSLS regressions where RMB invoicing is instrumented with the local currency use dummy. The F-statistic from the first stage regression is high, indicating that the regressions do not suffer from a weak instruments problem. With TSLS, the impact of RMB invoicing on firm exports is observed to be lower. Now, invoicing in RMB increases firm exports by 51.4% and exported quantity by 73.2%. Unlike OLS estimates, there is no statistically significant change in export prices when invoicing in RMB.

When comparing effect sizes across these models, the OLS estimates are found to exhibit substantial upward bias. This is to be expected, given the positive selection of varieties into RMB invoicing. Correcting for such selection with an IV reduces the magnitude of coefficients but retains their statistical and economic significance. Overall, the results indicate the export advantage that firms gain from invoicing in local currencies in competitive markets. Moreover, firms raise their export revenues when invoicing their products to China in the local currency by increasing the volume of sales rather than the adjustment of prices.

Table 1.7: Causal impact of RMB invoicing: IV approach

Dependent Variables:	log(exports)		log(quantity)		log(price)	
	OLS	TSLS	OLS	TSLS	OLS	TSLS
Model:	(1)	(2)	(3)	(4)	(5)	(6)
RMB	0.737***	0.415***	0.643***	0.549***	0.094**	-0.133
	(0.113)	(0.158)	(0.112)	(0.148)	(0.045)	(0.091)
Observations	171,013	171,013	171,013	171,013	171,013	171,013
R^2	0.939	0.939	0.953	0.953	0.962	0.962
IV coefficient, first stage	–	0.024***	–	0.024***	–	0.024***
F-statistic, first stage	–	307.85	–	307.85	–	307.85

Note: The dependent variable is either the log of exports or exported quantity (in kilos) or export prices for a given variety in China. Columns (1), (3) and (5) report OLS estimates from regressing these against a RMB invoicing dummy. In contrast, columns (2), (4) and (6) report second-stage regression results when RMB invoicing is instrumented with a dummy that takes the value of one if the variety is invoiced in local currencies in other destinations (excluding the US) in that year. All regressions include firm $\times$ time, firm $\times$ CN8 and CN8 $\times$ time fixed effects. Additional controls for the absolute and relative size of the variety in France's global exports (excluding China) are also included. Standard errors are clustered by firm and product. Significance codes: ***: 0.01, **: 0.05, *: 0.1

The IV estimation as shown in Section 1.6 may be sensitive to the precise definition of the local currency use channel. To test this, I replace the local currency use dummy with alternative proxies for this channel. These proxies are identical to those constructed in Table 1.6 and reflect the experience of the variety in being invoiced in local currency abroad (excluding China). The TSLS results from employing these variables as instruments for RMB adoption are shown in columns (2) and (3) of Table 1.8. For comparison, column (1) repeats the previous TSLS regression with local currency use dummy as the IV. The high values for the first stage F-statistic in columns (2) and (3) confirm that these proxies remain relevant instruments for RMB invoicing. Even when shifting to these proxies, the effect of RMB use on firm exports remains high and ranges from 37.7% to 99.4%. As before, the IV estimates are lower than the OLS coefficients seen in Table 1.7.

Table 1.8: Alternate proxies for the IV

Dependent Variable:	log(exports)		
Model:	(1)	(2)	(3)
RMB	0.415***	0.320**	0.690***
	(0.158)	(0.156)	(0.149)
Observations	171,013	171,013	171,013
R^2	0.939	0.939	0.939
Local currency IV	dummy	experience	destinations
IV coefficient, first stage	0.024***	0.019**	0.045***
F-statistic, first stage	307.85	134.17	678.5

Note: The dependent variable is the log of exports for a given variety in China. Columns report second-stage regression results when RMB invoicing is instrumented with different variables, all of which reflect the variety being invoiced in local currencies in other destinations. In column (1), the instrument is a dummy that takes the value of one if the variety is invoiced in local currencies in any other destination (excluding the US) in that year and is therefore identical to column (2) in Table 1.7. In columns (2) and (3), the instrument is a dummy that equals one for varieties exported in local currencies elsewhere for more than two years or to more than two destinations. All regressions include firm $\times$ time, firm $\times$ CN8 and CN8 $\times$ time fixed effects. Additional controls for the absolute and relative size of the variety in France's global exports (excluding China) are also included. Standard errors are clustered by firm and product. Significance codes: ***: 0.01, **: 0.05, *: 0.1

Even after instrumenting for RMB invoicing and the inclusion of a rich set of fixed effects, selection into RMB may still play a role. To further exclude this possibility, I next restrict the sample in two ways. First, I retain only those firms that were already exporting to China prior to 2011. This drops any firm that began exporting into China following the RMB reforms. The OLS and TSLS results from this modification are reported in Panel A of Table 1.9. Once again, the OLS estimates exhibit an upward bias when compared to the TSLS coefficients. After instrumenting for RMB adoption with the local currency dummy, firm exports are seen to grow by 57.1%. This increase is driven by a strong rise in exported quantity (80.9%) rather than changes in price.

In Panel B of Table 1.9, I further limit the sample by keeping only those varieties that were invoiced in RMB at least once over 2011-2017. Doing so reduces the number of observations by over 90% relative to Table 1.7 but still allows us to exploit the time variation in RMB use within and across the 4600 varieties which selected into RMB adoption during 2011-2017. While these varieties exhibit variation in RMB invoicing over time, they are consistently exported to other destinations in local currencies. Therefore the local currency IV remains strongly correlated with RMB adoption, as can be seen by the first-stage F statistic. Despite limiting the sample in this manner, RMB invoicing continues to show a high and positive impact on firm exports to China.

Table 1.9: TSLS results with restricted samples

Dependent Variables:	log(exports)		log(quantity)		log(price)	
Model:	OLS	TSLS	OLS	TSLS	OLS	TSLS
Panel A: Firms exporting to China prior to 2011						
RMB	0.734***	0.452***	0.649***	0.593***	0.085*	-0.141
	(0.115)	(0.167)	(0.115)	(0.156)	(0.045)	(0.091)
Observations	140,745	140,745	140,745	140,745	140,745	140,745
R^2	0.932	0.932	0.949	0.948	0.957	0.957
IV coefficient, first stage	–	0.025***	–	0.025***	–	0.025***
F-statistic, first stage	–	272.52	–	272.52	–	272.52
Panel B: Varieties exported in RMB at least once						
RMB	0.809***	0.759**	0.651***	0.624	0.158**	0.134
	(0.156)	(0.356)	(0.145)	(0.393)	(0.079)	(0.223)
Observations	14,408	14,408	14,408	14,408	14,408	14,408
R^2	0.954	0.953	0.959	0.959	0.958	0.958
IV coefficient, first stage	–	0.099***	–	0.099***	–	0.099***
F-statistic, first stage	–	108.27	–	108.27	–	108.27

Note: The dependent variable is either the log of exports or exported quantity (in kilos) or export prices for a given variety in China. In Panel A, the sample is restricted to firms that exported to China at least once prior to the RMB reforms in 2011. In Panel B, the sample is restricted to varieties that were exported at least once in RMB to China over 2011-2017. Columns (1), (3) and (5) report OLS estimates whereas, columns (2), (4) and (6) report second-stage regression results when RMB invoicing is instrumented with a dummy that takes the value of one if the variety is invoiced in local currencies in other destinations (excluding the US) in that year. All regressions include firm $\times$ time, firm $\times$ CN8 and CN8 $\times$ time fixed effects. Additional controls for the absolute and relative size of the variety in France's global exports (excluding China) are also included. Standard errors are clustered by firm and product. Significance codes: ***: 0.01, **: 0.05, *: 0.1

1.7 Conclusion

This paper exploits highly disaggregated customs data from France in order to examine the impact of China's RMB reforms on individual firms' invoicing behaviour. It is the first paper that analyzes RMB adoption at the firm-level and that provides novel evidence on the (positive) causal impact of RMB invoicing on export performance. In doing so, it contributes to the vibrant literature on the rich heterogeneity in invoicing currency choices across firms and existing research on the impact of RMB internationalization that so far relies on aggregate payments data at the country-pair level.

The results in this paper suggest that China's RMB internationalization policies have had positive but limited effects on the invoicing behaviour of exporting firms in France. The majority of firms continue invoicing in USD or EUR when exporting to China, especially those that newly enter the market. To expand currency usage, PBC promotion policies should therefore target the unique obstacles faced by small and new exporters eg. in accessing RMB-denominated trade finance and managing exchange rate fluctuations. The results also reveal that USD dominance is difficult to challenge, particularly its usage as a vehicle invoicing currency for exports to third markets. However, RMB may still emerge as an important regional currency for settling trade in Asia.

This points towards several avenues that remain open for future research. Further work can investigate the extent to which RMB invoicing changes when firms' buyers or suppliers are themselves trading with China. Information on the other side of the transaction is not available for extra-EU exports in the case of France, but such buyer-supplier customs data can shed light on RMB use within regional value chains centred around China. Since data on firms' invoicing currency only begins from 2011 in France, future research can also examine RMB adoption in other countries where customs data potentially report longer currency histories that predate China's RMB reforms. Such research can further add to our understanding of the RMB's evolving role in global trade and financial markets.

Chapter 2

Trade Liberalization Along the Firm Size Distribution: The Case of the EU-South Korea FTA

2.1 Introduction

Free trade agreements (FTAs) are regularly criticized for privileging the interests of the largest firms. Such concerns have contributed to public resistance against mega-regional FTAs such as the Transatlantic Trade and Investment Partnership (TTIP) and the EU-Canada Comprehensive Economic and Trade Agreement (CETA). Since then, policymakers have worked to include chapters in FTAs that are dedicated to supporting small enterprises. However, for such provisions to be effective, an improved understanding of the distributional effects of modern FTAs is required.

While there is ample research on selection effects from trade liberalization, see, e.g., the survey by Melitz and S. Redding (2014), evidence on the heterogeneous impact of lower tariffs or non-tariff barriers (NTBs) on *continuing* exporters remains scarce. The latter constitutes another, hitherto under-explored channel through which trade liberalization could further increase pre-existing inequality across firms, with implications for labour markets, social welfare and the design of FTAs. We address this gap in the literature by comparing the magnitude of FTA effects along the size distribution of incumbent exporters and examining these size-differentials over time and across sectors. Using customs data from France, we find robust evidence of size heterogeneity in export growth along the intensive margin of incumbent exporters.

Our main result is that the agreement boosted sales of incumbent exporters (i.e., at the intensive margin) in the top quartile of the size distribution by about 71 percentage points more than firms in the bottom quartile. The result is mostly driven by lower NTBs, but larger firms appear to react more strongly to lower trade costs too.[1] While

[1] In many models with rent-sharing, the wage distribution follows the size distribution Helpman, Itskhoki, and Redding (2010). Hence, our result suggests that the FTA may have increased wage inequality in France.

the result on the intensive margin is at odds with the standard Melitz (2003) model, our findings on exporter entry and the product-level extensive margin are in line with theory: the FTA does indeed increase the likelihood of export participation and the number of products exported by medium-sized firms.

Our key result therefore sheds doubt on the frequent assumption that all firms face identical (variable and fixed) trade costs and demand elasticities. Interestingly, leading models that relax one of these assumptions predict the opposite of what we observe. For example, models featuring linear demand systems such as Melitz and Ottaviano (2008) imply that more productive (and, hence, larger) exporters expand sales by less than their less productive peers when trade costs fall, as they face less elastic demand. Also, Arkolakis (2010) predicts that trade liberalization boosts the sales of larger exporters by lower rates than those of smaller firms because of increasing marginal foreign market penetration costs.

Our intensive margin result would be consistent with a configuration where the FTA lowers trade barriers more strongly for larger firms.[2] For instance, it is conceivable that detailed provisions of FTAs such as rules of origin reflect the interests of dominant firms rather than of smaller firms. Alternatively, our result would emerge if larger firms react more strongly to identical trade cost reductions. This would be the case if, unlike in Arkolakis (2010), marginal market access costs are decreasing in sales, or if taking advantage of the FTA entails recurring investment that firms with larger sales find easier to undertake.[3]

The EUKFTA is an excellent case to study the effects of trade liberalization along the firm size distribution. First, the agreement concerns two sizeable advanced economies and was the largest EU FTA in terms of joint market size when it entered into force in 2011.[4] Second, it is an ambitious agreement that mandated the reduction to zero of 94% of all EU tariff lines and 80% of South Korean tariff lines within the first year. This implies that the size of tariff cuts was primarily determined by the pre-existing level of MFN tariffs, alleviating concerns regarding endogeneity. Third, the EUKFTA is still considered the prototype of a deep new-generation trade agreement with ambitious language on NTBs, both at the sectoral (vertical) and the cross-sectoral (horizontal) level; see Mattoo, Rocha, and Ruta (2020).

In terms of methodology, we adopt a triple-difference approach that includes the most extensive set of three-way fixed effects to minimize omitted variable bias and other

[2] This could be the case for usually unobservable NTBs; tariffs, in contrast, do not vary across firms within products (but larger firms could cluster in product categories facing larger cuts).

[3] Such investment could be related to proving rules of origin which require firms to furnish extensive documentation.

[4] It remained the EU's biggest FTA until 2017, when the EU-Canada Comprehensive Economic and Trade Agreement (CETA) began to be provisionally applied.

sources of possible endogeneity.[5] To study the heterogeneous effects of NTB reductions on firms, we employ a novel 'umbrella' approach inspired by the gravity literature (Baier and Bergstrand, 2007). Essentially, this amounts to an events-study technique where the application of the FTA is summarized by an indicator variable. By definition, the indicator variable captures *all* trade effects attributable to NTBs, since tariff cuts are precisely observable so that their effects can be netted out. This strategy bypasses the need to measure the wide range of NTBs addressed by the FTA and therefore complements existing literature that uses specific proxies of NTBs (such as concerns raised by countries about technical barriers to trade or sanitary and phytosanitary measures).

Finally, the long panel dimension of French customs data enables us to deal with another methodological challenge – the anticipation of trade liberalization by firms. Official negotiations over the EUKFTA began in 2007, with the FTA entering into force in 2011. To deal with firm anticipation, we compare export performance after the inception of the FTA (2011-2016) to the period prior to official negotiations (2000-2006). In addition, we take an agnostic approach to constructing the control group of countries by including all export markets that are reported in the customs data (besides South Korea).

Our paper is related to several strands of research. First, it extends prior firm-level literature on the impact of trade liberalization by jointly studying the role of tariff and NTB reductions on firms' trading activities. Earlier work such as Iacovone and Javorcik (2010) on NAFTA and Bustos (2011) on MERCOSUR has examined the impact of tariff reductions on a variety of firm outcomes. NTBs and their impact on firms have also been discussed, although separately, in papers such as Fontagné and Orefice (2018) on technical barriers to trade (TBTs) and Fontagné, Orefice, Piermartini, and Rocha (2015) on sanitary and phytosanitary (SPS) measures.[6] Much of this literature has focused on the selection effect (i.e., the extensive margin of trade liberalization) while we have a special interest in *differential* size-effects along the intensive margin of continuing exporters.

Our findings also contribute to a small but growing literature that reports the advantages of large firms from NTB reductions; see Fontagné, Orefice, and Piermartini (2020) on border formalities, Carballo et al. (2016) on border entry timings, and Karpaty and Tingvall (2015) on corruption. This paper broadens the analysis to the wide range of NTBs that are addressed by a deep FTA, compares the impact of NTBs to tariffs, and studies the variation in NTB effects over time and across sectors. Al-

[5]Our regressions allow for multiple three-way fixed effects such as firm-product-destination, firm-destination-time and firm-product-time fixed effects.

[6]Our work also relates to the extensive gravity literature on the trade effects of FTAs, which uses aggregate data; see Head and Mayer (2014) or Yotov et al. (2008).

though our primary focus lies on FTA effects for incumbent exporters, we also provide an extension for the import side.

Third, we contribute to prior research on the trade effects of the EUKFTA. Lakatos and Nilsson (2017) use monthly EU-wide trade data at the 8-digit product level in a gravity-type model. They report an increase of 11.2% in the probability to export and a 10.7% increase in the value of EU exports. We are aware of only one existing study that uses firm-level data to evaluate EUKFTA: Kasteng and Tingvall (2019) use transaction-level import data for Swedish firms for one month (November 2016) to examine preference utilization rates.[7]

The rest of the paper is structured as follows. Section 2.2 provides the necessary background for the EUKFTA and discusses the tariff liberalization and NTB reductions envisioned by the agreement. It describes the customs database used and our measure of firm size. Section 2.3 sets out our empirical methodology for examining the impact of the agreement along the size distribution and shows baseline results. A range of robustness checks are reported in Section 2.4. Finally, concluding remarks and policy implications are presented in Section 2.5.

2.2 The EU-South Korea FTA

2.2.1 The EU-South Korea FTA as a Prototypical "New Generation" FTA

The EUKFTA was the first trade agreement signed by the EU with an Asian economy. Formal negotiations were launched in 2007 and after eight official rounds of talks, the agreement was signed in 2010. Following ratification in parliaments, the FTA was applied from July 2011 onwards. Since then, the EUKFTA has become a model for the EU's 'new generation' FTAs because of its unprecedented scope, depth and speed of liberalization. In that, it differs from earlier agreements and is an excellent example of what the literature refers to as a 'deep' trade agreement (see Dür, Baccini, and Elsig, 2014; Mattoo, Rocha, and Ruta, 2020). In particular, the commitments under EUKFTA extend beyond tariff reductions to so-called WTO-X provisions covering competition policy, intellectual property rights and capital mobility. The EUKFTA also features

[7] Using a New Quantitative Trade Theory (NQTT) model with multiple countries, multiple sectors and value chains following Caliendo and Parro (2015), European Commission (2017) examine the general equilibrium effects of the EUKFTA using GTAP data up to 2014. They find that the agreement boosted the EU's GDP by approximately 4.4 € billion and increased EU exports to South Korea by 42% relative to the benchmark scenario with no FTA.

provisions on transparency and regulatory stability aimed at supporting small firms.[8]

The agreement led to deep tariff cuts across the board. Upon its implementation in 2011, most industries experienced rapid liberalization with duties completely eliminated. In all, South Korea eliminated nearly 64% of its tariff lines immediately, with another 16% of tariff lines being already duty-free. Approximately 1.8% of tariff lines were phased out over ten years and longer, largely for relatively sensitive products in the agri-food and textiles sectors. As a consequence, for EU exporters, the simple average of South Korean duties fell from 12.1% to 6.2% upon entry in force, and, within five years, the agreement had eliminated 98.7% of duties in trade value for both agricultural and industrial goods (European Commission, 2010). In 2010, the simple average of the EU's applied MFN tariffs faced by South Korean exporters stood at 5.1%. With the FTA's implementation, this was reduced, essentially without phase-in, to approximately 0.5% in 2011.[9]

In our analysis, we use a complete global matrix of applied bilateral tariffs at the HS 6-digit product level that is drawn from Felbermayr, Teti, and Yalcin (2019). This database includes the phasing-out of tariffs from FTAs and fills in missing MFN tariffs by examining the nearest preceding or succeeding observation. Information on a given product's tariff staging category is drawn from the FTA tariff schedules, available through the WTO's RTA database.

In comparison to tariffs, the precise measurement of NTBs poses several challenges. This stems from the fact that NTBs group together all frictions to trade other than tariffs and tariff-rate quotas. They include impediments to trade that arise from geographic and historical factors such as distances, cultural norms, languages and institutional frameworks as well as 'behind-the-border' policy measures. The EUKFTA acted upon the latter through a range of provisions. Amongst other trade reforms, South Korea lowered the burden of third-party testing for EU electronics, recognized UNECE as the relevant standard-setting body for motor vehicles and agreed to policy coordination in SPS and TBT measures. The agreement also featured several 'horizontal' clauses that would benefit all sectors, e.g., by improving transparency, availability

[8]Dedicated SME chapters have been introduced in EU or US FTAs such as the EU-Japan Economic Partnership Agreement (EUJEPA) or the US-Mexico-Canada-Agreement (USMCA). They also feature in all agreements currently under negotiation.

[9]Figure B.1 in the Appendix B.1 provides details. These tariff cuts are important for the EU-South Korea relationship, given the role of goods trade in overall trade between their economies. In 2010, the year before the FTA's implementation, 79% of the EU's total exports to and 89% of the EU's total imports from South Korea comprised of goods.

of information and customs facilitation.[10] In our analysis, we work with an event studies approach to capture the comprehensive effects of NTBs.

We use French data to investigate the agreement. Within the EU, France is amongst the top trade partners of South Korea. In 2016, it accounted for approximately 8.85% of EU's total goods exports to South Korea, ranking fourth after Germany (39.40%), UK (11.91%) and Italy (9.06%). France has also widened its trade surplus in goods with respect to South Korea in recent years. This surplus stood at €1.57 billion in 2016, a 45% increase over the trade surplus of €1.08 billion in 2010, the year before the FTA went into effect. Turning to the composition of trade baskets, we note that French exports to South Korea are dominated by manufacturing industries such as machinery, transport, chemicals and plastics. At a more disaggregated level, manufactured goods such as cars and car parts, other aircraft and aircraft parts, packaged medicines and electronics capture substantially high shares in overall exports.

2.2.2 Data on Trade Flows

To examine the impact of the agreement on firms, we use customs data from France over the period of 2000-2016 (dataset DGDDI, 2018). These data provide information on export sales and import purchases of French trading firms (denoted by f), disaggregated by destination or source country (d) and product (p) over time (t).[11] Services trade is not included. Since each firm is assigned a unique identifier ('SIREN'), it is possible to follow its export and import activities over time. We aggregate transactions from the monthly to yearly level and products from the 8-digit Combined Nomenclature classification to the 6-digit HS 1992 Classification (to match the tariffs data). Due to changes in the reporting threshold in 2011, we follow Bergounhon, Lenoir, and Mejean (2018) by dropping observations where a firm's annual exports or imports amount to less than €1000. In all, the customs data cover 390,000 exporting firms and 413,500 importing firms trading in 5000 products with 194 countries.

[10] With data on tariffs and tariff elasticities, European Commission (2017) computes reductions in NTBs that would explain changes in trade flows not accounted for by tariff cuts. They report the highest NTB reduction for EU exports in electronic equipment (25.3%), raw materials (13%) and machinery and equipment (9.3%) sectors. NTBs faced by South Korea's exporters also fell significantly for metals (12.5%), raw materials (9.5%) and agricultural goods (7.8%). NTBs fell even for sectors that did not have dedicated provisions under the FTA. Therefore, their results highlight the role of 'horizontal' clauses that reduce trade frictions more broadly across sectors. Such NTB reductions are crucial as they drive the overwhelming majority of welfare gains in CGE evaluations of deep FTAs – especially when the initial level of tariffs applied on manufactured goods is relatively low, as is the case for EU and South Korea.

[11] The transaction-level customs data that support the findings of this study is covered by statistical secrecy and can be accessed only through a previous authorization of the French Custom Administration. The customs data is from the DGDDI (Direction Générale des Douanes et Droits Indirects – a directorate of the French Ministry of Finance). The authorization is granted by the "Comité du secret" of the CNIS (Conseil National de l'Information Statistique). The link to procedures for getting access to the data is: https ://www.comite-du-secret.fr/.

2.2.3 Measuring Size

We use the customs data described above to proxy a firm's size by the total of its trade with markets other than South Korea. In principle, the customs data could be merged with a balance sheet survey of firms that contains more conventional measures of size such as revenue, capital stock or employment.[12] However, the balance sheet data has no size information on firms with less than 25 employees. These firms account for more than half of French exporters (Fontagné, Orefice, and Piermartini, 2020). For the current analysis, retaining these firms is important as the objective is to study the differential impact of the FTA along the full size distribution.

Furthermore, using customs data instead of balance sheet information allows us to define size at the firm-product level, which is not feasible with balance sheet data. Defining firm size at the product level facilitates the estimation of size-specific tariff elasticities since tariffs vary at the product-level. For this reason, most of our analysis relies on a firm-product level measure of size. However, we provide sensitivity checks to examine the robustness of results regarding this choice.

Our size measure is therefore measured as the total trade (exports and imports) of a firm across destinations within an HS-6 digit product over the control period (2000-2006), using the GDP deflator (base year 2015) to adjust for price changes and excluding any trade with South Korea. This definition is based on a time window that ends five years before entry into force of the EUKFTA and even before negotiations on the agreement began, thereby taking into account the fact that size is endogenous to trade liberalization. Excluding trade with South Korea has the same advantage. Defined in this manner, our size measure is time-invariant. Such a trade-based proxy for size is also supported by prior literature (Melitz and S. Redding, 2014; Fontagné, Orefice, and Piermartini, 2020).[13]

2.2.4 Characteristics of French Exporters to South Korea

Our data include firms that either export to South Korea or import from it or trade both ways with the country. The data also feature firms that never trade with South Korea. Of all French exporting firms in the data, 4.87% exported to South Korea during the control period (2000-2006). That share stands at 2.16% when looking at exported firm-product combinations. How different are these firms from each other in terms of

[12] The Enquete Annuelle d'Enterprise (EAE) is a survey that provides balance sheet information of firms along with the SIREN identifier that enables matching.

[13] Any size measure (whether based on trade flows or balance sheet data) is unlikely to be a perfect measure of productivity as in Melitz (2003) models. Hence, our econometric strategy includes firm-product-time fixed effects to capture any time-varying supply-side shocks such as to firms' technologies and worker skills in a narrowly defined HS-6 digit product category.

our size measure? To examine this, we report summary statistics in Table 2.1. Within product classes, firms trading both ways with South Korea are (on average) more than 200 times larger than firms not trading with South Korea (comparing lines (i) and (iv) in Table 2.1).[14] Firms that export (but not import) to South Korea are still more than 35 times as large, with an average size of €32 million (in 2015 prices). This picture becomes even more pronounced when looking at median ($p50$) values instead of means, with the median size of exporters to South Korea being more than 50 times larger than firms that do not trade with South Korea at all. Within those groups, there is a substantial degree of skewness which tends to increase in the size of the groups and is thus smaller amongst firms trading with South Korea than in the full sample. Against the backdrop of existing literature, these results are as expected.[15]

Table 2.1: Size distribution and trade with South Korea (in million €)

		Number	Mean $\bar{\mu}$	$p25$	$p50$	$p75$	$\bar{\mu}/p50$
(i)	Two-way trade w/Korea	3105	200.875	1.107	8.349	51.513	24.060
(ii)	Exported to Korea	31,712	32.149	0.077	0.737	6.426	43.624
(iii)	Imported from Korea	19,362	10.163	0.035	0.249	1.845	40.815
(iv)	No trade w/Korea	4,385,575	0.859	0.004	0.013	0.064	66.077
(v)	All firm-product pairs	4,439,754	1.263	0.004	0.013	0.067	97.154

Note: This table reports summary statistics for our size measure calculated as the sum of (deflated) exports and imports of a firm in a given HS-6 digit product (excluding trade with South Korea) over the period 2000 to 2006 i.e. before negotiations for the EUKFTA began. It describes the size distribution of firm-product combinations which were i) exported to and imported from South Korea at least once over 2000-2006; ii) exported to but never imported from South Korea; iii) imported from but never exported to South Korea; iv) did not trade with South Korea and; v) traded with any destination over 2000-2006. $p25$, $p50$, and $p75$ denote the 25th, 50th, and 75th percentile of the size (sales) distribution.

French firms exporting to South Korea tend to be diversified across destinations as well as products. For instance in 2016, a large proportion of these firms exported not only within the EU but also to economies such as the US (72.2%), China (58.2%), and Japan (56.2%). Out of all firms selling to South Korea in 2016, 19.4% sold two and 24.6% sold more than two HS 6-digit products to South Korea. Hence, there is

[14] Amongst the two-way traders, 60% fall within three HS chapters - i) 84 (Nuclear reactors, boilers, machinery and mechanical appliances; parts thereof); ii) 85 (Electrical machinery and equipment and parts thereof; sound recorders and reproducers, television image and sound recorders and reproducers, and parts and accessories of such articles) and; iii) 90 (Optical, photographic, cinematographic, measuring, checking, precision, medical or surgical instruments and apparatus; parts and accessories thereof).

[15] Such size patterns arise in heterogeneous firms models with asymmetric countries. Given comparable high trading costs (both fixed and variable) resulting from geographical and cultural distance to France, only the most efficient firms select into far away markets (Chaney, 2008), and they tend to export their best-performing products (Mayer, Melitz, and Ottaviano, 2014). Similarly, the productivity (and size) premium of two-way traders over one-way traders has also been reported in prior literature such as Kasahara and Lapham (2013), who demonstrate the role of import-export complementarities in driving this gap.

ample variation across markets and products within French firms that export to South Korea. This feature of the data enables us to use a broad range of firm fixed effects in our regressions.

Next, quite in line with expectations and prior research, we find that multi-product firms, firms serving multiple destinations, and firms serving neighbouring markets like Japan, Taiwan or both are significantly more likely to export to South Korea in the control period (2000-2006). Using simple two-period linear probability panel models, Table B.4 in the Appendix shows that multi-product firms have a probability of exporting to South Korea that is by about 1.8% higher than that of other firms, multi-destination firms display a premium of 1.1%, and firms exporting to Japan and/or to Taiwan have a 7.4% higher likelihood. The latter observation suggests thinking of Japan and Taiwan as plausible alternative destinations for French exporters to South Korea. After the entry into force of the EUKFTA, the likelihood to export to South Korea appears significantly higher for all those firms, with a particularly strong effect amongst firms exporting to Japan and/or Taiwan. These findings reaffirm core predictions in the Melitz (2003) models regarding selection effects.

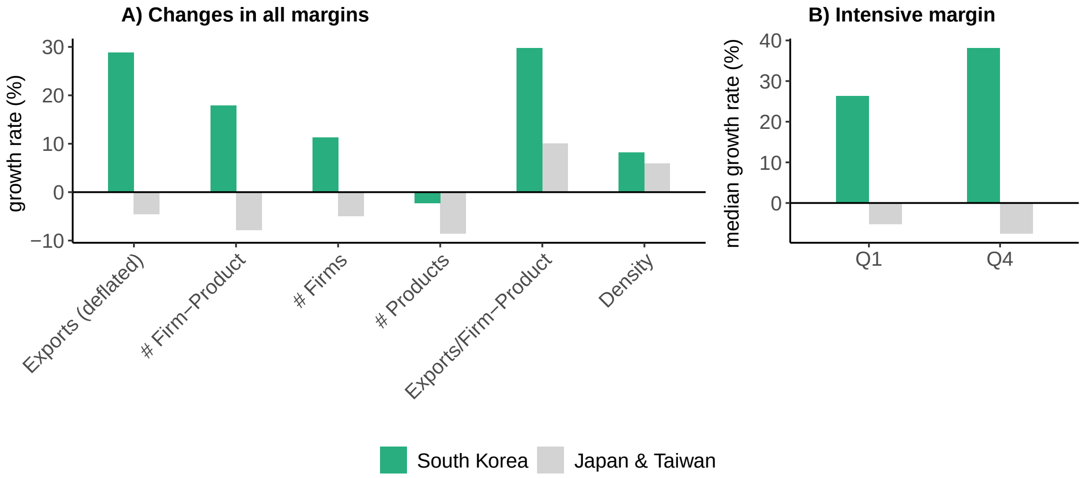
2.2.5 Effects of the FTA – A First Glance at the Data

To examine the contributions of various margins to overall growth in trade, we follow Bernard et al. (2009) and decompose exports to South Korea into i) unique number of firms; ii) unique number of products; iii) average exports per firm-product pair and; iv) density i.e. the fraction of all possible firm-product pairs for which exports are positive. We compute changes in these margins, where the margins are first averaged across years within two periods (2000-2006 and 2011-2016) and then differenced. We compare exports to South Korea with exports to Japan and Taiwan since they are similarly distant markets that imported comparable baskets of goods from France.[16]

Panel A in Figure 2.1 depicts changes for these various margins between the control and FTA period. We find that exports to South Korea posted a steep jump following the implementation of the FTA. The contrast with Japan and Taiwan is also striking. Exports to South Korea increased by approximately 29 percentage points, driven by increases in the number of exported firm-product combinations and the average sales per firm-product combination. In subsequent regressions, we expand the control group to include all other countries in the customs data set and introduce high dimensional fixed effects in order to account for a wide range of variables that can influence export outcomes such as demand shocks, macroeconomic conditions, firms' market knowledge and distribution networks.

[16] As shown in Table B.4 in Appendix B.2, there is a high degree of correlation between exporting to South Korea and exporting to either of these two countries prior to the adoption of the FTA.

Figure 2.1: Growth in exports from France



Note: Panel A shows the growth rate in various margins of French exports, where the margins are first averaged across years within the control (2000-2006) and FTA periods (2011-2016) and then differenced. Panel B shows the median growth rates in sales of firm-product combinations that were exported in both control and FTA periods to all three destinations (South Korea, Japan and Taiwan). The median growth is computed over all firm-product combinations within the bottom quartile ($Q1$) and the top quartile ($Q4$) of the size distribution. In both panels, export values are first adjusted by France's GDP deflator drawn from the World Bank Database.

In Appendix B.1, we further decompose the change in aggregate exports into the change in sales of continuously exporting firms, entrants and exiting firms (Figure B.2). We find that exports to South Korea in the post-FTA period were approximately €5.95 billion higher compared to the control period. This is primarily driven by a €7.69 billion increase in the sales of continuous firms. In comparison, firms that newly entered the South Korean market contributed €3.23 billion in additional exports, whereas firms that exited the market led to a decline in sales of €4.97 billion. In Japan and Taiwan, total exports of French firms shrank due to exiting firms.

These preliminary findings indicate that the EUKFTA provided a substantial boost to French exports to South Korea and that the export growth was overwhelmingly driven by continuous exporters i.e. firms that had already exported to South Korea in the period before the start of the EUKFTA negotiations. Looking at the set of continuously exported firm-product combinations, the top quartile of the size distribution saw an increase in exports to South Korea by 38% (at the median) whereas the bottom quartile grew by approximately 26% (see panel B of Figure 2.1).[17] In the following sections, we explore this skewness in export growth and compare the role of tariffs and NTB reductions in generating these growth differentials.

[17]This differential also exists for the simple average of growth rates for $Q1$ and $Q4$ exporters.

Finally, it is worth noting that tariff cuts are identical for all firms within product categories. Of course, it is perfectly possible that product classes populated by larger firms have experienced larger tariff cuts. Table B.2 in Appendix B.1 shows, however, that simple averages of tariff reductions within size classes are not biased in favor of large firms.[18]

2.3 Empirical Methodology

2.3.1 Specification

To guide our analysis, we use a simple demand function where exports of a French firm f in product p to a destination d at time t can be written as $X_{fpdt} = A_{pdt}(\mathcal{P}_{fpdt})^{-\eta_{fpd}}$. In this expression, buyers in d face the price $\mathcal{P}_{fpdt}$; A_{pdt} is a demand shifter that is common across firms but varies across products, destinations and time and; η_{fpd} is the demand elasticity which we take as time-invariant but specific to the destination and the firm-product combination.

The consumer price depends on tariffs and NTBs such that $\mathcal{P}_{fpdt} = \mathcal{P}_{pdt}\tau_{fpdt}t_{pdt}$ where $\tau_{fpdt} \geq 1$ is an iceberg factor that captures NTBs and $t_{pdt} \geq 1$ is an ad valorem tariff factor which does not vary across firms.[19] Following the gravity literature (see, e.g., Head and Mayer, 2014), we assume that NTBs are affected by FTAs such that – $\tau_{fpdt} = \tau^0_{fpdt}\exp(-\zeta_{fpd}FTA_{dt})$, where τ^0_{fpdt} corresponds to the base level of NTBs; FTA_{dt} is a dummy variable taking the value one if France has a free trade agreement with country d at time t, and ζ_{fpd} is the associated coefficient. Substituting and taking logs, we obtain the following:

$$\ln X_{fpdt} = \ln A_{pdt} - \eta_{fpd}\ln t_{pdt} - \eta_{fpd}\ln \mathcal{P}_{pdt} - \eta_{fpd}\ln \tau^0_{fpdt} + \eta_{fpd}\zeta_{fpd}FTA_{dt}. \quad (2.1)$$

The EUKFTA would affect this expression through changes in tariffs (t_{pdt}) and through NTBs as captured by FTA_{dt}. In this framework, the tariff elasticity η_{fpd} measures the effect of changes in trade costs (i.e., of tariffs and non-tariff barriers) whereas ζ_{fpd} measures the change in trade costs following the entry into force of an FTA. Provided with a sound estimate of η_{fpd}, it would be possible to back out ζ_{fpd}. Note, however, that clean identification is difficult, because components of Equation (2.1) such as A_{pdt},

[18] The evidence is more mixed when weighting tariff cuts by the size measure or firm-level exports to South Korea. Compared to the bottom quartile, medium-sized exporters tend to face smaller tariff cuts while the top exporters enjoy either similar or larger cuts. A better understanding of the structure of tariff cuts as a function of the size distribution is an interesting avenue for further research.

[19] If $\tilde{t} \in (0,1)$ is the ad valorem tariff rate, then $t = 1 + \tilde{t}$ is the tariff factor.

$\mathcal{P}_{pdt}$ or τ^0_{fpdt} are not readily observable. Including appropriate fixed effects is helpful, but risks making identification of ζ_{fpd} impossible.

In this paper, we are interested in the size-specific effects of trade policy. Focusing on the interaction between firm size measures and trade policy variables, we gain degrees of freedom in dealing with unobserved determinants of X_{fpdt}. Following our simple theoretical framework, we adopt a difference-in-differences approach that introduces high dimensional fixed effects that can control for demand shocks A_{fpdt}, the producing firm's costs (as reflected by the factory-gate price $\mathcal{P}_{fpdt}$) and non-actionable NTBs. Our corresponding specification is as follows:

$$\ln X_{fpdt} = \sum_{k=1}^{\mathcal{K}-1} \beta^k (\mathcal{I}_{dt} \times Size^k_{fp}) + \sum_{k=1}^{\mathcal{K}-1} \gamma^k (\ln t_{pdt} \times Size^k_{fp}) + \sum_{k=1}^{\mathcal{K}-1} \delta^k (\mathcal{I}_{dt} \times \ln t_{pdt} \times Size^k_{fp}) + \mathbf{Z} + \theta_{fpd} + \theta_{fpt} + \theta_{pdt} + \varepsilon_{fpdt}. \quad (2.2)$$

We aggregate exports to two periods – a control period (2000-2006) and an FTA period (2011-2016) – instead of working with yearly data, which raises issues related to the volatility of firm-level data.[20] The treated country is South Korea, with all remaining countries in the data forming the control group. Our choice of the control group of countries is hence both expansive and agnostic. The treatment dummy for the agreement $\mathcal{I}_{dt}$ takes the value of one for South Korea in the FTA period and zero otherwise.[21]

In our baseline regressions, we allocate exporters into $\mathcal{K}$ size bins (in robustness checks we also work with a continuous size measure). We interact these size bins with the dummy $\mathcal{I}_{dt}$, taking the category of the smallest exporters as the base category.[22] The associated series of coefficients β^k captures $\eta_{fpd}\zeta_{fpd}$ for the k-th size category relative to the base category. Hence, we cannot equate size specific effects $\eta^k_{fpd}\zeta^k_{fpd}$ to β^k. [23] However, it is clear that potential heterogeneity in estimated β^k coefficients can be due to η^k_{fpd}, ζ^k_{fpd}, or both.

[20]We use the French GDP deflator (base year as 2015) to account for changes in output prices within periods. We also estimate equation (1) on yearly data as a robustness check in Section 2.4. Also note that, in principle, we could collapse the data over the product dimension and study changes in exports at the firm-destination-time level. We do so as a robustness check in Table B.7 in Appendix B.2.

[21]In fact, $\mathcal{I}_{dt}$ is the product of the more general indicator variable FTA_{dt} and a South Korea dummy.

[22]Size bins are defined within HS-6 digit product. Due to the array of fixed effects present in the model, only $\mathcal{K}-1$ of the bin-specific interaction terms in our model are linearly independent, and so we can estimate bin-specific effects only for $\mathcal{K}-1$ bins. As a default, we omit the category of smallest exporters. Consequently, the specification cannot inform us about the absolute change in exports within a size bin but only about the change relative to the base category.

[23]With a continuous size measure, interpretation is more straightforward; see Section 2.2.3.

We also add size bin interactions with the applied tariff factor t_{pdt}, distinguishing between an average base effect (applying to any change in tariffs) and EUKFTA-specific tariff changes. With the inclusion of tariff controls in the specification and assuming that our fixed effects capture possible changes in preferences and supply-side determinants, the β^k coefficients provide a clean identification of NTB reductions by construction; i.e., they capture all effects of the FTA on firm exports across the size distribution net of tariffs. This allows us to interpret the β^k coefficients as the 'catch-all' effect of NTB reductions for a size class.

Our empirical methodology as outlined above therefore departs from prior literature by circumventing the need to define and construct proxies for the wide variety of horizontal and sector-specific NTBs that restrict cross-border trade. This is particularly relevant in the case of deep agreements such as the EUKFTA, whose provisions span a large number of behind-the-border issues. This approach also aligns with our main focus – that of examining size heterogeneity from NTB reductions in general and not on the narrower issue of the agreement's implementation, which requires computing the precise cuts in different NTBs such as SPS, TBT or red tape that were achieved by the FTA. Moreover, the β^k coefficients in this specification extend beyond policy-driven NTBs as, amongst other things, they also capture the trade effects of reductions in uncertainty.

Given that the control group includes all destinations excluding South Korea, we additionally control for other agreements signed by the EU in equation (2.2). This is incorporated in $\mathbf{Z}$ which is a vector of interactions between the various size bins and a dummy variable that takes the value of one in the second period for all other countries with which the EU implemented FTAs after the control period.[24] Moreover, $\mathbf{Z}$ includes interactions between these other FTAs, tariffs and size bins to account for any changes in tariff elasticities brought about by other agreements along the firm size distribution. Together, these terms account for any firm-specific demand or supply shocks affecting exports in the control group of countries. We prefer this approach over excluding the EU's new FTA partners from the control group as the latter may bias the estimates of fixed effects, particularly for products that may be heavily traded with those countries.

Our preferred specification includes the richest possible set of fixed effects (firm-product-destination, firm-product-time and product-destination-time) such that the β^k coefficients can still be identified.[25] These fixed effects control for variation in trade margins that could stem from factors other than the FTA such as demand-side shocks, changes in distribution networks, management practices or firm abilities amongst other in-

[24]This list of FTAs is drawn from the Design of Trade Agreements (DESTA) database by Dür, Baccini, and Elsig (2014) and reported in Table B.1 in Appendix B.1. In all, 16 agreements were implemented of which the deepest were the EUKFTA, as well as the EU-Georgia (2014) and EU-Moldova (2016) FTAs.

[25]We experiment with less exhaustive sets of fixed effects as well; see below.

fluences. A causal interpretation of β^k coefficients therefore relies on the relatively weak assumption that θ_{pdt} and θ_{fpt} fixed effects capture any omitted variables relating to demand-side and supply-side shocks, respectively. Moreover, the inclusion of firm-product-destination fixed effects addresses potential concerns regarding the endogeneity of tariff reductions, similar to the Baier and Bergstrand (2007) solution of including country-pair fixed effects in structural gravity estimations. Given the two-period structure of the model and the set of fixed effects, the underlying sample is a strongly balanced panel of firm-product-destination triplets. Hence, identification is based purely on variation over time in the intensive margin of firms' exports. Finally, the error term is clustered three-way by firm, product and destination.[26]

An important aspect of our econometric strategy is that the estimated coefficients only indicate the *relative* effects of NTB liberalization, i.e. relative to the chosen reference category. They do not reflect the aggregate impact of the agreement on French exports, for which one would need a structurally estimated model. Instead, our focus is on the *differential* impact of NTB reductions on firms' intensive margin along the size distribution.[27] Moreover, any unobservable variables relating to incumbent exporters to South Korea are captured by including firm-product-destination fixed effects. This fixed effect also drops firms that may have lobbied for the FTA but which did not export to South Korea in the control period. By measuring size using trade flows (excluding trade with South Korea) in the control period, we also shut down another potential channel for reverse causality.

2.3.2 Baseline Results: Impact of the FTA on the Intensive Margin

We report results on the intensive margin in Table 2.2 below. The most rigorous and our preferred specification following equation (2.2) is reported in column (1). In columns (2)-(4), we modify the set of fixed effects in order to demonstrate their role in capturing size heterogeneity along the intensive margin. Throughout these estimations, we compare exporters belonging to different quartiles of the size distribution to those in the bottom quartile ($Q1$, the excluded category). Standard errors are clustered by firm, product and destination.[28]

[26] Results are robust to this choice of clustering.

[27] An even stricter definition of the intensive margin would be at the 10-digit tariff line level, for sales by a given plant within a firm and to a given buyer in the foreign market. However, this is not feasible in our case due to limitations on the availability of plant and buyer-level information in the customs data as well as the level of aggregation of tariffs (HS 6-digit) in the database by Felbermayr, Teti, and Yalcin (2019).

[28] However, results do not change if standard errors are clustered by different dimensions of the data e.g. firm-product, firm-destination or product-destination. See Table B.6 in Appendix B.2 for further details.

Column (1) reports estimates of size-specific β, γ, and δ coefficients in the table's upper, middle and lower sections, respectively. The estimates of β coefficients show that the increase in exports due to the FTA is greatest for exporters in the top quartile ($Q4$) of the size distribution. In fact, $Q4$ exporters grew their sales to South Korea by approximately 66 percentage points more relative to those in the bottom quartile. Since we net out the effects of tariffs, we attribute this change to reductions in the costs of NTBs.[29] The regression result also reveals a monotonic pattern: the relative increase in exports due to the NTB reductions continuously falls as size shrinks. These coefficients are not only statistically significant from zero but also differ from each other based on Wald tests.

Moreover, column (1) reveals that $Q4$ exporters also react more strongly to tariff reductions (γ coefficients). For exporters belonging to the top size category, the absolute value of the tariff elasticity is by the amount 1.13 larger than for exporters in the bottom size bin. Exporters in other size categories do not appear to react differently to the smallest exporters to tariff cuts. The final three lines in Table 2.2 show estimates of δ coefficients. There is no evidence that changes of tariffs in the EUKFTA produced any different size-specific effects than changes of tariffs arising in other contexts.[30] Overall, the results in column (1) indicate that larger French exporters have increased exports to South Korea by more than smaller ones, with the boost from lower NTBs linearly declining with size. Only the largest firms appear to exhibit stronger behavioral responses to trade cost changes, so that, relative to the smallest exporters, the stronger responses of exporters in the second and third quartiles of the size distribution must be entirely due to larger effective reductions in the cost of NTBs.

Next, we turn to the role of fixed effects. Column (2) deviates from the main specification by assuming that demand shocks are common across products within destination-time. Therefore we replace θ_{pdt} by the less comprehensive θ_{dt} fixed effects. This change lowers the differential impact of NTB reductions on exporters. Even so, the statistical significance and monotonic ordering of β^k coefficients are retained. In column (3), we assume that supply shocks are common across products within firm-time by replacing θ_{fpt} with only θ_{ft} fixed effects. Doing so generates counter-intuitive estimates on tariff controls which turn positive, unlike the other regression results. This signals the importance of differentiating across products within firms, especially given the dominance of multi-product exporters in France's trade with South Korea. Our results also hold, albeit weakly for NTB reductions, when we replace θ_{fpd} with θ_{fd} fixed effects in column (4). This amounts to presuming that the capabilities of a firm in a given desti-

[29]This higher export growth due to NTB reductions is observed to operate through the quantity channel rather than prices, as can be seen in Table B.5 of Appendix B.2.

[30]Though the coefficient of -1.67 for firms in the $Q3$ bin is significant at the 90% confidence level, it is not robust to clustering choices. See Table B.6 in Appendix B.2.

Table 2.2: Impact of EUKFTA on firm-level outcomes by size quartiles

Dependent Variable:	ln(exports)				
	(1)	(2)	(3)	(4)	(5)
$\mathcal{I} \times Q4$	0.664***	0.554***	0.469***	0.203*	0.712***
	(0.144)	(0.112)	(0.111)	(0.110)	(0.150)
$\mathcal{I} \times Q3$	0.564***	0.436***	0.458***	0.198*	0.561***
	(0.138)	(0.110)	(0.128)	(0.116)	(0.143)
$\mathcal{I} \times Q2$	0.373**	0.339**	0.406***	0.234*	0.380**
	(0.154)	(0.130)	(0.132)	(0.125)	(0.155)
$\ln t \times Q4$	-1.13***	-0.868***	1.43***	-2.13***	
	(0.413)	(0.320)	(0.355)	(0.397)	
$\ln t \times Q3$	-0.280	-0.198	1.81***	-0.751***	
	(0.314)	(0.272)	(0.364)	(0.241)	
$\ln t \times Q2$	0.181	0.053	1.62***	-0.089	
	(0.337)	(0.246)	(0.339)	(0.200)	
$\mathcal{I} \times \ln t \times Q4$	-1.27	-0.871	-6.46	-0.750	
	(0.979)	(0.816)	(4.09)	(1.14)	
$\mathcal{I} \times \ln t \times Q3$	-1.67*	-1.16	-6.69	-2.17**	
	(0.858)	(0.762)	(4.07)	(1.02)	
$\mathcal{I} \times \ln t \times Q2$	0.704	-0.010	-5.40	0.171	
	(0.868)	(1.00)	(4.16)	(1.28)	
Fixed effects	$\theta_{fpd}, \theta_{fpt}, \theta_{pdt}$	$\theta_{fpd}, \theta_{fpt}, \theta_{dt}$	$\theta_{fpd}, \theta_{ft}, \theta_{pdt}$	$\theta_{fd}, \theta_{fpt}, \theta_{pdt}$	$\theta_{fpd}, \theta_{fpt}, \theta_{pdt}$
Observations	1,758,070	1,758,070	1,758,070	1,758,070	1,758,070
R^2	0.919	0.905	0.920	0.866	0.919

Note: The table above reports regression results for the impact of the EUKFTA along the intensive margin. $\mathcal{I}$ is a dummy variable that takes the value of one for French exports to South Korea from 2011 onwards and the value of zero otherwise. t denotes product-level ad valorem tariff factors. Column (1) contains our baseline results for the intensive margin following the specification in equation (2.2). In columns (2)-(4), we vary the set of fixed effects. Size is defined at the firm-product level and tariffs at the product-destination-time level. Only continuous exporters are retained i.e. firm-product combinations that report positive exports to a given destination in both control (2000-2006) and FTA (2011-2016) periods. Standard errors are clustered by firm, product and destination. Significance codes: ***: 0.01, **: 0.05, *: 0.1

nation are common across products. In this case, size heterogeneity in NTB elasticities is further reduced but sharpened for tariff elasticities in comparison to column (1). Overall, varying the combination of fixed effects largely preserves the size advantage of larger exporters from NTB reductions. Therefore, we proceed with the most conservative specification following equation (2.2) as it features all possible fixed effects to reduce issues arising from omitted variables.

Finally, in column (5) of Table 2.2, we keep all fixed effects but drop tariffs in order to examine the combined effect of the FTA (tariffs and NTBs) on incumbent exporters of differing sizes. As before, the positive impact of the FTA on the intensive margin is magnified for larger incumbents, with sales of top quartile exporters in South Korea growing by 71 percentage points more than of those of bottom quartile exporters. Comparing to column (1), we conclude that more than 90% of the size advantage of top quartile exporters is due to NTBs; their stronger reaction to tariff cuts playing only a minor role.

Our baseline result on the FTA's intensive margin effects is at odds with the workhorse Melitz (2003) model which implies that larger incumbent exports react with similar rates to trade cost changes than smaller ones. This result emerges from the assumption of identical trade cost structures for small and large firms, and/or identical reductions in trade costs, and/or identical elasticities of demand. Our result not only suggests that these assumptions may be problematic but that popular frameworks which depart from the Melitz (2003) assumptions, such as the one by Melitz and Ottaviano (2008) which allows for variable elasticities of demand, or by Arkolakis (2010) which endogenizes foreign market access costs have counterfactual implications too, since they predict advantages for smaller incumbent exporters. Before discussing possible model modification, though, it is necessary to check the effect of the EUKFTA on the extensive margins. This is why we now turn to the effect of the EUKFTA on market entry and product diversification.

2.3.3 Market Entry and Product Diversification

Although our focus is on the impact of NTB reductions on the intensive margin of exports, our data does permit us to examine two additional margins - firm entry into exporting and the diversification of export baskets following the agreement. We do so by retaining the two-period structure of our baseline model but moving the analysis to the firm-destination-time dimension as shown by Equation (2.3) below.

Now, the dependent variable Y_{fdt} corresponds to either a dummy variable for a firm's exporting status or the number of products exported to a given destination d at time t. Correspondingly, we move our size measure from the firm-product to the firm-wide level by aggregating imports and exports across all products traded by a firm in the control period with all countries except South Korea. In this case, tariffs are averaged across products at the destination-time level and interacted with size bins. The vector of controls **Z** includes interaction terms between size bins and a dummy that takes the value of one in the second period for all other countries with which the EU signed FTAs after 2006.[31]

$$Y_{fdt} = \sum_{k=1}^{\mathcal{K}-1} \beta^k (\mathcal{I}_{dt} \times Size_f^k) + \sum_{k=1}^{\mathcal{K}-1} \gamma^k (\ln t_{dt} \times Size_f^k) + \mathbf{Z} + \theta_{fd} + \theta_{ft} + \theta_{dt} + \varepsilon_{fdt} \quad (2.3)$$

These regressions include all possible fixed effects (firm-destination, firm-time and destination-time) such that the β^k coefficients can still be estimated. The results are reported in Table 2.3. In the case of firm entry, identification is based on entrants and exiters as the dependent variable does not vary for continuous exporters to a given destination.[32] In contrast, when examining adjustments to the product basket, only those firms are retained which exported in both periods to a given destination.

For firm entry in columns (1)-(2), we find the β^k coefficients to be negative. This implies that NTB reductions induced new firms into exporting to South Korea and these firms tended to be smaller than firms exiting the market following liberalization. This can also be seen when comparing density plots of firm sizes between continuously exporting firms to South Korea, entrants and exiters (see Figure B.3 in Appendix B.1). We also find that the intermediately sized exporters had higher tariff elasticities for the entry margin than smaller firms, confirming the predictions of the Melitz (2003) model on selection into exporting from the middle of the size distribution. In the case of the product margin reported in column (3), NTB reductions do not generate discernible size effects.

2.3.4 Possible Modeling Implications

To summarize our results across these margins, we find that NTB reductions strongly favour large firms along the intensive margin, whereas the effect is absent for the product margin. In the case of the firm entry margin, our findings closely follow the patterns of Melitz (2003) models. One approach to reconcile theory with data would be to allow for a correlation between exogenous cuts in variable trade costs and firm size. Another approach would be to assume the opposite as Arkolakis (2010), namely, that the marginal foreign market access costs are declining in market share and that the

[31]Coefficients on the interaction between $\mathcal{I}_{dt}$, $\ln t_{dt}$ and size bins cannot be identified due to multicollinearity with γ^k coefficients.

[32]The sample therefore reduces with the exclusion of continuous exporters.

Table 2.3: Market entry and product diversification

Dependent Variables:	Exporter (0,1)		ln(products)	
Model:	(1)	(2)	(3)	(4)
$\mathcal{I} \times Q4$	-0.048***	-0.050***	0.092	0.055
	(0.014)	(0.014)	(0.058)	(0.059)
$\mathcal{I} \times Q3$	0.019	-0.005	0.086	0.031
	(0.014)	(0.014)	(0.060)	(0.063)
$\mathcal{I} \times Q2$	0.044***	0.025	0.112*	0.079
	(0.015)	(0.016)	(0.060)	(0.065)
$\ln t \times Q4$		-0.039		-0.682*
		(0.145)		(0.351)
$\ln t \times Q3$		-0.488***		-1.00***
		(0.102)		(0.309)
$\ln t \times Q2$		-0.386***		-0.590*
		(0.105)		(0.340)
Fixed-effects	$\theta_{fd}, \theta_{ft}, \theta_{dt}$	$\theta_{fd}, \theta_{ft}, \theta_{dt}$	$\theta_{fd}, \theta_{ft}, \theta_{dt}$	$\theta_{fd}, \theta_{ft}, \theta_{dt}$
Observations	2,380,810	2,380,810	3,196,118	3,196,118
R^2	0.654	0.654	0.887	0.887

Note: The table above reports regression results for the impact of the EUKFTA on market entry and product margins. $\mathcal{I}$ is a dummy variable that takes the value of one for French exports to South Korea from 2011 onwards and the value of zero otherwise. Columns provide results following the specification in equation (2.3). Since the dependent variables here are defined at the firm-destination-time level, size is correspondingly computed at the firm level (aggregating across products) and tariffs are averaged across products within a given destination and time period. Standard errors are clustered by firm and destination. Significance codes: ***: 0.01, **: 0.05, *: 0.1

FTA lowers the level of entry costs for all firms. Such a structure could generate firm entry from the middle of the size distribution and, at the same time, higher growth rates by the largest firms.[33]

Such assumptions may be rationalized by the fact that larger firms have the capacity to plan and invest towards better leveraging of the benefits of the agreement eg. by hiring specialized consultants and lawyers to meet testing, certification and complex rules of origin requirements. Alternatively, large firms may be lobbying for favourable rules and consequently receiving larger NTB cuts or even larger tariff reductions in their product categories than smaller firms. The latter channel has been examined by recent work such as such as Blanga-Gubbay, Conconi, and Parenti (2020) who use detailed information from lobbying reports and develop a model where only large pro-FTA firms select into lobbying.[34]

2.4 Robustness Checks and Additional Results

In this section, we describe additional robustness checks to our key result: that larger incumbent exporters gained more from NTB reductions under the EUKFTA than smaller exporters. As there is no statistically significant change in tariff elasticities after the entry into force of the EUKFTA, the following tables report only the general tariff elasticities across size bins (while the regressions do include FTA-specific terms).

2.4.1 Truncating the Size Distribution

First, our baseline results may be affected by the presence of a few very large firms. To address this concern, we replicate our preferred specification in column (1) of Table 2.2 but drop the top 1%, 5% and 10% of firms. Size bins are redefined accordingly on the remaining sample. This allows us to test whether the size effect from NTB reductions in our baseline regressions is driven by the largest firms only. As Table 2.4 shows, NTB reductions continue to favour the largest firms. Similar to our main result, the advantage from size is consistently observed not only for NTB reductions but also for the average base effect of tariff cuts in the case of the top quartile firms. However, these size-specific tariff elasticities slightly diminish with the exclusion of an increasing number of large firms.[35]

[33] Working out the precise conditions under which this can happen is an avenue for future research.

[34] Due to lack of information on lobbying expenses of French firms, a similar exercise of linking lobbying efforts to size and trade flows is not feasible in our case.

[35] This pattern is also seen when we use a continuous measure of size; see Panel A of Table B.8 in Appendix B.2.

Table 2.4: Impact of NTB reductions and tariff cuts after excluding the largest firms

Dependent Variable:	ln(exports)			
Sample:	All	Drop top 1%	Drop top 5%	Drop top 10%
	(1)	(2)	(3)	(4)
$\mathcal{I} \times Q4$	0.664***	0.639***	0.628***	0.601***
	(0.144)	(0.142)	(0.154)	(0.160)
$\mathcal{I} \times Q3$	0.564***	0.558***	0.504***	0.420***
	(0.138)	(0.140)	(0.154)	(0.161)
$\mathcal{I} \times Q2$	0.373**	0.367**	0.397**	0.369**
	(0.154)	(0.155)	(0.162)	(0.162)
$\ln t \times Q4$	-1.13***	-0.987**	-0.851**	-0.773*
	(0.413)	(0.410)	(0.394)	(0.404)
$\ln t \times Q3$	-0.280	-0.207	-0.281	-0.161
	(0.314)	(0.327)	(0.347)	(0.354)
$\ln t \times Q2$	0.181	0.245	0.224	0.289
	(0.337)	(0.353)	(0.392)	(0.451)
$\mathcal{I} \times \ln t \times Q4$	-1.27	-1.43	-1.53*	-1.11
	(0.979)	(1.00)	(0.793)	(0.907)
$\mathcal{I} \times \ln t \times Q3$	-1.67*	-1.79*	-0.963	-0.971
	(0.858)	(0.912)	(1.11)	(1.26)
$\mathcal{I} \times \ln t \times Q2$	0.704	1.88*	-1.45*	-1.25
	(0.868)	(0.958)	(0.839)	(0.941)
Observations	1,758,070	1,696,706	1,545,340	1,390,912
R^2	0.918	0.917	0.919	0.924

Note: Regression results are based on equation (2.2), where the dependent variable is exports at the firm-product-destination level aggregated to two periods: 2000-2006 and 2011-2016. $\mathcal{I}$ is a dummy variable that takes the value of one for French exports to South Korea from 2011 onward and the value of zero otherwise. Product-level ad-valorem tariff factors are denoted by t. In each column, size bins are recalculated after dropping the top 1%, 5% and 10% of varieties from the sample. Additional controls include interactions between a $\mathcal{I}_{dt}$, tariffs and size bins as well as interactions between size bins and a dummy variable that takes the value of one for all other countries with which the EU implemented FTAs after 2006. All regressions include firm-product-time, product-destination-time and firm-product-destination fixed effects. Only continuous exporters are retained i.e. those firm-product combinations that have positive exports in a given destination for each of the two periods. Standard errors are clustered by firm, product and destination. Significance codes: ***: 0.01, **: 0.05, *: 0.1

2.4.2 Alternative Size Measures

In our baseline model, we work with size quartiles so that we do not have to assume any functional form linking effects and firm size measures. One obvious alternative would be a linear specification, where we interact the EUKFTA dummy ($\mathcal{I}_{dt}$) with the log of a continuous measure of size (where size continues to be defined as in our baseline regression). As Table 2.5 shows, we find that the size advantage for NTB reductions holds. Increasing the size of an exporter by one percent increases additional exports to Korea by 0.079 percent. Also, tariff cuts continue to have a stronger effect for larger firms, but the relative relevance of NTBs still dominate.[36]

One could also work with two size categories instead of four as in our baseline by interacting the FTA indicator with a dummy for exporters in the top half of the size distribution. Comparing firms above and below the median size, we observe that NTB reductions boosted exports by approximately 33.9 percentage points more in the upper half of the size distribution relative to the lower half. This pattern of monotonicity is approximately maintained when the effects of NTB reductions under the agreement are estimated at every decile of the size distribution, relative to the bottom decile.[37]

Our findings may also depend on how we define the "size" of the exporter. We examine this in Table 2.6 where each column corresponds to a different proxy of size. For reasons of comparison, column (1) replicates column (1) of Table 2.2 using the baseline measure of size defined as the pre-treatment (2000-2006) sum of exports and imports at the firm-product level across all countries except South Korea; see Section 2.2.3. Column (2) defines size based on intra-EU trade only, as this provides us with a proxy of the domestic performance of French exporters. Column (3) defines size on extra-EU trade to capture the advantage obtained by selling overseas. Finally, column (4) switches to a definition of size that sums exports and imports of a firm over all the products traded by the firm in the pre-treatment period.

In columns (2) and (3), the size hierarchies are confirmed, indicating that experience in both domestic and foreign markets is relevant for leveraging NTB reductions.[38] Using our intra-EU trade measure, we find that the top quartile exporters increase sales to South Korea by approximately 36.8 percentage points more than the bottom quartile due to NTB cuts. This is a lower premium compared to column (1) of Table 2.2, since skewness of the size distribution reduces substantially with this measure.

[36] As in the baseline, the main effects for size and tariffs cannot be estimated due to the presence of firm-product-time and product-destination-time fixed effects.

[37] We show these results in Figure B.4 in Appendix B.2. Note that estimates are far less precisely estimated as standard errors grow when the distribution is split into an increasing number of bins. Therefore, we prefer the comparison across quartiles in the baseline model.

[38] Statistical significance is maintained when we move from quartiles to the log value of these alternative size proxies; see Panel B of Table B.8 in Appendix B.2.

Table 2.5: Impact of EUKFTA using log size measure and median dummy

Dependent Variable:	ln(exports)			
Size measure:	continuous		median (0,1)	
	(1)	(2)	(3)	(4)
$\mathcal{I} \times size$	0.079***	0.088***	0.339***	0.375***
	(0.009)	(0.010)	(0.064)	(0.061)
$\ln t \times size$	-0.243***		-1.05***	
	(0.065)		(0.342)	
Observations	1,758,070	1,758,070	1,758,070	1,758,070
R^2	0.919	0.919	0.919	0.919

Note: The table above reports regression results for the impact of the EUKFTA along the intensive margin following the specification in equation (2.2). Firm size is defined within a product class and tariffs at the product-destination-time level. Columns (1)-(2) report results when using the log value of our default size measure. Columns (3)-(4) report results when using *median* dummy that takes the value of one for exporter sizes above the median and zero otherwise. Only continuous exporters are retained i.e. firm-product-destination triplets with positive exports in both control (2000-2006) and FTA (2011-2016) periods. Standard errors are clustered by firm, product and destination. Significance codes: ***: 0.01, **: 0.05, *: 0.1

Interestingly, size heterogeneity in tariffs disappears when we employ a firm-wide measure that aggregates over all the products. This indicates that the proper estimation of size-specific tariff elasticities requires size to vary at the product-level. For larger exporters to benefit more from tariff cuts, it is crucial that they are large in the product categories affected by the cuts and hence, the size obtained from selling other goods does not help. One conclusion from this result is that it may well be costs associated to abiding by product-specific rules of origin that drive the size patterns as observed in column (1).

Finally, we can also consider the length of exporters' experience as a relevant proxy for their productivity. To test this, we replace the size bins from equation (2.2) with $experience$, a dummy based on the number of years the exporter was active in foreign markets over the control period (2000-2006). Table 2.7 reports the results of this robustness check. The positive and statistically significant coefficients on $\mathcal{I} \times experience$ confirm our prior that experienced exporters gained more from NTB cuts delivered by the EUKFTA. Consistent exporters with more than three years of experience also react more to tariff reductions more generally.

Table 2.6: Impact of EUKFTA with varying definitions of size

Dependent Variable:	ln(exports)			
Size measure:	Firm-Product Global	Firm-Product Intra-EU	Firm-Product Extra-EU	Firm Global
	(1)	(2)	(3)	(4)
$\mathcal{I} \times Q4$	0.664***	0.368***	0.645***	0.505***
	(0.144)	(0.091)	(0.135)	(0.137)
$\mathcal{I} \times Q3$	0.564***	0.250***	0.541***	0.437***
	(0.138)	(0.085)	(0.131)	(0.134)
$\mathcal{I} \times Q2$	0.373**	0.288***	0.275**	0.212
	(0.154)	(0.091)	(0.138)	(0.141)
$\ln t \times Q4$	-1.13***	-1.90***	0.256	-0.129
	(0.413)	(0.430)	(0.327)	(0.199)
$\ln t \times Q3$	-0.280	-0.887***	0.531*	-0.160
	(0.314)	(0.269)	(0.300)	(0.217)
$\ln t \times Q2$	0.181	-0.194	0.486	0.091
	(0.337)	(0.236)	(0.362)	(0.202)
Observations	1,758,070	1,564,004	1,652,294	1,758,070
R^2	0.918	0.917	0.916	0.918

Note: Regression results are based on equation (2.2) where the dependent variable is exports at the firm-product-destination level aggregated to two periods:2000-2006 and 2011-2016. $\mathcal{I}$ is a dummy variable that takes the value of one for French exports to Korea from 2011 onwards and the value of zero otherwise. t denotes product-level ad valorem tariff factors. In each column, firm size is defined differently. Using data only from the control period (2000-2006) and excluding trade with South Korea, these size measures are: column (1) global trade within the firm-product pair (baseline measure); column (2) intra-EU trade in the firm-product pair; column (3) extra-EU trade in the firm-product pair and; column (4) global trade of the firm across products. Since regressions include firm-product-destination fixed effects, only continuously exported varieties are retained i.e. varieties that have positive exports in a given destination for each of the two periods. Standard errors are clustered by firm, product and destination. Significance codes: ***: 0.01, **: 0.05, *: 0.1

Table 2.7: Exporting experience over 2000-2006

Dependent Variable:		ln(exports)	
Experience:	>1 year	>3 years	7 years
	(1)	(2)	(3)
$\mathcal{I} \times experience$	0.436**	0.171**	0.115**
	(0.180)	(0.080)	(0.051)
$\ln t \times experience$	-0.911	-1.12***	-1.11***
	(0.612)	(0.329)	(0.237)
Observations	1,758,070	1,758,070	1,758,070
R^2	0.918	0.918	0.918

Note: Regressions are based on equation (2.2), where the dependent variable is exports at the firm-product and destination level aggregated to two periods: 2000-2006 and 2011-2016. Experience is a dummy that takes the value of one for firm-product combinations that were exported more than once (column 1), more than three years (column 2) and in all years (column 3) in the control period (2000-2006). Standard errors are clustered by firm, product and destination. Significance codes: ***: 0.01, **: 0.05, *: 0.1

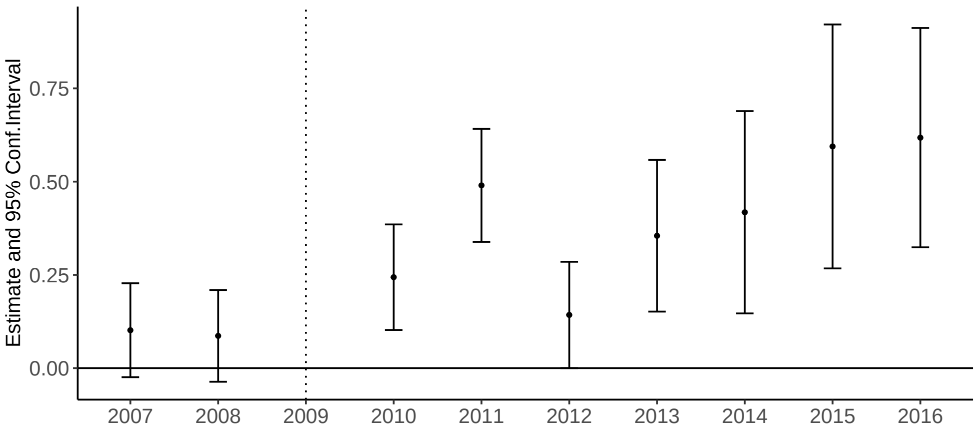
2.4.3 The Impact of NTB Reductions over Time

Our two-period baseline model reveals that larger firms benefit substantially more from NTB reductions but cannot reveal whether this size advantage grows, diminishes or remains stable over time. To address this issue, we exploit the long time dimension of the French customs data. Correspondingly, the specification in equation (2.2) is expanded by replacing the $\mathcal{I}_{dt}$ dummy with a South Korea dummy (Kor_d) and a series of year dummies. This modification allows us to examine the evolution of the impact of NTB reductions in South Korea on large firms relative to smaller ones. As the agreement was signed in 2010 and implemented in 2011, we take 2009 as the reference year.

To ensure that our proxy for size (measured over 2000-2006) remains exogenous, we estimate the specification for data ranging from 2007 to 2016. As we are interested primarily in adjustments at the intensive margin, only those firm-product-destination triplets are retained that registered positive exports throughout this period. As before, we include interactions between tariffs and size bins as well as interactions between size bins, year dummies and other countries with which the EU entered into FTAs over the sample period.

Figure 2.2 shows the resulting coefficient estimates and 95% confidence intervals for the adjustment in exports of the top quartile of varieties relative to the bottom quartile stemming from a reduction in NTBs under the EUKFTA. We observe a clear break following the signing of the agreement in 2010, with large firms posting substantially higher sales growth from NTB reductions compared to smaller firms. This size advantage therefore kicks in prior to entry into force of the agreement in 2011. The coefficient is even higher closer to the end of the sample period, indicating that the size advantage may have magnified.

Figure 2.2: Export growth from NTB reductions - $Q4$ exporters relative to bottom $Q1$ exporters



Note: This graph shows the adjustment in exports of the top quartile ($Q4$) of exporters relative to the bottom quartile, from a reduction in NTBs under the EUKFTA. Following the specification provided by equation (2.2), it plots coefficients and 95% confidence intervals on the interaction $Kor_d \times Year_t \times Q4$, where Kor_d is a dummy for South Korea and $Q4$ is a dummy variable that takes the value of one for varieties in the top quartile of the size distribution and zero otherwise. The chosen reference year is 2009, the year before the EUKFTA was signed. A set of firm-product-year, product-destination-year and firm-product-destination fixed effects are included. Standard errors are clustered by firm, destination and year.

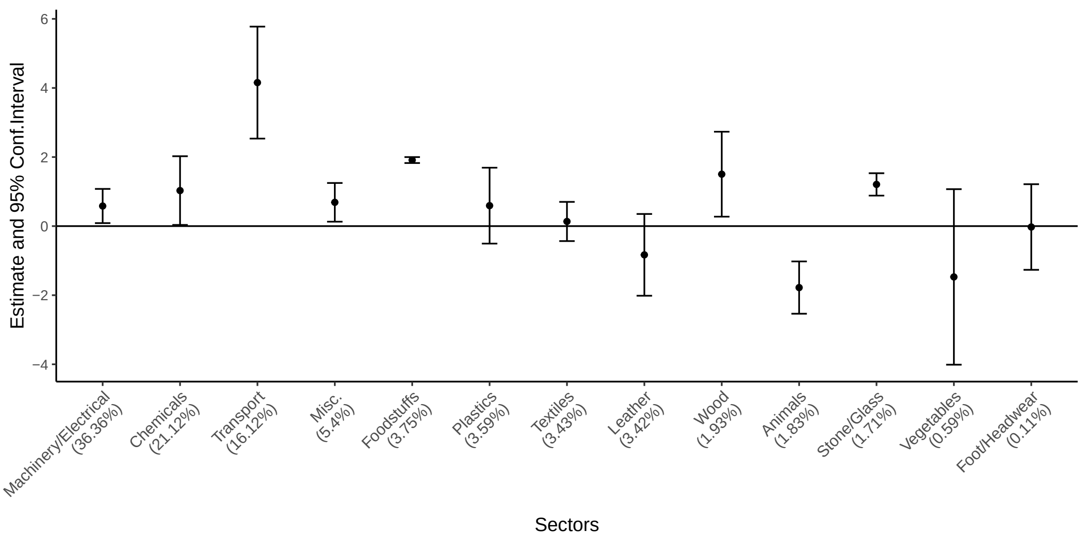
2.4.4 Heterogeneity across Sectors

Till now, our analysis has focused on identifying treatment effects of NTB reductions averaged across products. However, the size differential from NTB reductions may be driven by certain sectors. Therefore as a robustness check, we examine this channel by splitting the sample and estimating our baseline regression separately for each goods sector.[39] Figure 2.3 shows the estimated values and 95% confidence intervals of the coefficients on $\mathcal{I}_{dt} \times Q4$ interactions for each sector, where the sectors are arranged in decreasing order of their shares in France's exports to South Korea over the control period, 2000-2006.

[39]Based on the goods sector classification for the Harmonized System (HS) provided by WITS.

We observe substantial heterogeneity in estimates across sectors that was masked by our earlier results. For the five sectors machinery/electrical, chemicals, transport equipment, miscellaneous goods, and foodstuffs, which together account for more than 80% of export sales to South Korea, top quartile exporters enjoy a statistically significant size advantage over bottom quartile exporters.[40] The wood and stone/glass sectors, although small in France's export basket, behave similarly. Among the remaining sectors, only the animal products sector appears to display the opposite pattern whereby smaller exporters benefit more from NTB reductions in comparison to larger ones. Overall, Figure 2.3 confirms that size advantage from NTB reductions is not driven by outlier sectors alone but emerges as a recurring feature in the data.[41]

Figure 2.3: Heterogeneity across sectors - $Q4$ exporters relative to bottom $Q1$ exporters



Note: Following the specification provided by equation (2.2), this graph plots coefficients and 95% confidence intervals on the interaction term $\mathcal{I}_{dt} \times Q4$ for every sector. The regression for each sector includes similar interaction terms for the second and third quartiles. Standard errors are clustered by product and destination. Sectors are arranged from left to right by the decreasing order of their shares in France's exports to South Korea (in brackets) during the control period.

[40] Size dispersion across exporters within these sectors is also high, as can be seen in Table B.3 in Appendix B.1.

[41] When looking at tariff elasticities (γ_k), coefficients are statistically different between $Q4$ and $Q1$ exporters at the 90% confidence interval for only some of the sectors eg. miscellaneous products, wood and leather.

2.4.5 Phasing-In of Tariffs

An alternative approach to examining the impact of NTB reductions on the intensive margin of exports is shown in Table 2.8. By exploiting the tariff schedule of South Korea under the FTA, we split the sample into exports of goods that were already duty-free in South Korea in 2010 (MFN=0), those that became duty-free upon entry into force of the agreement (EIF) and goods whose tariffs were set to be gradually phased out by South Korea over three, ten or more than ten years. By estimating equation (2.2) for these categories, we can ascertain whether the FTA had a differential impact depending on the length of phasing out periods.[42]

Looking at the first column in Table 2.8, we observe that the exports of already duty-free goods grow more for larger firms than smaller firms following the FTA. Since South Korea applied no tariffs on these products prior to the agreement, we can be certain that the β^k coefficients capture size-specific elasticities from NTB reductions. The effect is particularly strong for the top quartile firms. The size advantage for already duty-free products is also robust to using a continuous measure of exporter size (see Panel C of Table B.8 in Appendix B.2). The size premium in export growth from NTB cuts is also present for goods with short tariff phase-outs as seen in columns (2) and (3) and only disappears for sensitive agri-foods where tariffs (and potentially NTBs) are cut only over ten years or longer. The lower panel of Table 2.8 reports general (i.e., not EUKFTA-specific) size-dependent tariff elasticities. The results confirm that larger firms tend to react more strongly to changes in trade costs.[43]

2.4.6 Imports from South Korea

So far, we have studied the agreement's impact on France's exports to South Korea. However, our data allow us to check whether the size advantage from NTB reductions applies to France's imports from South Korea as well. Now, the dependent variable is the log value of a firm's import purchases at the product-country-time level and tariffs correspond to duties applied by the EU on its trade partners. In Figure 2.4, we replicate the dynamic diff-in-diff regression discussed in Section 2.4.3 for imports to examine the evolution of the size effect over time. Overall, the evidence for a size differential in import growth is less clear, and this differential is not observed to be statistically significant when looking at the yearly coefficients in Figure 2.4. Therefore, the advantage of large firms from NTB reductions under the EUKFTA appears to be

[42] In this case, the interactions between the $\mathcal{I}_{dt}$ dummy, tariffs and size bins drop out due to multicollinearity and are therefore excluded in the estimations reported in Table 2.8.

[43] The elasticities remain identified even if we focus on products for which South Korea applied an MFN tariff of zero before the FTA as import tariffs set by other export destinations of France differ from South Korean tariffs.

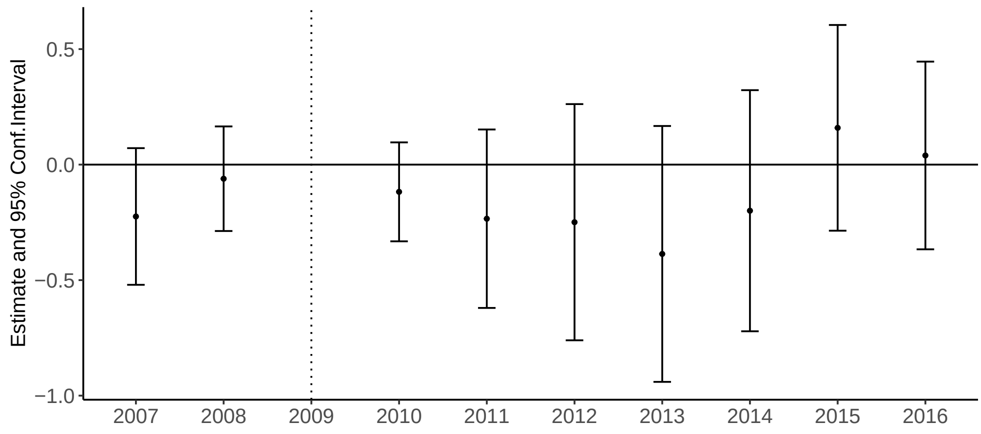
Table 2.8: Impact of EUKFTA across tariff staging categories

Dependent Variable:	ln(exports)				
Product categories:	MFN=0	EIF	3 years	10 years	11+ years
	(1)	(2)	(3)	(4)	(5)
$\mathcal{I} \times Q4$	1.48***	0.644***	0.741**	0.041	-0.410
	(0.304)	(0.176)	(0.333)	(0.240)	(0.546)
$\mathcal{I} \times Q3$	1.02***	0.539***	0.745**	-0.048	-0.022
	(0.346)	(0.174)	(0.341)	(0.207)	(0.498)
$\mathcal{I} \times Q2$	1.01**	0.292	0.284	-0.036	0.678
	(0.397)	(0.185)	(0.399)	(0.262)	(0.670)
$\ln t \times Q4$	-4.6***	-0.708	-2.78***	-0.137	-0.933
	(1.430)	(0.482)	(0.777)	(1.030)	(1.070)
$\ln t \times Q3$	-3.29**	0.033	-1.52**	0.542	0.583
	(1.430)	(0.414)	(0.752)	(1.010)	(1.230)
$\ln t \times Q2$	-3.42**	0.648	-1.89**	1.13	0.171
	(1.420)	(0.550)	(0.867)	(1.190)	(1.670)
Observations	191,936	893,520	252,892	224,251	43,156
R^2	0.924	0.919	0.908	0.916	0.926

Note: Regression results are based on equation (2.2) where the dependent variable is exports at the firm-product-destination level aggregated to two periods: 2000-2006 and 2011-2016. Using the tariff schedule of South Korea, the sample is split into exports of goods that were already duty-free in 2010 (MFN=0) (column (1)), those that became duty-free upon entry into force of the agreement (EIF) (column (2)), and goods whose tariffs were set to be gradually phased out by South Korea over three, ten or more than ten years (columns (3), (4), and (5)). $\mathcal{I}_{dt}$ is a dummy that takes the value 1 for South Korea in the post-treatment window (2011-2016) and zero otherwise. All regressions include firm-product-time, product-destination-time and firm-product-destination fixed effects. Only continuous exporters are retained i.e. firm-product combinations that report positive exports to a given destination in both control (2000-2006) and FTA (2011-2016) periods. Standard errors are clustered by firm, product and destination. Significance codes: ***: 0.01, **: 0.05, *: 0.1

driven via exports rather than imports. For tariffs, this is not overly surprising as the EU's level of protection before the FTA was much lower than the South Korean one. Possibly, a similar pattern prevailed regarding NTBs. Also note that the asymmetry between exports and imports is unlikely to be driven by exchange rate movements as the inclusion of destination-time fixed effects into all our specifications effectively nets out the influence of currency revaluations.

Figure 2.4: Import growth from NTB reductions: Top quartile ($Q4$) importers relative to bottom quartile ($Q1$)



Note: This graph shows the adjustment in import purchases of the top quartile of exporters relative to the bottom quartile, from a reduction in NTBs under the EUKFTA. Following the dynamic counterpart of the specification provided in equation (2.2), it plots coefficients and 95% confidence intervals on the interaction $Kor_d \times Year_t \times Q4$, where $Q4$ is a dummy variable that takes the value of one for importers in the top quartile of the size distribution and zero otherwise. The chosen reference year is 2009, the year before the EUKFTA was signed. A set of firm-product-year, product-destination-year and firm-product-destination fixed effects are included. Standard errors are clustered by firm, destination and year.

2.5 Conclusion

This paper sheds light on the effects of an important new generation trade agreement along the firm size distribution. We exploit French firm-level customs data for the period 2000 to 2016 and employ a differences-in-differences strategy to identify treatment effects for different quartiles of the size distribution. We find a new and robust stylized fact: Exporters of different sizes react differently to NTB reductions. More precisely, we document that French exporters with larger pre-FTA sizes expand their exports to South Korea by larger rates than firms further down the size distribution.

This effect is driven chiefly by NTBs, i.e., the summary effects of the FTA net of tariff concessions. It suggests that the NTB provisions of the FTA are not just about reducing the fixed costs of market access for firms, but also – and maybe predominantly – about lowering the variable trade costs for more efficient firms by more than for the less efficient ones. Interestingly, larger firms also seem to react more strongly to reductions in tariff reductions (which are independent of size), a result that is not specific to the EUKFTA and that only partly explains the overall advantage of large firms.

Our main finding confirms a widely-held prior that FTAs benefit larger firms more than smaller ones. Based on this presumption, many modern FTAs include special provisions that aim to support small and medium sized enterprises (SMEs). For example, since the release of the EU's 'Trade for All' strategy in 2015, the EU has included such SME provisions in all new trade agreements. These typically include commitments for the EU and its partners to provide information on the contents of the trade agreement on a dedicated website that has a database searchable by tariff code, with information on tariffs, import requirements, rules of origin, etc. In addition, such chapters provide for SME Contact Points on each side to facilitate bilateral cooperation between governments so that the specific needs of SMEs are addressed.

Such provisions can be readily justified on political economy and on equity grounds. They may be necessary to win the support of SMEs to conclude and ratify modern FTAs. Governments may also wish to spread the gains from trade more widely, as their incidence across firms affects the distribution of profits and wages. Whether SME chapters are required to enhance efficiency depends, however, on the details of the mechanism that gives rise to our empirical observation.

If for exogenous reasons, larger firms face lower iceberg trade costs and if those are complementary to politically induced variable trade costs, our result would simply reflect the technological superiority of larger firms without providing a rationale for policy intervention. Similarly, if higher sales in a foreign market require repeated lumpy payments – for example, as additional warehouses need be maintained or a bigger sales organization needs to be financed – larger firms have a natural advantage. The situation could be different if economies of scale or externalities are involved. This could be the case if learning-by-exporting externalities become stronger with size or if, contrary to Arkolakis (2010), marginal foreign market access costs are decreasing in market shares.

Whether such mechanisms are present and whether they justify political interventions on efficiency grounds, and if yes, which ones, however, depends on details. The case is clearest if the FTA itself contains elements that make it hard for small firms to benefit from the agreement. For example, it is known that rules-of-origin are costly to document and abide by. Simplifying them could benefit smaller exporters more than the usual SME chapters while also increasing overall efficiency. Similarly, if size differentials reflect lobbying activity by large firms, it would be advisable to install safeguard against such attempts during the negotiation process. To make further progress, one would need to develop and test structural models that embed such mechanisms.

Moreover, further empirical research should test the effectiveness of existing SME provisions exploiting the fact that an increasing share of FTAs contain such language. Also, it would be important to test for size-specific effects of NTBs and tariffs across the exporter size distribution in other trade agreements and in other countries. Our focus on a representative European agreement – the one between the EU and South Korea – and on French exporters could be the start for a broader and more comprehensive research agenda.

Chapter 3

Brothers in Arms: The Value of Coalitions in Sanctions Regimes

3.1 Introduction

The 'war by other means' theory proposed by Blackwill and Harris (2016) argues that economic policies such as sanctions are effective instruments for the pursuit of geopolitical objectives. In their view, sanctions can serve as strategic substitutes for military intervention and reduce the risk of armed conflict. This unique form of (economic) statecraft is made feasible by sanctioning states' central position in the global economy and their influence over international flows of capital, commodities and services. In this paper, we bring the Blackwill and Harris (2016) theory to the data in two key respects. First, we examine the welfare costs incurred from employing sanctions for geopolitical purposes and their distribution across sanctioning and sanctioned states. Second, we provide novel evidence on how the deterrent force of these sanctions relies upon the structure of sanctioning coalitions i.e. alliances of countries that coordinate their sanctions packages.

To magnify the impact of sanctions, countries have made considerable efforts to coordinate their actions. Substantial diplomatic capital is spent toward enlarging coalitions or preserving them when measures escalate in severity. Negotiations surrounding sanctions packages also tend to be highly sensitive as they test the unity of existing alliances. Even within the EU, divisions emerged over the extent of Russia sanctions since the bloc's members have varying degrees of dependence on Russian gas and bear different economic costs from imposing sanctions (Chowdhry et al., 2020). Not only do coalitions increase the "moral suasion" of sanctions (Hufbauer, Schott, and Elliott, 1990) and raise the economic costs incurred by targeted regimes, they also reduce opportunities for sanctions-busting i.e. the circumvention of sanctions by directing trade toward third parties. Building and maintaining coalitions is therefore difficult but vital to the effective implementation of sanctions.

To analyze the economic impact of sanctions coalitions, we rely upon a computable general equilibrium (CGE) model. This is a Caliendo and Parro (2015)-type model of the world economy based on Hinz and Monastyrenko (2022) and similar to Chowdhry et al. (2020). It allows us to compute changes in welfare that would be experienced by sanctioned and sanctioning states as well as third parties under different hypothetical setups of sanctions coalitions. By comparing welfare changes across various constellations of sanction alliances, we highlight the value of coordinating sanctions packages.

To examine these issues, we focus on the sanctions regimes against Iran in 2012 and Russia in 2014 as they have been unprecedented in terms of depth and severity. In the case of Iran, we analyze the wave of sanctions that occurred in 2012 following concerns related to the country's nuclear program. Amongst these sanctions, the hardest-hitting measures included an embargo against Iranian oil and natural gas and the isolation of Iran from the SWIFT system and global financial markets. These sanctions were eased in 2016 as part of the "Joint Comprehensive Plan of Action" (JCPOA) deal. However, the withdrawal of the United States from the JCPOA in 2018 triggered a reinstatement of sanctions against Iran.

In addition to Iran, we also examine the series of sanctions imposed against Russia in 2014 following its annexation of Crimea. These sanctions were initially limited to targeted travel bans, visa restrictions and asset freezes on Russian and Crimean officials. However, sanctions were toughened following the shooting down of a civilian airplane in the contested Donbas region in July 2014. After this incident, new trade and financial sanctions were imposed. These measures included restrictions on exports of dual-use and sensitive technologies, restrictions on access to loans and capital markets for major Russian banks, energy companies and defense equipment manufacturers and the addition of more Russian entities on the sanctions list. Together, these policies aimed to severely restrict economic activity in Russia. In August 2014, Russia retaliated by banning imports of agri-food products from sanctioning states. These sanctions regimes have continued and escalated even further in 2022 following Russia's invasion of Ukraine.

We conduct several simulations in order to compare the welfare outcomes under different hypothetical sanctions coalitions. In our first counterfactual scenario, we show that welfare losses incurred by Iran and Russia are higher when sanctions are imposed collectively rather than unilaterally by coalition members. In the second scenario, we find that China's potential cooperation in the sanctions package against Russia and Iran would significantly increase the punitive force of these sanctions without imposing high welfare costs on existing sanctioning states or China itself. In the third scenario, we expand on the second scenario and evaluate which countries would need to be brought on board as prospective allies to amplify the deterring effect of sanctions.

In the case of Iran sanctions, the most important third-party countries which would increase the punitive impact of sanctions are China, India, UAE, Brazil, Azerbaijan and Russia. For Russian sanctions, the leading potential coalition partners are China, South Korea, Belarus, Turkey, Brazil and India. Finally, in the fourth scenario, we compute how burden-sharing within coalitions could mitigate the adverse welfare impact experienced by sanctioning states.

The remainder of the paper is structured as follows: In Section 3.2, we provide an overview of the growing literature on the economic impact of sanctions and the role of coalition-building in sanctions regimes. Section 3.3 describes the various data sources used for estimating the gravity model and counterfactual scenarios. These gravity estimations are reported in Section 3.4 while simulation results are described in Section 3.5. Finally, Section 3.6 concludes with policy recommendations.

3.2 Related Literature

Sanctions can stabilize or destabilize governments, lead to power shifts within countries and change the political relationship between sender and target (Escribà-Folch, Wright, and Wright, 2015). However, sanctions' effectiveness in achieving political goals is contested (Pape, 1997; Pape, 1998; Hufbauer, Schott, Elliott, and Oegg, 2007). Moreover, authoritarian regimes often redistribute the burden of sanctions within their countries while bolstering regime support via the "rally-around-the-flag" effect (Grauvogel and Von Soest, 2014).

Sanction effectiveness crucially hinges upon two things: First, sanctions have to be threatened with credibility, and second, the sanctioning coalition has to be broad enough to display this credibility and to exert economic pressure. In the political science literature, research on (military) coalitions and alliances generally restricts itself to the context of conflict. There are however, some papers that document important findings on how alliances work and affect trade outcomes. These works have studied the relationship between trade and changing military partnerships and show that intensively trading country-pairs are less likely to be involved in military disputes (Polachek, 1980; Pollins, 1989). Military alliances are also observed to have a large positive effect on bilateral trade, with the effect being higher in bipolar than in multipolar systems (Gowa and Mansfield, 1993). Furthermore, alliances that include a major power trade more than those without major power participation (Mansfield and Bronson, 1997). Overall, these papers reflect the role of alliances in preserving trade flows.

More recent studies in this literature investigate the role of third parties and their relationship to sanction-senders and targets. For instance, Early (2012) analyzes 96 episodes of US sanctions to show that third parties will cooperate (sanction bust) when the costs from sanctions are low (high). Peksen and Peterson (2016) provide further discussion on the role of third parties as alternative markets for targeted nations. The paper shows that sanction-senders are more likely to threaten or impose sanctions when the target has limited opportunities to redirect lost trade to third parties.

In the future, more sanctions can be expected to enforce compliance with environmental policy objectives. In this regard, coalitions and trade sanctions have also been examined in the context of climate change negotiations. For instance, Lessmann, Marschinski, and Edenhofer (2009) use a model of coalition stability and find that trade sanctions can raise participation in International Environmental Agreements. Nordhaus (2015) finds that climate coalitions remain unstable without sanctions against non-participants and proposes a Climate Club with minor trade penalties to induce stable coalitions and high levels of mitigation. Hagen and Schneider (2017) also examine the stability of climate change coalitions in the presence of sanctions by incorporating the possibility of retaliation by targeted nations. In doing so, they find an interesting threshold effect. Coalitions are seen to be stable only when they reach a specific size. Conversely, when the coalition size falls below the threshold, the effect of retaliation dominates and decreases incentives to becoming a coalition member. Therefore, the risk of retaliation to sanctions significantly reduces the willingness to form coalitions.

One aspect that the literature in economics and political science has shown clearly is that sanctions typically have negative consequences not only for targets (see e.g. Afesorgbor and Mahadevan, 2016), but also for senders, with costs distributed unequally across countries within sanctioning coalitions. While the EU itself increasingly uses sanctions (Attia and Grauvogel, 2019) the sanctions policy of the US has affected the EU and its member states asymmetrically through direct costs (Chowdhry et al., 2020).

Economically, sanctions (see e.g. Crozet and Hinz, 2020; Neuenkirch and Neumeier, 2015; Neuenkirch and Neumeier, 2016), blockades and embargoes (see e.g. Etkes and Zimring, 2015; Heilmann, 2016) affect financial and/or trade flows and can result in welfare losses, trade destruction or trade diversion. Accounting for third-country effects is thus of great importance as even "bilateral" sanctions lead to trade diversion and hence impact third parties through global value chains. However, a shortcoming in the related political science literature is that analysis is mainly focused on a two-country world, which limits explanatory power (Early and Cilizoglu, 2020; Felbermayr, Kirilakha, et al., 2020). An analysis with a multi-country multi-sector trade model to account for trade diversion and sectoral effects is thus desired.

3.3 Data

For estimating the gravity model, data on trade flows is taken from UN Comtrade. Furthermore, the CEPII Gravity dataset is used for information pertaining to variables such as joint membership of countries in the WTO, free trade area or currency union. Finally, the CGE model is estimated and calibrated using standard data sources. The main input for simulations for the model are derived from the GTAP 10 database (Aguiar et al., 2019). This data supplies the model with information on consumption shares, input coefficients, bilateral trade shares, trade balances and bilateral tariffs. Furthermore, the data is concorded to 65 GTAP sectors and 141 countries or regions.

3.4 Impact of the 2012 Iran and 2014 Russia Sanctions

3.4.1 Model of the world economy

We build a Caliendo and Parro (2015)-type model of the world economy based on Hinz and Monastyrenko (2022) and similar to Chowdhry et al. (2020) in order to run a variety of simulations. We provide a brief overview of the model setup in Appendix C.1. Conveniently, the model yields a sectoral gravity equation that — including a time dimension — can be estimated as follows:

$$X_{odst} = \exp\left(\boldsymbol{\beta} z_{odst} + \Gamma_{ost} + \Gamma_{dst} + \Gamma_{osd}\right) \tag{3.1}$$

The dependent variable is the value of trade (exports or imports) from origin (o) to destination (d) in sector (s) in a given year (t). Equation (3.1) includes fixed effects Γ_{ost}, Γ_{dst}, and Γ_{ods} to purge all origin $\times$ sector $\times$ time and destination $\times$ sector $\times$ time-specific factors, as well as unobserved time-invariant and sector-specific bilateral characteristics. In addition, the specification features z_{odst} which is a vector of time-varying bilateral trade frictions. This includes the incidence of sanctions, as well as customary important policy variables like joint membership in the WTO, an FTA, or a currency union. Correspondingly, $\boldsymbol{\beta}$ is the vector of the respective sector-specific coefficients. The equation is estimated with a Poisson pseudo-maximum likelihood (PPML) procedure where the standard errors are clustered three-way by origin, destination and year.

Sanctions impact on aggregate trade flows

Before estimating the disaggregated sectoral gravity model as described by Equation (3.1), we first examine the impact of Iran and Russia sanctions on aggregate trade flows. Results from this exercise are reported in Table 3.1. Looking at Iran, we find that sanctions reduced overall exports flows by $(\exp(0.3641) - 1) \times 100 = 43.92\%$. Sanctions have an even stronger effect on imports, with trade flows dropping by 48.04%. In the case of Russia, sanctions policies decrease export flows by 36.71% whereas import are reduced by 22.89%. Together, these coefficients indicate the severe impact of sanctions imposed on Iran and Russia on aggregate trade.

Table 3.1: Impact of the Iran and Russia sanctions on aggregate international trade

Dependent Variable:	Trade value
Sanctions on flows to Iran	-0.3641***
	(0.0597)
Sanctions on flows from Iran	-0.3923***
	(0.0505)
Sanctions on flows to Russia	-0.3127***
	(0.0695)
Sanctions on flows from Russia	-0.2061***
	(0.0634)
WTO	0.3239*
	(0.1821)
Common currency	0.1408*
	(0.0763)
FTA	0.0713*
	(0.0391)
Fixed-effects	
origin × year	Yes
destination × year	Yes
origin × destination	Yes
Fit statistics	
Observations	188,795
Pseudo R^2	0.9921

Note: Clustered (origin, destination & year) standard-errors in parentheses
*Signif. Codes: ***: 0.01, **: 0.05, *: 0.1*

Sanctions impact by sector

Having analyzed the impact of the Iran and Russia sanctions on aggregate trade flows, we next turn to estimating Equation (3.1) at the sectoral level. We use these estimated percentage changes in exports and imports due to the sanctions by sector to inform the simulations in Section 3.5.[1] These sector-wise coefficients are reported in Figure 3.1 and Figure 3.2.

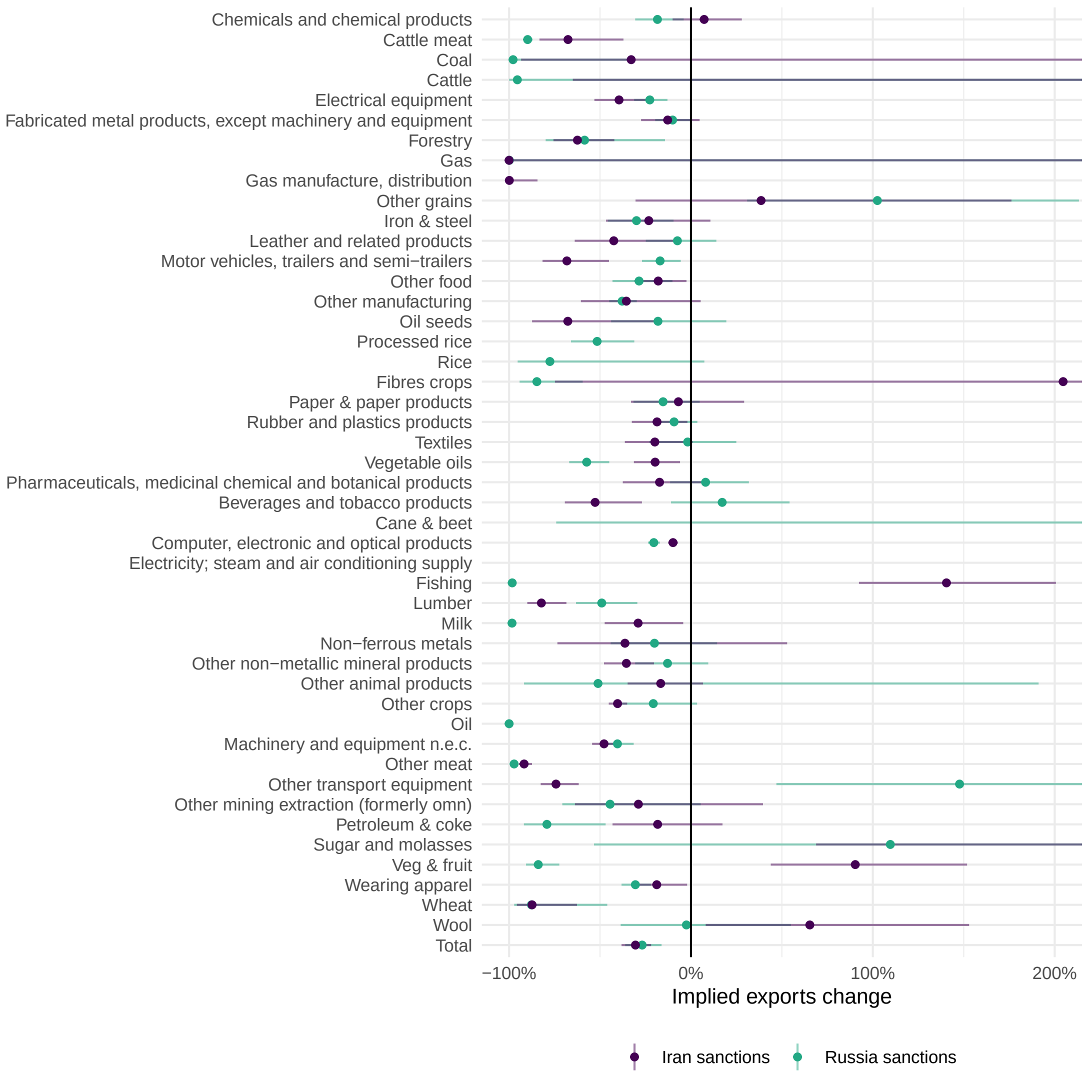
Looking at the export side in Figure 3.1, we note that sanctions reduced exports to Russia and Iran across a broad range of sectors as the majority of the estimated coefficients are negative and statistically significant. Some of the sectors hit hardest in Iran include gas, meat, lumber, transport equipment and motor vehicles. In the case of exports to Russia, the contraction is deepest in energy products such as oil, gas, coal and petroleum. Following Russia's embargo on imports of agri-food products from sanctioning states in 2014, we also see high and negative coefficients for exports to Russia in goods such as milk, fishing, rice, vegetables and fruit. The few sectors such as other grains, electricity and fiber crops that report positive coefficients have wide confidence intervals and are largely not statistically significant.

Turning toward the impact of sanctions on imports in Figure 3.2, we again observe negative coefficients for the majority of sectors. The steepest decline in imports from Iran is in wheat, fiber crops, textiles and electrical equipment. For imports sourced from Russia, the detrimental effect of sanctions is highest for agri-food products including cane and beet, cattle, other meat, fiber crops, vegetable oils and beverages.

Together these figures reveal that sanctions against Iran and Russia hit trade flows in most sectors but by varying degrees. This heterogeneity in sectoral responses to sanctions is masked by gravity estimations at the aggregate level. Therefore, subsequent simulations with the CGE model will draw upon these sectoral elasticities to generate more precise counterfactuals for the role of coalitions in sanctions regimes.

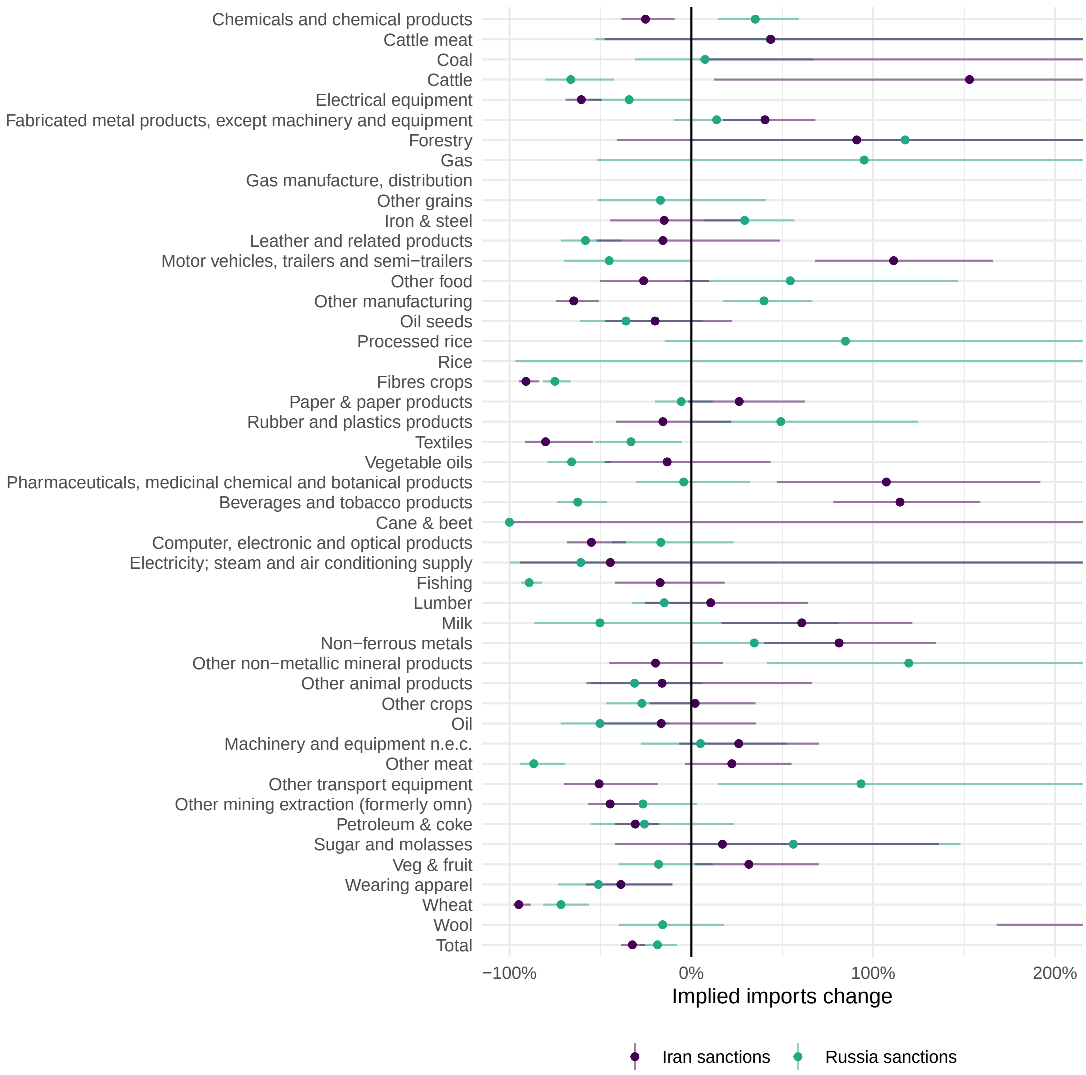
[1] Note that we only use those estimated changes where the coefficient is statistically significant at the 95% interval.

Figure 3.1: Sanctions impact on exports to Russia and Iran by sector



Note: Figure above displays coefficients and their 95% confidence intervals derived from estimating the sectoral gravity model as outlined in equation (3.1). The estimates capture the impact of sanctions on each sector's exports to Iran and Russia.

Figure 3.2: Sanctions impact on imports from Russia and Iran by sector



Note: Figure above displays coefficients and their 95% confidence intervals derived from estimating the sectoral gravity model as outlined in equation (3.1). The estimates capture the impact of sanctions on each sector's imports from Iran and Russia.

Table 3.2: Welfare change

Case	Scenario	Welfare change (%)
Iran	Multilateral	-2.65
Iran	Unilateral	-2.36
Russia	Multilateral	-8.94
Russia	Unilateral	-7.53

Note: The table above displays the collective welfare loss incurred by all members of the current sanctions coalition against Iran or Russia. In the multilateral scenario, members coordinate their sanctions whereas in the unilateral scenario, each country imposes its sanctions in isolation.

3.5 Simulating the Value of Coalitions

While gravity estimations in Section 3.4 revealed the negative impact of sanctions on trade flows at the aggregate or sectoral level, they do not account for the full economic costs associated with sanctions. To do so requires running simulations in a general equilibrium model. Therefore, we now compute a series of counterfactual scenarios that evaluate different setups of sanctions coalitions with the help of the CGE model.

The model computes changes in welfare when moving from the baseline case where no country imposes sanctions on Iran or Russia to the counterfactual scenarios which feature different hypothetical sanctions coalitions. This change in welfare is then compared to the benchmark welfare loss from sanctions, calculated as the difference between welfare in the baseline (no sanctions) and welfare in the status-quo i.e. under the current state of sanctions coalitions against Russia and Iran.

3.5.1 Individual contributions of countries

In the first set of scenarios, we examine the change in welfare costs from the baseline if each member of the current sanctioning coalition unilaterally imposed sanctions on Russia or Iran. Each of these scenarios corresponds to a complete breakdown of existing sanctions coalitions as each sanctioning country acts in isolation and is therefore the only sanctioning state. Doing so allows us to compute the 'contribution' of each coalition member toward the sanctions regime in the status quo. Defining the counterfactuals in such a manner has an additional advantage. It allows us to contrast the combined welfare effect of countries coordinating their sanctions against the effect (summed across countries) of imposing sanctions on a unilateral basis.

Results from these counterfactuals are reported in Table 3.2. In both cases, the benchmark welfare loss incurred by Iran and Russia under the existing coalition framework exceeds the total welfare loss incurred if the coalition members were to impose sanc-

Figure 3.3: Iran sanctions

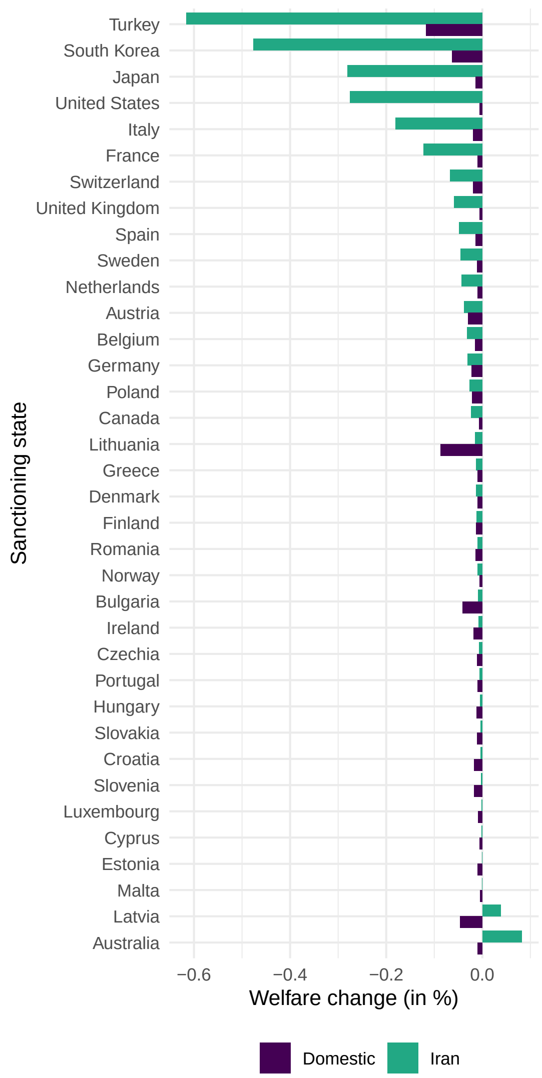
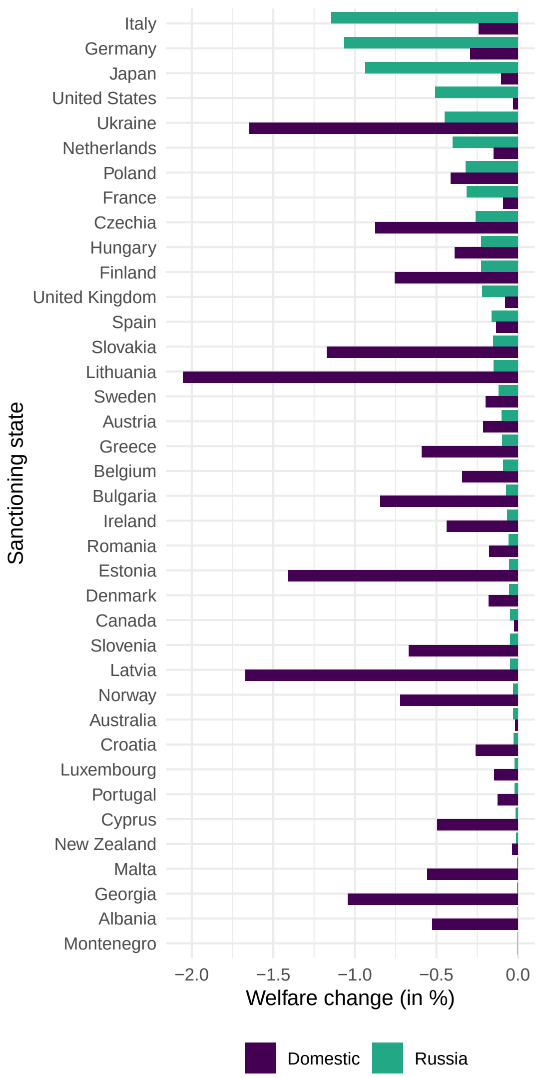


Figure 3.4: Russia sanctions



Note: Figures above display each country in the current sanctions coalition against Iran or Russia and the welfare change it experiences domestically and that which it imposes on the sanctioned state.

tions unilaterally. This difference between the benchmark and unilateral scenarios demonstrates the impact of coordinating sanctions across countries. For Iran, the welfare loss from coordinated sanctions is 0.29 pp higher whereas for Russia, its 1.41 pp greater.

How costly is it for sanction senders to impose these welfare losses on targeted nations? To investigate this issue, we next compare the welfare loss incurred by Iran and Russia against the welfare loss incurred domestically for each of the sanctioning states. Note that these domestic welfare losses stem from the increase in cross-border frictions which raises the operating costs for businesses trading with sanctioned states. The costs are further magnified in the presence of supply chains and for countries dependent upon inputs sourced from the sanctioned state. In Figure 3.3 and Figure 3.4, we plot the welfare changes experienced by sanctioning states domestically and that incurred by the sanctioned state. The simulations produce several interesting outcomes. First, countries differ substantially in their capacity to inflict welfare loss on the sanctioned state. For Iran, the coalition members which exert the highest coercive force (welfare loss on Iran) are Turkey (-0.62%), South Korea (-0.48%), Japan (-0.28%), US (-0.28%), Italy (-0.18%), France (-0.12%) and Switzerland (-0.07%). Looking at Russia, the punitive impact is highest for coalition members such as Italy (-1.14%), Germany (-1.06%), Japan (-0.93%), US (-0.51%) and Ukraine (-0.45%).

Considering the magnitude of domestic welfare losses, we observe substantial skewness across countries in their contribution toward the sanctions regime. Seen in this alternative respect, the top five contributors to the Iran sanctions are Turkey (-0.12%), Lithuania (-0.09%), South Korea (-0.06%), Latvia (-0.05%) and Bulgaria (-0.04%). In the case of Russia, the leading contributors are Lithuania (-2.05%), Latvia (-1.67%), Ukraine (-1.65%), Estonia (-1.41%) and Slovakia (-1.17%). For these nations, sanctions against Russia are extremely costly. Moreover, these domestic welfare losses are comparable between the unilateral and multilateral sanctions scenarios.[2]

Third, we note that the US is the most effective in imposing the Iran and Russia sanctions in terms of welfare cost borne at home vis-à-vis welfare loss imposed on the target. Closely following the US are Japan, France, the UK, Italy and Germany. In comparison, nations such as Lithuania, Estonia and Latvia incur high costs of sanctions that translate only into marginal welfare loss for Russia. Several of these countries are also ranked low in terms of their effectiveness in imposing sanctions against Iran. Therefore, coalition members differ substantially not only in their economic expenditure or contribution towards the sanctions regime but also in their effectiveness.

3.5.2 Impact of non-cooperating China

In this scenario, we evaluate the implicit welfare cost associated with China's non-alliance. Here, the counterfactual consists of China joining the existing sanctioning coalition and therefore imposing new export and import sanctions against Russia and Iran. The resulting shift in welfare costs borne by Russia and Iran as well as the sanctioning states then reflects the hidden costs of China's non-cooperation in the status quo. Results from these two counterfactuals are reported in Table 3.3 and Table 3.4.

[2] For further details, see Figure C.1 and Figure C.2 in Appendix C.2.

In each table, we report the benchmark welfare change (i.e. under the existing sanctions coalition) and the welfare change from China joining the coalition. These changes are computed from a baseline scenario where no country sanctions Russia or Iran. In both cases, China's involvement in the sanctions regime greatly deepens the welfare loss incurred by the sanctioned state. Iran's welfare reduces by an additional 6.19 pp and Russia's by 2.31 pp.

At the same time, China itself incurs minimal additional welfare loss from joining the Iranian (0.06 pp) or Russian (0.1 pp) sanctions coalitions. Moreover, existing coalition members do not experience any significant changes in their welfare from China sanctioning Iran or Russia. Taken together, these counterfactuals indicate that China can substantially raise the coercive power of the sanctions regimes without facing significant domestic welfare costs or imposing such costs on current coalition members. Interestingly, an important dimension of the implicit cost of China's non-cooperation in the status quo is borne by the rest of the world (RoW). This group could experience a 0.01 pp increase in their welfare (on average) from China sanctioning Iran and a 0.02 pp increase from China sanctioning Russia.

Table 3.3: Iran sanctions with China

	Benchmark	incl. China
Iran	-2.65 %	-8.84 %
Current coalition	-0.02 %	-0.02 %
China	0.01 %	-0.05 %
Rest of the world	0.01 %	0.02 %

Table 3.4: Russia sanctions with China

	Benchmark	incl. China
Russia	-8.94 %	-11.25 %
Current coalition	-0.52 %	-0.52 %
China	0.02 %	-0.08 %
Rest of the world	0.11 %	0.13 %

Note: The tables above display welfare changes from sanctions in the benchmark scenario and a scenario in which China joins the existing sanctions coalitions.

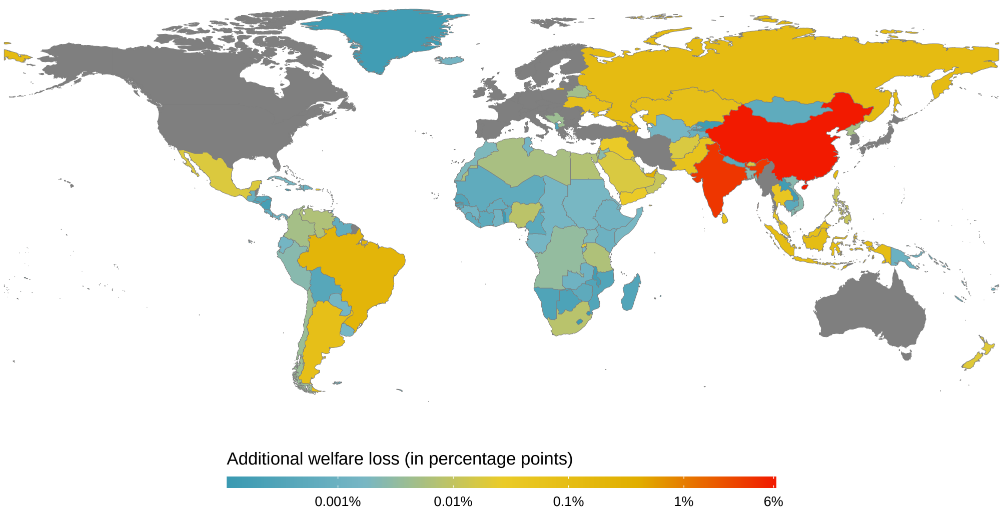
3.5.3 Ideal coalition partners

In the next set of scenarios, we examine the issue of which countries would need to join the existing sanctions coalitions in order to make sanctions more costly for targeted nations when compared to the benchmark scenario. To do so, we construct a succession of counterfactuals. In each counterfactual, we expand the current sanctioning coalition by including one additional country that does not impose sanctions against Russia or Iran in the benchmark case. This presumes that the addition of new members in the sanctions coalition does not cause other coalition members to depart.

This generates a series of counterfactuals, one for each country in the world that did not sanction Russia and/or Iran in 2019. Comparing the welfare loss incurred under these various counterfactuals with the welfare loss under the benchmark scenario allows us to compute the additional coercive power of each third-party country to the sanctions coalition. Based on this, we create a ranked list of nations that would be 'ideal' coalition partners to be approached if the existing coalition decides to strengthen the sanctions regimes against Russia and Iran. The results from these counterfactuals are depicted in Figure 3.5 and Figure 3.6. In both maps, we plot the additional welfare loss incurred by the targeted nation from each country joining (one at a time, with replacement) the existing sanctions regime against Iran or Russia.

In the case of Iran sanctions, the most important third-party countries which would increase the punitive impact of sanctions are China (6.19 pp), India (3.72 pp), UAE (0.36 pp), Azerbaijan (0.26 pp), Brazil (0.22 pp) and Russia (0.12 pp). For Russian sanctions, the leading potential coalition partners that would increase the welfare loss for Russia are China (2.3 pp), South Korea (1.05 pp), Belarus (0.77 pp), Turkey (0.3 pp) and India (0.14 pp). Coordinating sanctions with these countries would reduce opportunities for sanctions-busting by targeted nations and increase the effectiveness of sanctions regimes.

Figure 3.5: New coalition partners: Welfare loss imposed on Iran

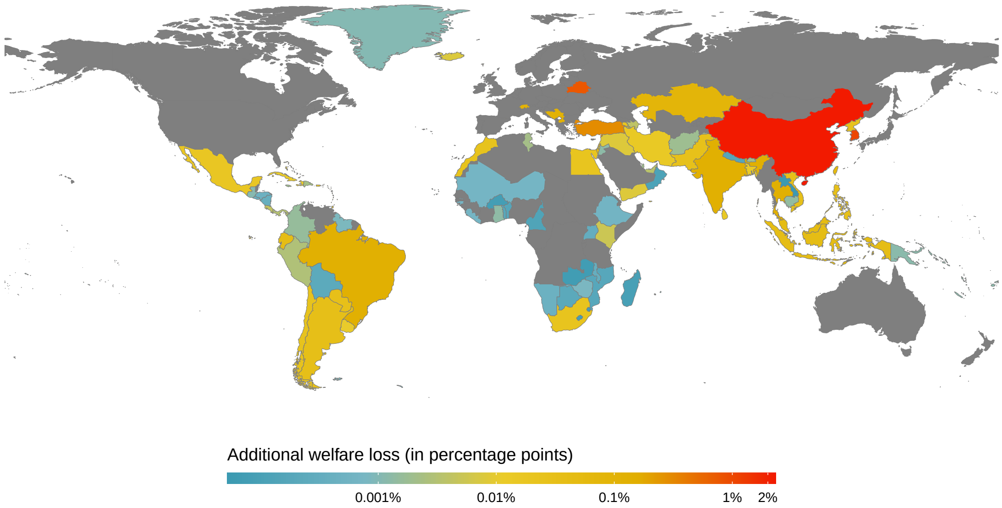


Note: The map above displays the additional welfare loss incurred by Iran from each new country joining the current sanctions coalition. Iran and countries which already sanction Iran are depicted in grey.

3.5.4 Burden-sharing

The counterfactual results reported in Section 3.5.1 show that sanctions impose uneven domestic welfare costs on coalition members. Therefore, in the final set of sce-

Figure 3.6: New coalition partners: Welfare loss imposed on Russia



Note: The map above displays the additional welfare loss incurred by Russia from each new country joining the current sanctions coalition. Russia and countries which already sanction Russia are depicted in grey.

narios, we examine the potential for burden-sharing within the coalition. Calls for such burden-sharing mechanisms have been raised previously by countries at the UN, given the increasing frequency and severity of sanctions.[3] In mitigating the adverse impact of sanctions and their asymmetric incidence across countries, burden-sharing policies can also stabilize sanctions coalitions and incentivize new countries to join.

Here, we investigate one potential mechanism by which sanctioning states can offset the economic costs incurred from the Iran and Russia sanctions. For each of these sanctions regimes, we evaluate the change in welfare if coalition members were to deepen their economic cooperation. In terms of the model, this increased cooperation is translated into a specific reduction in non-tariff barriers (NTBs) that applies to all trade flows within the coalition.

By increasing access to coalition partners' markets, such NTB reductions can potentially counteract the decline in trade between sanctioning states and the targeted nation. Such burden-sharing proposals have been highlighted in the international relations literature including Boudreau (1998) who argues that leading UN members should facilitate new trading arrangements as a form of mutual support amongst sanctioning states. Increasing market integration within the sanctions coalition is also a

[3]See "Calls for burden-sharing mechanism to ease sanctions effects on third states, as Sixth Committee continues discussion of report of Charter Committee", United Nations Press Release GA/L/3075, October 1998. Link: https://bit.ly/3pZgmMh.

Figure 3.7: burden-sharing of Iran sanctions

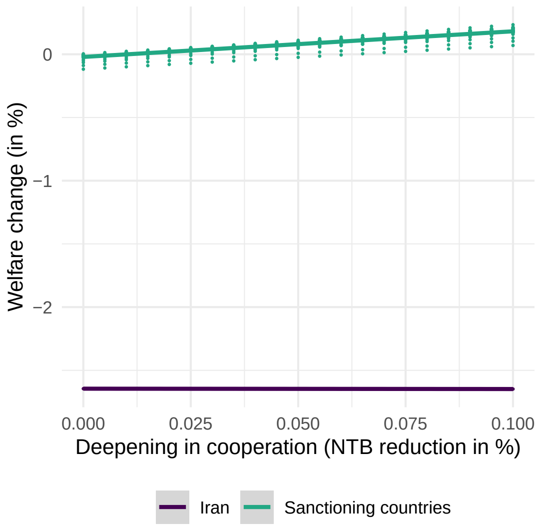
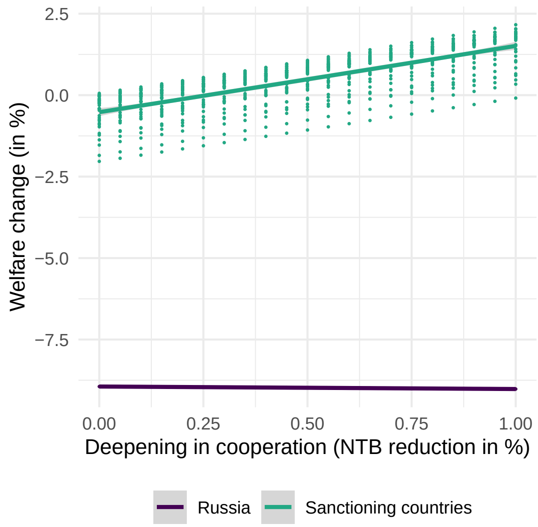


Figure 3.8: burden-sharing of Russia sanctions



Note: The figures above display welfare changes experienced by the current coalition members (green) and targeted state (purple) under various scenarios featuring deepening economic cooperation or reductions in NTBs within the coalition (horizontal axis).

more feasible burden-sharing mechanism given the difficulties in establishing direct compensation or transfers of sanctions-induced economic costs.

The resulting welfare impact of deepening economic cooperation between sanctioning countries is shown in Figure 3.7 and Figure 3.8. In each graph, we report welfare changes across several simulated counterfactuals where each counterfactual corresponds to a specific reduction in NTBs. It allows us to examine the depth of NTB reduction required to compensate the welfare losses of all sanctioning members fully. In the case of Iran sanctions, close to 0.1% of NTB reductions are required across all coalition states in order to fully offset the cost of sanctions. In the case of Russian sanctions, the NTB reduction needed is higher. All sanctioning states must lower NTBs by at least 1% in order to fully compensate every member for the welfare loss from sanctions. In line with expectations, this deepening cooperation amongst the Iran or Russia sanctions coalitions has no clear impact on the welfare loss incurred by the targeted nation. Together, these results lend credence to the proposals of burden-sharing through closer economic cooperation. They reveal that reducing trade frictions within the coalition can minimize the expense of imposing sanctions and therefore serve to keep these coalitions viable.

3.6 Conclusion

This paper provides novel empirical results concerning the impact of coalitions on the economic cost and deterrent power of sanctions. To do so, we examine various hypothetical geometries of sanctions coalitions against Iran and Russia and compute the welfare losses they impose on sanctioning and sanctioned states. The welfare losses are calculated by running simulations with a Caliendo and Parro (2015)-type CGE model that uses sector-specific trade elasticities (sanctions coefficients) drawn from structural gravity estimations.

The simulations provide convincing evidence that coalitions serve two important purposes. First, they can magnify the coercive force of sanctions regimes by raising the welfare losses incurred by targeted nations. Second, they can reduce the welfare losses borne by individual sanctioning states, such as when the coalition moves towards deeper economic cooperation. These twin objectives of raising the punitive force of sanctions whilst lowering domestic welfare losses depend not just on the size of the coalition but also the constellation of nations that belong to it. For instance, leading developing economies such as China, India and Brazil are 'ideal' prospective allies for increasing the cost of sanctions on targeted nations. The cost of not having these members in the existing sanctions coalition is particularly high in the case of China. Counterfactuals show that China's non-cooperation in sanctions against Iran and Russia in the status-quo greatly reduces the deterrence capability of sanctions.

Results in this paper also suggest that tighter economic cooperation between sanctioning states in the form of lower NTBs can offset the welfare losses from sanctions. Moreover, the NTB reductions required to compensate for these welfare losses is not high. Taken together, these results have profound policy implications. They suggest that if members offer even marginal improvements in market access to those in the sanctions coalition, the welfare cost of engaging in such 'economic warfare' can be neutralized. Beyond lowering the domestic cost of sanctions, deeper cooperation between existing coalition members provides an additional advantage. It can act as a burden-sharing mechanism that reduces the skewness in how welfare losses are distributed across coalition members. Simulation results show that these welfare costs from Iran and Russia sanctions are disproportionately borne by small states such as Latvia, Lithuania and Estonia. Given this inequity in economic burdens, deeper economic cooperation within the coalition can make sanctions regimes more viable and stable in the long run. In conclusion, this paper contributes to the vibrant literature on sanctions by demonstrating the vital importance of coordinating sanctions packages through coalitions.

Chapter 4

Après-ski: The Spread of Coronavirus from Ischgl through Germany*

4.1 Introduction

By mid-May 2020, the highly contagious SARS-CoV-2 virus infected about 4.5 million people worldwide and led to almost 300,000 fatalities.[1] The outbreak prompted governments to impose lockdowns affecting nearly 3 billion people worldwide, in an unprecedented attempt to 'flatten the curve' of infections so that healthcare systems are not overwhelmed. In Germany, even though restrictions were phased in from March 9th to 23rd 2020, the number of confirmed cases increased to approximately 175,000 with almost 8,000 deaths by mid-May 2020.[2] However, the spread *within* Germany was far from homogeneous — the two southernmost states, Bayern and Baden-Württemberg, were amongst the most affected, and even within these states there was a lot of variation.

Figure 4.1 depicts the spatial distribution of confirmed COVID-19 cases per 100,000 inhabitants in each of the 401 'Kreise' (counties)[3] using data provided by the Robert-Koch Institute, the German federal government agency and research institute responsible for disease control and prevention.[4] The left-hand side map indicates that as early as March 13th 2020, 356 out of 401 counties already reported some confirmed cases. By May 9th, infections had increased across the country with counties in southern and eastern Germany experiencing significantly higher case burdens, as shown by the histogram on the right-hand side.

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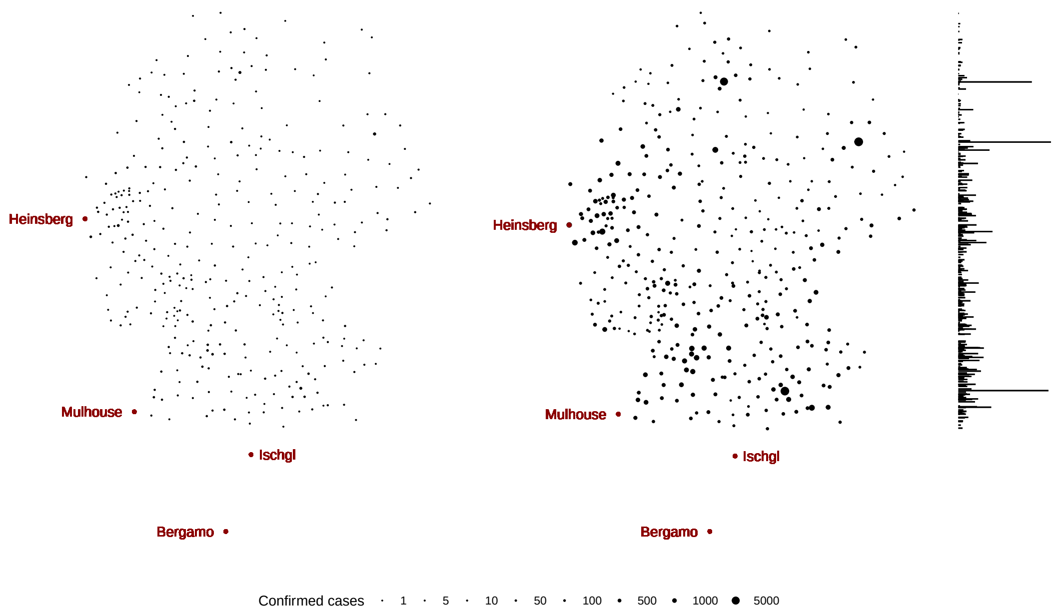
[1]See e.g. https://ourworldindata.org/coronavirus-data.

[2]See https://experience.arcgis.com/experience/478220a4c454480e823b17327b2bf1d4/page/page_-1/.

[3]Strictly speaking, in Germany there are 294 so-called 'Landkreise' (rural counties) and 107 'kreisfreie Städte' (cities not belonging to any 'Kreis').

[4]See https://www.rki.de/DE/Content/InfAZ/N/Neuartiges_Coronavirus/Fallzahlen.html.

Figure 4.1: Confirmed cases in Germany on March 13th and May 9th, 2020



Note: Map on the left-hand side shows confirmed cases on March 13th 2020, map on the right-hand side shows those on May 9th 2020. The histogram to the very right shows the change by latitude, binned by county.

The county with the lowest case incidence rate (CIR) was Mansfeld-Südhartz in North-Eastern Saxony-Anhalt (0.03%). Tirschenreuth in Bavaria, the most affected county, had a CIR 52 times higher (1.53%).[5] Across counties, the standard deviation of the CIR was almost as large as its mean. A similar dispersion was observed for the case fatality rate (CFR), which was reported as zero for 26 counties.[6]

Which factors explain this spatial distribution? In this paper, we explore whether tourists visiting super-spreader locations, in particular the resort town of Ischgl in neighbouring Austria, brought home the virus from trips in February and March 2020, as hypothesized by German and international media outlets.[7] Another earlier hotspot in Germany, the county of Heinsberg, located in the Carnival-celebrating Rhineland region, may have contributed to the diffusion of the virus. The transmission may have occurred from other neighbouring countries as well such as the highly affected town

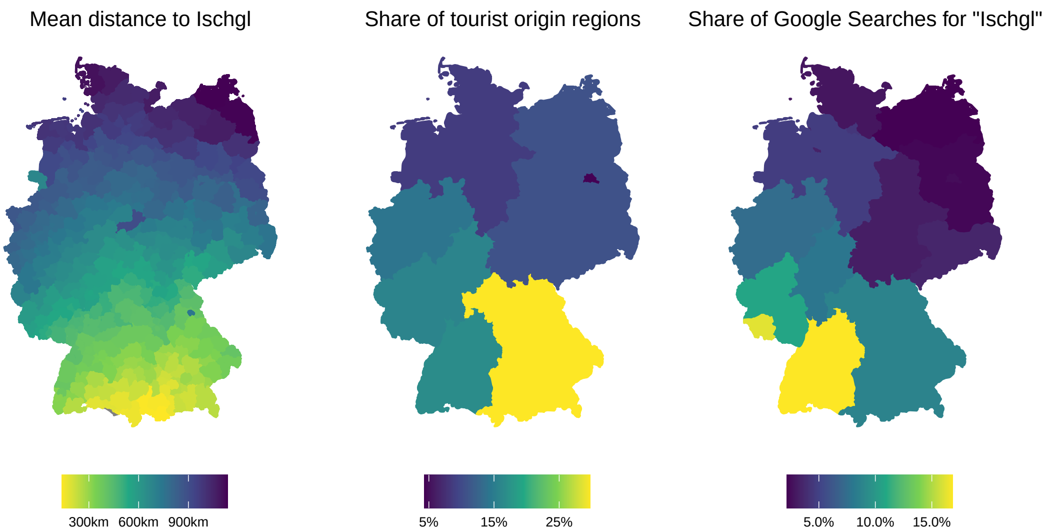
[5]Case incidence rate (CIR) is defined as the number of infected individuals divided by population size.

[6]Case fatality rate (CFR) is defined as the number of confirmed deaths divided by the number of confirmed cases.

[7]See e.g. "A Corona Hotspot in the Alps Spread Virus Across Europe", March 31st, 2020, Der Spiegel (https://www.spiegel.de/international/world/ischgl-austria-a-corona-hotspot-in-the-alps-spread-virus-across-europe-a-32b17b76-14df-4f37-bfcf-39d2ceee92ec).

of Mulhouse in the French border region of 'Grand Est' and Bergamo, the northern Italian city.

Figure 4.2: Road distances, tourist origins and Google searches



Note: The map on the left-hand side shows each counties' mean distance to Ischgl (own computation based on OpenStreetMap data). The map in the center depicts the share of each region in the total number of German tourists who visited Ischgl over the winter season of 2018-2019 (own computation based on data shared by Statistik Austria). Note that there are only seven regions provided in the raw data: Bavaria, Baden-Württemberg, Nordrhein-Westphalia, Berlin, Middle Germany (Saarland, Rhineland-Palatinate, Hesse), Eastern Germany (Brandenburg, Saxony-Anhalt, Thuringia, Saxony, Mecklenburg-Vorpommern) and Northern Germany (Lower Saxony, Bremen, Hamburg, Schleswig-Holstein). The map on the right-hand side shows the share of Google searches for the town of Ischgl from all 16 German states (own computation based on data from Google Trends).

We evaluate these claims by exploiting the exogenous variation in the road distances of German counties from these important clusters of infections — Ischgl, Heinsberg, Mulhouse and Bergamo. The road distance to Ischgl proxies well the probability of tourists visiting the Austrian ski and après-ski hotspot, as can be seen in Figure 4.2. The reported share of Ischgl tourists from origin regions in Germany in the previous year's skiing season (winter of 2018–2019), as well as the share of Google search queries in the same time period for the town of Ischgl from German states, are highly correlated with the road distance (-0.76 and -0.83, respectively). Whereas the distribution of Google searches for 'Ischgl' shows a gradual decline of search intensity with growing distance — underlining the town's Germany-wide tourist appeal. The same distributions for search queries for 'Mulhouse', 'Heinsberg' and 'Bergamo' are markedly different (see figure D.1 in appendix D.1): Whereas for the former ones, searches are highly concentrated in the near proximity of the respective towns, for the latter they are very evenly spread across all of Germany.

By estimating negative binomial regressions, we compute the elasticity of cases and mortality from COVID-19 with respect to distance from these initial European hotspots. The primary aim of our analysis is to explain the substantial spatial heterogeneity in the first wave of COVID-19 infections across German counties. By observing the spatial heterogeneity over the first wave, we indirectly evaluate the efficacy of early lockdown measures in halting the diffusion of the virus.

To guide our empirical analysis, we present a stylized two-period model where the mobility of persons drives infection transmissions. This simple model yields an insightful and testable proposition: the (absolute value of the) elasticity of COVID-19 cases with respect to distance from a super-spreader location is lower (higher) when individuals are more (less) mobile. We evaluate this proposition by examining the evolution of estimated distance elasticities over time. Finally, we demonstrate the significance of Ischgl as 'Ground Zero' for the outbreak in Germany during the first COVID-19 wave by performing a back-of-the-envelope counterfactual scenario with a hypothetical location for the town.

Crucially, all our regressions control for a host of possible confounding variables — including the relative latitude of a county. In robustness checks, we also examine how road distance to Ischgl affects case loads across counties that belong to the same latitude decile. Hence, our results do not simply capture the general effects of a county's distance to the South, e.g. Lombardy in Northern Italy, the European region hit hardest and earliest by the pandemic. We also control for testing by health authorities to account for the spatial pattern in the likelihood of detecting COVID-19 cases.

Our results paint a clear picture for the first wave of COVID-19: Cases increased strictly proportionally with population, but the share of the population infected was, amongst other factors, a function of the road distance to the major Austrian ski resort Ischgl. Were all German counties as far away from Ischgl as Vorpommern-Rügen, Germany would have 45% fewer COVID-19 cases. In contrast, distance to other hotspots was not important. Catholic culture appears to increase the number of cases — likely through Carnival celebrations in late February 2020.[8] We fail to find evidence for a host of socio-demographic determinants such as trade exposure to China, the share of foreigners, the age structure, or a work-from-home index. In line with expectations, fatality rates did not depend on distance to Ischgl. However, case fatality rates increased strongly with the share of population above 65 years and tended to fall in the number of available hospital beds. Finally, distance to Ischgl did not become irrelevant over time for observed cases, suggesting that early lockdown measures were effective in reducing mobility and avoiding further diffusion of the virus across German counties.

[8]Carnival is a typical Catholic tradition. The German South and South-West are predominantly Catholic, the North and North-East predominantly Protestant, but there is substantial variation within those regions as well.

Studying the diffusion of the virus across space is of utmost importance to guide the pandemic response which has so far largely been framed and implemented at national levels. Yet, with substantial heterogeneities in the number of infections — both in absolute and per capita numbers — a more fine-grained approach may be required that can take into consideration the specificity of the diffusion. Our analysis also highlights that international tourism is a powerful channel for the spread of contagious diseases. Timely travel bans can therefore limit transmission paths and control the cross-border spillover of infections. Popular destinations such as Ischgl have a critical role to play in such containment strategies since they can rapidly turn into super-spreader locations.

Declared a global pandemic by the WHO on March 11th 2020, the SARS-CoV-2 virus and its associated disease COVID-19 present an enormous challenge to the world economy. Outside of China where the virus was first detected, several European countries such as Italy, Spain and the UK have been hit particularly hard by the outbreak. Within Europe, Germany was treated as an exception during the first wave due to its low case fatality rates (4.41%) in comparison to Italy (13.93%), Spain (11.84%) and the UK (14.91%).[9]

The absence of proven treatments and vaccines during the first wave necessitated quarantine measures which curtailed human mobility and halted economic activity such as industrial production, retail sales and tourism. Although there is a great degree of uncertainty, the economic costs are expected to be high. For 2020, the International Monetary Fund expects global GDP to fall by 4.4%, more than in the world economic and financial crisis of 2009.[10] In the Eurozone, the Fund expects an output contraction of 8.3%. The World Trade Organization (WTO) projected the volume of international trade in goods to fall by 9.2% in 2020, with substantial downside risks as the resurgence of the virus requires new lockdowns.[11]

Since the outbreak of disease, economists have worked on several strands of research. The literature is moving fast; here we present only a few characteristic papers. Macroeconomists have introduced optimizing behavior by economic agents into the basic epidemiological SIER (Susceptible-Infected-Exposed-Recovered) model to examine the economic consequences of pandemics under different policy choices (Eichenbaum, Rebelo, and Trabandt, 2020; Farboodi, Jarosch, and Shimer, 2020; Krueger, Uhlig, and Xie, 2020). Behavioral economists have started to examine the long-run effects of this crisis on preferences (Kozlowski, Veldkamp, and Venkateswaran, 2020). Trade economists are studying the diffusion of health-related shocks through trade networks (Sforza and Steininger, 2020). Economic historians are investigating past pandemics to search for patterns that may inform current policy making (Barro, Ursua, and Weng,

[9] Figures as of May 9th, 2020
[10] World Economic Outlook, October 2020.
[11] See WTO press release dated October 6th, 2020 accessible via https://bit.ly/3i1yXlr.

2020), whereas econometricians are working to fill data gaps in order to properly calibrate macroeconomic models (Stock, 2020).

Our paper is most closely linked to the emerging literature on the geographical dispersion of the SARS-CoV-2 virus. J. Harris (2020) shows how the subway system was critical for the propagation of infections in New York City and identifies several distinct hotspot zip codes from where the virus subsequently spread. Jia et al. (2020) also examine the geographical distribution of COVID-19 cases by using detailed mobile phone geo-location data to compute population outflows from Wuhan to other prefectures in China. Pluemper and Neumayer (2020), in a related research endeavor also using German county-level data, find a positive association between the wealth of a district and a negative association with social deprivation in the initial phase of the pandemic up until mid-April 2020, which turned positive for the former and disappeared for the latter afterwards. Also related is Cuñat and Zymek (2020), who combine the SIR model with a structural gravity framework to simulate the spread of contagion in the UK. Our work contributes to the literature by (i) using exogenous variation in the distance to a super-spreader location to identify the role of tourism in the spatial diffusion of COVID-19 in the first wave and; (ii) providing a very simple test for the effectiveness of lockdown measures.

The remainder of this paper is structured as follows. Section 4.2 provides the relevant context to this analysis by describing the circumstances of the outbreak in Ischgl, Heinsberg, the French region of Grand Est and Bergamo. Section 4.3 outlines a simple theoretical model which underpins our empirical analysis. In Section 4.4, we describe our empirical strategy, the datasets used and the construction of key variables. Section 4.5 presents the main regression results followed by a counterfactual analysis in Section 4.6. Finally, Section 4.7 concludes.

4.2 Context

By mid-May 2020, there were around 16,000 confirmed cases of COVID-19 in Austria. The largest cluster of infections, comprising more than 20% of total cases, was located in the alpine province of Tyrol that is home to approximately 8% of Austria's population. The province's capital city, Innsbruck, was the first to report COVID-19 infections in the country, on February 25th, 2020. In Tyrol, the ski resort town of Ischgl is considered to be one of the epicentres, where the virus spread within après-ski bars, restaurants and shared accommodation.

A highly popular destination for international tourists, Ischgl was first flagged as a risk zone by Iceland on March 5th, 2020 after infection tracing revealed it as an important origin for COVID-19 cases. By March 8th, Norway's testing results also revealed that 491 of its 1198 cases had acquired the infection in Tyrol.[12] Despite these early warnings, skiing in Ischgl continued for nine more days. It was only on March 13th, 2020 that the town was placed under quarantine measures. On the same day, Germany's leading centre for epidemiological research, Robert Koch Institute (RKI), also designated Ischgl as a high-risk area — alongside Italy, Iran, Hubei Province in China, North Gyeongsang Province in South Korea, and the Grand Est region in France.

As the caseload of infections increased, Austrian authorities finally announced a lockdown in Tyrol on March 19th, 2020. This substantial delay in response is likely to have exacerbated the spread of the pandemic in Austria and other European countries, given the timing of the ski season and the location of the province which is bordered by Italy, Germany and Switzerland. As of March 20th, one-third of all cases in Denmark and one-sixth of those in Sweden were traced to Ischgl.[13] With data on mobile phone usage, a software company was also able to trace the movement of tourists after their visit to Ischgl. Returning skiers went to several destinations such as Munich, Frankfurt, Berlin, Cologne and Hamburg.[14]

In Germany, the states of Bavaria, Baden-Württemberg and North Rhine-Westphalia (NRW) reported the highest number of confirmed cases of the disease in the first wave. Together, they accounted for about two-thirds of Germany's total 175,000 COVID-19 cases as of mid-May 2020. Besides Ischgl, the district of Heinsberg in NRW has emerged as another important cluster that may have intensified the outbreak in Germany. The virus was reported to have spread there through Carnival celebrations, with an attendant testing positive on February 25th, 2020.

The northeastern French region of Grand Est was also heavily affected by the pandemic. Close to France's border with Germany, the spread of infections in the area was largely traced to a mass church gathering in Mulhouse.[15] Given the region's proximity to the hard-hit German state of Baden-Württemberg and the regular cross-border movement of German and French workers, we incorporate the town of Mulhouse in the Grand Est region into our analysis. We also include the distance to Bergamo in our regressions. The city is located in the densely inhabited region of Lombardy, which be-

[12] See "How an Austrian ski paradise became a COVID-19 hotspot", March 20th, 2020, Euractiv,(https://www.euractiv.com/section/coronavirus/news/ischgl-oesterreichisches-skiparadies-als-corona-hotspot/)

[13] See "Austrian Ski Region Global Hotspot for Epidemic", March 19th, 2020, Financial Times, (https://www.ft.com/content/e5130f06-6910-11ea-800d-da70cff6e4d3)

[14] See "So bahnte sich das Virus seinen Weg von Ischgl nach Deutschland", March 29th, 2020, Die Welt, (https://www.welt.de/wirtschaft/article206879663/Corona-Pandemie-So-hat-Ischgl-das-Virus-in-die-Welt-getragen.html)

[15] See e.g. "Special Report: Five days of worship that set a virus time bomb in France", March 30th, 2020, Reuters, (https://www.reuters.com/article/us-health-coronavirus-france-church-spec/special-report-five-days-of-worship-that-set-a-virus-time-bomb-in-france-idUSKBN21H0Q2)

came the epicenter of the COVID-19 outbreak in Italy. Therefore, these four locations — Ischgl, Heinsberg, Mulhouse and Bergamo — constitute interesting candidates as 'super-spreader locations' for studying the transmission of infections within Germany.

4.3 Theoretical Model

In this section, we sketch a stylized two-period model where the virus is transmitted through mobility of the population. This simplified setup is only used to guide our econometric strategy and does not form the basis of any structural estimations.

Let there be two rounds of infections. In the first round, people can be infected by visiting a super-spreader location such as Ischgl. Let P_i be the (time-invariant) population of county i and I_i^0 the number of infected individuals at the end of period 0. Let $f\left(D_i^0\right)$ denote the likelihood that an individual from county i has visited the super-spreader location in period 0 and has become infected, with f being a function of county i's distance to the super-spreader location. Let $f : [1, \infty) \to [0, 1]$ be a continuous and twice differentiable function with $f' < 0$ and $f'' < 0$.[16]

Hence,

$$
\begin{aligned}
& I_i^0 = P_i f\left(D_i^0\right) \\
\Leftrightarrow \quad & \iota_i^0 = I_i^0 / P_i = f\left(D_i^0\right),
\end{aligned}
$$

with $\iota_i^0 \in [0, 1]$ being the initial infection rate in county i.

In the second round, individuals randomly meet within Germany. If an infected person comes into contact with a susceptible person, the latter is also infected. Thus, in the absence of outside mobility between counties, new infections in period 1 would be given by

$$
\begin{aligned}
& \iota_i^1 - \iota_i^0 = \gamma \iota_i^0 (1 - \iota_i^0) \\
\Leftrightarrow \quad & I_i^1 = I_i^0 + \gamma P_i \iota_i^0 (1 - \iota_i^0),
\end{aligned}
$$

where $\gamma \in [0, 1]$ is the probability that an infection occurs when a susceptible individual meets an infected one.

However, individuals tend to move — within and across counties.[17] Let M_{ij} denote those individuals from county i that meet other individuals from county j, with

[16]The underlying intuition being that a person is less likely to visit Ischgl when the road distance is greater.

[17]For simplicity we assume there is no mobility outside of Germany.

$\sum_j M_{ij} = P_i$.[18] Assuming symmetry in mobility between counties, i.e. $M_{ij} = M_{ji}$, we have

$$I_i^1 = I_i^0 + \gamma M_{ii} \left(1 - \iota_i^0\right) \iota_i^0 + 2\gamma \sum_{j \neq i} M_{ji} \left(1 - \iota_i^0\right) \iota_j^0. \tag{4.1}$$

The elasticity of the infection rate with respect to the distance to the super-spreader location is given by $\delta^t \equiv \frac{\partial I_i^t}{\partial D_i^0} \frac{D_i^0}{I_i^t}$. Assuming $\frac{\partial M_{ii}}{\partial D_i^0} = 0$, it can be shown that

Proposition. *If any $M_{ij} > 0 \; \forall \; i, j$, then $\delta^0 < \delta^1$.*

Proof. See appendix D.2.

As both δ^0 and δ^1 are negative, we expect the elasticity of infections with respect to distance from the super-spreader location to be greater (i.e. closer to zero) with mobility than without mobility. When there is no inter-county geographical mobility after period 0, then $M_{ij} = 0$ for all $j \neq i$, the elasticity is larger in absolute terms than when mobility is allowed; when even intra-county mobility is not permitted, then the elasticity is time-invariant: $\delta^1 = \delta^0$ as $I_i^1 = I_i^0$.

The intuition for this result is simple: as mobility between and within counties spreads the virus further over time, the role of distance to Ischgl in explaining the spatial variation of infections decreases. We assume mobility between counties i and j, M_{ij}, to be exogenous to i's and j's distance to Ischgl, believing this to be a rather innocuous assumption.[19]

4.4 Empirical Model and Data

4.4.1 Model and Hypotheses

As reflected in our stylized model, we are interested in understanding the number of COVID-19 patients (I_i^0 and I_i^1) and fatalities registered in a county. For this reason, the appropriate econometric strategy is to estimate a count data model, such as a Poisson or negative binomial model. In this context, we expect the variation of our dependent variable to exceed that of a true Poisson since (i) counts will not be independent in a pandemic; and; (ii) there may be unobserved heterogeneity. Therefore, we employ a negative binomial model in which the variance is assumed to be a function of the

[18] Note that one could assume gravity-type micro-foundations with frictions to interactions between i and j, e.g. à la Anderson (2011).

[19] Essentially, the spatial distribution of counties, mobility costs between them, and their population sizes are assumed to be independent of the counties' distance to Ischgl.

mean (NB-2 model; see Cameron and Trivedi, 2013). Since the NB-2 model nests the simple Poisson model, one can test for over-dispersion.[20] One handy feature of the negative binomial model is that its coefficients can be interpreted exactly as in a linear model in which the dependent variable is logarithmic.

We exploit the variation in cases and deaths across the numerous counties as of May 9th, 2020, and estimate elasticities with respect to road distances from Ischgl, Heinsberg, Mulhouse and Bergamo. We run cross-sectional regressions which are specified as follows:

$$\text{cases}_i = exp(\alpha + \sum_{k\in\{0,1,2,3\}} \delta_k \log(D_i^k) + \gamma \mathbf{Z}_i + \varepsilon_i) \tag{4.2}$$

$$\text{deaths}_i = exp(\alpha + \rho \log(\text{lagged cases}_i) + \sum_{k\in\{0,1,2,3\}} \delta_k \log(D_i^k) + \gamma \mathbf{Z}_i + \varepsilon_i) \tag{4.3}$$

The main coefficients of interest in the above regressions are the set of δ coefficients which capture the elasticity of COVID-19 cases or deaths with respect to the road distance of any given county i from Ischgl, Heinsberg, Mulhouse or Bergamo. In equation (4.2), these distance elasticities enable us to test our first hypothesis — namely, that COVID-19 cases decay as distance from an infection cluster increases. Therefore, as per our theoretical model, we expect the δ coefficients to be negative.

Equation (4.3) takes the number of COVID-19 deaths as the dependent variable and introduces log cases lagged by 18 days as an additional explanatory variable. We control for cases with a lag since the mean time between the onset of symptoms and death is estimated at 17.8 days (Verity et al., 2020). This allows us to test our second hypothesis — that distance to super-spreader locations should not matter for the number of deaths in a county, controlling for the number of infections in a county. Proximity to any of the hotspots may have affected the incidence rate but should not determine the medical severity of cases and therefore the fatalities.

Our third and final hypothesis is that the distance of a county from these towns is more crucial for spreading infections in the initial phase of the epidemic — in the absence of restrictions on the movement of people. With time, COVID-19 expands its reach to more locations and the role of these initial clusters may become less relevant. A test of this hypothesis can be conducted by introducing time variation in the number of cases and deaths at the county-level. By repeatedly estimating equation (4.2) for each day within this period, we obtain a time series of coefficients for the distance vari-

[20] As frequently observed with negative binomial models, as in our exercise, it does not matter substantially whether dispersion is assumed constant across observations (NB-1 model) or is a function of the expected mean (NB-2 model).

ables. These time series can then be examined graphically in order to determine when and for how long distance to initial infection clusters mattered in the propagation of COVID-19.

Clearly, distance to Ischgl correlates with other potential determinants of infections. Hence, while we trivially have no issues with reverse causality, our exercise is potentially subject to substantial omitted variable bias. In our exercise, we have no other way to deal with this problem than to load the vector $\mathbf{Z_i}$ with a rich and well-design array of control variables. The most important is geographical latitude, relative to the southernmost point of Germany. This rules out that the coefficient δ_0 simply captures proximity to Italy. The control also captures climatic variation, as well as other factors, e.g. cultural practices, that tend to have a north-south gradient and may influence infection rates. Moreover, we add further county-specific characteristics such as population and population density, GDP per capita, share of population that is older than 65 years, shares of Protestants and Catholics, share of foreigners, a work-from-home index that captures the prevalence of home office work, exposure to trade with China and the number of hospital beds in a county. All these controls may exhibit non-zero correlation with distance to Ischgl. For example, the share of Catholics is much higher in the South than in the North and Catholic festivities, e.g. Carnival, may propagate infections.

The $cases_i$ variable in equations (4.2) and (4.3) refer to diagnosed cases rather than to a full count of the infected population, or a random draw. There could be many more undetected cases in the German population than diagnosed ones. For instance, Li et al. (2020) find that in early stages of pandemics, the number of infected people was six higher than official statistics revealed. To deal with this issue, we control for the number of tests per county. Interestingly, there is substantial variation across counties in the share of population tested.

Despite these efforts to contain omitted variable bias, we adopt a cautious reading of our results and refrain from interpreting them as causal. Nonetheless, our evidence on the spatial determinants of the COVID-19 spread in Germany reveals interesting correlations and strong indications of a link between the COVID-19 burden of a county and its distance from a super-spreader location.

Finally, in Appendix D.3 we analyze the sensitivity of our results to the choice of distance measures by switching from road distance to travel time and great circle distances. We also introduce fixed effects for deciles of counties' latitude and great circle distance to Ischgl to study the East-West spread of the virus, as well as ensuring the analysis does not simply pick up North-South variations. Note that by subtracting the log of population from both sides of equation (4.2) and the log of lagged number of cases from both sides of equation (4.3), one can interpret the estimated coefficients as

Table 4.1: Summary statistics

Variable	Mean	Std. dev.	Median	Max	Min
Number of confirmed cases, current	423.03	591.92	278.00	6261.00	13.00
Number of confirmed cases, 18 day lag	364.91	515.48	240.00	5295.00	13.00
Case incidence rate (CIR), in %	0.21	0.16	0.17	1.52	0.03
Deaths	21.66	30.10	12.00	261.00	0.00
Case fatality rate (CFR), in % &	5.63	3.92	5.06	24.05	0.00
Population (in thousands)	201.23	231.06	149.07	3421.83	34.08
Number of tests (in thousands)	257.90	176.51	208.50	500.48	24.86
Road distance to Ischgl (in km)	609.75	237.28	610.95	1134.26	138.69
Road distance to Heinsberg (in km)	428.40	184.31	433.05	805.37	0.27
Road distance to Mulhouse (in km)	521.01	211.44	507.22	1069.63	56.81
Road distance to Bergamo (in km)	782.55	228.83	768.56	1319.22	326.32
Population / Area	517.74	676.47	195.78	4531.17	36.47
Log of relative latitude	3.35	1.75	3.32	7.51	0.22
GDP per capita (in thousand Euros)	37.16	16.14	33.11	172.44	16.40
Share of foreigners	0.07	0.05	0.06	0.31	0.01
Share of 65+	0.21	0.02	0.21	0.29	0.15
Share of Catholics	0.32	0.24	0.29	0.88	0.02
Share of Protestants	0.30	0.17	0.26	0.72	0.04
Work-from-Home Index	0.53	0.04	0.52	0.67	0.46
Trade with China measure	6338.35	4079.31	5321.70	30228.97	470.53
Number of hospital beds	1255.41	1598.54	851.50	20390.00	42.00

Note: Epidemological data refer to May 9, 2020; other data to year of 2019 or latest available year. Case fatality rate calculated on the basis of reported cases 18 days earlier.

elasticities (or semi-elasticities) of case incidence rates (CIRs) or of case fatality rates (CFRs), respectively. Therefore, as an additional robustness check, we estimate models with CIR and CFR as dependent variables using OLS.

4.4.2 Data

We use publicly available data on COVID-19 cases in Germany provided by the Robert Koch Institute (RKI). The RKI database reports confirmed cases as well as fatalities from COVID-19, although it should be noted that these numbers may under-represent the actual spread of infections due to limitations in testing. A valuable feature of the RKI dataset for our purposes is its level of geographic disaggregation. Information is available not just at the country-wide or Bundesländer (state) level, but at the county-level in Germany. The data spans from March 10th to May 9th, 2020 and relates to the first wave of COVID-19 infections experienced by Germany.[21] In this paper, we work with cumulative confirmed cases and COVID-19 related deaths as of May 9th, 2020.

[21]The RKI data have been criticized for inaccurate timing of reported cases. For example, there are differences between weekends and weekdays. Since we do not exploit daily variation in cases, our estimations should be largely free from this problem. Besides these issues, German administrative data are generally perceived as being of high quality.

We merge this database with information on the county-level from the Regionaldatenbank Deutschland. We include data on the local population,[22] which allows us to control for the demographic structure of each county, given the higher risk of hospitalization and fatalities from COVID-19 amongst older populations. We also control for another population characteristic, namely religious affiliation, that may indicate whether Carnival gatherings — largely a Catholic festival — may have contributed to the spread.

To control for the levels of economic activity, we utilize GDP per capita at the county-level for the latest available year, 2018. We further include a variable that describes the regional intensity of jobs that can be performed from home, the 'work-from-home' index at the county-level computed by Alipour, Falck, and Schüller (2020). The ability to work from home, and thus avoid public spaces and offices, may have played an important role in determining the local spread of the virus (Fadinger and Schymik, 2020). As another possible channel for the transmission of the virus within Germany, we incorporate the exposure of counties to international trade with China, where the outbreak was first reported. The trade (export and import) exposure measures are taken from Dauth et al. (2017).[23]

The number of confirmed cases may also be dependent on the testing capacity and healthcare infrastructure of the county. However, there is no reliable data available as of the time of writing on the number of tests conducted daily in each county. Given this limitation, we use the number of tests performed in each of the 16 German Bundesländer. This information is provided by the RKI. For healthcare capacity, which may impact the prevalence of testing and the possibility of adequate treatment, we use the number of hospital beds in each county as an indicator. This is again drawn from Regionaldatenbank Deutschland database for the year 2018.

In order to examine the impact the three hotspots had on the spread of the virus, we exploit each of the county administrative centers' distance to the towns of Ischgl, Heinsberg and Mulhouse. We compute road distance and travel times based on the shortest path in road networks with data from the OpenStreetMap project. In a robustness exercise we additionally use the great circle distance between the respective locations; see Table D.1 in Appendix D.3.

[22] As of December 31st, 2017.

[23] The measure is constructed from national sector-level import and export data and regional sector-level employment shares.

4.5 Regression Results

In this section, we analyze regression results based on specifications described in equations (4.2) and (4.3) and assess the evolution of estimated coefficients such as distance elasticities over time.

4.5.1 COVID-19 Cases

Table 4.2 reports results for confirmed COVID-19 cases, where we introduce a richer set of controls with each successive regression. Starting with column (1), we find that the coefficient on population is statistically identical to 1, implying that cases rose proportionately with population size.[24] Counties with bigger populations did not have higher case rates (infections per number of inhabitants). This finding is robust across all our specifications.

The coefficient for the number of tests is positive and statistically significant — i.e. counties located in states that conducted more tests reported more confirmed cases. The estimated coefficient is large; it suggests that an increase in the number of tests by 1% correlates with an increase in the number of cases by 0.441%. This implies that increasing the number of tests by 10% reveals about 12 more cases of infected persons in the median county.[25] This emphasizes the vital importance of testing in understanding the spread of infections and its role in the policy response. In all columns of Table 1, we also report the θ parameter which indicates the extent of over-dispersion in the data. If the θ parameter were to approach infinity, the negative binomial distribution would approach a Poisson distribution. However, the parameter is seen to be finite across specifications. Hence, our choice of negative binomial regressions over Poisson estimation is indeed valid. Not surprisingly, infection data exhibits over-dispersion.

In column (2), we introduce road distance to Ischgl as an additional explanatory variable. In doing so, we find that the pseudo-R^2 increases by 6 percentage points or 9%, indicating the relevance of this variable for the overall fit of the model. The resulting coefficient implies that a county whose road distance to Ischgl is by 1% lower than that of another county has a count of infections that is higher by 0.68%. However, Ischgl may not be the only cluster from where the virus may have spread through Germany. To examine this possibility, column (3) introduces the road distances to other clusters — Heinsberg, Mulhouse and Bergamo — as controls. By additionally controlling for the latitude of each county, we exploit precisely the variation in road distance and not the geographical location of a county on the North-South axis. As such, latitude has no

[24] For this reason, we can interpret the coefficients in our regressions as also measuring the effect on the log CIR of counties.

[25] $(1.1^{0.441} - 1) \times 279 = 11.98$; adding further covariates reduces the importance of tests by about half.

Table 4.2: Count of Confirmed Cases: Negative Binomial Regressions

	Dependent variable:				
	Number of confirmed cases				
	(1)	(2)	(3)	(4)	(5)
log(Population)	0.991^{***}	1.073^{***}	1.053^{***}	1.070^{***}	1.070^{***}
	(0.048)	(0.053)	(0.055)	(0.053)	(0.054)
log(Number of tests)	0.441^{***}	0.286^{***}	0.258^{***}	0.190^{***}	0.188^{***}
	(0.039)	(0.044)	(0.040)	(0.043)	(0.043)
log(Distance to Ischgl)		-0.679^{***}	-0.840^{*}	-0.787^{*}	-0.795^{*}
		(0.075)	(0.460)	(0.459)	(0.458)
log(Distance to Heinsberg)			-0.138^{***}	-0.064	-0.077
			(0.050)	(0.098)	(0.096)
log(Distance to Mulhouse)			-0.053	-0.019	-0.032
			(0.201)	(0.212)	(0.211)
log(Distance to Bergamo)			-0.256	-0.294	-0.250
			(0.757)	(0.713)	(0.722)
log(Latitude)			0.174	0.243	0.236
			(0.169)	(0.204)	(0.216)
log(Population / Area)				0.038	0.001
				(0.045)	(0.047)
Share of Catholics				0.712^{**}	0.734^{**}
				(0.297)	(0.295)
Share of Protestants				0.166	0.187
				(0.259)	(0.252)
Share of 65+				-1.183	-0.739
				(2.343)	(2.222)
Share of Foreigners				-0.547	-0.810
				(1.116)	(1.142)
log(GDP p.c.)					0.069
					(0.122)
Work-from-Home Index					1.078
					(1.216)
log(China Trade)					-0.006
					(0.070)
Pseudo R2	0.66	0.73	0.74	0.76	0.76
Observations	401	401	401	401	401
θ	3.202^{***}	3.925^{***}	(0.287)	(0.306)	(0.308)
	(0.219)	(0.273)	(0.287)	(0.306)	(0.308)

Note: Constant not reported. Robust standard errors: *p<0.1; **p<0.05; ***p<0.01.

measurable effect on the case load.[26] The coefficient on the distance to Ischgl remains statistically significant. Proximity to Ischgl also appears to be more important than proximity to the other hotspots. For the purpose of illustration, compare the city of Munich, which is about 190 km away from Ischgl, to Hamburg, 935 km away. Everything else equal, Hamburg should have 73.7% fewer COVID-19 cases than Munich.[27] The high elasticity implies a fast decay of infections as one moves away from Ischgl.

In column (4), we control for a wide range of county-level variables that could also predict infections. Notably, the distance elasticity for Heinsberg is no longer statistically significant whereas the distance elasticity to Ischgl is stable. Examining the demographic characteristics, factors such as population density, share of the elderly (65 years and older) and foreign residents in total population are not significant determinants of the spread. In contrast, a 1% point increase in the share of Catholics is associated with a 0.712% increase in cases — probably attesting to the role of carnival celebrations in February 2020, which are typical for Catholic regions but not for Protestant ones, in propagating the virus. To illustrate the importance of this correlation: increasing the share of Catholics in the county with the smallest share (0.02, county of Weimar in Thuringia) to the share observed in the most Catholic county (share of 0.88, county of Freyung-Grafenau in Bavaria) almost doubles the case count.[28]

Our baseline specification additionally controls for economic factors and is reported in column (5). In comparison to the minimalist specification reported in column (2), controlling for demographic and economic factors increases rather than decreases the distance elasticity to Ischgl; adding additional socio-economic controls keeps it approximately constant. A 1% reduction in road distance to Ischgl corresponds to a 0.8% increase in the number of confirmed cases.[29]

Looking at the coefficient on a county's trade exposure to China, where the virus first appeared, we observe that the transmission of the virus in Germany was not driven by the strength of economic ties to China. Our results therefore undermine possible claims that the participation of local firms in global production chains involving China

[26] If latitude is included in specification (1), its coefficient is observed to be negative (coefficient of -0.45) and highly statistically significant; if distance to Ischgl is added (without the distances to other super-spreader locations, the coefficient on latitude remains negative but turns statistically insignificant while distance to Ischgl appears significant (with a coefficient of -0.40).

[27] $100\% \times [(935/190)^{-0.84} - 1] = -73.7\%$. In fact, Hamburg had about 20% fewer infections in May 2020, but 24% more population in comparison to Munich.

[28] $[\exp(0.712 \times 0.86) - 1] \times 100\% = 84.4\%$. The case incidence rate is 0.10% in Weimar and 0.24% in Freyung-Grafenau.

[29] Table D.1 column (2) in the Appendix uses travel time instead of road distance as a measure of distance; the pseudo-R^2 goes down slightly, but our main results remain intact. Importantly, when distance variables are constructed using great circle distances, distance to Ischgl is no longer statistically significant, but log latitude changes sign and becomes large in absolute value. This is not surprising, as latitude almost perfectly predicts geodesic distance to Ischgl (coefficient of correlation $\rho = 0.989$); latitude highly correlates to travel time, too, but the ρ is somewhat lower at 0.972. We view this as supportive of our identification strategy which relies on road distance conditional on latitude.

may have led to the import of the virus and therefore propagated contagion.[30] We also find that the 'Work-from-Home' (WFH) Index is not a significant factor in the diffusion process. This runs counter to the results reported by Fadinger and Schymik (2020) – possibly because we control for WFH at the more disaggregated county (NUTS-3) level as opposed to the NUTS-2 level. Rather, infections are seen to be dependent on population size and the proximity to local hotspots. All together, the models have relatively high values for pseudo-R^2, which offers a rough measure of the variation in infection rates that our models are able to explain.

For the sake of checking robustness, Table D.2 in the Appendix reports regressions analogous to those in Table 4.2, but with the dependent variable being the case incidence rate and the estimation method being OLS. This regression design is more restrictive than our preferred one, but we generally find that our findings are confirmed. In our most comprehensive regression, Hamburg is predicted to have a CIR that is 0.23 percentage points lower than Munich's (in the data, Hamburg's CIR is 0.27% and Munich's 0.46%).

4.5.2 COVID-19 Fatalities

Having examined confirmed infection cases, in Table 4.3 we address the observed spatial heterogeneity in COVID-19 deaths across counties. All regressions contain the log of confirmed infections 18 days prior as a major predictor of the death count.[31] As in Table 4.2, regressions also include the log of the number of tests conducted in a county and the log of population.

In all specifications, the coefficient on log lagged cases is observed to be statistically significant and greater than 1, implying that deaths are increasing more than proportionately to the number of reported cases in a county. An underlying issue of congestion in healthcare facilities may explain this relationship. Importantly this relation is not driven by population: across all specifications, we find that more populous counties tend to have lower number of fatalities, holding the case load constant. But note that the two variables are strongly correlated, as the previous section has shown. The number of tests has no measurable influence on death counts.

Without adding the controls introduced in column (1), distance to Ischgl has a large, negative effect on the dependent variable; however, this would be a meaningless result

[30]While the first patient in Germany contracted COVID-19 from a visiting colleague from China in January 2020, this case was well identified. The systematic spread of the virus however, began in Germany much later from March 2020 onwards and is not observed to be linked to exposure to China. See https://www.spiegel.de/wissenschaft/medizin/corona-virus-erster-fall-in-deutschland-bestaetigt-a-19843b8d-8694-451f-baf7-0189d3356f99.

[31]As noted above, Verity et al., 2020 find that the average time between a confirmed infection and a death is approximately 18 days.

Table 4.3: Count of Deaths: Negative Binomial Regressions

	Dependent variable:				
	Number of deaths				
	(1)	(2)	(3)	(4)	(5)
log(Lagged number of confirmed cases)	1.301^{***}	1.394^{***}	1.395^{***}	1.443^{***}	1.441^{***}
	(0.062)	(0.064)	(0.064)	(0.065)	(0.065)
log(Population)	-0.372^{***}	-0.487^{***}	-0.487^{***}	-0.490^{***}	-0.387^{***}
	(0.074)	(0.078)	(0.079)	(0.088)	(0.097)
log(Number of tests)	−0.052	−0.003	−0.003	0.038	0.016
	(0.043)	(0.047)	(0.047)	(0.050)	(0.050)
log(Distance to Ischgl)		0.549	0.548	0.448	0.523
		(0.350)	(0.353)	(0.351)	(0.354)
log(Distance to Heinsberg)		0.091^{**}	0.091^{**}	0.067	0.064
		(0.037)	(0.037)	(0.042)	(0.043)
log(Distance to Mulhouse)		0.005	0.004	−0.043	−0.049
		(0.175)	(0.176)	(0.173)	(0.176)
log(Distance to Bergamo)		−0.149	−0.151	−0.187	−0.018
		(0.660)	(0.661)	(0.682)	(0.683)
log(Latitude)		−0.086	−0.085	0.040	−0.073
		(0.187)	(0.190)	(0.194)	(0.195)
log(GDP p.c.)			−0.008	0.071	0.142
			(0.117)	(0.142)	(0.139)
log(Population / Area)				-0.118^{**}	−0.076
				(0.050)	(0.051)
Share catholics				−0.098	−0.068
				(0.273)	(0.273)
Share protestants				0.009	0.021
				(0.315)	(0.316)
Share population 65+				6.705^{***}	7.248^{***}
				(1.637)	(1.642)
Share foreigners				2.711	2.560
				(1.666)	(1.695)
log(Number of hospital beds)					-0.144^{***}
					(0.049)
Pseudo R2	0.76	0.77	0.77	0.78	0.78
Observations	401	401	401	401	396
θ	4.098^{***}	4.341^{***}	4.341^{***}	4.717^{***}	4.829^{***}
	(0.419)	(0.454)	(0.454)	(0.503)	(0.520)

Note: Constant not reported. Robust standard errors: *p<0.1; **p<0.05; ***p<0.01.

as it only reflects the geography of case counts. Once we control for confirmed cases, distances to the super-spreader locations cease to have a negative effect; if at all, there is a positive effect which is, however, only marginally statistically significant. This is reasonable since health outcomes are likely to depend more on the individual case or county's demographic and economic characteristics than on the distance to a ski resort.[32]

However, mortality rises sharply with the share of the elderly in county populations (see columns (4) and (5)), conforming with medical findings that case fatality ratios are higher for older age groups (Verity et al. (2020)). For the purpose of illustration, comparing the county with the smallest share of elderly (0.15, county of Vechta, Lower Saxony) to the county with the greatest (0.29, county of Dessau-Roßlau in Saxony-Anhalt), model (5) predicts more than a doubling of the death count.[33]

Variables such as the share of Catholics that had an important effect in Table 4.2, are no longer significant. This indicates that the capacity of the health system does not depend on a county's predominant religious group. Also, the share of foreigners is not significant. In contrast, healthcare infrastructure, as proxied by the number of hospital beds, turns out as a statistically significant predictor of COVID-19 morbidity. A 10% increase in number of beds in a county lowers deaths by approximately 1.36%.[34] Thus, access to quality medical care is imperative for minimizing the loss of human life due to the pandemic.

While this finding warrants further investigation, we would like to stress that the number of beds is predetermined in our specification, so we do not face the issue of reverse causality. Moreover, the effect is estimated conditional on a number of variables that explain both fatalities and the number of hospital beds, such as density (population per area, capturing the urban/rural divide) or GDP per capita. Also note that mobility from counties with few beds to others with more beds would attenuate the effect; hence, we are likely to identify a lower bound of the true effect.

For robustness, Table D.3 in the Appendix reports regressions analogous to those in Table 4.3, but with CFR as the dependent variable and OLS as the estimation method. Results are broadly robust. For example, increasing the number of beds by 10% in a county, lowers the case fatality rate by 0.092% points;[35] the median CFR being 4.18%.

[32]Note that, in the robustness checks (CFR model) presented in Table D.3 distance to Ischgl is never statistically significant.

[33]$\exp[6.705 \times (0.29 - 0.15)] - 1$.

[34]$1.1^{-0.144} - 1$.

[35]$-0.963 \times \ln(1.1)$.

4.5.3 Super-spreader Effects Over Time

So far, we have focused on examining a cross-section of the RKI database by running regressions on a snapshot of COVID-19's impact across counties as of May 9th, 2020. Now, we move towards analysing the time dimension as well. Our question here is: Did the role of super-spreader locations like Ischgl diminish over time during the first wave of infections? This is addressed clearly by Figure 4.3. It depicts the evolution of the 'daily distance elasticities' that are computed by repeatedly estimating our baseline specification for confirmed cases. To the extent that tourists returning from Ischgl explain an initial distribution of infections but subsequent mobility spreads the virus further, one would expect the measured elasticity to decline in absolute value. This corresponds to the proposition of our theoretical model. If the lockdown (phased in from March 9th to March 23rd 2020) has been effective in restricting mobility, our model predicts that distance elasticity will remain highly negative as initial exposure continues to be important.

Strikingly, we observe distinct phases in the behaviour of the Ischgl elasticity that broadly corresponds with the timeline of Germany's lockdown. Over the initial period, this elasticity reduces in absolute value as individuals continue to be mobile. Once mobility is severely restricted with the imposition of a lockdown, it remains stable, significantly different from zero and strongly negative. Thus, distance from Ischgl is a relevant predictor of cases not just over varying specifications as pointed out in Table 4.2, but also over time.[36]

The same exercise is carried out for other control variables that were observed to be significant in Table 4.2 to construct Figure 4.4. It shows that the positive relationship between cases and testing capacity is consistent and statistically significant over time. In the case of population size, results are in alignment with the cross-sectional regression as elasticities remain close to 1. How well does the baseline model explain the variation in cases across counties? As shown in Figure 4.5, the pseudo-R^2 is high and improves substantially with time up to 0.80 on March 25th 2020, and has only slightly fallen from there. Again, if the infection had spread geographically after the containment measures, we would expect a sizeable decline in R^2 from our model; however, we do not observe this pattern. We conclude that restrictions in mobility after March 23rd 2020 helped contain the virus imported from Ischgl in those counties where it first arrived.

[36] The Mulhouse elasticity is never statistically significant whereas the Heinsberg elasticity becomes statistically insignificant by March 28th 2020. The elasticity of the distance to Bergamo is initially positive but not statistically significant, with large standard errors.

Figure 4.3: Distance coefficents

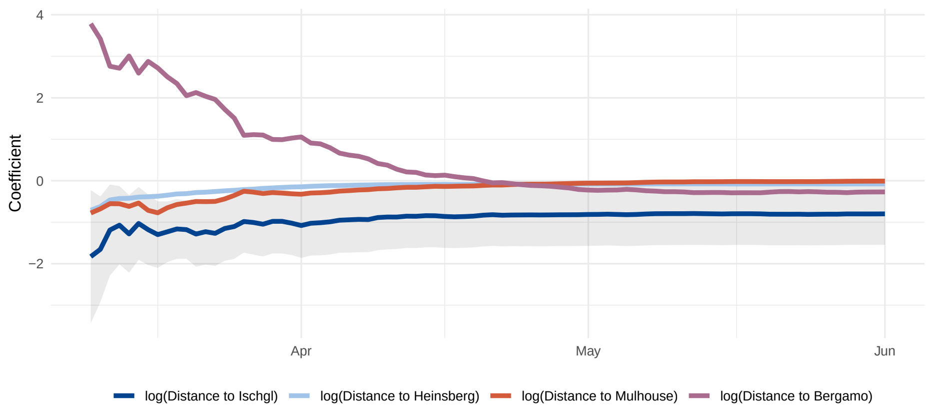


Figure 4.4: Control variable coefficients

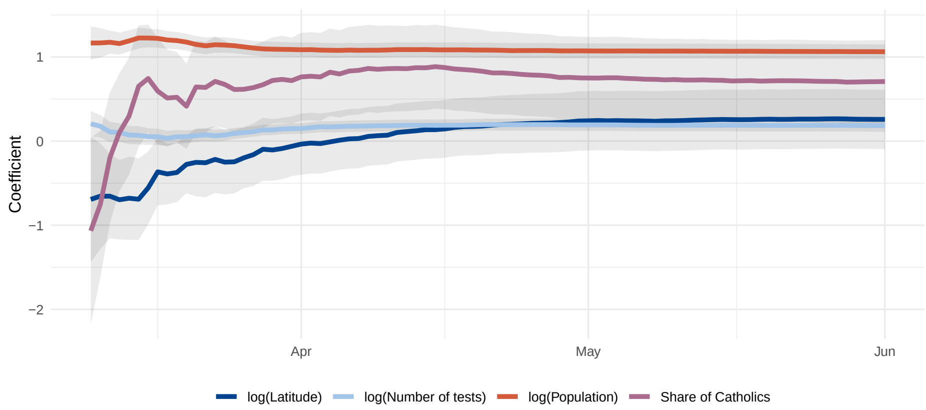
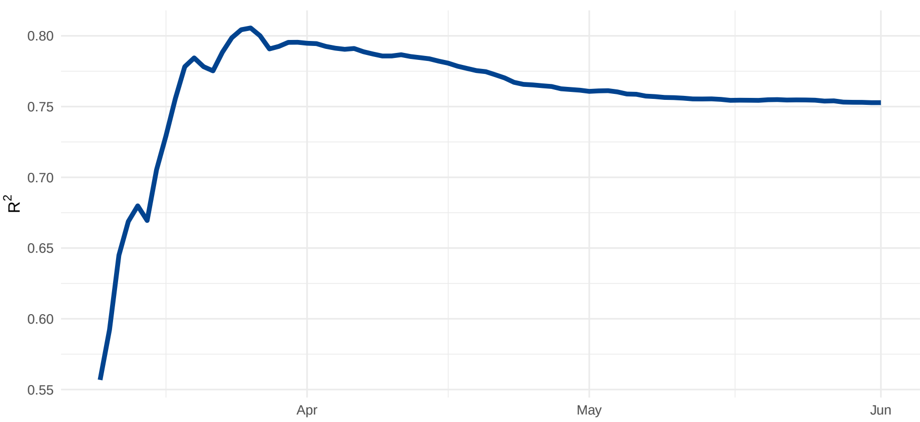


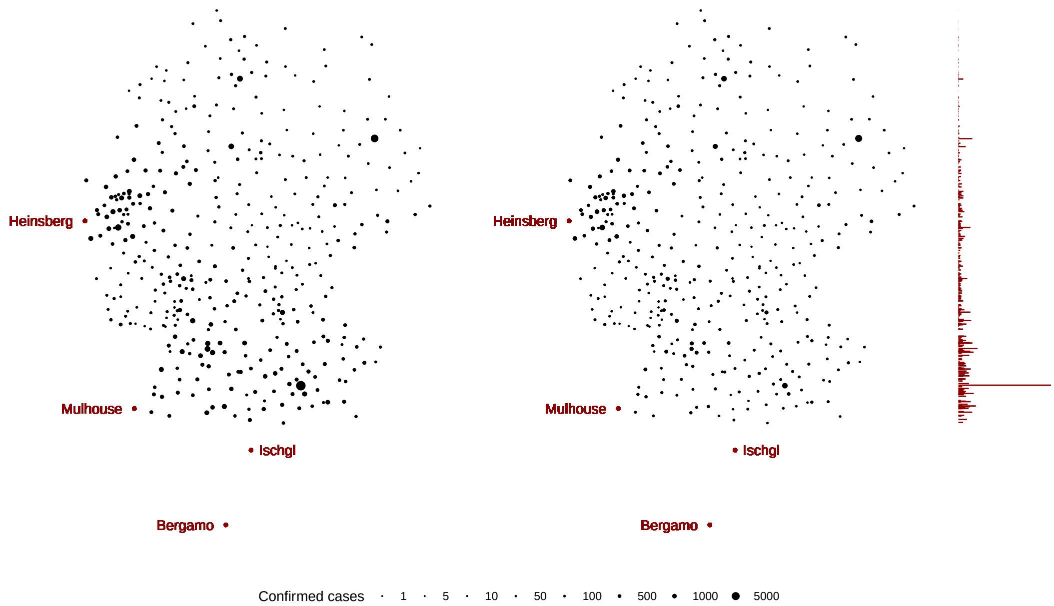
Figure 4.5: Pseudo-R^2 by date of data



4.6 Counterfactual Scenario

The previous section described how the prevalence of infections in Germany during the first COVID-19 wave was related to counties' geographic proximity to a super-spreader location such as Ischgl. To further gauge the impact of proximity to Ischgl on case counts, we now perform a simple back-of-the-envelope counterfactual exercise. We predict the number of confirmed cases were Ischgl located 1,134 km away from all counties, the distance at which the Kreis Vorpommern-Rügen, the northeastern-most county, is actually located from Ischgl. This assumes a situation in which no county is located close to the resort town, and hence simulates a situation in which fewer German tourists may have returned from their ski trip infected with the virus.

Figure 4.6: Predicted confirmed cases on May 9th vs. back-of-the-envelope counterfactual



Note: Map on the left-hand side shows predicted confirmed cases on May 9th, map on the right-hand side shows those predicted in counterfactual scenario for May 9th. The histogram to the very right shows the difference by latitude, binned by county.

Using the baseline negative binomial regression, we compare the predicted number of confirmed cases against the number of cases with the new, hypothetical location of Ischgl. The experiment leads to the total number of cases in Germany (as of May 9th 2020) dropping from the predicted level of 172,275 to the counterfactual level of 94,304 i.e. a 45 % reduction. This back-of-the-envelope calculation validates our prior findings and offers a compelling demonstration of the spatial aspects of the virus transmission. Figure 4.6 below presents maps for the predicted and counterfactual scenarios, with a histogram that captures the differences in number of cases by latitude. The south, in reality located relatively closely to Ischgl, would have seen far fewer cases.

4.7 Conclusion

This paper studies the geographical distribution of COVID-19 cases and fatalities across the 401 German counties. It tests the hypothesis that returning visitors from super-spreader locations like Ischgl, a popular ski resort in Tyrol, Austria, have played a major role in spreading the disease. Indeed, distance to Ischgl turns out to be an important predictor for case incidence rates, but not for case fatality rates. Were all German counties situated as far from Ischgl as the most distant county of Vorpommern-Rügen, Germany would have 45% fewer COVID-19 cases. Distance to Ischgl does not become irrelevant over time, suggesting that lockdown measures have avoided further diffusion of the virus across German counties. In contrast, distances to other hotspots are unimportant.

Catholic culture, likely capturing local Carnival festivities in late February 2020, appears to increase the number of cases while other socio-demographic determinants such as trade exposure to China, the share of foreigners, the age structure, GDP per capita, or a work-from-home index do not add any explanatory power. Case fatality rates increase strongly in the share of population above 65 years and fall in the number of available hospital beds.

We view our results as evidence towards confirming the role of super-spreader locations for the diffusion of a pandemic. Additionally, we find evidence for the efficacy of the lockdown measures put in place in reducing the spread of the virus. Further improvements of the analysis will be possible as more data become available, for example on testing strategies at the county-level.

Appendix A

Appendix to Chapter 1

A.1 String Distance Measures

A.1.1 Levenshtein distance

Consider a hypothetical firm exporting five different CN8 products to China, two of which are invoiced in RMB. In this scenario, each element of the firm's China product vector (CPV) takes the value of one, whereas elements in the RMB product vector (RPV) take the value of one or zero depending on whether the corresponding product is invoiced in RMB or not. Then the (normalized) Levenshtein distance between these two vectors is 3/4. The numerator reflects the number of changes that are required to transform the RPV into the CPV and the denominator reflects the normalization (maximum length of product vectors minus one).

$$CPV_{ft} = \begin{pmatrix} 1 \\ 1 \\ 1 \\ 1 \\ 1 \end{pmatrix} \quad RPV_{ft} = \begin{pmatrix} 1 \\ 0 \\ 1 \\ 0 \\ 0 \end{pmatrix}$$

A.1.2 Bray-Curtis similarity index

Consider the same hypothetical firm exporting five different CN8 products to China, two of which are invoiced in RMB. In contrast to before, the CPV and RPV are now defined based on product export shares, as shown below.

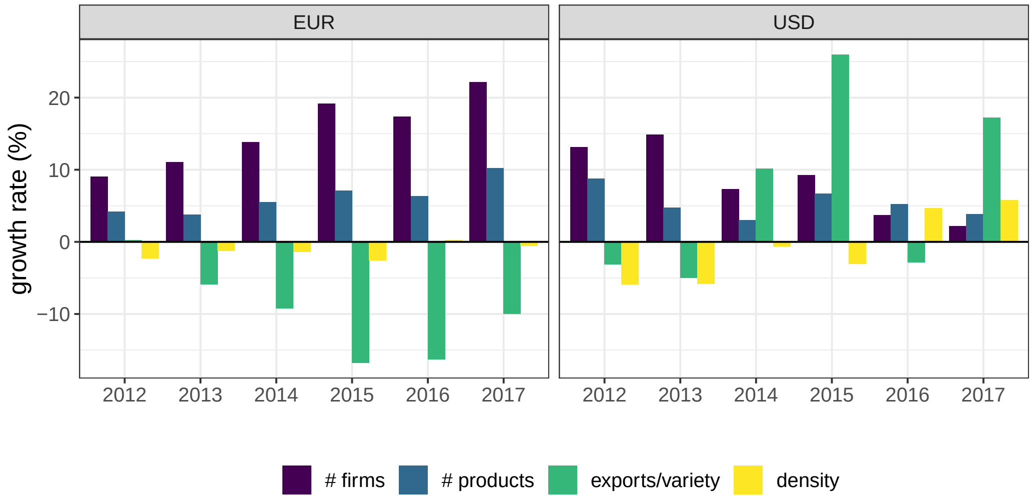
$$CPV_{ft} = \begin{pmatrix} 0.1 \\ 0.5 \\ 0.2 \\ 0.1 \\ 0.1 \end{pmatrix} \quad RPV_{ft} = \begin{pmatrix} 0.7 \\ 0 \\ 0.3 \\ 0 \\ 0 \end{pmatrix}$$

Then the corresponding Bray-Curtis similarity index is calculated as follows:

$$BC = 1 - \frac{|0.1-0.7|+|0.5-0|+|0.2-0.3|+|0.1-0|+|0.1-0|}{(0.1+0.5+0.2+0.1+0.1)+(0.7+0+0.3+0+0)} = 1 - 0.7 = 0.3$$

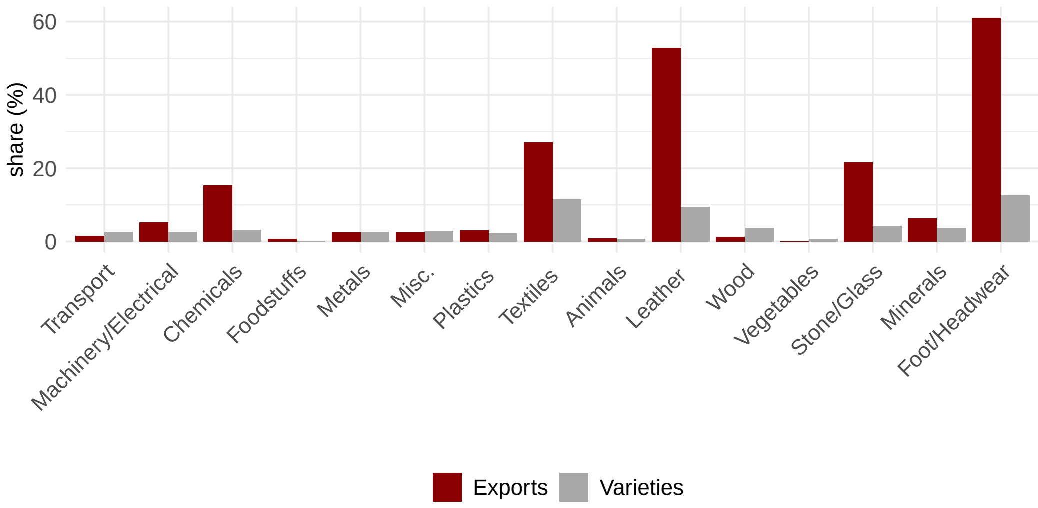
A.2 Further Descriptive Statistics

Figure A.1: Margin decomposition of growth in EUR and USD-denominated exports



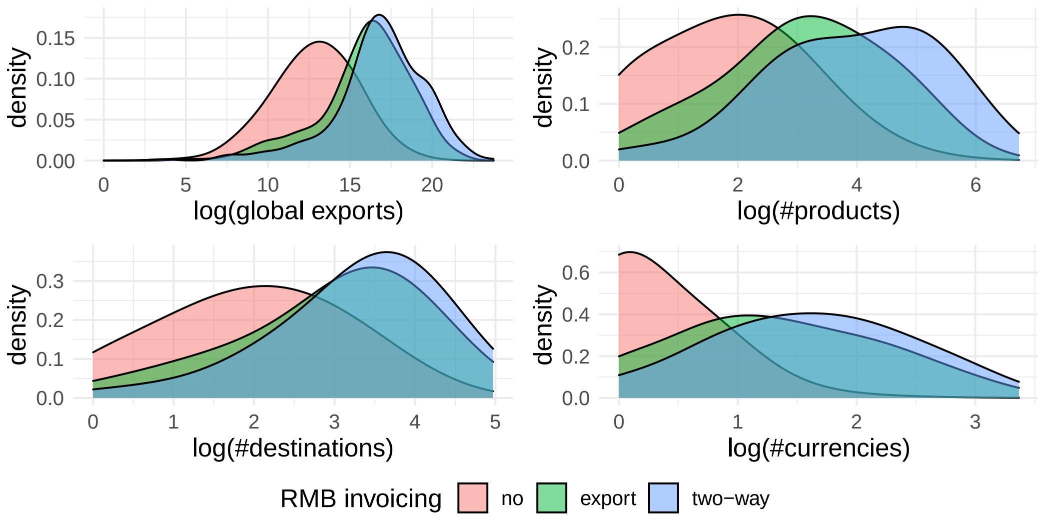
Note: Figures above display the growth in EUR-denominated and USD-denominated exports to China (relative to 2011) along multiple margins following the decomposition strategy proposed by Bernard et al. (2009).

Figure A.2: RMB penetration in exports to China by sector



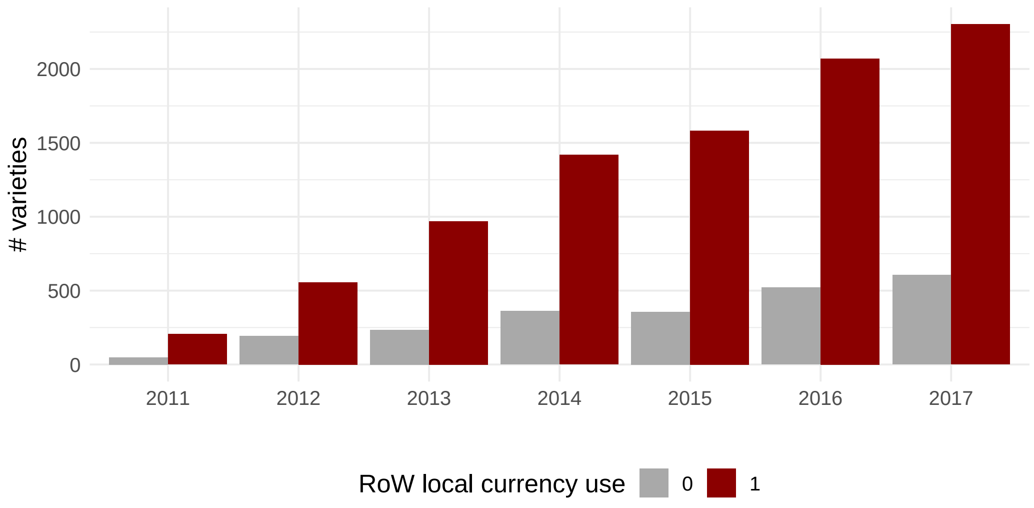
Note: The figure shows the share of total export value (red) or varieties (grey) in a given sector that is invoiced in RMB over 2011-2017. Sectors are arranged from left to right in decreasing order of their share in France's aggregate exports to China over the same time period.

Figure A.3: Kernel density distributions: RMB vs non-RMB invoicing firms



Note: The figure displays several kernel density plots for firms that i) never traded in RMB (red); ii) only exported in RMB (green) and; iii) exported and imported in RMB (blue).

Figure A.4: Invoicing strategies in other markets for varieties invoiced in RMB in China



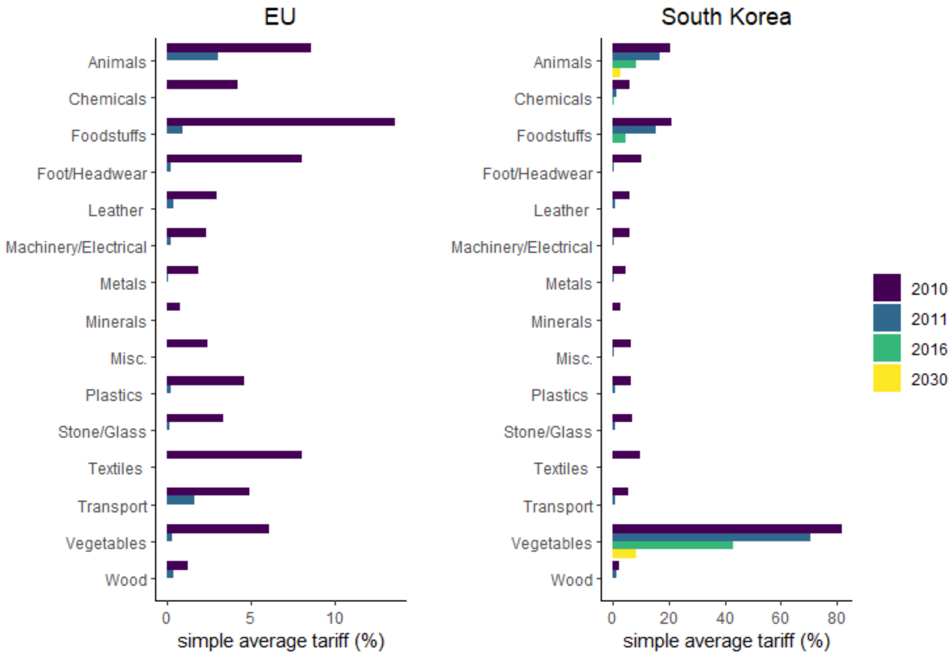
Note: The figure shows the number of RMB-invoiced varieties exported to China in any given year which are also invoiced in local currencies when exported to other extra-EU destinations excluding the US (in red) or not (in grey).

Appendix B

Appendix to Chapter 2

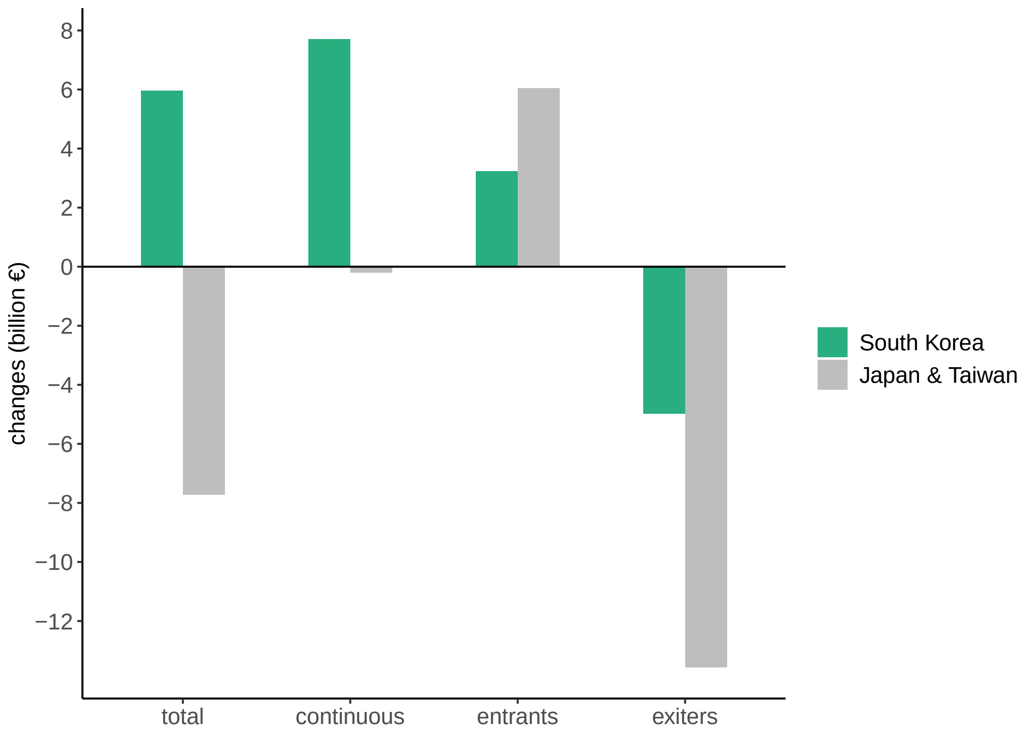
B.1 Further Descriptive Statistics

Figure B.1: Tariff schedules of the EU and South Korea



Note: The graph depicts changes in average applied tariffs imposed by the EU on South Korean products (left) and by South Korea on EU products (right). These changes are examined over the course of the agreement's transition period, from 2011 to 2030. In 2010, prior to the entry into force of the FTA, applied tariffs corresponded to MFN duties. Considering the differences in tariff levels between the EU and South Korea, the horizontal axes have different scales.

Figure B.2: Change in exports by type of firm and destination (billion €)



Note: The figure above shows the absolute change in France's exports between the control (2000-2006) and post-FTA period (2011-2016) and compares this change for South Korea (in green) with that of Japan and Taiwan. Changes are further decomposed into changes in sales of continuous, entrants and exiting firms to that destination. Export values are adjusted by France's GDP deflator drawn from the World Bank Database.

Table B.1: EU FTAs entering into force over 2006-2016

Agreement	EIF	depth index	rasch depth
EU Enlargement	2007	5	0.85
CARIFORUM EU EPA	2008	7	1.58
Albania EU SAA	2009	7	1.26
Cote d'Ivoire EU EPA	2009	3	0.26
EU Montenegro SAA	2010	6	1.37
European Economic Area (EEA)	2011	5	0.67
EU Korea	2011	7	2.03
EU Enlargement	2013	5	0.90
EU Serbia SAA	2013	7	1.42
Central America EC	2013	6	1.76
Colombia EC Peru	2013	7	1.89
EU Georgia	2014	7	2.03
Bosnia and Herzegovina EC SAA	2015	4	1.06
EU Kosovo SAA	2015	5	1.18
EU Moldova	2016	7	2.11
EU SADC EPA	2016	4	0.54

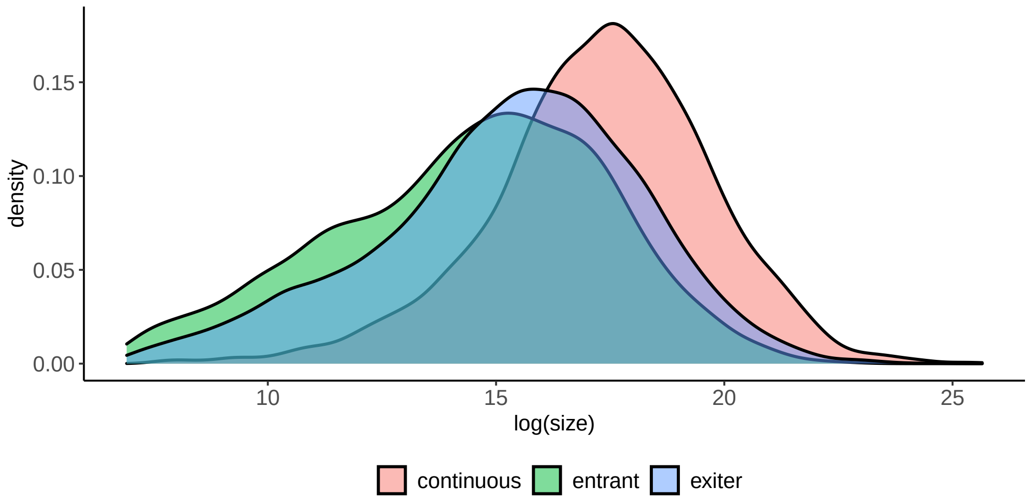
Note: This table lists trade agreements between the EU and other trade partners that entered into force over 2006-2011. Their respective depth indices are drawn from Dür, Baccini, and Elsig (2014).

Table B.2: Percentage point reductions in tariffs for French exporters to South Korea

	Simple average	Weighted average	Weighted average
Weights:	–	Exporter size	Korea sales
Q4	7.79	6.26	6.18
Q3	7.47	1.78	3.46
Q2	9.00	1.84	1.70
Q1	6.68	3.27	6.63

Note: This table provides the percentage point change in the simple average and weighted average tariffs faced by continuous French exporters to South Korea, between the control (2000-2006) and FTA period (2011-2016). The averages are computed within quartiles of the size measure as defined in Section 2.2.3. The weights are taken as the exporter size or the sales of the exporter to South Korea in the control period.

Figure B.3: Size distributions by type of firm



Note: This graph plots the kernel densities of size defined at the firm level for firms that i) exported to South Korea in both periods (continuous); ii) that began exporting to South Korea during the post-FTA period (entrant) and; iii) that exited South Korea in the post-FTA period (exiter).

Table B.3: Size dispersion across sectors

Sector	C.V.
Transport	12.72
Foodstuffs	11.82
Machinery/Electrical	11.55
Minerals	10.73
Leather	9.92
Misc.	8.61
Chemicals	8.50
Metals	8.01
Plastics	7.79
Wood	6.68
Stone/Glass	6.49
Vegetables	5.56
Foot/Headwear	5.25
Textiles	4.96
Animals	3.87

Note: This table reports the coefficient of variation (C.V.) for the baseline size measure across goods sectors. The coefficient of variation for a sector is calculated as the ratio of standard deviation to the mean of size for all firm-product combinations within that goods sector.

B.2 Robustness Checks

Table B.4: Characteristics of firms exporting to South Korea

Dependent Variable:	Exporter to Korea (0,1)		
Model:	(1)	(2)	(3)
multi-product	0.018***		
	(0.0007)		
multi-product × FTA	0.006**		
	(0.002)		
multi-destination		0.011***	
		(0.0007)	
multi-destination × FTA		0.010***	
		(0.003)	
exporter to Japan/Taiwan			0.074***
			(0.003)
exporter to Japan/Taiwan × FTA			0.025***
			(0.008)
R^2	0.576	0.576	0.581

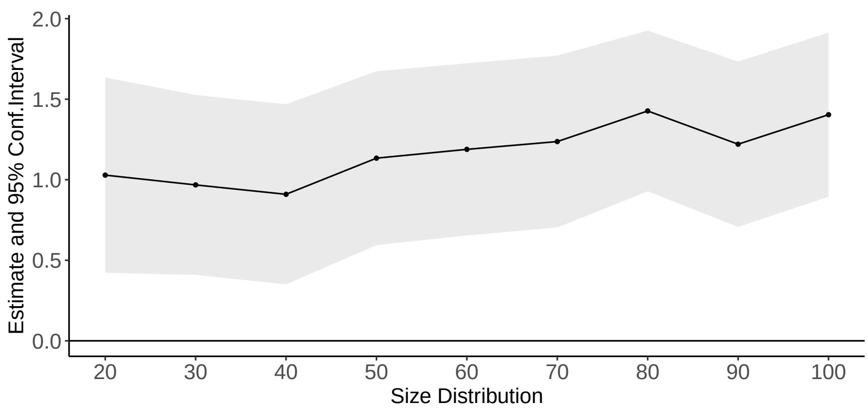
Note: Number of observations $N = 1,685,204$. This table reports coefficients from linear probability models. The dependent variable is set equal to 1 if the firm exported to South Korea in the given year and 0 otherwise. To compare the coefficients between the control and post-FTA periods, we interact the explanatory variables with an FTA dummy that equals 1 over 2011-2016. Multi-product and multi-destination are dummy variables at the firm-year level. In column (3), the explanatory variable is a dummy taking the value of 1 if the firm exported to Japan or Taiwan in the given year. All regressions include firm and year fixed effects. Standard errors are clustered by firm and year. Significance codes: ***: 0.01, **: 0.05, *: 0.1

Table B.5: Impact of EUKFTA: Decomposition of export revenues

Dependent Variables:	ln(exports)	ln(quantity)	ln(price)
Model:	(1)	(2)	(3)
$\mathcal{I} \times Q4$	0.664***	0.687***	-0.023
	(0.144)	(0.153)	(0.086)
$\mathcal{I} \times Q3$	0.564***	0.689***	-0.125
	(0.138)	(0.154)	(0.089)
$\mathcal{I} \times Q2$	0.373**	0.434***	-0.061
	(0.154)	(0.161)	(0.102)
$\ln t \times Q4$	-1.13***	-0.780*	-0.352**
	(0.413)	(0.427)	(0.163)
$\ln t \times Q3$	-0.280	0.170	-0.450**
	(0.314)	(0.333)	(0.184)
$\ln t \times Q2$	0.181	0.474	-0.293
	(0.337)	(0.411)	(0.211)
Observations	1,758,070	1,758,070	1,758,070
R^2	0.919	0.945	0.967

Note: Regressions are estimated for export revenues, export quantity and export prices (calculated as the ratio of revenue to quantity) at the firm-product-destination level aggregated to two periods: control (2000-2006) and post-FTA (2011-2016). EUKFTA is a dummy that takes the value of 1 for South Korea in the 2011-2016 period. Regressions include firm-product-time, product-destination-time and firm-product-destination fixed effects. Only continuous exporters are retained i.e. firms that have positive exports in a given product-destination for each of the two periods. Standard-errors are clustered three-way by firm, product and destination. Significance codes: ***: 0.01, **: 0.05, *: 0.1

Figure B.4: Impact of NTB reductions on intensive margin: Comparison across deciles



Note: Following the specification provided by equation (2.2), this graph plots coefficients on the interaction between the $\mathcal{I}_{dt}$ (EUKFTA dummy) with size deciles. The base category comprises of exporters in the bottom decile of the size distribution. The 95% confidence intervals are constructed using standard errors clustered by firm, product and destination.

Table B.6: Impact of EUKFTA: Different clustering methods

Dependent Variable:	ln(exports)			
Model:	(1)	(2)	(3)	(4)
$\mathcal{I} \times Q4$	0.664***	0.664***	0.664***	0.664***
	(0.141)	(0.137)	(0.172)	(0.144)
$\mathcal{I} \times Q3$	0.564***	0.564***	0.564***	0.564***
	(0.136)	(0.134)	(0.172)	(0.138)
$\mathcal{I} \times Q2$	0.373**	0.373**	0.373**	0.373**
	(0.152)	(0.151)	(0.181)	(0.154)
$\ln t \times Q4$	-1.13***	-1.13***	-1.13***	-1.13***
	(0.405)	(0.393)	(0.364)	(0.413)
$\ln t \times Q3$	-0.280	-0.280	-0.280	-0.280
	(0.308)	(0.318)	(0.325)	(0.314)
$\ln t \times Q2$	0.181	0.181	0.181	0.181
	(0.335)	(0.329)	(0.346)	(0.337)
$\mathcal{I} \times \ln t \times Q4$	-1.27	-1.27	-1.27	-1.27
	(0.998)	(0.808)	(1.47)	(0.979)
$\mathcal{I} \times \ln t \times Q3$	-1.67*	-1.67*	-1.67	-1.67*
	(0.908)	(0.903)	(1.38)	(0.858)
$\mathcal{I} \times \ln t \times Q2$	0.704	0.704	0.704	0.704
	(0.891)	(1.47)	(1.28)	(0.868)
Clustering	$p-d$	$f-d$	$f-p$	$f-p-d$
Observations	1,758,070	1,758,070	1,758,070	1,758,070
R^2	0.918	0.918	0.918	0.918

Note: Regressions are estimated for exports at the firm-product-destination level aggregated to two periods: control (2000-2006) and post-FTA (2011-2016). EUKFTA is a dummy that takes the value of 1 for South Korea in the 2011-2016 period. Regressions include firm-product-time, product-destination-time and firm-product-destination fixed effects. Only continuous exporters are retained i.e. firms that have positive exports in a given product-destination for each of the two periods. Standard-errors are clustered by destination and product in column (1), firm and destination in column (2), firm and product in column (3) and firm, product and destination in column (4). Significance codes: ***: 0.01, **: 0.05, *: 0.1

Table B.7: Impact of EUKFTA: Aggregating exports to the firm-destination-time level

Dependent Variable:		ln(exports)	
Model:	(1)	(2)	(3)
$\mathcal{I} \times \ln(size)$	0.046***		
	(0.008)		
$\mathcal{I} \times median$		0.103**	
		(0.050)	
$\mathcal{I} \times Q4$			0.240**
			(0.101)
$\mathcal{I} \times Q3$			0.233**
			(0.106)
$\mathcal{I} \times Q2$			0.185*
			(0.106)
$\ln t \times \ln(size)$	-0.302***		
	(0.094)		
$\ln t \times median$		-1.84***	
		(0.385)	
$\ln t \times Q4$			-2.62***
			(0.768)
$\ln t \times Q3$			-1.91***
			(0.581)
$\ln t \times Q2$			-0.724
			(0.598)
Base category	–	$< 50^{th} ptile$	$Q1$
R^2	0.879	0.879	0.879

Note: $N = 798{,}129$. This table examines the differential impact of NTB reductions on exports, aggregated across products within firms. Thus, the dependent variable is at the firm-destination-time level. It plots coefficients on the interaction between the EUKFTA dummy with various size bins. The regressions include firm-time, destination-time and firm-destination fixed effects. Only continuous exporters are retained i.e. firms that have positive exports in a given destination in both control and post-FTA periods. Additional controls interactions between size bins and a dummy variable that takes the value of 1 in the second period for all other countries with which the EU implemented FTAs after 2006. The 95% confidence intervals are constructed using two-way clustered (firm & destination) standard errors.

Table B.8: Robustness checks with continuous size measure

Dependent Variable:	ln(exports)			
Panel A: Truncating size distribution				
Sample:	Full sample	Drop top 1%	Drop top 5%	Drop top 10%
$\mathcal{I} \times \ln(size)$	0.079***	0.070***	0.070***	0.075
	(0.009)	(0.010)	(0.015)	(0.026)
$\ln t \times \ln(size)$	-0.243***	-0.251***	-0.239***	-0.273
	(0.065)	(0.062)	(0.072)	(0.095)
Observations	1,758,070	1,696,706	1,545,340	1,390,912
R^2	0.919	0.918	0.919	0.924

Panel B: Alternate size measures				
Size measure:	Firm-Product Global	Firm-Product Intra-EU	Firm-Product Extra-EU	Firm Global
$\mathcal{I} \times \ln(size)$	0.079***	0.045***	0.086***	0.075***
	(0.009)	(0.011)	(0.017)	(0.006)
$\ln t \times \ln(size)$	-0.243***	-0.328***	-0.075	-0.070
	(0.065)	(0.071)	(0.082)	(0.063)
Observations	1,758,070	1,564,004	1,652,294	1,758,070
R^2	0.919	0.918	0.917	0.919

Panel C: Tariff staging					
Product categories:	MFN=0	EIF	3 years	10 years	11+ years
$\mathcal{I} \times \ln(size)$	0.171***	0.074***	0.068***	0.042*	-0.058**
	(0.015)	(0.012)	(0.013)	(0.022)	(0.022)
$\ln t \times \ln(size)$	-0.519***	-0.258***	-0.277*	-0.118	-0.173***
	(0.109)	(0.094)	(0.158)	(0.077)	(0.048)
Observations	191,936	893,520	252,892	224,251	43,156
R^2	0.924	0.919	0.908	0.916	0.925

Note: This table replicates the robustness checks reported in Section 2.4 for a continuous measure of size instead of size bins. Panel A shows how the magnitude of the size advantage varies when the size distribution is truncated, as discussed in Section 2.4.1. Panel B shows the impact of using alternative measures of exporter size as discussed in Section 2.4.2. Finally, Panel C reports the size coefficients for products belonging to different tariff staging categories in South Korea's tariff schedule; see Section 2.4.5 for detailed discussion.

Table B.9: Count models: Market entry and product diversification

Dependent Variables:	Exporter (0,1)		# Products	
	(1)	(2)	(3)	(4)
$\mathcal{I} \times Q4$	-0.269*	-0.287*	0.053	0.021
	(0.161)	(0.164)	(0.060)	(0.066)
$\mathcal{I} \times Q3$	0.046	-0.068	0.017	-0.036
	(0.164)	(0.169)	(0.062)	(0.069)
$\mathcal{I} \times Q2$	0.233	0.135	0.080	0.034
	(0.188)	(0.193)	(0.068)	(0.075)
$\ln t \times Q4$		-0.353		-0.553
		(0.422)		(0.376)
$\ln t \times Q3$		-2.35***		-0.930**
		(0.437)		(0.414)
$\ln t \times Q2$		-2.04***		-0.804*
		(0.503)		(0.442)
Estimation	Poisson	Poisson	Neg. Binomial	Neg. Binomial
Fixed effects	$\theta_{fd}, \theta_{ft}, \theta_{dt}$	$\theta_{fd}, \theta_{ft}, \theta_{dt}$	$\theta_{fd}, \theta_{ft}, \theta_{dt}$	$\theta_{fd}, \theta_{ft}, \theta_{dt}$
Observations	1,817,127	1,817,127	3,196,118	3,196,118
Pseudo R^2	0.119	0.120	0.340	0.340
Over-dispersion			32.398	32.402

Note: The table above reports regression results for the impact of the EUKFTA for the market entry and product margins. $\mathcal{I}$ is a dummy variable that takes the value of one for French exports to South Korea from 2011 onwards and the value of zero otherwise. Columns provide results following the specification in equation (2.3). Since the dependent variables here are defined at the firm-destination-time level, size is correspondingly computed at the firm level (aggregating across products) and tariffs are averaged across products within a given destination and time period. Significance codes: ***: 0.01, **: 0.05, *: 0.1

Appendix C

Appendix to Chapter 3

C.1 Model Setup

The KITE model builds on Caliendo and Parro (2015) and its implementation is similar to that of Aichele, Felbermayr, and Heiland (2014) and Hinz and Monastyrenko (2022). There are N countries, indexed o and d, and J sectors, indexed j and k. Production uses labor as the sole factor, which is mobile across sectors but not across countries. All markets are perfectly competitive. Sectors are either wholly tradable or non-tradable.

There are L_d representative households in each country that maximize their utility by consuming final goods C_d^j in the familiar Cobb-Douglas form

$$u(C_d) = \prod_{j=1}^{J} C_d^{\alpha_d^j} \quad \text{with} \quad \sum_{j=1}^{J} \alpha_d^j = 1.$$

where α_d^j is the constant consumption share on industries j's goods. Household income I_d is derived from the supply of labor L_d at wage w_d and a lump-sum transfers of tariff revenues. Intermediate goods $\omega^j \in [0,1]$ are produced in each sector j using labor and *composite* intermediate goods from all sectors. Let $\beta_d^j \in [0,1]$ denote the cost share of labor and $\gamma_d^{k,j} \in [0,1]$ with $\sum_k \gamma_d^{k,j} = 1$ the share of sector k in sector j's intermediate, such that

$$q_d^j(\omega^j) = z_d^j(\omega^j) \left[l_d^j(\omega^j) \right]^{\beta_d^j} \left[\prod_{k=1}^{J} m_d^{k,j}(\omega^j)^{\gamma_d^{k,j}} \right]^{1-\beta_d^j}$$

where $z_d^j(\omega^j)$ is the overall efficiency of a producer, $l_d^j(\omega^j)$ is labor input, and $m_d^{k,j}(\omega^j)$ represent the composite intermediate goods from sector k used to produce ω^j. With constant returns to scale and perfectly competitive markets, unit costs are

$$c_d^j = \frac{\Upsilon_d^j w_d^{\beta_d^j}}{z_d^j(\omega^j)} \left[\prod_{k=1}^{J} (P_d^k)^{\gamma_d^{k,j}} \right]^{1-\beta_d^j}$$

where P_d^k is the price of a composite intermediate good from sector k, and the constant $\Upsilon_d^j = \prod_{k=1}^J (\gamma_d^{k,j} - \beta_d^j \gamma_d^{k,j})^{-\gamma_d^{k,j} + \beta_d^j \gamma_d^{k,j}} (\beta_d^j \gamma_d^j)^{-\beta_d^j \cdot \gamma_d^j}$. Hence, the cost of the input bundle depends on wages and the prices of *all* composite intermediate goods in the economy. Producers of composite intermediate goods supply Q_d^j at minimum costs by purchasing intermediate goods ω^j from the lowest cost supplier across countries, so that

$$Q_d^j = \left[\int r_d^j(\omega^j)^{1-1/\sigma^j} d\omega^j \right]^{\sigma^j/(\sigma^j-1)}.$$

$\sigma^j > 0$ is the elasticity of substitution across intermediate goods within sector j, and $r_d^j(\omega^j)$ the demand for intermediate goods ω^j from the lowest cost supplier such that

$$r_d^j(\omega^j) = \left(\frac{p_d^j(\omega^j)}{P_d^j} \right)^{-\sigma^j} Q_d^j$$

where P_d^j is the unit price of the composite intermediate good

$$P_d^j = \left[\int p_d^j(\omega^j)^{1-\sigma^j} d\omega^j \right]^{1/(1-\sigma^j)}$$

and $p_d^j(\omega^j)$ denotes the lowest price of intermediate good ω^j in d across all possible origin locations, i.e.

$$p_d^j = \min_o \left\{ p_{od}^j \right\}. \tag{C.1}$$

Composite intermediate goods are used in the production of intermediate goods ω^j and as the final good in consumption as C_d^j, so that the market-clearing condition is written as

$$Q_d^j = C_d^j + \sum_{k=1}^{J} \int m_d^{j,k}(\omega^j) d\omega^j \tag{C.2}$$

Trade in goods is costly, such that the offered price of ω^j from o in d is given by

$$p^j_{od} = \phi^j_{od} \cdot \frac{c^j_o}{z^j_o(\omega^j)} \tag{C.3}$$

where ϕ^j_{od} denote generic bilateral sector-specific trade frictions. These can take a variety of forms — e.g. tariffs, non-tariff barriers, but also sanctions. In that case, we can specify

$$\phi^j_{od} = \tau^j_{od} \cdot \kappa^j_{od},$$

where $\tau^j_{od} \geq 1$ represent sector-specific ad-valorem tariffs and $\kappa^j_{od} \geq 1$ other iceberg trade costs. Tariff revenue $(\tau^j_{od} - 1)$ is collected by the importing country and transferred lump-sum to its households.

Ricardian comparative advantage is induced à la Eaton and Kortum (2002) through a country-specific idiosyncratic productivity draw z^j from a Fréchet distribution.[1]

The price of the composite good is then given as

$$P^j_d = A^j \left[\sum_{o=1}^{N} \lambda^j_o (c^j_o \phi^j_{od})^{-\theta^j} \right]^{-1/\theta^j} \tag{C.4}$$

which, for the non-tradable sector towards *all* non-domestic sources collapses to

$$P^j_d = A^j (\lambda^j_d)^{-1/\theta^j} c^j_d \tag{C.5}$$

where $A^j = \Gamma(\xi^j)^{1/(1-\sigma^j)}$ with $\Gamma(\xi^j)$ being a Gamma function evaluated at $\xi^j = 1 + (1-\sigma^j)/\theta^j$. Total expenditures on goods from sector j in country d are given by $X^j_d = P^j_d Q^j_d$. The expenditure on those goods originating from country o is called X^j_{od}, such that the share of j from o in d is $\pi^j_{od} = X^j_{od}/X^j_d$. This share can also be expressed as

$$\pi^j_{od} = \frac{\lambda^j_o (c^j_o \phi^j_{od})^{-\theta^j}}{\sum_{h=1}^{N} \lambda^j_h (c^j_h \phi^j_{hd})^{-\theta^j}} \tag{C.6}$$

Total expenditures on goods from sector j are the sum of the firms' and households' expenditures on the composite intermediate good, either as input to production or for

[1] The productivity distribution is characterized by a location parameter λ^j_o that varies by country and sector inducing *absolute* advantage, and a shape parameter θ^j that varies by sector determining *comparative* advantage.

final consumption

$$X_d^j = \sum_{k=1}^{J}(1-\beta_d^k)\gamma_d^{j,k}\sum_{o=1}^{N}X_o^k\frac{\pi_{od}^k}{\tau_{od}^k\zeta_{od}^k} + \alpha_d^j I_d \tag{C.7}$$

with $I_d = w_d L_d + R_d + B_d$, i.e., labor income, tariff revenue and the aggregate trade balance. Sectoral trade balance is simply the difference between imports and exports

$$B_d^j = \sum_{o=1}^{N} X_{od}^j - X_{do}^j \tag{C.8}$$

and the aggregate trade balance $B_d = \sum_{j=1}^{J} B_d^j$, and $\sum_{d=1}^{N} B_d = 0$, with B_d being exogenously determined. The total trade balance can then be expressed as

$$\sum_{j=1}^{J}\sum_{o=1}^{N} X_d^j \frac{\pi_{od}^j}{\tau_{od}^j\zeta_{od}^j} - B_d = \sum_{j=1}^{J}\sum_{o=1}^{N} X_o^j \frac{\pi_{do}^j}{\tau_{do}^j\zeta_{do}^j}. \tag{C.9}$$

A counterfactual general equilibrium for alternative trade costs in the form of $\hat{\phi}_{od}^j = \phi_{od}^{j\prime}/\phi_{od}^j$[2] can be solved for in changes following Dekle, Eaton, and Kortum (2008).

[2] I.e. where any variable $\hat{x}$ denotes the relative change from a previous value x to a new one x'.

C.2 Additional Simulation Results

Figure C.1: Domestic costs of Iran sanctions

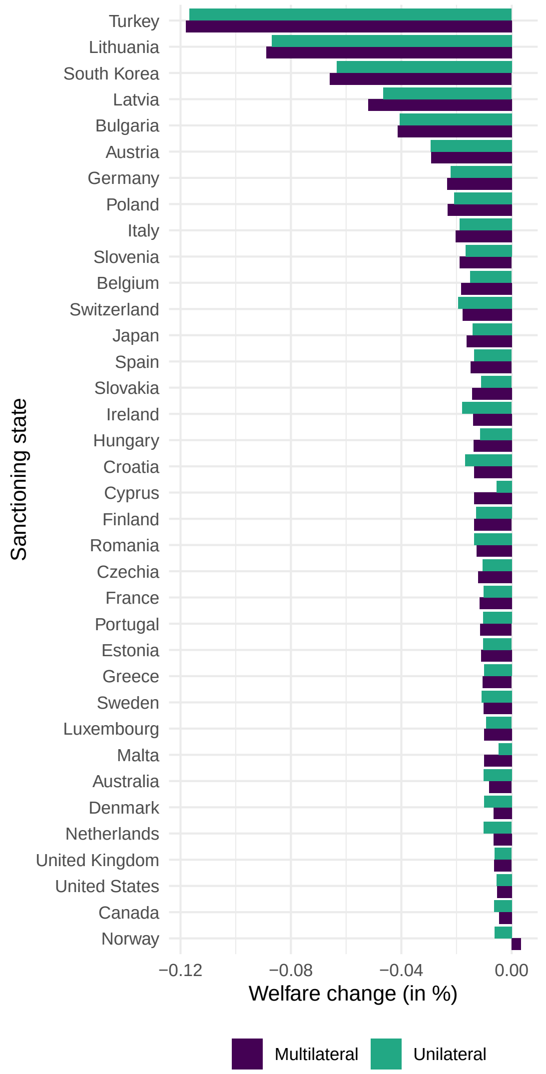
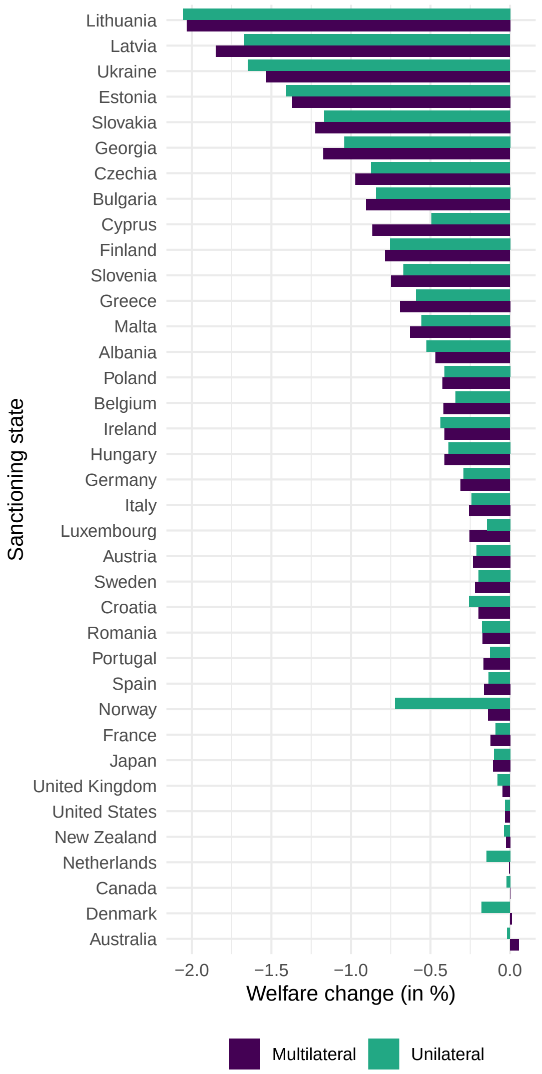


Figure C.2: Domestic costs of Russia sanctions



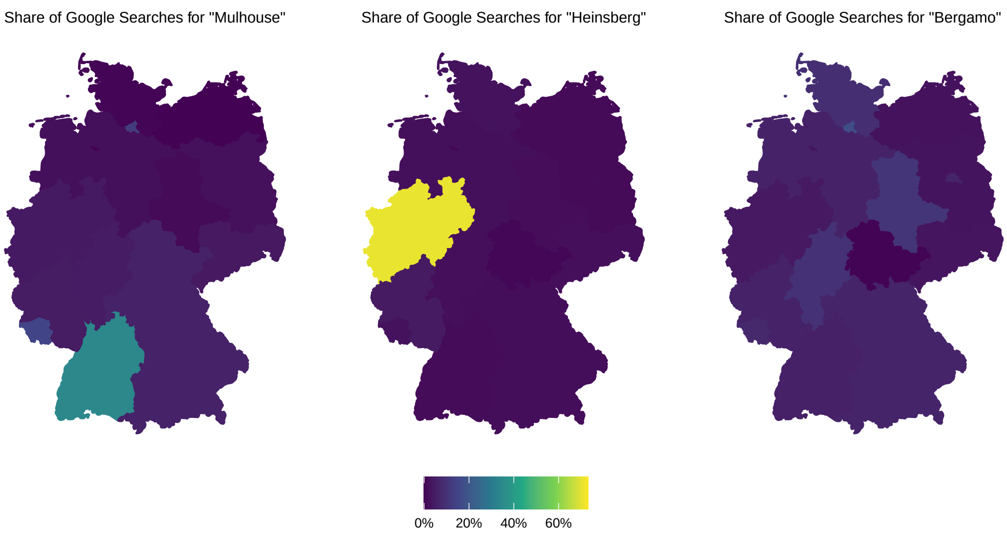
Note: Figures above display each country in the current sanctions coalition against Iran or Russia and the welfare change it experiences domestically either in a unilateral or a multilateral sanctions regime.

Appendix D

Appendix to Chapter 4

D.1 Further Figures

Figure D.1: Google searches for Mulhouse, Heinsberg and Bergamo



Note: The maps show the share of Google searches about the towns of Mulhouse, Heinsberg and Bergamo from all 16 German states (own computation based on data from Google Trends).

D.2 Proof of Proposition

Consider

$$\begin{aligned} & \delta^0 < \delta 1 \\ \Leftrightarrow \quad & \frac{\partial I_i^0}{\partial D_i}\frac{D_i}{I_i^0} < \frac{\partial I_i^1}{\partial D_i}\frac{D_i}{I_i^1}. \end{aligned}$$

Cancelling D_i, and applying the chain rule yields

$$\frac{\partial I_i^0}{\partial D_i}\frac{1}{I_i^0} < \frac{\partial I_i^1}{\partial I_i^0}\frac{\partial I_i^0}{\partial D_i}\frac{1}{I_i^1}.$$

Dividing by $\frac{\partial I_i^0}{\partial D_i}$ flips the sign because it is negative, yielding

$$\frac{I_i^1}{I_i^0} > \frac{\partial I_i^1}{\partial I_i^0}. \tag{D.1}$$

We now turn to equation (4.1) that describes the number of new infections in county i in period 1:

$$I_i^1 = I_i^0 + (1 - \iota_i^0)M_{ii}\iota_i^0 + 2\left(1 - \iota_i^0\right)\sum_{j \neq i} M_{ji}\iota_j^0 \tag{D.2}$$

Dividing by I_i^0 gives us the left-hand side of condition (D.1)

$$\frac{I_i^1}{I_i^0} = 1 + (1 - \iota_i^0)\frac{M_{ii}}{P_i} + 2\frac{1 - \iota_i^0}{P_i}\sum_{j \neq i}\frac{M_{ij}}{P_i}\iota_j^0 \tag{D.3}$$

Taking the derivative of (D.2) with respect to I_i^0 yields

$$\frac{\partial I_i^1}{\partial I_i^0} = 1 + (1 - \iota_i^0)\frac{M_{ii}}{P_i} - \frac{M_{ii}}{P_i}\iota_i^0 - 2\sum_{j \neq i}\frac{M_{ij}}{P_i}\iota_j^0 \tag{D.4}$$

Inserting equations (D.4) and (D.3) into condition (D.1) and rearranging yields

$$
\begin{aligned}
1+(1-\iota_i^0)\frac{M_{ii}}{P_i}+2\frac{1-\iota_i^0}{\iota_i^0}\sum_{j\neq i}\frac{M_{ij}}{P_i}\iota_j^0 &> 1+(1-\iota_i^0)\frac{M_{ii}}{P_i}-\frac{M_{ii}}{P_i}\iota_i^0-2\sum_{j\neq i}\frac{M_{ij}}{P_i}\iota_j^0 \\
2\frac{1-\iota_i^0}{\iota_i^0}\sum_{j\neq i}M_{ij}\iota_j^0 &> -M_{ii}\iota_i^0-2\sum_{j\neq i}M_{ij}\iota_j^0 \\
2\frac{1}{I_i^0}\sum_{j\neq i}M_{ij}\iota_j^0 &> -\frac{M_{ii}}{P_i}\iota_i^0
\end{aligned}
$$

Since the left side is always positive, and the right side is always negative, this proves that $\delta^0 < \delta^1$.

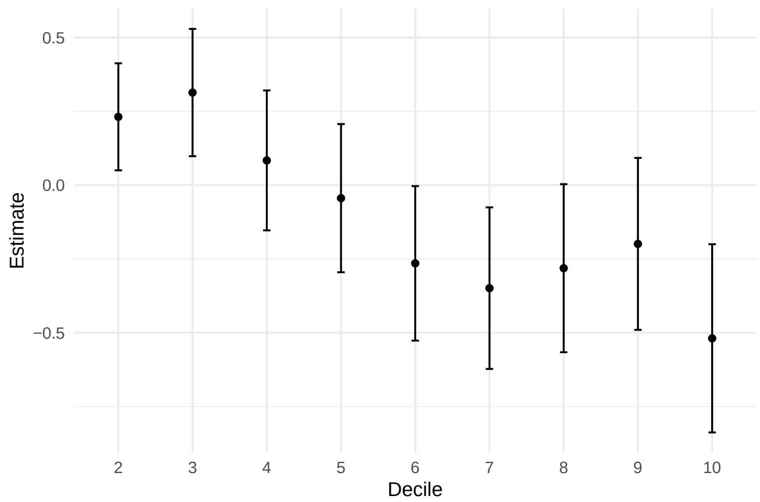
■

D.3 Robustness Checks

In this appendix, we perform a number of robustness checks to determine whether distance elasticities are sensitive to variable definitions or model choice. In all prior regressions, we used continuous measures of distance. However, we can divide the measure into bins in order to test whether the relationship between case counts and distance from Ischgl is non-linear. We therefore alter our baseline specification by introducing a series of dummies for the various deciles of road distance to Ischgl. The estimated coefficients then capture cases relative to the first decile i.e. relative to counties that are nearest to Ischgl. Figure D.2 below plots this sequence of coefficients and reveals close to a linear relationship. To explain with an example, counties belonging to the 10th decile that are farthest away from Ischgl have approximately 0.5% fewer cases in comparison to the reference group of counties closest to Ischgl.

Table D.1 below compares our baseline negative binomial specification for confirmed cases in column (1) with regressions that employ alternative measures of distance. We find that Ischgl dominates over Heinsberg and Mulhouse as a super-spreader location even when switching from road distance to travel time. The results for other controls closely follow the pattern observed in Table 4.2. While the elasticities on population size, testing and share of Catholics are highly significant and comparable across specifications, the coefficients on other demographic and economic factors remain largely insignificant. There in no marked improvement in the model's Pseudo R^2 either when estimating with alternative definitions of distance. Switching to a great circle distance, which should not matter for the spread of the disease, yields a much smaller and statistically insignificant elasticity. Note that now relative latitude to Ischgl has a negative, albeit statistically insignificant coefficient, as it is highly collinear to the great circle

Figure D.2: Robustness check: Distance coefficient by decile



distance. All other coefficients remain largely unchanged.

In column (4) we include fixed effects for great circle distance decile. We therefore compare counties with very similar great circle distance to Ischgl, yet varying road distances. The results are still very similar to those in column (1). The same holds true for column (5). Here we include latitude decile fixed effects, hence comparing counties at similar latitudes. While the distance to Ischgl elasticity is somewhat smaller, the effect remains highly significant.

The following robustness checks relate to the choice of the dependent variable and the estimation strategy. In Table D.2, we move towards analyzing CIR as opposed to the number of cases. With CIR as our outcome variable, we are now no longer in a count-model and can estimate regressions with simple OLS. Consistent with prior findings, we observe that distance to Ischgl is a significant predictor for infections. In a similar vein, we move from count models for fatalities in Table 4.3 to estimating OLS regressions for CFR in Table D.3. This change does not undermine our main results. While testing capacity and share of the elderly influence CFR, distances of counties from super-spreader locales do not.

Table D.1: Robustness check: Distance measure

	Dependent variable:				
	Number of confirmed cases				
	(1)	(2)	(3)	(4)	(5)
log(Population)	1.068^{***}	1.070^{***}	1.044^{***}	1.090^{***}	1.120^{***}
	(0.053)	(0.055)	(0.055)	(0.060)	(0.055)
log(Number of tests)	0.186^{***}	0.192^{***}	0.207^{***}	0.206^{***}	0.228^{***}
	(0.044)	(0.045)	(0.044)	(0.052)	(0.050)
log(Distance to Ischgl)	-0.879^{***}	-0.817^{***}	-0.153	-0.800^{**}	-0.586^{***}
	(0.293)	(0.249)	(0.204)	(0.402)	(0.189)
log(Distance to Heinsberg)	-0.080	-0.041	-0.036	-0.124	-0.123
	(0.094)	(0.157)	(0.110)	(0.090)	(0.093)
log(Distance to Mulhouse)	-0.086	-0.091	0.003	0.071	0.017
	(0.111)	(0.127)	(0.108)	(0.124)	(0.093)
log(Latitude)	0.209	0.082	-0.237	0.073	
	(0.232)	(0.179)	(0.192)	(0.218)	
log(Population / Area)	0.001	0.006	0.003	0.028	0.022
	(0.047)	(0.049)	(0.049)	(0.050)	(0.048)
Share of Catholics	0.733^{**}	0.812^{**}	0.754^{**}	0.643^{**}	0.515^{*}
	(0.295)	(0.336)	(0.324)	(0.278)	(0.294)
Share of Protestants	0.187	0.168	0.212	0.164	0.064
	(0.252)	(0.262)	(0.262)	(0.284)	(0.276)
Share of 65+	-0.769	-0.577	-1.172	-0.188	0.007
	(2.220)	(2.247)	(2.308)	(1.979)	(1.986)
Share of Foreigners	-0.845	-0.714	-0.842	-1.031	-1.190
	(1.152)	(1.193)	(1.196)	(1.141)	(1.224)
log(GDP p.c.)	0.066	0.047	0.064	-0.011	0.100
	(0.123)	(0.122)	(0.121)	(0.127)	(0.121)
Work-from-Home Index	1.137	0.987	1.326	1.148	0.199
	(1.204)	(1.204)	(1.245)	(1.197)	(1.263)
log(China Trade)	-0.002	0.006	0.041	0.010	-0.014
	(0.069)	(0.068)	(0.068)	(0.069)	(0.069)
Distance measure	Road	Travel time	Great circle	Road	Road
Fixed effects	-	-	-	Great circle decile	Latitude decile
Pseudo R2	0.76	0.75	0.75	0.77	0.77
Observations	401	401	401	401	401
θ	4.404^{***}	4.347^{***}	4.261^{***}	4.701^{***}	4.556^{***}
	(0.308)	(0.303)	(0.297)	(0.330)	(0.319)

Note: Constant and great circle decile fixed effects in column (4) not reported. Robust standard errors: *p<0.1; **p<0.05; ***p<0.01.

Table D.2: Case Incidence Rate, OLS Regressions

	Dependent variable:				
	Number of confirmed cases / Population x 100.000				
	(1)	(2)	(3)	(4)	(5)
Number of tests	0.090***	0.057***	0.050***	0.043***	0.043***
	(0.008)	(0.008)	(0.009)	(0.011)	(0.011)
log(Distance to Ischgl)		−0.134***	−0.224***	−0.213**	−0.213**
		(0.016)	(0.084)	(0.084)	(0.084)
log(Distance to Heinsberg)			−0.026**	−0.014	−0.018
			(0.012)	(0.014)	(0.014)
log(Distance to Mulhouse)			−0.025	−0.029	−0.032
			(0.042)	(0.045)	(0.045)
log(Distance to Bergamo)			0.167	0.166	0.182
			(0.158)	(0.157)	(0.160)
log(Latitude)			−0.013	0.010	0.006
			(0.043)	(0.045)	(0.046)
log(Population / Area)				−0.010	−0.018
				(0.010)	(0.012)
Share of Catholics				0.102*	0.110**
				(0.052)	(0.054)
Share of Protestants				−0.024	−0.019
				(0.057)	(0.058)
Share of 65+				−0.157	−0.060
				(0.369)	(0.382)
Share of Foreigners				0.142	0.073
				(0.278)	(0.288)
log(GDP p.c.)					0.011
					(0.031)
Work-from-Home Index					0.300
					(0.296)
log(China Trade)					0.001
					(0.014)
Observations	401	401	401	401	401
R^2	0.252	0.359	0.367	0.386	0.389

Note: Constant not reported. Robust standard errors: $^{*}p<0.1$; $^{**}p<0.05$; $^{***}p<0.01$.

Table D.3: Case Fatality Rate, OLS Regressions

	Dependent variable:				
	Number of deaths / Confirmed cases 18 days ago				
	(1)	(2)	(3)	(4)	(5)
Log(Lagged Number of confirmed cases)	1.335***	1.864***	1.861***	2.239***	2.227***
	(0.411)	(0.463)	(0.464)	(0.477)	(0.475)
log(Population)	−1.902***	−2.539***	−2.540***	−2.689***	−2.007***
	(0.489)	(0.537)	(0.538)	(0.609)	(0.655)
log(Number of tests)	−0.135	0.060	0.054	0.305	0.161
	(0.280)	(0.305)	(0.308)	(0.356)	(0.358)
log(Distance to Ischgl)		2.107	2.131	1.504	1.970
		(2.274)	(2.290)	(2.263)	(2.273)
log(Distance to Heinsberg)		0.492*	0.491*	0.352	0.281
		(0.252)	(0.253)	(0.298)	(0.306)
log(Distance to Mulhouse)		0.172	0.181	−0.226	−0.251
		(1.180)	(1.188)	(1.186)	(1.193)
log(Distance to Bergamo)		−1.214	−1.208	−1.074	−0.235
		(4.318)	(4.325)	(4.202)	(4.222)
log(Latitude)		0.155	0.137	0.827	0.162
		(1.209)	(1.226)	(1.272)	(1.300)
log(GDP p.c.)			0.077	0.518	1.114
			(0.768)	(0.971)	(1.039)
log(Population / Area)				−0.589*	−0.344
				(0.332)	(0.329)
Share catholics				−0.947	−1.038
				(1.801)	(1.828)
Share protestants				−0.321	−0.424
				(1.991)	(2.020)
Share population 65+				43.298***	46.561***
				(10.891)	(11.013)
Share foreigners				16.140	14.921
				(11.118)	(11.256)
log(Number of hospital beds)					−0.963***
					(0.343)
Observations	401	401	401	401	396
R^2	0.060	0.077	0.077	0.127	0.136

Note: Constant not reported. Robust standard errors: $^{*}p<0.1$; $^{**}p<0.05$; $^{***}p<0.01$.

BIBLIOGRAPHY

Afesorgbor, Sylvanus Kwaku and Renuka Mahadevan (2016). "The impact of economic sanctions on income inequality of target states". *World Development*, 83, 1–11.

Aguiar, Angel, Maksym Chepeliev, Erwin L Corong, Robert McDougall, and Dominique van der Mensbrugghe (2019). "The GTAP Data Base: Version 10". *Journal of Global Economic Analysis*, 4 (1), 1–27.

Aichele, Rahel, Gabriel Felbermayr, and Inga Heiland (2014). *Going deep: The trade and welfare effects of TTIP*. Tech. rep.

Alfaro-Urena, Alonso, Isabela Manelici, and Jose P Vasquez (2020). "The effects of joining multinational supply chains: New evidence from firm-to-firm linkages". *Available at SSRN 3376129*.

Alipour, Jean-Victor, Oliver Falck, and Simone Schüller (2020). "Germany's capacities to work from home". *CESifo Working Paper*, (8227).

Amiti, Mary, Oleg Itskhoki, and Jozef Konings (2020). *Dominant currencies: How firms choose currency invoicing and why it matters*. Tech. rep. National Bureau of Economic Research.

Anderson, James E (2011). "The Gravity Model". *Annual Review of Economics*, 3 (1), 133–160.

Antràs, Pol and Davin Chor (2018). *On the measurement of upstreamness and downstreamness in global value chains*. Routledge.

Arkolakis, Costas (2010). "Market penetration costs and the new consumers margin in international trade". *Journal of Political Economy*, 118 (6), 1151–1199.

Attia, Hana and Julia Grauvogel (2019). "Easier in than out: The protracted process of ending sanctions". *GIGA Focus Global*, (05).

Autor, David, David Dorn, Lawrence F Katz, Christina Patterson, and John Van Reenen (2020). "The fall of the labor share and the rise of superstar firms". *The Quarterly Journal of Economics*, 135 (2), 645–709.

Bahaj, Saleem and Ricardo Reis (2020). "Jumpstarting an international currency".

Baier, Scott and Jeffrey Bergstrand (2007). "Do free trade agreements actually increase members' international trade?" *Journal of International Economics*, 71 (1), 72–95.

Barbiero, Omar (2020). "The valuation effects of trade". *SSRN 3513170*.

Barro, Robert, Jose Ursua, and Joanna Weng (2020). "The Coronavirus and the great influenza pandemic. Lessons from the "Spanish Flu" for the Coronavirus's potential effects on mortality and economic activity". *NBER Working Paper 26866*.

Bergounhon, Flora, Clémence Lenoir, and Isabelle Mejean (2018). *A guideline to French firm-level trade data*. Tech. rep.

Bernard, Andrew, Bradford Jensen, Stephen Redding, and Peter Schott (2009). "The margins of US trade". *American Economic Review*, 99 (2), 487–93.

Blackwill, Robert and Harris (2016). *War by other means: Geoeconomics and statecraft*. Harvard University Press.

Blanga-Gubbay, Michael, Paola Conconi, and Mathieu Parenti (2020). "Globalization for sale". *CEPR Discussion Paper No. DP14597*.

Boudreau, Donald G (1998). "UN sanctions burden sharing: Article 50 of the UN charter and assistance to third states". *Diplomacy and Statecraft*, 9 (1), 24–52.

Boz, Emine, Camila Casas, Georgios Georgiadis, Gita Gopinath, Helena Le Mezo, Arnaud Mehl, and Tra Nguyen (2020). "Patterns in invoicing currency in global trade".

Boz, Emine, Gita Gopinath, and Mikkel Plagborg-Møller (2019). "Dollar invoicing and the heterogeneity of exchange rate pass-through". *AEA Papers and Proceedings*. Vol. 109, 527–32.

Bustos, Paula (2011). "Trade liberalization, exports, and technology upgrading: Evidence on the impact of MERCOSUR on Argentinian firms". *American Economic Review*, 101 (1), 304–40.

Caliendo, Lorenzo and Fernando Parro (2015). "Estimates of the Trade and Welfare Effects of NAFTA". *Review of Economic Studies*, 82 (1), 1–44.

Cameron, Colin and Pravin Trivedi (2013). *Regression analysis of count data*. Cambridge, UK: Cambridge University Press.

Carballo, Jerónimo, Alejandro Graziano, Georg Schaur, and Christian Volpe Martincus (2016). *Endogenous border times*. Tech. rep. IDB Working Paper Series.

Chaney, Thomas (2008). "Distorted gravity: The intensive and extensive margins of international trade". *American Economic Review*, 98 (4), 1707–21.

Chowdhry, Sonali, Gabriel Felbermayr, Julian Hinz, Katrin Kamin, Anna-Katharina Jacobs, and Hendrik Mahlkow (2020). *The economic costs of war by other means*. Kiel Policy Brief 147. URL: http://hdl.handle.net/10419/225500.

Corsetti, Giancarlo, Paolo Pesenti, Richard Clarida, and Jeffrey Frankel (2007). "The simple geometry of transmission and stabilization in closed and open economies". *NBER international seminar on macroeconomics*. Vol. 2007. 1. The University of Chicago Press Chicago, IL, 65–129.

Crowley, Meredith A, Oliver Exton, and Lu Han (2020). "The looming threat of tariff hikes: Entry into exporting under trade agreement renegotiation". *AEA Papers and Proceedings*. Vol. 110, 547–51.

Crozet, Matthieu and Julian Hinz (2020). "Friendly fire: The trade impact of the Russia sanctions and counter-sanctions". *Economic Policy*, 35(101), 97–146.

Cuñat, Alejandro and Robert Zymek (2020). "The (structural) gravity of epidemics".

Dauth, Wolfgang, Sebastien Findeisen, Jens Südekum, and Nicole Wössner (2017). "The adjustment of labor markets to robots". *CEPR Discussion Paper*, (12306).

Dekle, Robert, Jonathan Eaton, and Samuel Kortum (2008). "Global rebalancing with gravity: Measuring the burden of adjustment". *IMF Economic Review*, 55 (3), 511–540.

Di Giovanni, Julian, Andrei A Levchenko, and Isabelle Mejean (2020). *Foreign shocks as granular fluctuations*. Tech. rep. National Bureau of Economic Research.

Döhring, Björn et al. (2008). *Hedging and invoicing strategies to reduce exchange rate exposure- A Euro-area perspective*. Tech. rep. Directorate General Economic and Financial Affairs (DG ECFIN), European Commission.

Dür, Andreas, Leonardo Baccini, and Manfred Elsig (2014). "The design of international trade agreements: Introducing a new dataset". *The Review of International Organizations*, 9 (3), 353–375.

Early, Bryan R (2012). "Alliances and trade with sanctioned states: A study of US economic sanctions, 1950–2000". *Journal of Conflict Resolution*, 56 (3), 547–572.

Early, Bryan R and Menevis Cilizoglu (2020). "Economic sanctions in flux: Enduring challenges, new policies, and defining the future research agenda". *International Studies Perspectives*, 21 (4), 438–477.

Eaton, Jonathan and Samuel Kortum (2002). "Technology, geography, and trade". *Econometrica*, 70 (5), 1741–1779.

Eichenbaum, Martin, Sergio Rebelo, and Mathias Trabandt (2020). "The macroeconomics of epidemics". *NBER Working Paper 26882*.

Eichengreen, Barry and Guangtao Xia (2019). "China and the SDR: Financial liberalization through the back door". *Quarterly Journal of Finance*, 9 (03), 1950007.

Escribà-Folch, Abel, Joseph Wright, and Joseph Wright (2015). *Foreign pressure and the politics of autocratic survival*. Oxford Studies in Democratization.

Etkes, Haggay and Assaf Zimring (2015). "When trade stops: Lessons from the Gaza blockade 2007–2010". *Journal of International Economics*, 95 (1), 16–27.

European Commission (2010). *EU-South Korea FTA: Quick reading guide*. Tech. rep.

— (2017). *Evaluation of the implementation of the free trade agreement between the EU and its Member States and the Republic of Korea*. Tech. rep.

— (2021). *EU-China Comprehensive Agreement on Investment: Key facts and figures*. Tech. rep.

Fadinger, Harald and Jan Schymik (2020). "The effects of working from home on Covid-19 infections and production - A macroeconomic analysis for Germany". *Covid Economics*, (9), 107–134.

Farboodi, Maryam, Gregor Jarosch, and Robert Shimer (2020). "Internal and external effects of social distancing in a pandemic". *Covid Economics*, (9), 22–58.

Felbermayr, Gabriel, Aleksandra Kirilakha, Constantinos Syropoulos, Erdal Yalcin, and Yoto V Yotov (2020). "The global sanctions database". *European Economic Review*, 129, 103561.

Felbermayr, Gabriel, Feodora Teti, and Erdal Yalcin (2019). "Rules of origin and the profitability of trade deflection". *Journal of International Economics*, 121, 103248.

Fontagné, Lionel, Houssein Guimbard, and Gianluca Orefice (2020). "Product-level trade elasticities: Worth weighting for".

Fontagné, Lionel and Gianluca Orefice (2018). "Let's try next door: Technical barriers to trade and multi-destination firms". *European Economic Review*, 101, 643–663.

Fontagné, Lionel, Gianluca Orefice, and Roberta Piermartini (2020). "Making small firms happy? The heterogeneous effect of trade facilitation measures". *Review of International Economics*, 28 (3), 565–598.

Fontagné, Lionel, Gianluca Orefice, Roberta Piermartini, and Nadia Rocha (2015). "Product standards and margins of trade: Firm-level evidence". *Journal of International Economics*, 97 (1), 29–44.

Fontagné, Lionel, Angelo Secchi, and Chiara Tomasi (2018). "Exporters' product vectors across markets". *European Economic Review*, 110, 150–180.

Gaubert, Cecile and Oleg Itskhoki (2021). "Granular comparative advantage". *Journal of Political Economy*, 129 (3), 871–939.

Georgiadis, Georgios, Helena Le Mezo, Arnaud Mehl, and Cedric Tille (2021). "Fundamentals vs. policies: Can the US dollar's dominance in global trade be dented?"

Goldberg, Linda S and Cedric Tille (2016). "Micro, macro, and strategic forces in international trade invoicing: Synthesis and novel patterns". *Journal of International Economics*, 102, 173–187.

Gowa, Joanne and Edward Mansfield (1993). "Power politics and international trade". *American Political Science Review*, 87 (2), 408–420.

Grauvogel, Julia and Christian Von Soest (2014). "Claims to legitimacy count: Why sanctions fail to instigate democratisation in authoritarian regimes". *European Journal of Political Research*, 53 (4), 635–653.

Hagen, Achim and Jan Schneider (2017). *Boon or bane? trade sanctions and the stability of international environmental agreements*. Tech. rep.

Harris, Jeffrey (2020). "The subways seeded the massive Coronavirus epidemic in New York City". *NBER Working Paper*, (27021).

Head, Keith and Thierry Mayer (2014). "Gravity equations: Workhorse, toolkit, and cookbook". *Handbook of International Economics*. Ed. by Gita Gopinath, Elhanan Helpman, and Kenneth Rogoff. Vol. 4. Handbook of International Economics. Elsevier, 131–195.

Heilmann, Kilian (2016). "Does political conflict hurt trade? Evidence from consumer boycotts". *Journal of International Economics*, 99 (C), 179–191.

Helpman, Elhanan, Oleg Itskhoki, and Redding (2010). "Inequality and unemployment in a global economy". *Econometrica*, 78 (4), 1239–1283.

Hinz, Julian and Evgenii Monastyrenko (2022). "Bearing the cost of politics: Consumer prices and welfare in Russia". *Journal of International Economics*, 103581. ISSN: 0022-1996. DOI: https://doi.org/10.1016/j.jinteco.2022.103581. URL: https://www.sciencedirect.com/science/article/pii/S0022199622000137.

Hufbauer, Gary, Jeffrey Schott, and Kimberly Elliott (1990). *Economic sanctions reconsidered: History and current policy*. Vol. 1. Peterson Institute.

Hufbauer, Gary, Jeffrey Schott, Kimberly Elliott, and Barbara Oegg (2007). "Economic sanctions reconsidered. Washington, DC". *Peterson Institute for International Economics*.

Iacovone, Leonardo and Beata Javorcik (2010). "Multi-product exporters: Product churning, uncertainty and export discoveries". *The Economic Journal*, 120 (544), 481–499.

Ito, Takatoshi, Satoshi Koibuchi, Kiyotaka Sato, and Junko Shimizu (2010). *Why has the Yen failed to become a dominant invoicing currency in Asia? A firm-level analysis of Japanese exporters' invoicing behavior*. Tech. rep. National Bureau of Economic Research.

Ito, Takatoshi, Satoshi Koibushi, Sato Kiyotaka, Shimzu Junko, and Yoshimi Taiyo (2021). *The Dollar, the Yen, or the RMB? A survey data analysis of invoicing currencies among Japanese overseas subsidiaries*. Tech. rep.

Jia, Jayson S, Xin Lu, Yun Yuan, Ge Xu, Jianmin Jia, and Nicholas A Christakis (2020). "Population flow drives spatio-temporal distribution of COVID-19 in China". *Nature*, 1–11.

Karpaty, Patrik and Patrik Tingvall (2015). "Offshoring and home country R&D". *The World Economy*, 38 (4), 655–676.

Kasahara, Hiroyuki and Beverly Lapham (2013). "Productivity and the decision to import and export: Theory and evidence". *Journal of International Economics*, 89 (2), 297–316.

Kasteng, Jonas and Patrik Tingvall (2019). *Who Uses the EU's free trade agreements? – A Transaction-level analysis of the EU–South Korea Free Trade Agreement*. Tech. rep. The National Board of Trade Sweden.

Kozlowski, Julian, Laura Veldkamp, and Venky Venkateswaran (2020). "Scarring body and mind: The long-term belief-scarring effects of Covid-19". *Covid Economics*, (8), 1–26.

Krueger, Dirk, Harald Uhlig, and Taojun Xie (2020). "Macroeconomic dynamics and reallocation in an epidemic". *Covid Economics*, (5), 21–55.

Lakatos, Csilla and Lars Nilsson (2017). "The EU-Korea FTA: Anticipation, trade policy uncertainty and impact". *Review of World Economics*, 153 (1), 179–198.

Lessmann, Kai, Robert Marschinski, and Ottmar Edenhofer (2009). "The effects of tariffs on coalition formation in a dynamic global warming game". *Economic Modelling*, 26 (3), 641–649.

Li, Ruiyun, Sen Pei, Bin Chen, Yimeng Song, Tao Zhang, Wan Yang, and Jeffrey Shaman (2020). "Substantial undocumented infection facilitates the rapid dissemination of novel Coronavirus (SARS-CoV-2)". *Science*, 368 (6490), 489–493.

Lyonnet, Victor, Julien Martin, and Isabelle Mejean (2016). "Invoicing currency and financial hedging".

Mansfield, Edward and Rachel Bronson (1997). "Alliances, preferential trading Arrangements, and international trade". *American Political Science Review*, 91 (1), 94–107.

Mattoo, Aaditya, Nadia Rocha, and Michele Ruta (2020). "Overview: The evolution of deep trade agreements". *Handbook of Deep Trade Agreements*. Ed. by Aaditya Mattoo, Nadia Rocha, and Michele Ruta. World Bank.

Mayer, Thierry, Marc Melitz, and G Ottaviano (2014). "Market size, competition, and the product mix of exporters". *American Economic Review*, 104 (2), 495–536.

Melitz, Marc (2003). "The impact of trade on intra-industry reallocations and aggregate industry productivity". *Econometrica*, 71 (6), 1695–1725.

Melitz, Marc and G Ottaviano (2008). "Market size, trade, and productivity". *Review of Economic Studies*, 75, 295–316.

Melitz, Marc and S Redding (2014). "Heterogeneous firms and trade". *Handbook of International Economics*. Vol. 4. Elsevier, 1–54.

Messer, Todd (2020). "Foreign currency as a barrier to trade: Evidence from Brazil". *Available at SSRN 3888401*.

Neuenkirch, Matthias and Florian Neumeier (2015). "The impact of UN and US economic sanctions on GDP growth". *European Journal of Political Economy*, 40, 110–125.

— (2016). "The impact of US sanctions on poverty". *Journal of Development Economics*, 121, 110–119.

Nordhaus, William (2015). "Climate clubs: Overcoming free-riding in international climate policy". *American Economic Review*, 105 (4), 1339–70.

Pape, Robert A (1997). "Why economic sanctions do not work". *International security*, 22 (2), 90–136.

— (1998). "Why economic sanctions still do not work". *International Security*, 23 (1), 66–77.

Peksen, Dursun and Timothy M Peterson (2016). “Sanctions and alternate markets: How trade and alliances affect the onset of economic coercion”. *Political Research Quarterly*, 69 (1), 4–16.

Pluemper, Thomas and Eric Neumayer (2020). “The COVID-19 pandemic predominantly hits poor neighborhoods, or does it? Evidence from Germany”. *medRxiv*. DOI: 10.1101/2020.05.18.20105395.

Polachek, Solomon W. (1980). “Conflict and trade”. *Journal of Conflict Resolution*, 24 (1), 55–78. ISSN: 00220027. DOI: 10.1177/002200278002400103.

Pollins, Brian M (1989). “Conflict, cooperation, and commerce: The effect of international political interactions on bilateral trade flows”. *American Journal of Political Science*, 33 (3), 737–761.

Rauch, James E (1999). “Networks versus markets in international trade”. *Journal of international Economics*, 48 (1), 7–35.

Rodrik, Dani (2020). *Globalization in the new normal*. Tech. rep. Distinguished ECFIN Lecture, European Commission.

Sforza, Alessandro and Marina Steininger (2020). “Globalization in the time of COVID-19”. *CESifo Working Paper*, (8184).

Stock, James (2020). “Data gaps and the policy response to the novel Coronavirus”. *Covid Economics*, (3), 1–11.

Verity, Robert, Lucy C Okell, Ilaria Dorigatti, Peter Winskill, Charles Whittaker, Natsuko Imai, Gina Cuomo-Dannenburg, Hayley Thompson, Patrick GT Walker, Han Fu, et al. (2020). “Estimates of the severity of Coronavirus disease 2019: A model-based analysis”. *The Lancet Infectious Diseases*.

Yotov, Yoto V., Roberta Piermartini, José-Antonio Monteiro, and Mario Larch (2008). *An advanced guide to trade policy analysis: The Structural gravity model*. World Trade Organization and UNCTAD.

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München, 15. March 2022

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