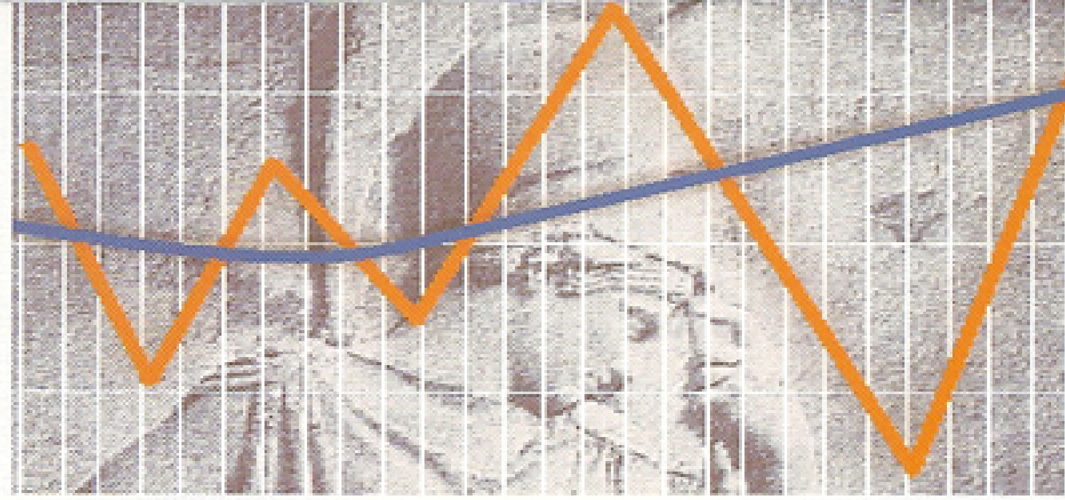

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ifo Beiträge zur Wirtschaftsforschung

Essays on International Trade and Development

Benedikt Heid

ifo Institut
Leibniz-Institut für Wirtschaftsforschung
an der Universität München e.V.

Herausgeber der Reihe: Hans-Werner Sinn
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Preface

This volume was prepared by Benedikt Heid while he was working at the ifo Institute and the University of Bayreuth. It was completed in December 2013 and accepted as a doctoral thesis by the Department of Economics at the Ludwig-Maximilians-Universität München.

It includes six self-contained chapters which deal with the following topics in empirical international trade and development economics: the expansion of firms' export destinations across space and time (chapter 1), the extension of structural gravity models for developed countries to include unemployment (chapter 2) and for Latin American and Caribbean developing countries to additionally include informal employment (chapter 3) as well as the relation between foreign direct investment, trade, and informal employment as illustrated by the maquiladora industry in Mexico (chapter 4), the interaction between migration and trade and their effects on unemployment (chapter 5), and the dynamics of democracy and income (chapter 6).

Keywords: democracy; dynamic panel estimators; export destination choice; firm-level customs data; fixed effects instrumental variable panel estimators; gravity equation; income; informality; informal sector; international trade; *maquiladoras*; *maquilas*; Mexico; MFA/ATC quota removal; migration; offshoring; preferential trade agreements; spatial correlation; trade and labor markets; structural estimation; trade costs; unemployment.

JEL codes: F12; F13; F14; F15; F16; F22; F23; C23; C26; C33; D72; E21; O10; O17; O24.

Für meinen Vater

Es ist nicht genug zu wissen – man muss auch anwenden.
Es ist nicht genug zu wollen – man muss auch tun.
—Johann Wolfgang von Goethe

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First and foremost, I would like to express my gratitude to my advisors, Gabriel Felbermayr and Mario Larch. During the last years, you gave me the opportunity to learn from both of you, both from your lectures, your scholarly writing, and countless discussions. Thank you also for your constant encouragement, patience, and trust in me. Both of you epitomize the ideal combination of academic rigor and policy relevance economists should strive for. Your passion for economics and economic policy will always be a guiding principle for me. Thank you for all the hours of your precious time you shared with me. Furthermore, I would like to thank Carsten Eckel for joining my committee.

I am also indebted to my coauthors, Alejandro Riaño, Fabrice Defever, Julian Langer, and Mario Larch (again). I learnt so much through working with you, discussing new ideas, models, specifications, referee reports, and what it actually means what we are doing. I could not hope for better collaborators, and I am looking forward to continuing to work with you in the future.

I would also like to thank my colleagues, both at the ifo Institute as well as the University of Bayreuth, without whom this experience would have been incomplete. Thank you for numerous discussions and your help in so many ways. Special thanks go to Steffen Sirries with whom I not only shared an office and our common interest in econometrics and economics but also in good food, wine, and spirits, and the good life in general.

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Muchísimas gracias to my girlfriend Laura Márquez Ramos for your love, support, and for enduring both distance and long work hours. Since we met at a conference, you are a central part of my life.

Finally, I would like to express my deepest gratitude to my parents without whom nothing of this would have been possible. Thank you so much for your unconditional support during all those years. We were, are, and will always be the “Family of the Year”.

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vorgelegt von
Benedikt Sebastian Heid

Referent:	Prof. Gabriel Felbermayr, PhD
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Introduction

Variatio delectat. This truism is not only reflected by the fact that trade economists have made it a corner stone of most of their models but also by the vast array of topics typically covered in international trade. There is no denying that the range of the topics covered by the present dissertation is also rather broad. Nevertheless, its six chapters are unified by two recurring themes, one topical, the other methodological: Chapters 2, 3, 4, and 5 deal with the interaction of labor markets and international trade in developed and emerging economies. Chapter 2 develops an estimable gravity model of international trade with unemployment generated by search frictions and estimates it for a set of OECD countries. Chapter 3 extends this model to incorporate informal labor markets, a typical feature of emerging economies, and applies it to a set of Latin American and Caribbean countries. Chapter 4 zooms in on Mexico's experience with the rise of foreign-owned processing plants, so-called *maquiladoras*, and its labor market effects. Chapter 5 returns to developed countries and studies the interaction between trade, unemployment, and migration. The latter chapter is also the link between the topical and methodological overarching themes of this dissertation as it uses dynamic panel estimators which allow to control for unobserved time-invariant heterogeneity and true state dependence. Chapter 1 also uses a dynamic panel estimator to analyze the spread of firms' export destinations across space and time. Finally, Chapter 6 applies the same methods to analyze the determinants of democracy.

All chapters are self-contained and include their own introductions, conclusions, and appendices and can thus be read independently. To guide the time-pressed reader, I present the main contributions and results of each chapter in the following.

Standard models of international trade abstract from firm dynamics. When trade costs fall, firm exports instantaneously adjust to the new optimal level. In addition, standard models assume that firms export to all potential export markets when trade costs are not infinite. Empirically, Eaton et al. (2004) observe that about a third of all firms only export to one market. Helpman et al. (2008b) show that this also translates into aggregate trade flows as in their sample of 158 countries, about half of all possible country pairs do not trade with each other. They develop a gravity model which rationalizes the observed zero trade flows by assuming that the productivity distribution of firms is bounded from above. When firms have to cover fixed costs to serve a particular market, it can be an equilibrium outcome that some countries are not served by any firm from a particular country, as even the most productive firm from this country cannot recover the market-specific export fixed costs as its productivity is too low. While Helpman et al. (2008b) brought trade models more in line with observed aggregate trade flow data, the model is still at odds with export behavior at the firm level: It implies a country hierarchy of export destinations, i.e. if a country is served by a low productivity firm, the same country has to be served by all firms with a higher productivity from the same country. In addition, the model still implies that a fall in trade costs leads to an instantaneous adjustment of firm export behavior to the new equilibrium. Lawless (2009) shows that aggregate data hide the substantial entry and exit dynamics in particular export markets at the firm level. Also, firms do not stick to a clear hierarchy of markets. Chapter 1, which is joint work with Fabrice Defever and Mario Larch, contributes to this strand of literature and presents evidence that about one third of firms do not stick to a country hierarchy. More importantly, we show that this heterogeneity features a common pattern: Firms' export destinations are clustered in space—even more than standard gravity models predict. In addition, firms tend to spread their export destinations in a spatial way: When a firm has exported to a particular country in a particular year, it will tend to export to a country which shares a common border with its previous export destinations in the next year, even when controlling for standard determinants of export destination choices like distance and market size. This behavior has been incorporated in models of international trade (see Morales et al., 2011; Nguyen, 2012, and Albornoz et al., 2012) but empirical evidence so far has relied either on structural models (Morales et al., 2011) or has not taken into account econometric pitfalls like true state dependence and heterogeneity of export destinations at the firm level (Albornoz et al., 2012 and Lawless, 2013). Using the removal of import quotas on textile and apparel exports to European Union countries as well as the United States and Canada in 2005 as a quasi-natural experiment, we present causal evidence for this phenomenon of 'spatial exporters' or 'extended gravity' (see Morales et al., 2011).

Chapter 2, which is joint work with Mario Larch, widens the perspective and moves from a positive analysis of the exporting behavior of individual firms using detailed firm-product-level customs data from one country to a normative, i.e. welfare analysis of trade liberalization using aggregate trade flow data between OECD countries. The cornerstone of empirical quantitative welfare analysis in international trade is the gravity equation: It posits that the trade flow between two countries increases proportionally to the market sizes of the two trading partners and decreases with increasing distance between the two countries, similar to Newton's law of universal gravitation in classical mechanics. Since its first application to trade flows by Tinbergen (1962) and Linnemann (1966), it has become increasingly used in international trade (see Head and Mayer, 2014 for a recent overview).[1] The main reason for this is its empirical success: A simple descriptive regression of (log) trade flows on (log) distance and the (log) GDPs of the exporting and importing country for 28 OECD countries in 2006 explains 84% of the variation in trade flows.[2] This has spurred the interest of trade economists to come up with a thorough theoretical foundation for the gravity equation.[3] A first attempt has been made by Anderson (1979); the model however did not have a major impact on the subsequent literature. Nearly a quarter of a century later, Anderson and van Wincoop (2003) again used a model with Armington (1969) preferences where goods where differentiated across countries to come up with a workable theory for the gravity equation. At the same time, Eaton and Kortum (2002) developed a Ricardian-type model from which they derived a gravity equation, complementing the demand-side driven approach by Anderson and van Wincoop (2003). The crucial point of Anderson and van Wincoop (2003), though often neglected in empirical papers, is that the theory behind the gravity equation can be used for a welfare analysis of counterfactual scenarios like e.g. abolishing a border between two countries, or the inception of a preferential trade agreement. Importantly, the gravity model takes into account the general equilibrium (i.e. income and third-country) effects which have an impact on the welfare analysis of (counterfactual) trade liberalization episodes. Arkolakis et al. (2012) show that these frameworks all imply the same estimate for the welfare gain from moving from autarky to the observed level of trade, and that the change in the import share of GDP (joint with the elasticity of trade with respect to trade costs) is a sufficient statistic for the welfare gains from trade.

However, all the frameworks covered by Arkolakis et al. (2012) assume perfect labor markets, i.e. full employment. We show in Chapter 2 that when one relaxes this assumption and introduces labor market frictions, the welfare formula of Arkolakis et al. (2012) has to

[1] Ravenstein (1885, 1889) applied a similar gravity equation to migration flows even earlier.
[2] I use the data from Head et al. (2010) which are also used in Chapter 2.
[3] The following section is based on Head and Mayer (2014).

be augmented by the net employment change brought about by trade liberalization.[4] We then present a simple model of trade with search-generated unemployment and investigate how the incorporation of labor market frictions affects the estimation of gravity equations as well as the calculation of the effects of counterfactual trade liberalization scenarios.[5] We apply our quantitative framework to investigate the welfare, GDP, and (un)employment effects of existing preferential trade agreements between 28 OECD countries as well as the international spill-over effects of labor market reforms in the United States and Germany.[6] We find that accounting for labor market frictions increases the welfare gains by more than 50 percent in comparison with welfare gains implied by standard gravity frameworks assuming a perfect labor market when we employ commonly used values for the elasticities in our model. The additional change in welfare is brought about by a change in unemployment when trade is liberalized. When trade costs fall, imports of foreign varieties become cheaper, leading to a lower consumer price index in the corresponding country. When labor markets are characterized by search frictions, firms have to incur costs to post vacancies in order to find workers. The lower price level translates one-to-one into lower recruiting costs for domestic firms, therefore more created vacancies and ultimately to lower unemployment.[7] We also find that unilateral improvements in labor market institutions in one country (e.g. the recent Hartz reforms in Germany) reduce the unemployment rate not only in the improving country but also in all of its trading partners due to positive spill-over effects of the labor market reform. This is consistent with reduced-form empirical evidence by Felbermayr et al. (2013). In addition, we present a novel way to estimate the elasticity of substitution as well as the matching elasticity using cross-country trade data.

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[4] Other recent examples of quantitative trade models which imply some sort of modified gravity equation and which are also not captured by the welfare equivalence of Arkolakis et al. (2012) are Waugh (2010) and Fieler (2011). Waugh (2010) argues that trade costs are higher for countries with a lower income per capita to reconcile bilateral trade data with international price data. Fieler (2011) finds that taking into account non-homothetic preferences across countries may improve the empirical fit of trade flows between countries with different levels of income per capita compared to standard gravity models which assume that preferences are homothetic.

[5] Our labor market model is similar to Felbermayr et al. (2013); however, we go beyond their analysis as we investigate the implications of their framework for gravity models and structurally estimate and use it for a quantitative counterfactual evaluation of trade and labor market policies.

[6] Eaton et al. (2013) look at the relation between the observed changes in manufacturing output and the unemployment rate during the financial crisis between 2007 and 2011 using the model presented in Dekle et al. (2007) which implies a gravity equation. However, they assume that the economy is in full employment in equilibrium such that unemployment only arises in their counterfactual analysis where they assume that wages are nominally rigid. Also, they do not investigate the impact of labor market frictions on the welfare equivalence from Arkolakis et al. (2012) nor on the estimation of gravity equations.

[7] Felbermayr et al. (2011a) and Felbermayr et al. (2013) on the one hand and Helpman and Itskhoki (2010) on the other use a similar mechanism in a one- and two-sector model, respectively.

Whereas Chapter 2 focuses on the welfare and employment effects of trade liberalization in developed countries, Chapter 3 turns towards the emerging economies in Latin America and the Caribbean. In principle, one could apply the quantitative framework from Chapter 2 also to this set of countries to study the effects of trade liberalization. However, their labor markets are remarkably different to those in OECD countries as large parts of their labor force is employed in the informal sector. Irrespective of the variety of definitions used, informal employment comprises between 25 to more than 70 percent of the labor force in Latin American countries.[8] The informal sector is characterized by low productivity, small scale establishments. Informal workers are often self-employed, or, when they work as employees, do not possess a written labor contract, or do not have access to social security or health insurance (see ILO, 2010). Due to its low productivity, wages in the informal sector are considerably lower. Informal establishments are also characterized by no strict distinction between private and firm accounts, and often, workers are family members or close relatives (see de Laiglesia and Jütting, 2009 and de Mel et al., 2009). Sometimes, informal workers are paid in kind instead of receiving a monetary wage. Therefore, informal sector employment has generally been seen as detrimental for the welfare of workers.

Empirical evaluations of the impact of trade liberalization on welfare and informal employment until now have focused on single country case studies using a small open economy assumption, contrary to structural gravity models where general equilibrium effects are at the center stage of the analysis.[9] I extend the model from Chapter 2 to incorporate an informal sector whose productivity is linked to its overall size, reflecting the concept of surplus labor or disguised unemployment as discussed by Lewis (1954), one of the first formal analyses of informal employment. Workers can choose between working in the formal and informal sector, taking up the idea of Maloney (2004) that workers may voluntarily choose to work in the informal sector. In the formal sector, workers face the risk of becoming unemployed, whereas this risk does not exist in the informal sector, as workers can always become self-employed. I then use this framework to analyze the welfare and employment effects of preferential trade agreements using a sample of 13 Latin American and Caribbean countries and compare its results to standard frameworks which assume full employment such as Anderson and van Wincoop (2003) or a unified labor market with search frictions such as

[8] For an overview of informality, its different definitions as well as the situation in Latin America and the Caribbean, see Gasparini and Tornarolli (2009).

[9] Examples for these country studies are Goldberg and Pavcnik (2003) for Brazil and Colombia; Fiess et al. (2010) for Argentina, Brazil, Colombia, and Mexico; Bosch et al. (2012) for Brazil; Coşar et al. (2011) for Colombia, and Arias et al. (2013) for Brazil and Mexico.

the framework presented in Chapter 2. I find that the welfare effects of trade liberalization are quantitatively and qualitatively different to those from a framework with full employment and the framework from Chapter 2. I find that on average, preferential trade agreements decrease welfare, reduce the size of the informal sector, and increase the unemployment rate as now more workers are searching for jobs in the formal sector.

In Mexico, informal employment is especially rampant, with 30 to 50% of the labor force employed in the informal sector, depending on the specific definition of informality (see Heid et al., 2011). Policy makers in Mexico see foreign direct investment, especially in the form of greenfield investments, as a way to generate more (formal) employment, especially for Mexico's low skilled workers (see Martin, 2000). Since the 1980s, Mexico has experienced an increase in its *maquiladora* sector. *Maquila* plants, or *maquiladoras* for short, are (predominantly U.S.-owned) export processing plants whose defining characteristics are that they import intermediate inputs (again mainly from the United States), assemble final goods by taking advantage of the low labor cost in Mexico, and export essentially all output again back to the United States. This business model was encouraged by an episode of trade and investment liberalization during the 1980s and increased further in the wake of NAFTA.[10] What are the welfare and labor market consequences of this rise of the *maquiladoras*? Interestingly, the literature on *maquiladoras* (see e.g. Feenstra and Hanson, 1997; Mollick, 2008; Mollick, 2009, and Bergin et al., 2009) abstracts from the decisive feature of the Mexican labor market: the large informal sector.[11] In Chapter 4, which is joint work with Mario Larch and Alejandro Riaño, we evaluate the rise of the *maquiladoras* during the 1990s using a quantitative trade model which is tailor-made to reproduce the key stylized facts of the Mexican economy: A production structure which is characterized by a domestically-owned standard manufacturing sector and a foreign-owned *maquiladora* sector which imports intermediates from abroad, is relatively more skill-intensive and sends its profits outside Mexico, as well as a labor market which is characterized by the possibility of informal employment for low-skilled workers which do not obtain a job in the formal part of the economy. Specifically, we combine a multi-sector model of heterogeneous firms in the spirit of Bernard et al. (2007) with a model of heterogeneous firms featuring search-generated

[10] NAFTA is the North Atlantic Free Trade Agreement between Canada, the United States, and Mexico which came into force on January 1, 1994. For a general overview on NAFTA, see Lederman and Servén (2005) and the papers in the same issue of the *World Bank Economic Review*.

[11] Verhoogen (2008) studies trade and wage inequality in Mexican manufacturing but does exclude *maquila* plants. Waldkirch et al. (2009) study the employment effects of FDI but also do not consider *maquiladoras*; neither studies informal employment.

labor market frictions in the vein of Felbermayr et al. (2011a). We treat Mexico as a small open economy following the modeling strategy of Demidova and Rodríguez-Clare (2009) who generalize the small open economy setup under monopolistic competition from Flam and Helpman (1987) to a heterogeneous firms framework. We calibrate our model to key moments in the data and simulate an exogenous increase in U.S. demand for goods produced by *maquiladoras* similar to the increase in demand observed in the 1990s. We find that the shift in relative demand towards *maquila* goods leads to an increase in *maquiladora* employment at the expense of standard manufacturing employment. Along standard Stolper-Samuelson arguments, we find that the skill premium declines. Interestingly, the accompanying labor reallocation leads to a net decline in low-skilled employment which ultimately leads to an increase in the share of workers employed in the informal sector. In combination with a decrease in the average productivity in the standard manufacturing sector, this leads to lower welfare. Hence our study shows that while Mexican exports have surged, the rise of the *maquiladoras* might be considered a mixed blessing for Mexican workers.

While informal labor markets and their consequences are mainly a phenomenon of emerging economies such as Mexico, migration is a pervasive feature affecting also developed economies. Particularly, immigration into developed economies has become increasingly important: two thirds of the increase in the total number of immigrants worldwide between 1960 and 2000 is due to inflows into Western Europe and the United States (for a detailed overview of global migration trends, see Özden et al., 2011). This increase in immigration has highlighted the importance of studies which shed light on the impact of immigration on the labor market. Labor economists have tended to focus on identifying the causal impact of immigration on wages or employment by using case studies like e.g. the Mariel boat lift (see Card, 1990) or by identifying labor demand responses for finely defined labor markets (see e.g. Borjas, 1999). Parallel to this literature, empirical trade economists have tried to identify the impact of trade liberalization on the level of unemployment (see Dutt et al., 2009 and Felbermayr et al., 2011b) by using cross-country panel regressions. Since at least Mundell (1957), it is well known that international trade can be a substitute for migration, at least in a standard two goods, two factors trade model without trade costs. Hence, goods trade has the same effect as if factors could wander freely between countries. Empirical evidence has suggested that, to the contrary, trade and migration may rather be complements than substitutes (see Gould, 1994 and Felbermayr et al., 2010b), so the literature might be summarized somewhat tongue-in-cheek that it agrees at least on the fact that immigration and trade are not statistically independent of each other. If so, then there arises the need to check whether the results of previous studies do not suffer from an omitted variable bias introduced by omitting either trade or immigration from their empirical specifications. Chapter 5 tackles the latter part by revisiting the impact of trade openess on unemployment while controlling for immigration. It is joint work with Mario Larch. Specifically, we use dynamic panel estimators to analyze the determinants of unemployment rates as introduced by Nickell et al. (2005) and subsequently used by Felbermayr et al. (2011b) to study the impact of trade openness on unemployment. This approach allows us to control for country-specific unobserved time-invariant heterogeneity as well as true state dependence of the unemployment rate, reflecting the high persistence in unemployment rates. We augment the regression from Felbermayr et al. (2011b) by including net inflows of migrants into a country and apply it to a panel data set of 24 OECD countries between 1997 and 2007. We also present an alternative specification which uses a Romer and Frankel (1999) type instrument to control for endogenous migrant inflows. Across our different empirical strategies and robustness checks, we find a robust insignificant effect of immigration on the unemployment rate. Interestingly, we find no significant effect of trade openness on unemployment, either, contrary to the original findings in Felbermayr et al. (2011b).

Controlling for state dependence in the presence of unobserved heterogeneity is also an important issue in the literature dealing with the influence of the level of income on the probability of a country having a democratic political system. This is the subject of Chapter 6 which is joint work with Julian Langer and Mario Larch. The relation between income and democracy is of major interest for both development economists and political scientists alike. Following Lipset (1959), a major proponent of "modernization theory", it has been increasingly accepted that higher levels of income per capita lead to the emergence of democratic regimes. This broad consensus is based on a large body of empirical evidence in favor of modernization theory which uses a variety of econometric approaches.[12] However, Acemoglu et al. (2008) argue that the relation between income and democracy breaks down when controlling for unobserved heterogeneity and state dependence. Using the difference GMM (Generalized Method of Moments) dynamic panel estimator from Arellano and Bond (1991), Acemoglu et al. (2008) show that in a regression with the level of democracy as

[12] Amongst others, Barro (1999), Gundlach and Paldam (2009), Corvalan (2010), Benhabib et al. (2011), Boix (2011), Treisman (2011), and Moral-Benito and Bartolucci (2012) find evidence consistent with modernization theory.

dependent variable, the lagged level of income per capita turns out to be insignificant. Hence previous studies, which have not used dynamic panel estimators, erroneously inferred that there exists an income-democracy nexus. Our paper demonstrates that this conclusion does not hold up to closer scrutiny. While Acemoglu et al. (2008) make a first step in taking into account both effects of unobserved heterogeneity and true state dependence, they do not go far enough. As is well documented by Arellano and Bover (1995) and Blundell and Bond (1998), the difference GMM dynamic panel estimator suffers from a potentially large small sample bias even if the autoregressive parameter of the lagged dependent variable is only moderately large. As political regimes tend to be stable over long periods of time, leading to a high autocorrelation in any measure of democracy, it seems natural to apply the system GMM estimator of Blundell and Bond (1998) to the data set of Acemoglu et al. (2008) as it does not suffer from the small sample bias. Our paper is the first to apply this type of estimator in this literature. We show that our findings are robust to using an alternative measure of democracy as well as to using different external instruments. In addition, we apply methods proposed by Roodman (2009b) to prevent a proliferation of instruments which have not been applied previously in this literature.

Chapter 1

Spatial Exporters*

1.1 Introduction

Firm exports exhibit a geographical pattern. Not only do different firms serve different numbers of countries but also the spatial distribution of those countries differs across firms. Standard gravity models predict that firms are more likely to export to larger countries and to countries that are closer to the country of origin of the firm. These standard gravity forces generate some degree of unconditional spatial concentration of export destinations of firms. Recently, the literature has highlighted that this observed spatial correlation is larger than what the standard gravity model would predict, a fact which has been labeled 'extended gravity' (see Morales et al., 2011, and Albornoz et al., 2012) or 'spatial exporters' (see Defever et al., 2011).

In this paper, we provide causal evidence for 'extended gravity' or 'spatial exporters', i.e. time-varying firm-specific heterogeneity in export destinations shaped by firms' previous export experience in spatially close countries. We take into account unobserved time-invariant heterogeneity at the firm-country level which may arise because firms can differ in their ability to serve specific markets, e.g. due to differences in language skills of their sales force. We also control for true state dependence at the firm-destination level which captures market-specific sunk costs of exporting (see Das et al., 2007). We show that the probability that a firm

* This chapter is based on joint work with Mario Larch and Fabrice Defever. It is a revised version of CESifo Working Paper No. 3672, 2011. A previous version of this paper has been circulated under the title "Spatial Exporter Dynamics".

exports to a country increases by about 2 percentage points for each additional prior export destination with a common border with this country.

One reason for observing spatial exporter patterns may be the crucial need for gathering local information from trading partners over time. Different local information which has been acquired through previous export experience may then lead to different trade networks across firms.[1] When demand is uncertain but correlated across markets, firms may enter new destinations gradually to learn about profits in proximate markets from their previous export experience (see Albornoz et al., 2012; Nguyen, 2012). Also, when firms have to adapt products to specific markets, adaptation costs may be reduced if a firm already has entered markets which are relatively similar (see Morales et al., 2011). As a consequence, when trade barriers fall, firms will expand their export destinations not randomly but following a spatial pattern.

These channels highlight that one has to take into account two different aspects of the firm's problem: i) when to enter a new destination, and ii) where to go. When destination choices of a firm for different destinations are uncorrelated, the decision problem is simple: Every market entry decision can be analyzed on its own. Hence, the two problems of when and where to export can be separated.[2] However, if destination choices are correlated, these two decisions become intrinsically related. Empirically, this leads to a dynamic discrete choice problem. As explained by Morales et al. (2011), this problem is formulated in a straight-forward way theoretically but quickly leads to an empirically de facto unsolvable problem because it involves computing the expected profits for every possible combination of time paths of entries into destinations.[3] Complementary to the structural empirical approach

[1] For instance, an exporting firm may gain access to a new export market via a multinational retailer which already serves a third country. As the network of subsidiaries of wholesalers and of multinational firms tends to expand spatially (see Basker, 2005 and Defever, 2012), this mechanism also implies a spread of exports to contiguous countries. In addition to geography, cultural closeness can also generate a similar pattern through networks of ethnically related firms. For instance, networks may reduce search costs as firms may learn about potential suitable suppliers within their ethnic community (see for instance Rauch, 2001). Recently, Chaney (2011) has developed a model describing trade patterns as an international network. Firms tend to build on their network for finding new trading partners, similar to social interactions between individuals (see Jackson and Rogers, 2007).

[2] For instance, Das et al. (2007) structurally estimate the parameters of a firm's dynamic problem of when to start and stop exporting, irrespective of the specific export market choice.

[3] Therefore, Morales et al. (2011) do not solve this dynamic problem explicitly. Instead, they resort to moment inequality estimators to obtain bounds on the parameters of interest in their structural empirical model. Their estimates based on firm-level export data for Chilean manufacturing firms in the chemicals sector show that startup costs of accessing a new country are significantly determined by the countries to which a firm had previously exported. Albornoz et al. (2012) and Nguyen (2012) focus their analysis on the timing of entry only and assume a hierarchy between countries in terms of profitability and a constant correlation of profits across all export destinations. Together, these assumptions elude the question of where to go. Lawless (2013) shows that entry decisions of firms are correlated with their export status in

suggested by Morales et al. (2011), we use reduced form regressions exploiting a quasi-natural experiment.

We present evidence for 'spatial exporters' relying on the removal of binding import quotas under the MultiFiber Arrangement/Agreement on Textiles and Clothing (MFA/ATC) regime in 25 EU countries, the United States, and Canada in 2005 to study the export destination choice of a sample of Chinese textile and apparel exporters which never exported to these countries before 2005. This exogenous shock has generated a large entry of firms in a set of potential new destinations and a substantial redistribution of quota rents towards new entrants into these markets (see Khandelwal et al., 2013). We can then study firms' subsequent export destination choices in other countries which were not directly affected by the lifting of the MFA quotas. As the timing of the lifting of the MFA quotas was exogenous to the firms, it helps us to overcome the endogeneity problem introduced by the dynamic nature of the firm's export destination choice.

As a first step, we use the lifting of the MFA quotas as a quasi-natural experiment to study the export destination choice in non-MFA countries employing a differences-in-differences estimator where we define as the treatment group the countries which are contiguous to a previously restricted MFA country. In order to exploit all the available information about firms' export history (not only firms' experience in previously restricted MFA countries), we use the quasi-natural experiment as an instrument to study the effect of previous export experience of firms on subsequent destination choices. Finally, using a dynamic panel estimator, we account for the endogeneity, the persistence, and true state dependence in export destination choices.

Our empirical strategy gauges the relative importance of the time-varying cross-country correlation of a firm's export destination choices resulting from its export history due to both geographical proximity of previous export destinations as well as cultural closeness measured by common language, common colonizer, or similar income levels. As we use reduced form regressions we do not rely on a specific channel imposed by an underlying structural model. Rather, we quantify the effects of any correlation across destination markets resulting from a firm's export history on the probability to export to a specific country, irrespective of whether it arises from the demand or supply side.

previous geographically close export destinations. However, she does not control for true state dependence nor firm-specific country fixed effects as we do.

Our paper provides causal evidence of the spatial correlation of export decisions at the firm level that has been put upfront by recent theoretical developments on export dynamics (see Albornoz et al., 2012; Nguyen, 2012, Morales et al., 2011, and Chaney, 2011). It could also contribute to explain the pattern of zero bilateral trade flows observed empirically (see Evenett and Venables, 2002). Understanding exporting firm behavior is also crucial from a policy perspective. If across-country path dependence in firm destination choices is important, it also has ramifications for trade liberalization policies. Then, reducing trade barriers between two countries can lead to more trade with other countries nearby than standard gravity forces would predict, even though they did not lower their trade barriers. This gives rise to externalities across countries.[4] Therefore, our research highlights an additional reason for potential efficiency increases in trade liberalization through policy coordination between countries.

The remainder of the paper is organized as follows: Section 1.2 describes the data set and our identification strategy. Section 1.3 presents our baseline empirical results. We start with a differences-in-differences (diff-in-diff) approach which investigates the impact of the lifting of the MFA quotas on the probability of exporting to countries which are contiguous to a previously restricted MFA country. We then investigate the impact of actual export experience in previous markets on a firm's destination choice. As our regressor of interest in the latter specification is potentially endogenous, we present instrumental variable regressions where we use the lifting of the MFA quotas as an instrument. Finally, we present dynamic panel specifications which use an alternative set of internal instruments to control for our potentially endogenous regressor of interest as well as the endogenous lagged dependent variable. Section 1.4 presents evidence at the firm-product-couple level. Section 1.5 presents robustness checks with respect to including lagged export values, competitor's success, excluding trade agents, state owned firms, foreign owned firms or processing trade firms. The last section concludes.

[4] For instance, Borchert (2008) finds that the growth of Mexican exports to Latin America was higher for products with a large reduction in the preferential U.S. tariff under NAFTA. Similarly, Molina (2010) identifies a strong positive effect of RTAs in promoting exports outside the bloc of liberalized countries. While it is difficult to explain these findings with standard trade models, they can easily be rationalized in the presence of firm-specific cross-country correlations in export destination choices.

1.2 Data and identification

1.2.1 Sample and dependent variable

To investigate the importance of spatial exporters, we use transaction level customs panel data on the universe of Chinese exporters for the years 2000 to 2006. We only keep products which fall in the Harmonized System (HS) chapters of textile and clothing products, i.e. chapters 50 to 63, as these are the products covered by the MFA regime. We aggregate all transactions of a firm in a country in one year into one observation. The sample is restricted to continuous exporters, i.e. firms that export at least to one country every year.[5] Specifically, we investigate the export destination choice between 150 non-MFA member countries of firms which did not export in any of the MFA restricted countries during the years 2000 to 2004.[6] Hence, our sample includes both firms that enter the MFA member countries after 2004 as well as those who export to other countries between 2000 and 2006. Overall, our sample is composed of 1,295 continuous exporters which never entered the MFA restricted countries before 2005.

Our dependent variable is the firm specific vector of export status $\mathbf{y}_{it} = (y_{i1t}, \ldots, y_{ijt}, \ldots, y_{i\mathcal{J}t})$ which indicates whether a firm i exports to a specific destination j in year t, which also defines the unit of observations. $\mathcal{J}$ is the number of non-MFA countries in our sample. In Table A.2 we present the descriptive statistics for our dependent variable.[7] 1.2 percent of our observed destination choices turn out to be positive. Hence, serving a foreign market is a rare event.

In order to shed light on the entry of firms into different markets in our sample of firms, we follow Eaton et al. (2011) and first assume that firms follow a common hierarchy, meaning that a firm that sells to the $k+1$st most popular export destination necessarily sells to the kth most popular destinations as well. We present the top seven export destinations of the Chinese exporters in our sample, excluding the MFA-restricted countries. In Table 1.1 we report the number of firms exporting to each of the seven most popular destinations, as well

5 This allows us to abstract from selection into exporting at the firm-extensive margin. See Das et al. (2007) for a structural model of selection into exporting.

6 The previously restricted MFA countries are the 25 EU countries as of 2005, the United States, and Canada. A comprehensive list of all non-MFA countries in our sample can be found in Appendix A.1.

7 As we use two lags in our dynamic panel specifications and have to skip one additional lag in order to ensure exogeneity of the instrument with the second lag, we use four years for all our specifications for comparability.

as the unconditional empirical probability of Chinese exporters selling there. We clearly see that common gravity variables, like distance and country size, matter.

Again following Eaton et al. (2011), in Table 1.2 we report strings of the top-seven destinations that obey a hierarchical structure, alongside the number of firms selling to each string. For example, the export string JPN means that the firm exports to Japan but to no other destination among the top 7 non-MFA destinations. Similarly, the string JPN-KOR means that the firm exports to Japan and South Korea but no other destination among the top 7 non-MFA destinations, and so forth. Overall, 66 percent (861/1295) of all firms in our sample adhere to the hierarchy given by the top seven non-MFA export destinations. Hence, about a third of the firms export to a different set of countries, implying a substantial amount of heterogeneity across firms in terms of the set of export destinations they serve. The column labeled "Independence" in Table 1.2 reports, based on the unconditional probabilities presented in Table 1.1, the number of firms selling to each hierarchical string assuming independence across destination choices of a firm. If a firm chose export destinations independently, the number of firms sticking to the common hierarchy would be 770, implying that only 59 percent (770/1295) would follow a common hierarchy. In the data, we observe 861 firms which stick to the common hierarchy, i.e. 12 percent more than what independence would imply. Hence, in our empirical specification we will have to take into account that export destinations within firms are clustered spatially, and that there is considerable heterogeneity in export destinations across firms. We therefore allow for time-invariant firm-specific attractiveness of export destinations.

1.2.2 Identification strategy

Under the MultiFiber Arrangement/Agreement on Textiles and Clothing (MFA/ATC) regime, restrictions were upheld on many products even after China acceded to the WTO on December 11th, 2001. On January 1st, 2005 the removal of import quotas lead to the entry of a large number of firms in the then 25 EU countries, the United States, and Canada.[8] Figure 1.1 shows the average number of exporters into these markets across all restricted HS-6 products. While around 100 to 150 firms had been exporting a restricted MFA product while the import restrictions were still upheld, this number jumped to more than 300 in 2005.

[8] See Harrigan and Barrows (2009), Brambilla et al. (2010), Upward et al. (2011), and Khandelwal et al. (2013).

Table 1.1: Chinese textile and apparel firms exporting to the seven most popular non-MFA destinations in 2006

Export destination	Number of exporters	Fraction of exporters
Japan (JPN)	973	0.751
South Korea (KOR)	328	0.253
Singapore (SGP)	81	0.063
Australia (AUS)	70	0.054
Vietnam (VNM)	62	0.048
Thailand (THA)	57	0.044
Malaysia (MYS)	46	0.036
All Chinese exporters*	1,295	

Notes: *in our sample. Table shows the seven most popular export destinations of the 1,295 textile and apparel firms in our sample excluding the 27 MFA/ATC restricted export destinations for the year 2006. The table follows closely Table I in Eaton et al. (2011). We describe the construction of the sample in detail in Section 1.2.1.

Table 1.2: Chinese textile and apparel firms exporting to strings of top-seven non-MFA destinations in 2006

	Number of Exporters	
Export String[a]	Data	Independence
JPN	676	565
JPN-KOR	175	191
JPN-KOR-SGP	8	13
JPN-KOR-SGP-AUS	1	1
JPN-KOR-SGP-AUS-VNM	0	0
JPN-KOR-SGP-AUS-VNM-THA	0	0
JPN-KOR-SGP-AUS-VNM-THA-MYS	1	0
Total	861	770

Notes: [a]The export string JPN means exporting to Japan but no other destination among the top 7 non-MFA destinations; JPN-KOR means exporting to Japan and South Korea but no other destination among the top 7, and so forth. The table follows closely Table II in Eaton et al. (2011). We describe the construction of the sample in detail in Section 1.2.1.

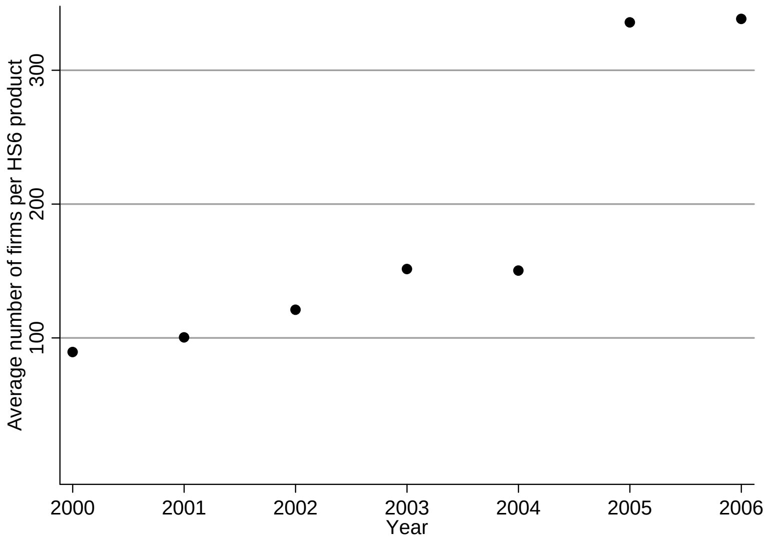
One possible reason behind the large and rapid entry of firms into MFA countries in 2005 can be seen in the fear that safeguard mechanisms could potentially re-introduce quotas. Actually, the EU countries, the United States, and Canada had product-specific safeguard mechanisms which were not phased out until 2008. The possible use of these safeguard measures was likely and it was unclear which products would be affected. This is corroborated by Figure 1.1 which shows that the average number of exporters across products did not increase in 2006 so that there is no evidence of a gradual entry of firms into the previously restricted MFA countries, at least on average. This can be explained by the new and transitional license system for textile exports that has been reintroduced in 2005 by the Chinese government. The intention was to limit the growth of Chinese exports of MFA products for the years 2006 to 2008. Looking back, the restrictions imposed in 2005 were by and large ineffective. However, the new restrictions had an impact on the growth of Chinese textile exports for 2006 to 2008.[9]

The lifting of the MFA quotas in 2005 exogenously changed the potential profitability of exporting to the previously restricted MFA countries. New entrants could reap part of the quota rents which previously accrued to those firms with an export license, leading to the increase in the number of firms in the EU, the United States, and Canada. If firms are 'spatial exporters', this change should have influenced the subsequent export destination choices in non-MFA countries. The same firms which quickly entered the previously restricted MFA countries *for the first time* could then potentially learn about other profitable export opportunities in countries which are geographically or culturally related to the previously restricted MFA countries.

1.3 Specifications

We have now described our identification strategy in general terms. It is compatible with several complementary empirical specifications which rely on different assumptions about the data-generating process. Specifically, we will use a differences-in-differences (diff-in-diff)

[9] China and the EU agreed in June 2005 to re-impose quotas on some products. Despite the implementation of a new license system China did not restrict the number of the licenses nor the volume of exports. As a reaction, EU retailers ordered large amounts of Chinese textile products before the quota implementation. Only two months after the signing of this agreement import quotas were exhausted and 75 million items of textile and clothing products were stuck in European ports (see Brambilla et al., 2010; Buckley, 2005, and Wikipedia, 2013). A diplomatic solution was reached at the beginning of September 2005 putting an end to a situation the UK press called the "Bra Wars" (see e.g. White and Gow, 2005 and Wikipedia, 2013).



Notes: Yearly average number of firms exporting to one EU country, the United States or Canada for HS 6-digit products for which the quota fill rate was higher than 90 percent.

Figure 1.1: Average number of exporting firms to one EU country, the United States or Canada per restricted MFA 6-digit product

strategy, (panel) instrumental variable regressions and dynamic panel estimations. This multitude of specifications provides robust evidence for spatial exporters. We will next discuss in turn our specifications and the corresponding results.

1.3.1 Differences-in-differences

Viewing the removal of the quota restrictions as a quasi-natural experiment, it seems natural to start with a differences-in-differences (diff-in-diff) specification.

MFA restrictions were removed January 1st, 2005. This lifting opened up new potential export markets but was not influenced by the decisions of individual firms and thus exogenous at the firm level. Beginning from this date, firms in our sample were able to enter the previously restricted MFA countries for the first time. There they could potentially acquire information about contiguous export markets. Therefore, firms should export more to destinations which are contiguous to MFA countries after the removal of the MFA restrictions. Hence, our treatment indicator C_j is defined at the country-level.[10] It is a dummy variable indicating whether a country j is contiguous to a MFA-restricted country. This also renders our treatment exogenous to the firm's choices, as the set of MFA restricted countries is the same for all firms. Similar to Morales et al. (2011), we assume a one year lag to quantify

[10] We therefore use standard errors clustered at the country-level following the recommendation for differences-in-differences estimates by Bertrand et al. (2004).

'spatial exporters', reflecting the fact that the learning or product adaptation processes of the firm take time. Hence, we define the year 2006 as our post-treatment period. $y2006_t$ is the corresponding dummy variable for the year 2006. The treatment effect, δ, measures whether firms export more frequently to countries that are contiguous to previously restricted MFA countries in 2006 and is captured by the interaction term of $y2006_t$ and C_j.

Specifically, our first empirical specification is therefore given by

$$y_{ijt} = \delta(y2006_t \times C_j) + \theta_{ij} + \theta_t + \epsilon_{ijt}, \tag{1.1}$$

where y_{ijt} is a dummy variable indicating whether a firm i exported to country $j \in \mathcal{J}$ in year t, where $\mathcal{J}$ is the set of non-MFA countries. We also introduce θ_{ij}, a firm-destination fixed effect, and θ_t, a year fixed effect. ϵ_{ijt} is a remainder error term. Note that this regression is equivalent to a diff-in-diff specification as the year and firm-destination fixed effects control for the treatment period as well as the treatment group dummies. We estimate specification (1.1) with ordinary least squares which leads to a linear probability model due to our binary dependent variable.[11]

The firm-destination fixed effects capture all country-firm characteristics that do not change over the considered time period. This includes time-constant destination-specific variables generally known to influence bilateral trade flows from the gravity literature such as market size, overall remoteness of a country (multilateral resistance terms), and trade costs. Crucially, it also controls for time-constant firm-specific heterogeneity such as productivity, quality, labor costs, and assortative matching of workers. For example, a firm might employ managers with specific language skills which influence the firm's export destination choice.[12] θ_t captures the general time trend in the empirical probability of exporting to a country.

We expect δ to be positive if firms are spatial exporters. δ is identified by firms which start to export to a country in 2006 which is contiguous to a MFA-restricted country. A positive effect can stem from two sources: 1.) A country j which is contiguous to a

[11] As we are only interested in average effects and not in predictions for individual firms and given the high number of fixed effects, we stick to the linear probability model, see Winkelmann and Boes (2009). As we also control for lagged endogenous variables in later specifications, we can extend our regression framework by using a linear dynamic panel estimator in a straight-forward way, simplifying the interpretation and comparison of results across our different specifications.

[12] In a strict sense, some gravity variables will change over time (such as market size and the multilateral resistance terms). However, note that we only consider one post-treatment year (2006). Hence, to bias our results the gravity variables would have to be considerably different in 2006 and at the same time this change would have to be correlated according to the same spatial pattern as our treatment.

previously MFA-restricted country now is more attractive as a potential export destination as firms can now reap, in addition to the direct profits of selling in j, the expected profit from gaining some information about the previously restricted MFA countries themselves, *irrespective of whether the firm has exported to a MFA-restricted country or not.* 2.) Firms which actually did export to a MFA-restricted country in 2005 *for the first time* and gained knowledge about potential business opportunities in contiguous country j. Note that firms which stop exporting to country j in 2006 decrease the estimate of δ (and may even render the coefficient negative).

Table 1.3 reports estimates of the diff-in-diff specification as given in equation (1.1). Specifications I to VI give the estimated treatment effects for exporting to a contiguous MFA country one year after the lift of the quota restrictions for different definitions of contiguity. A firm's destination choice can be correlated not only in markets which are geographically proximate to its previous export destinations but also in markets which share some other form of closeness. Specifically, we define contiguity according to whether the countries share a common border, a common language, a common colonizer, a common income group, or are located on the same continent using data provided by CEPII, see Mayer and Zignago (2011). Therefore, our concept of space is general and can refer to geographic as well as cultural cross-country correlation in export destination choices.

Appendix A.2 gives a detailed description of the construction of our contiguity variables. Table A.2 contains summary statistics for all variables. In our sample, e.g. 1.2 percent of all observations are countries which share a common border with an export destination of the same firm the year before.

Looking at specification I, contiguity is defined according to whether countries share a common border. The coefficient estimate of 0.003 implies an average increase of 0.3 percentage points in the probability of choosing a new export destination that is contiguous to a previously restricted MFA country in 2006. This effect may sound small. We therefore compare this marginal effect to the observed empirical probability of a firm exporting to a particular country in our sample reported. We report these empirical probabilities in Table A.1. For example, this implies about a 14 percent (0.003/0.022) increase in the probability of a firm exporting to Russia in 2006, as Russia shares a common border with Finland, an MFA country.[13]

[13] Note that we do not compare our estimates to the unconditional observed frequency of exporting to a country (the mean of our dependent variable, 0.012), as this frequency ignores the spatial correlation of

Specifications II to V run separate regressions where we construct our contiguity measure according to whether countries share the same language (specification II), whether countries have common colonial ties (specification III), whether countries are in the same income group (specification IV), or whether countries are located on the same continent (specification V). Evidently, especially space in the geographic sense (common border and common continent) plays a significant role in firms' export location choice. We do not find evidence for other definitions of contiguity, like common language, common colonizer or common income group, as important determinants for spatial exporters.

In column VI, we include all different contiguity measures at the same time to gauge the relative importance of the different measures. The marginal effects are hardly affected by conditioning on all other contiguity measures. Also the significance stays by and large the same. This hints at the orthogonality of the different contiguity measures and lends credibility to the treatment effects given in columns I to V.

In the specification given in equation (1.1) we do not condition on whether the firm has exported to a previously restricted MFA country. Hence, we identify a combination of the effects 1.) and 2.) mentioned before. Whereas 1.) increases the profitability of a destination only due to the option value of going to a MFA restricted country and therefore for all firms in our sample without any action from the firm[14], 2.) directly measures actually occurred spatial exporting only for firms that did export to an MFA restricted country first and afterwards to a contiguous one.

While Table 1.3 provides a first step towards evidence for spatial exporters, an interesting question is to identify how past learning from a country for the first time affects future export decisions, i.e. focus on the second source from above. This is what we do next.

1.3.2 Fixed effects regression taking into account firm-level history

Until now, we only focused on those countries which were contiguous to previously restricted MFA countries and neglected the impact of a firm's previous export history. In order to capture spatial exporting which takes into account firm-level history, we construct our

exports due to standard gravity forces such as country size and distance between origin and destination countries. Russia is the first country in our list of most frequent export destinations which shares a common border with an MFA country. Also note that the empirical probabilities given in Table A.1 are slightly different to those reported in Table 1.1 as we use all years in our regression data set to calculate the empirical probabilities.

[14] Note that this effect is heterogeneous across firms as it depends on a firm's export history.

Table 1.3: Diff-in-diff

	I	II	III	IV	V	VI
$y2006_t \times C_j$ defined according to...						
common border	0.003***					0.002***
	(0.001)					(0.001)
common language		0.000				0.000
		(0.000)				(0.000)
common colonizer			-0.001			-0.000
			(0.000)			(0.000)
common income group				-0.000		-0.000
				(0.000)		(0.000)
common continent					0.001***	0.001**
					(0.000)	(0.000)
Observations	777,000	777,000	777,000	777,000	777,000	777,000
# of firms	1,295	1,295	1,295	1,295	1,295	1,295

Notes: The dependent variable is y_{ijt} which is a dummy variable indicating whether a firm i exported to country j in year t. All regressions include firm-destination fixed effects, as well as year dummies and a constant (all not reported). Standard errors are in parentheses. All regressions use robust standard errors clustered at the country level to take into account that the regressor only varies at the country level following the suggestion for differences-in-differences estimates by Bertrand et al. (2004). *, ** and *** denote significance at the 10%-, 5%- and 1%-level, respectively.

contiguity measure, $N_{ij,t-1}$, which measures the number of countries which are contiguous to country j and to which firm i has exported in $t-1$ for each firm i and destination j. As the set of the previous export destinations is firm-specific, so are the contiguity variables. Specifically, $N_{ij,t-1} = \mathbf{w}_j'\mathbf{y}^*_{i,t-1}$, where $\mathbf{y}^*_{i,t-1}$ is the $(\mathcal{N} \times 1)$ vector of the export indicators for firm i in $t-1$ whose typical element $y_{i\ell,t-1}$ is 1 if firm i exported to country ℓ in year $t-1$, and zero otherwise. For the construction of our explanatory variable, $N_{ij,t-1}$, we use a set of $\mathcal{N} = 177$ countries, *including* the previously restricted MFA countries. In our regression sample, however, we continue to investigate the choice between $\mathcal{J} = 150$ non-MFA countries as in the previous section. $\mathbf{w}_j$ is the jth row of $\mathbf{W}$, a $(\mathcal{N} \times \mathcal{N})$ contiguity matrix. The typical entry $w_{\ell m}$ of $\mathbf{W}$ is 1 if countries ℓ and m are contiguous, and zero otherwise.[15]

As with C_j, we measure $N_{ij,t-1}$ by defining contiguity in terms of the countries sharing a common border, sharing a common language, sharing a common colonizer, being in a common income group, or being located on the same continent. For example, $N_{ij,t-1} = 2$ measured in terms of common border means that for firm i, country j shares a common border with two countries to which firm i has exported in $t-1$.

To take into account whether a firm actually has exported to a country in the previous year, we run the following regression:

$$y_{ijt} = \delta \mathbb{I}(N_{ij,t-1} > 0)_{ijt} + \theta_{ij} + \theta_t + \epsilon_{ijt}, \tag{1.2}$$

where $\mathbb{I}$ is the indicator function taking value one if $N_{ij,t-1} > 0$. In this regression, δ now quantifies the effect of actual experience in a previous export destination on future export decisions to contiguous countries. We expect δ to be positive if previous export experience from contiguous countries matters. Note that in contrast to $y2006_t \times C_j$, $\mathbb{I}(N_{ij,t-1} > 0)_{ijt}$ varies at the firm-level.

Table 1.4 gives the result for specification (1.2) and is organized in the same way as Table 1.3. Column I shows that the probability of exporting to a country increases by 1.4 percentage points if the firm previously exported to an export destination with a common

[15] In principle, one could also think about using $\mathbf{y}^{MFA}_{i,t-1}$ to construct $N_{ij,t-1}$, whose dimension is $(\mathcal{N} \times 1)$ and whose typical element $y^{MFA}_{i\ell,t-1}$ is 1 if firm i exported to country ℓ in $t-1$, and this country is an MFA country, and zero otherwise. By using $\mathbf{y}^*_{i,t-1}$ instead of $\mathbf{y}^{MFA}_{i,t-1}$ to construct $N_{ij,t-1}$, we also count previous export destinations of a firm which are not previously restricted MFA countries. We reran all our specifications using this alternative regressor but results hardly changed. Note, however, that focusing on $\mathbf{y}^{MFA}_{i,t-1}$ would potentially bias our coefficient estimates as $\mathbf{y}^{MFA}_{i,t-1}$ sets all those elements of $\mathbf{y}^*_{i,t-1}$ equal to 0 which identify positive non-MFA country export flows.

border. Is this effect large or small? We again compare this marginal effect to the empirical probability of a firm exporting to a particular country in our sample reported in Table A.1. Given these empirical probabilities, this implies e.g. a 20 percent increase in the probability of a firm exporting to Singapore when it has previously exported to Malaysia.[16] This effect is larger than the effect identified in Table 1.3 because we now focus on source 2.), i.e. the effect of actual export experience in contiguous countries.

Again, the effect of sharing a common border is the largest and most significant effect. Also sharing a common language or colonial ties are significant, albeit with smaller magnitudes. For example, the probability of exporting to Australia increases by about 4 percent (0.002/0.054) if the firm has previously exported to Great Britain (or some other English-speaking country). Similarly, the probability of exporting to India increases by about 11 percent (0.002/0.019) if the firm has previously exported to Great Britain with which it shares a language. In column VI we again find that effects are quantitatively very similar when conditioning on all different dimensions of spatial exporters jointly.

Similarly, we can also estimate the impact of an increase in the number of previous contiguous export destinations by omitting the indicator function from equation (1.3), i.e.:

$$y_{ijt} = \delta N_{ij,t-1} + \theta_{ij} + \theta_t + \epsilon_{ijt}. \tag{1.3}$$

Table 1.5 reports the estimates. Results are virtually unchanged, with sharing a common border remaining the regressor with the largest point estimate. The slight change in the specification implies that the probability of exporting to a country that shares a common border with a previous export destination increases by 1.2 percentage points if the firm actually exports to one additional contiguous country in the previous year.

A problem of regressions (1.2) and (1.3) is that, contrary to regression (1.1), now the regressor of interest is potentially endogenous as firms may anticipate that they may learn from previous export destinations and potentially choose their export destinations accordingly. We will therefore present (panel) instrumental variable regressions in the next subsection.

[16] Note that Japan and South Korea, our most frequent export destinations, do not have a common border with any country (the Democratic People's Republic of Korea is not included in our data set). We therefore chose Singapore, the third most frequent export destination. Malaysia shares a common border with Singapore.

Table 1.4: Fixed effects regression taking into account firm-level history—dummy

	I	II	III	IV	V	VI
$\mathbb{I}(N_{ij,t-1} > 0)_{ijt}$ defined according to...						
common border	0.014***					0.014***
	(0.003)					(0.003)
common language		0.002***				0.002***
		(0.001)				(0.001)
common colonizer			0.002**			0.001
			(0.001)			(0.001)
common income group				0.001		0.000
				(0.001)		(0.001)
common continent					0.001	-0.000
					(0.001)	(0.001)
Observations	777,000	777,000	777,000	777,000	777,000	777,000
# of firms	1,295	1,295	1,295	1,295	1,295	1,295

Notes: The dependent variable is y_{ijt} which is a dummy variable indicating whether a firm i exported to country j in year t. All regressions include firm-destination fixed effects, as well as year dummies and a constant (all not reported). Standard errors are in parentheses. All regressions use robust standard errors clustered at the firm level to take into account the potential autocorrelation in the export destination choice at the firm level. *, ** and *** denote significance at the 10%-, 5%- and 1%-level, respectively.

Table 1.5: Fixed effects regression taking into account firm-level history—N

	I	II	III	IV	V	VI
$N_{ij,t-1}$ defined according to...						
common border	0.012***					0.010***
	(0.003)					(0.003)
common language		0.001**				0.000
		(0.001)				(0.000)
common colonizer			0.003***			0.002*
			(0.001)			(0.001)
common income group				0.002**		0.001
				(0.001)		(0.001)
common continent					0.001*	0.000
					(0.001)	(0.001)
Observations	777,000	777,000	777,000	777,000	777,000	777,000
# of firms	1,295	1,295	1,295	1,295	1,295	1,295

Notes: The dependent variable is y_{ijt} which is a dummy variable indicating whether a firm i exported to country j in year t. All regressions include firm-destination fixed effects, as well as year dummies and a constant (all not reported). Standard errors are in parentheses. All regressions use robust standard errors clustered at the firm level to take into account the potential autocorrelation in the export destination choice at the firm level. *, ** and *** denote significance at the 10%-, 5%- and 1%-level, respectively.

1.3.3 Instrumental variable regressions

In order to account for the potential endogeneity of our regressor $\mathbb{I}(N_{ij,t-1} > 0)_{ijt}$, we instrument it by the exogenous regressor of interest from regression (1.1), $y2006_t \times C_j$. The exogeneity of our instrumental variable is again justified by the fact that the instrument is a country-specific variable and is not influenced by firm decisions. Still, our instrument is relevant as the instrument and the potential endogenous regressor are correlated by construction: C_j indicates countries contiguous to (previously) MFA restricted countries and $N_{ij,t-1}$ is positive if a firm exports to at least one country. As the MFA restricted countries in sum make up a large share of the world market, it is very likely that $N_{ij,t-1} > 0$ if $C_j = 1$. In addition, the regression results from the diff-in-diff specifications clearly show the relevance of the proposed instrument. For our estimation, we use the two-stage least-squares within panel instrumental variables estimator which includes the full set of firm-country fixed effects as used in the previous specification.

Comparing the results from Table 1.4 which assumes that $\mathbb{I}(N_{ij,t-1} > 0)_{ijt}$ is exogenous with the instrumental variable regressions that allow $\mathbb{I}(N_{ij,t-1} > 0)_{ijt}$ to be endogenous given in Table 1.6 shows that there is no qualitative change in our results. However, the size of the effect of contiguity is approximately seven times larger. Again, sharing a common border has the largest effect (point estimate of 0.104) and only the geographical contiguity measures turn out to be statistically significant. Results also remain largely unchanged when including all contiguity measures simultaneously (see column VI in Table 1.6).

Table 1.7 reproduces Table 1.5 but instruments $N_{ij,t-1}$ with $y2006_t \times N_j$.[17] Comparing results shows that the effects of geographical contiguity (common border and common continent) are about seven times larger. Hence, our estimate in specification I implies that the probability of exporting to a country that shares a common border with a previous export destination increases by 8 percentage points if the firm actually exports to one additional contiguous country in 2005.[18]

[17] We use $y2006_t \times N_j$ as this has the same kind of variation at the country level as our potentially endogenous regressor, $N_{ij,t-1}$. In principle, we could also again instrument by $y2006_t \times C_j$, or even use $y2006_t \times N_j$ in our diff-in-diff specification. These choices hardly matter for our results. These estimates are available from the authors upon request.

[18] We also experimented with the year 2004 and 2005 to construct our instrument, finding similar but larger effects. The estimate for common border for defining the treatment period to begin in the years 2004 and 2005 are 0.311 and 0.225 for $\mathbb{I}(N_{ij,t-1} > 0)_{ijt}$ and 0.268 and 0.187 for $N_{ij,t-1}$, respectively. This is consistent with the argument that by focusing on previous years, we would get an upward biased estimate of the effect of contiguity as exporting to contiguous countries would be confounded by other reasons. By using the lifting of the MFA restrictions, we likely minimize these other effects.

Table 1.6: Instrumental variable regressions—dummy

	I	II	III	IV	V	VI
$\mathbb{I}(N_{ij,t-1} > 0)_{ijt}$ defined according to...						
common border	0.104***					0.089**
	(0.029)					(0.045)
common language		0.004				0.005
		(0.004)				(0.012)
common colonizer			-0.128			-0.008
			(0.833)			(0.768)
common income group				0.043		0.154
				(0.280)		(0.364)
common continent					0.009***	-0.001
					(0.002)	(0.033)
Observations	777,000	777,000	777,000	777,000	777,000	777,000
# of firms	1,295	1,295	1,295	1,295	1,295	1,295

Notes: The dependent variable is y_{ijt} which is a dummy variable indicating whether a firm i exported to country j in year t. All regressions include firm-destination fixed effects, as well as year dummies and a constant (all not reported). We use the two-stage least-squares within panel instrumental variables estimator where we instrument the endogenous regressor by $y2006_t \times C_j$. Standard errors are in parentheses. All regressions use robust standard errors clustered at the firm level to take into account the potential autocorrelation in the export destination choice at the firm level. *, ** and *** denote significance at the 10%-, 5%- and 1%-level, respectively.

Table 1.7: Instrumental variable regressions—N

	I	II	III	IV	V	VI
$N_{ij,t-1}$ defined according to...						
common border	0.080***					0.075***
	(0.029)					(0.028)
common language		-0.000				0.002
		(0.001)				(0.002)
common colonizer			-0.027			-0.016
			(0.057)			(0.100)
common income group				-0.003		-0.004*
				(0.002)		(0.002)
common continent					0.005***	0.003
					(0.001)	(0.003)
Observations	777,000	777,000	777,000	777,000	777,000	777,000
# of firms	1,295	1,295	1,295	1,295	1,295	1,295

Notes: The dependent variable is y_{ijt} which is a dummy variable indicating whether a firm i exported to country j in year t. All regressions include firm-destination fixed effects, as well as year dummies and a constant (all not reported). We use the two-stage least-squares within panel instrumental variables estimator where we instrument the endogenous regressor by $y2006_t \times N_j$. Standard errors are in parentheses. All regressions use robust standard errors clustered at the firm level to take into account the potential autocorrelation in the export destination choice at the firm level. *, ** and *** denote significance at the 10%-, 5%- and 1%-level, respectively.

Even though we rely on panel data for our regressions so far, we have, until now, ignored the persistence and state dependence in the export status of firms. We turn to this issue in the next section.

1.3.4 Dynamic panel results taking into account state dependence

At least since Roberts and Tybout (1997) and Das et al. (2007) it is well known that whether a firm has exported in the previous period is highly correlated with its current export status. This evidence is provided at the firm level, irrespective of the variation of export destinations within a firm across time. Hence, it is based on persistence at the firm level export status, not at the firm-destination level. In principle, it is possible that this persistence is also evident at the firm-destination level. And indeed in our data set, the correlation between our dependent variable and its one year lag is 0.75.

One can distinguish between two major sources of this observed persistence. First, there maybe some unobserved time-invariant firm-destination component which determines whether a firm enters a specific destination. Second, there can be true state dependence, i.e. the previous export history of a firm in a specific country drives future export destination choices. In other words, export history in export destination choice matters.

Whereas the first persistence is captured in our specification by the firm-destination fixed effect θ_{ij}, we did not properly account for potential true state dependence in our estimations so far. As has been demonstrated by Nickell (1981), fixed effect estimators are biased in the presence of true state dependence. How does this affect our estimates? In our setting, consider a firm which exports to both Singapore and Malaysia in 2005 and 2006. Then, when not including lags of the dependent variable, our regressor of interest explains the firm's exporting behavior in Malaysia by its previous export experience in Singapore and vice versa.[19] To control for this confounding factor, avoid the Nickel bias, and account for the high persistence in our dependent variable, we employ the dynamic panel estimator from Blundell and Bond (1998).

[19] Note that for firms which continuously export to both destinations in all years included in the sample, this will be captured by the firm-destination fixed effects. However, firm-destination fixed effects will not cover this persistence for intermittent exporters.

Specifically, our dynamic panel specification including lags of the dependent variable is given by

$$y_{ijt} = \phi_1 y_{ij,t-1} + \phi_2 y_{ij,t-2} + \delta \mathbb{I}(N_{ij,t-1} > 0)_{ijt} + \theta_{ij} + \theta_t + \epsilon_{ijt}. \tag{1.4}$$

We include two lags of the dependent variable as Roberts and Tybout (1997) show that typically two lags have a significant and decaying impact on the export decision of a firm.[20]

Table 1.8 presents our dynamic panel estimates for specification (1.4), i.e. using dummy variables to indicate contiguity between a destination and previous export destinations. The table is organized in the same way as the previous tables but includes also the estimates for the two lags of the dependent variable. As can clearly be seen, we find true state dependence in all our specifications even at the firm-destination level. Our result that sharing a common border is the largest and most significant contiguity effect is corroborated by the dynamic panel estimates. Note that the dynamic panel estimator allows us to treat our contiguity variable as predetermined, consistent with the fact that lagged values of our regressor of interest can not be changed by the firm in the current period but future values may be adjusted by the firm, as stressed by the mechanisms in Morales et al. (2011), Albornoz et al. (2012), and Nguyen (2012). Sharing a common language, colonial ties or being in the same income group are all significant but have smaller effects than common border.

Column VI presents results when we include all regressors at the same time. Sharing a common border still has a similar impact on the probability of exporting to a country compared to the specification in column I. The same holds for the two countries sharing a common language or being in the same income group. Interestingly, sharing a common colonizer has a significant and positive effect in column III. This effect vanishes, however, in column VI. Being on the same continent even turns out to have an albeit small but significantly negative effect. Note, however, that a country which is located on the same continent very likely also shares a common border or a common language with a previous export destination. In other words, there is a high correlation between our different contiguity measures conditioning on true state dependence and firm-destination fixed effects. We again compare our estimated marginal effect to the empirical probability of a firm exporting to a

[20] While most applications of dynamic panel estimators only include one lag, Cameron and Trivedi (2005) show that the dynamic setting can easily be extended to more lags. We also experimented with including only one lag. However, these specifications where clearly rejected by model specification tests such as the autocorrelation tests or Sargan test.

particular country from Table A.1. Given the empirical probabilities, this implies e.g. a 33 percent (0.023/0.070) increase in the probability of a firm exporting to Singapore when it has previously exported to Malaysia.

We use the Sargan test and a test for the first and second order autocorrelation of the residuals to test our specifications. The bottom three lines of Table 1.8 report p-values of the respective tests. While we find evidence for first order autocorrelation in the residuals across all specifications, we do not find evidence for second order autocorrelation, implying that the moment conditions used for the dynamic panel estimator are valid. We re-run our model assuming homoskedastic error terms in order to calculate a Sargan overidentification test, as this test is only valid under homoskedasticity. In most specifications also the Sargan test does not reject our model specification. Only in specifications VI the Sargan test rejects the validity of our internal instruments. Overall, our results suggest a proper model specification.

We again can use the number of contiguous export destinations as an alternative regressor. Hence, the dynamic panel specification in this case is given by

$$y_{ijt} = \phi_1 y_{ij,t-1} + \phi_2 y_{ij,t-2} + \delta N_{ij,t-1} + \theta_{ij} + \theta_t + \epsilon_{ijt}. \tag{1.5}$$

Results, which are reported in Table 1.9, are hardly affected by this different measure of contiguity. Again, we find strong evidence for true state dependence, and again sharing a common border has the largest impact on the destination choice. Our specification tests for first and second order autocorrelation again do not invalidate our regressions. However, the Sargan test does reject the validity of our internal instruments in specifications IV-VI. Note however, that this test is only valid under homoskedastic errors, which is normally violated in trade data (see for example Santos Silva and Tenreyro, 2006).

1.4 Multi-product firms

Until now our analysis considered an export destination as contiguous if the firm previously exported any product to a contiguous market. It is well known that a substantial fraction of firms produce and export multiple products, and that multi-product firms make up for the majority of sales in a given industry, see Arkolakis and Muendler (2010) and Bernard et al. (2010a). In our sample, 56 percent of firms export in more than one HS-6 product category. If there exists within-firm correlation of export destination choices between products, then

Table 1.8: Dynamic panel estimates—dummy

	I	II	III	IV	V	VI
$\mathbb{I}(N_{ij,t-1} > 0)_{ijt}$ defined according to...						
common border	0.024***					0.023***
	(0.004)					(0.004)
common language		0.003***				0.003***
		(0.001)				(0.001)
common colonizer			0.003***			0.001
			(0.001)			(0.001)
common income group				0.003***		0.002***
				(0.001)		(0.001)
common continent					-0.001	-0.005***
					(0.001)	(0.001)
$y_{ij,t-1}$	0.344***	0.344***	0.343***	0.343***	0.344***	0.347***
	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)
$y_{ij,t-2}$	0.077***	0.078***	0.076***	0.078***	0.076***	0.074***
	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)
Observations	777,000	777,000	777,000	777,000	777,000	777,000
# of firms	1,295	1,295	1,295	1,295	1,295	1,295
AR(1)	0	0	0	0	0	0
AR(2)	.821	.822	.889	.803	.929	.950
Sargan	.383	.604	.170	.406	.061	.008

Notes: The dependent variable is y_{ijt} which is a dummy variable indicating whether a firm i exported to country j in year t. All regressions include firm-destination fixed effects, as well as year dummies and a constant (all not reported). Standard errors are in parentheses. All regressions use robust standard errors and treat the lags of the dependent variable as well as the regressors of interest as predetermined. We use the two-step system GMM estimator from Blundell and Bond (1998) and, due to the two-step estimation, we use the Windmeijer (2005) finite sample correction for the standard errors. *, ** and *** denote significance at the 10%-, 5%- and 1%-level, respectively. The values reported for AR(1) and AR(2) are the p-values for first and second order autocorrelated disturbances in the first differences equations. The row for the Sargan reports the p-values for the null hypothesis of validity of the overidentifying restrictions and can only be computed assuming homoskedasticity. To report this statistic, we re-estimate the model accordingly.

Table 1.9: Dynamic panel estimates—N

	I	II	III	IV	V	VI
$N_{ij,t-1}$ defined according to…						
common border	0.023***					0.013***
	(0.004)					(0.004)
common language		0.002***				-0.001**
		(0.000)				(0.000)
common colonizer			0.004***			0.000
			(0.001)			(0.001)
common income group				0.006***		0.005***
				(0.000)		(0.001)
common continent					0.004***	0.002***
					(0.000)	(0.000)
$y_{ij,t-1}$	0.344***	0.348***	0.343***	0.338***	0.342***	0.356***
	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)
$y_{ij,t-2}$	0.077***	0.081***	0.079***	0.081***	0.084***	0.098***
	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)
Observations	777,000	777,000	777,000	777,000	777,000	777,000
# of firms	1,295	1,295	1,295	1,295	1,295	1,295
AR(1)	0	0	0	0	0	0
AR(2)	.825	.673	.738	.582	.535	.212
Sargan	.346	.057	.081	.010	.003	0

Notes: The dependent variable is y_{ijt} which is a dummy variable indicating whether a firm i exported to country j in year t. All regressions include firm-destination fixed effects, as well as year dummies and a constant (all not reported). Standard errors are in parentheses. All regressions use robust standard errors and treat the lags of the dependent variable as well as the regressors of interest as predetermined. We use the two-step system GMM estimator from Blundell and Bond (1998) and, due to the two-step estimation, we use the Windmeijer (2005) finite sample correction for the standard errors. *, ** and *** denote significance at the 10%-, 5%- and 1%-level, respectively. The values reported for AR(1) and AR(2) are the p-values for first and second order autocorrelated disturbances in the first differences equations. The row for the Sargan reports the p-values for the null hypothesis of validity of the overidentifying restrictions and can only be computed assuming homoskedasticity. To report this statistic, we re-estimate the model accordingly.

a firm may enter a new export market with a product when it has previously sold a different product in a contiguous market.

There are both supply and demand side reasons which can explain this type of economies of scope. When costs for product adaptation are lower for other products within a firm once they have been incurred for a specific market and product, the additional cost of adapting the product for a similar market may be lower. In addition, when a firm sells its products under a single brand in order to benefit from brand loyalty of consumers, successful exports in one product category provide information about likely profitable exports across the whole product mix of a firm's brand.

To take into account these effects, we modify our dynamic panel specification given in equation (1.4) as follows:

$$\begin{aligned} y_{ijt} = \quad & \phi_1 y_{ij,t-1} + \phi_2 y_{ij,t-2} + \delta_1 \mathbb{I}(N^{sameproduct}_{ij,t-1} > 0)_{ijt} \\ & + \delta_2 \mathbb{I}(N^{otherproducts}_{ij,t-1} > 0)_{ijt} + \theta_{ij} + \theta_t + \epsilon_{ijt}, \end{aligned} \tag{1.6}$$

where i now denotes the firm-product couple at the HS6-digit product category and no longer a single firm, and where $N^{sameproduct}_{ij,t-1}$ is the number of contiguous export destinations where the firm has exported the same product before and $N^{otherproducts}_{ij,t-1}$ is the number of contiguous export destinations where the firm has previously exported products from other HS-6 digit categories. Note that θ_{ij} now captures all unobserved time-invariant firm-product-destination characteristics.

As we now focus on firm-product couples, we use all firm-product couples which have never entered the previously restricted countries before 2005. In our sample, there are 6,573 firm-product couples of 1,965 firms, implying that a firm exports about 3.3 products on average.[21] In our previous regression, we kept only those firms that never exported any product into the previously restricted MFA countries. As firms may have entered into the previously restricted MFA countries only with a subset of their products, we now keep all other firm-product couples where we do not observe exports into the previously restricted MFA countries before 2005. Hence, there are more firms in our multi-product sample than in the previous regressions.[22]

[21] Descriptive statistics of the firm-product couple level data set can be found in Table A.3 in the Appendix.

[22] Imagine a firm which has exported panties to an MFA country in 2004 but not bras. In our firm level regressions, this firm is dropped from the sample. However, in our multi-product regressions we will keep the bra observations.

In Table 1.10, we present the results for the multi-product specification. Even at the firm-product couple level, we find a very similar pattern of true state dependence in the export status with significant but decaying effects of the two lags of the dependent variable. Overall we find hardly any evidence that exporting to a country is more likely after a previous entry into contiguous export destinations across products, the only exception being the common border coefficient in specifications I and VI. We find that the probability of choosing a country increases by 1.8 percentage points when a firm previously has exported the same HS-6 product to a contiguous country, and by 0.2 percentage points if it has exported other HS-6 products. For the other contiguity measures, our results indicate no (economically) significant effect of across product learning for sharing a common language with at least one previous export destination or being in the same income group. Interestingly, we find small significant negative effects for common colonizer and common continent. This may hint at a potential for diversification in a firm's export portfolio by selling different products to different contiguous countries when they share a colonial past or are located on the same continent. Note that our results for the same HS-6 product are in line with the effects found at the firm-level in Section 1.3.4. As found in the firm level regressions, when including all different contiguity measures at the same time, we find very similar marginal effects (see specification VI).

Concerning the specification tests, we find that the tests for autocorrelation in the disturbances in first differences indicate a well-specified model. However, contrary to the firm-level regressions, the Sargan test now rejects the validity of the overidentifying restrictions. Remember, however, that this test assumes homoskedasticity and that the total number of observations has increased by a factor of more than five. With nearly four million observations based on 6,573 firm-product couples, the amount of heteroscedasticity is substantially higher by construction as compared to the firm-level regressions. This may very well explain the rejection of the overidentifying restrictions by the Sargan test based on the assumption of homoskedasticity.

In Table 1.11 we present multi-product regressions with the number of contiguous export destinations as an alternative regressor:

$$\begin{aligned} y_{ijt} = \quad & \phi_1 y_{ij,t-1} + \phi_2 y_{ij,t-2} + \delta_1 N_{ij,t-1}^{sameproduct} \\ & + \delta_2 N_{ij,t-1}^{otherproducts} + \theta_{ij} + \theta_t + \epsilon_{ijt}. \end{aligned} \tag{1.7}$$

Table 1.10: Multi-product firms: dynamic panel estimates—dummy

		I	II	III	IV	V	VI
$\mathbb{I}(N_{ij,t-1} > 0)_{ijt}$ defined according to...							
common border	$\mathbb{I}(N^{sameproduct}_{ij,t-1} > 0)_{ijt}$	0.018***					0.018***
		(0.002)					(0.002)
	$\mathbb{I}(N^{otherproducts}_{ij,t-1} > 0)_{ijt}$	0.002***					0.004***
		(0.001)					(0.001)
common language	$\mathbb{I}(N^{sameproduct}_{ij,t-1} > 0)_{ijt}$		0.002***				0.001***
			(0.000)				(0.000)
	$\mathbb{I}(N^{otherproducts}_{ij,t-1} > 0)_{ijt}$		-0.000**				-0.000
			(0.000)				(0.000)
common colonizer	$\mathbb{I}(N^{sameproduct}_{ij,t-1} > 0)_{ijt}$			0.002***			0.001**
				(0.001)			(0.001)
	$\mathbb{I}(N^{otherproducts}_{ij,t-1} > 0)_{ijt}$			-0.001**			-0.001
				(0.000)			(0.000)
common income group	$\mathbb{I}(N^{sameproduct}_{ij,t-1} > 0)_{ijt}$				0.001***		0.001*
					(0.000)		(0.000)
	$\mathbb{I}(N^{otherproducts}_{ij,t-1} > 0)_{ijt}$				-0.000		-0.000
					(0.000)		(0.000)
common continent	$\mathbb{I}(N^{sameproduct}_{ij,t-1} > 0)_{ijt}$					-0.000	-0.003***
						(0.000)	(0.000)
	$\mathbb{I}(N^{otherproducts}_{ij,t-1} > 0)_{ijt}$					-0.003***	-0.003***
						(0.000)	(0.000)
$y_{ij,t-1}$		0.309***	0.315***	0.311***	0.310***	0.325***	0.338***
		(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
$y_{ij,t-2}$		0.076***	0.081***	0.076***	0.077***	0.093***	0.106***
		(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
Observations		3,943,800	3,943,800	3,943,800	3,943,800	3,943,800	3,943,800
# of firm-product couples		6,573	6,573	6,573	6,573	6,573	6,573
# of firms		1,965	1,965	1,965	1,965	1,965	1,965
AR(1)		0	0	0	0	0	0
AR(2)		.693	.964	.657	.705	.219	.016
Sargan		0	0	0	0	0	0

Notes: The dependent variable is y_{ijt} which is a dummy variable indicating whether a firm-product couple i exported to country j in year t. All regressions include firm-product-destination fixed effects, as well as year dummies and a constant (all not reported). Standard errors are in parentheses. All regressions use robust standard errors and treat the lags of the dependent variable as well as the regressors of interest as predetermined. We use the two-step system GMM estimator from Blundell and Bond (1998) and, due to the two-step estimation, we use the Windmeijer (2005) finite sample correction for the standard errors. *, ** and *** denote significance at the 10%-, 5%- and 1%-level, respectively. The values reported for AR(1) and AR(2) are the p-values for first and second order autocorrelated disturbances in the first differences equations. The row for the Sargan reports the p-values for the null hypothesis of validity of the overidentifying restrictions and can only be computed assuming homoskedasticity. To report this statistic, we re-estimate the model accordingly.

By and large, results are very similar when compared to Table 1.10. Common border for the same product has again the largest marginal effect. We again do not find evidence for across product learning. Interestingly, we now find a small but significant effect of having exported to a common continent.

To sum up, we hardly find evidence for across product learning of spatial exporters. This probably hints at only small economies of scope for multi-product firms when entering new export markets with several products, at least across markets.

1.5 Robustness checks

We now discuss several effects that could influence our results and which are unrelated to the cross-country correlation in export destination choices of firms due to spatial exporters. Specifically, we investigate the role of lagged export values at the firm level, the impact of competitors' success in previous contiguous export destinations, trading agents, state-owned firms, foreign-owned firms, and processing trade. Regression results pertaining to these robustness checks can be found in Table 1.12. All robustness checks use specification VI from Table 1.9 as a starting point.

Lagged export values: In addition to learning from its previous export experience, a firm may also exhibit increasing returns to scale via a learning by doing mechanism in textile and apparel production. Since the quotas of the MFA represent an artificial quantity restriction, removing it should result in a large increase in the volume of export sales. As our regressor of interest is correlated with a firm's export volume by construction and this might bias our results, we include the lagged export value as an additional control variable in column I. As can be seen from column I in Table 1.12, contiguity between export destinations still has a significant positive impact on a firm's exporting decision even when controlling for the lagged export value.

Competitors' success: Krautheim (2012) theoretically investigates the importance of spillover effects from competing firms on exporting fixed costs. The number of exporting firms of the same product or the number of export markets already entered by close competitors may influence a firm's ability to export to a specific destination. Wen (2004) shows that Chinese firms producing in the same industry tend to cluster geographically across Chinese regions. We therefore use the sum of the number of previously entered contiguous export

Table 1.11: Multi-product firms: dynamic panel estimates—N

		I	II	III	IV	V	VI
$N_{ij,t-1}$ defined according to...							
common border	$N^{sameproduct}_{ij,t-1}$	0.017***					0.010***
		(0.002)					(0.002)
	$N^{otherproducts}_{ij,t-1}$	0.001***					-0.000
		(0.000)					(0.000)
common language	$N^{sameproduct}_{ij,t-1}$		0.002***				-0.000
			(0.000)				(0.000)
	$N^{otherproducts}_{ij,t-1}$		-0.000***				-0.000***
			(0.000)				(0.000)
common colonizer	$N^{sameproduct}_{ij,t-1}$			0.004***			0.001**
				(0.000)			(0.000)
	$N^{otherproducts}_{ij,t-1}$			0.000			-0.000***
				(0.000)			(0.000)
common income group	$N^{sameproduct}_{ij,t-1}$				0.005***		0.003***
					(0.000)		(0.000)
	$N^{otherproducts}_{ij,t-1}$				0.000***		0.000***
					(0.000)		(0.000)
common continent	$N^{sameproduct}_{ij,t-1}$					0.004***	0.002***
						(0.000)	(0.000)
	$N^{otherproducts}_{ij,t-1}$					0.000***	0.000***
						(0.000)	(0.000)
$y_{ij,t-1}$		0.310***	0.309***	0.312***	0.312***	0.308***	0.349***
		(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
$y_{ij,t-2}$		0.078***	0.075***	0.078***	0.085***	0.080***	0.124***
		(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
Observations		3,943,800	3,943,800	3,943,800	3,943,800	3,943,800	3,943,800
# of firm-product couples		6,573	6,573	6,573	6,573	6,573	6,573
# of firms		1,965	1,965	1,965	1,965	1,965	1,965
AR(1)		0	0	0	0	0	0
AR (2)		.876	.597	.816	.605	.948	0
Sargan		0	0	0	0	0	0

Notes: The dependent variable is y_{ijt} which is a dummy variable indicating whether a firm-product couple i exported to country j in year t. All regressions include firm-product-destination fixed effects, as well as year dummies and a constant (all not reported). Standard errors are in parentheses. All regressions use robust standard errors and treat the lags of the dependent variable as well as the regressors of interest as predetermined. We use the two-step system GMM estimator from Blundell and Bond (1998) and, due to the two-step estimation, we use the Windmeijer (2005) finite sample correction for the standard errors. *, ** and *** denote significance at the 10%-, 5%- and 1%-level, respectively. The values reported for AR(1) and AR(2) are the p-values for first and second order autocorrelated disturbances in the first differences equations. The row for the Sargan reports the p-values for the null hypothesis of validity of the overidentifying restrictions and can only be computed assuming homoskedasticity. To report this statistic, we re-estimate the model accordingly.

Table 1.12: Dynamic panel estimates—N—robustness checks

	I	II	III	IV	V	VI
	lagged export value	competitors' success	drop trading agents	drop state owned firms	drop foreign owned firms	drop processing trade firms
$N_{ij,t-1}$ defined according to...						
common border	0.013***	0.011**	0.014***	0.017***	0.007	0.006
	(0.004)	(0.004)	(0.004)	(0.004)	(0.006)	(0.007)
common language	-0.001*	-0.000	-0.001*	-0.001*	-0.001	-0.001
	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.001)
common colonizer	0.000	0.001	0.000	0.000	0.002	0.002
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
common income group	0.005***	0.003***	0.005***	0.005***	0.002*	0.004***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
common continent	0.002***	0.002***	0.002***	0.001**	0.002*	0.002*
	(0.000)	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
firmvalue$_{i,t-1}$	-0.000					
	(0.000)					
$N_{-ij,p,t-1}$ defined according to...						
common border		-0.000*				
		(0.000)				
common language		-0.000*				
		(0.000)				
common colonizer		-0.000***				
		(0.000)				
common income group		0.000***				
		(0.000)				
common continent		0.000***				
		(0.000)				
$y_{ij,t-1}$	0.357***	0.480***	0.363***	0.368***	0.370***	0.344***
	(0.013)	(0.012)	(0.014)	(0.015)	(0.024)	(0.026)
$y_{ij,t-2}$	0.101***	0.218***	0.100***	0.096***	0.115***	0.077**
	(0.013)	(0.013)	(0.014)	(0.014)	(0.023)	(0.025)
Observations	770,000	770,000	738,750	727,800	88,800	160,200
# of firms	1,295	1,295	1,236	1,213	148	547
AR(1)	0	0	0	0	0	0
AR(2)	.156	0	.065	.034	.718	.252
Sargan	0	0	0	0	.033	0

Notes: The dependent variable is y_{ijt} which is a dummy variable indicating whether a firm i exported to country j in year t. All regressions include firm-destination fixed effects, as well as year dummies and a constant (all not reported). Standard errors are in parentheses. All regressions use robust standard errors and treat the lags of the dependent variable as well as the regressors of interest as predetermined. We use the two-step system GMM estimator from Blundell and Bond (1998) and, due to the two-step estimation, we use the Windmeijer (2005) finite sample correction for the standard errors. *, ** and *** denote significance at the 10%-, 5%- and 1%-level, respectively. The values reported for AR(1) and AR(2) are the p-values for first and second order autocorrelated disturbances in the first differences equations. The row for the Sargan reports the p-values for the null hypothesis of validity of the overidentifying restrictions and can only be computed assuming homoskedasticity. To report this statistic, we re-estimate the model accordingly.

destinations over all competitors in the same Chinese prefecture, $N_{-ij,p,t-1}$, to control for these spillover effects. We can construct this control variable using again all of our different contiguity measures. As can be seen from column II in Table 1.12, controlling for spillover effects from close competitors hardly affects our results.

Trading agents: The raw data contains a number of trading agents ("intermediary firms") which mediate trade for other firms but do not directly engage in production. Including these firms could cause problems as their behavior is probably different from that of manufacturing firms. To exclude the possibility that our results are driven by these trading agent business networks, we exclude trading firms which are identified by certain keywords in their names. Ahn et al. (2011) use the Chinese characters for "importer", "exporter", and "trading" to identify "intermediary firms". By contrast, we follow Upward et al. (2011) and use a more comprehensive list of keywords which are typically used by various kinds of trading agents in China. These trading companies represent about 4 percent of our observations. Column III in Table 1.12 shows that dropping trading agents does not change our conclusions.

State-owned firms: Khandelwal et al. (2013) argue that state-owned firms seem to have been more likely to obtain a license before the MFA quota restrictions were lifted. This makes them potentially different from privately-owned firms. We therefore re-run our regressions excluding state-owned firms. Again, our results shown in column IV of Table 1.12 hold up excluding state-owned firms.

Foreign-owned firms: We exclude all foreign-owned firms and processing trade exports as the choice of destinations of Chinese firms could be influenced by the foreign headquarters location or by the location of other foreign direct investments realized by the parent company. While the qualitative results reported in column V in Table 1.12 are similar, our results lose some of their significance. This may well be due to the large drop in the number of firms and observations to about a tenth of the full sample.

Processing trade: Our data set allows us to distinguish between processing and ordinary exports. The former refers to exports that are assembled in an export processing zone and use a high share of imported intermediate inputs. Note that foreign owned firms often engage in processing exports but not necessarily so. Processing exports may be special with respect to the export locations choice because they could be influenced by a third foreign party. In addition, Chinese processing trade firms may have less liberty in their export destination choice. Excluding processing trade export transactions leads again to a substantial drop in the number of observations to around a fifth of the original sample. Column VI in Table 1.12 shows that our results are again qualitatively similar but lose some of their statistical significance.

1.6 Conclusion

How do firms choose new export destinations? While there are many factors that are important for this decision, one empirical regularity strikes out: Firms tend to choose new export markets that are geographically close to their prior export destinations more often than standard gravity models would predict.

We quantify the effect of this spatial pattern using Chinese customs data and the quasi-natural experiment of the end of the import quota restrictions on Chinese textile exports which generates an exogenous set of potential new destinations (25 EU countries, the US, and Canada). We use the sample of firms which have never exported to the 27 previously restricted MFA countries before 2005 to identify the effect of previous export history in contiguous countries on the probability of exporting to one of the 150 countries which were not covered by the MFA import restrictions. This allows us to quantify the importance of 'extended gravity' or 'spatial exporters', i.e. the time-varying firm-specific heterogeneity in export destinations shaped by firms' previous export experience in spatially close countries taking into account unobserved time-invariant heterogeneity at the firm-country level as well as true state dependence.

Our baseline results show that the probability to export to a country increases by about 2 percentage points for each prior export destination with a common border with this country. For example, this implies a 33 percent increase in the probability of a firm exporting to Singapore, one of the top export destinations in our data set of non-MFA countries, when it has previously exported to Malaysia, a country which shares a common border with Singapore. Our results are robust across multiple specifications (differences-in-differences, instrumental variables, and dynamic panel estimators). We also conduct a battery of robustness checks which control for lagged export values, competitor's success, multi-product firms, the role of direct transactions, trading agents, state-owned firms, foreign-owned firms, and processing trade.

Chapter 2

International Trade and Unemployment: A Quantitative Framework*

2.1 Introduction

The quantification of the welfare effects of trade liberalization is one of the core issues in empirical international trade. All empirical frameworks for evaluating welfare effects of trade policies so far assume perfect labor markets with full employment. For example, Arkolakis et al. (2012) have shown that an ex post analysis of the welfare effects (measured in terms of real income) of a move from autarky to the observed level of trade liberalization is possible by using only data on the observed import share in a country and an estimate of the trade elasticity. If we relax the assumption of full employment, then real income is given by the real wage bill of all employed workers, i.e., $e_j L_j w_j / P_j$, where e_j is the share of the labor force L_j which is employed times the wage w_j which is paid to a worker in terms of the price level P_j. Hence assuming a constant labor force, any change in welfare $\hat{W}_j$ can be decomposed into a change in net employment and the real wage, i.e.,

$$\hat{W}_j = \hat{e}_j \widehat{\left(\frac{w_j}{P_j}\right)}, \tag{2.1}$$

where hats denote changes. In Arkolakis et al. (2012), $\hat{e}_j = 1$ by assumption, and the change in real wages is given by $\hat{\lambda}_{jj}^{1/\varepsilon}$, the change in the share of domestic expenditures,

* This chapter is based on joint work with Mario Larch. It is a revised version of CESifo Working Paper No. 4013, 2012. A previous version of this paper has been circulated under the title “Gravity with Unemployment”.

$\hat{\lambda}_{jj}$, raised to some power of ε, the elasticity of imports with respect to variable trade costs. Assuming full employment allows Arkolakis et al. (2012) to conduct a very simple ex post analysis of the welfare effects of moving from autarky to the observed level of trade integration. As $\lambda_{jj} = 1$ under autarky, one can calculate the welfare gains from trade from the observed domestic expenditure share when an estimate of the trade elasticity is available. When we allow for unemployment, however, this is not feasible any longer as we do not observe the counterfactual employment level under autarky. When we are interested in an ex ante evaluation of any counterfactual trade policy besides autarky, we additionally need estimates of trade cost parameters to get an estimate of the counterfactual domestic consumption share, which typically are obtained from estimating gravity models, regardless of whether we assume perfect or imperfect labor markets.

In the following, we present a simple quantitative framework for bilateral trade flows based on Armington (1969) preferences and recently developed models of international trade with search and matching labor market frictions. Our framework allows us to derive sufficient statistics for the welfare effects of trade liberalization similar to those of Arkolakis et al. (2012) but augmented by the aggregate employment change. The additional insights of incorporating labor market frictions into a quantitative trade model come at minimal cost: We only require knowledge of the elasticity of the matching function. Hence, our framework is easily applied to all topics where trade flow effects are inferred, such as free trade agreements, currency unions, borders and ethnic networks.

We apply our framework to a sample of 28 OECD countries from 1950 to 2006 in order to evaluate two scenarios. First, we calculate the effects of introducing preferential trade agreements (PTAs) starting from a counterfactual world without any PTAs. Second, we evaluate the effects of a hypothetical labor market reform in the United States. We find that, on average, introducing PTAs as observed in 2006 increases GDP about four percent more when accounting for employment effects arising from imperfect labor markets. Countries with only small increases in GDP, however, experience negative employment effects. On average, welfare effects are eight percent larger when allowing for imperfect labor markets. When we use commonly assumed values for the elasticities in our model instead of our estimates, we find that accounting for labor market frictions increases the welfare gains by more than 50 percent. In our framework, changes in trade costs or labor market policies affect labor market outcomes through changes in relative prices and income. When trade costs fall, imports of foreign varieties become cheaper, leading to a lower consumer price index in the corresponding country. When labor markets are characterized by search frictions, firms have to incur costs to post vacancies in order to find workers. The lower price level translates one-to-one into lower recruiting costs for domestic firms.[1] Firms *ceteris paribus* create more vacancies so that more workers find a job and unemployment is reduced. Hence, standard methods neglecting labor market effects considerably underestimate the welfare gains from trade liberalization.

Our second counterfactual experiment analyzes a hypothetical improvement of labor market institutions in the United States. As expected, GDP and welfare increase in the United States but also improve for its trading partners due to positive spillover effects of the labor market reform. A unilateral labor market reform which for example increases the matching efficiency will increase the number of successful matches between workers and firms and thus rise employment, GDP, and welfare in the corresponding country. As workers spend part of their income on foreign varieties, the increase in income leads to higher import demand for all trading partners. This translates into lower unemployment in the trading partners, leading to a positive correlation between changes in unemployment rates across countries.

In Section 2.2 we present our quantitative framework and show how to estimate trade cost parameters and elasticities. We then derive expressions for the counterfactual trade and employment levels for welfare evaluations of trade and labor market policy changes using the estimated trade cost parameters and elasticities. As an illustration of our approach, Section 2.3 evaluates the effects of preferential trade agreements and labor market reforms for a sample of 28 OECD countries. Section 2.4 concludes.

Our paper is related to several literatures, notably the gravity literature which models bilateral trade flows. Within our framework, changes in employment and GDP directly affect bilateral trade flows which can be described by a gravity equation. It captures the key stylized facts that trade increases with market size and decreases with distance. The empirical success of the gravity equation spurred a great deal of interest in its theoretical underpinnings. Anderson (1979) and Bergstrand (1985) address the role of multilateral price effects for trade flows. A more recent contribution by Eaton and Kortum (2002) develops a quantifiable Ricardian model of international trade to investigate the role of comparative advantage and geography for bilateral trade flows. Anderson and van Wincoop (2003) refine the gravity

[1] Felbermayr et al. (2011a) and Felbermayr et al. (2013) on the one hand and Helpman and Itskhoki (2010) on the other use a similar mechanism in a one- and two-sector model, respectively.

equation's theoretical foundations by including average trade barriers to capture multilateral resistance and highlight the importance of proper empirical comparative static analysis. Fieler (2011) introduces non-homothetic preferences into the Ricardian framework of Eaton and Kortum (2002) to rationalize the fact that bilateral trade is large between rich countries and small between poor countries. Waugh (2010) provides a complementary framework with asymmetric trade costs to explain the cross-country-pair differences in bilateral trade volumes and income levels. Anderson and Yotov (2010) elaborate on the incidence of bilateral trade costs in the Anderson and van Wincoop (2003) framework. These theoretical developments allow to employ the gravity equation to infer the GDP and welfare effects of counterfactual trade liberalization scenarios accounting for general equilibrium effects, which is a core issue in empirical work on international trade.

Despite this multitude of theoretical foundations for the gravity equation, to date all of them assume perfect labor markets. Crucially, this implies that changes in real welfare ignore changes in the total number of employed workers due to trade liberalization or labor market reforms. A different strand of the theoretical trade literature stresses various channels through which trade liberalization affects (un)employment. Brecher (1974), Davis (1998), and Egger et al. (2012) focus on minimum wages to analyze the interactions between trade and labor market policies. A binding minimum wage prevents downward wage adjustments when a country opens up to trade. Instead, firms adjust the number of employed workers. Others have stressed labor market frictions arising due to fair wages or efficiency wages (Amiti and Davis, 2012; Davis and Harrigan, 2011, and Egger and Kreickemeier, 2009). Fair wages or efficiency wages lead firms to pay wages above the market clearing level in order to ensure compliance of workers. When trade is liberalized, average productivity of firms increases, which leads to an increase of the fair or efficiency wage due to rent-sharing as well as an increase in unemployment. Finally, search-theoretic foundations of labor market frictions are introduced into trade models (Davidson et al., 1988, 1999; Felbermayr et al., 2011a; Helpman et al., 2010a, and Helpman and Itskhoki, 2010). In these models, workers search for jobs and firms for workers. Once a firm-worker match is established, they bargain over the match-specific surplus. Trade and labor markets interact via relative prices of hiring workers and goods prices which affect search and recruitment efforts. While our framework relies on a search-theoretical foundation of labor market frictions, we employ different approaches to divide the rent between workers and firms like minimum wages, efficiency wages, and bargaining.[2]

Theoretically, the effects of trade liberalization on (un)employment are ambiguous, but Dutt et al. (2009) as well as Felbermayr et al. (2011b) provide reduced-form evidence that more open economies have lower unemployment rates on average. In contrast to these reduced-form approaches, our structural quantitative framework accounts for country-specific general equilibrium effects and allows to quantify employment, GDP, and welfare effects of policies.[3]

2.2 A quantitative framework for trade and unemployment

2.2.1 Goods market

The representative consumer in country j is characterized by the utility function U_j. We assume that goods are differentiated by country of origin, i.e. we use the simplest possible way to provide a rationale for bilateral trade between similar countries based on preferences à la Armington (1969).[4] In Appendix B.2, we demonstrate that our framework and counterfactual analysis are isomorphic to a Ricardian model of international trade along the lines of Eaton and Kortum (2002). The quantity of purchased goods from country i is given by q_{ij}, leading to the following utility function

$$U_j = \left[\sum_{i=1}^{n} \beta_i^{\frac{1-\sigma}{\sigma}} q_{ij}^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}}, \tag{2.2}$$

[2] Cuñat and Melitz (2010) and Cuñat and Melitz (2012) study the effect of differences in labor market frictions on patterns of comparative advantage. However, their model does neither feature trade costs, the center piece of gravity analysis, nor does it consider unemployment.

[3] A recent literature studies the labor market effects of trade liberalization using structural dynamic models (Kambourov, 2009; Artuç et al., 2010; Coşar et al., 2011; Menezes-Filho and Muendler, 2011; Coşar, 2013; Dix-Carneiro, 2013, and Helpman et al., 2013). However, all these studies focus on single countries and hence abstract from the interdependencies of trade flows between countries, a decisive feature of our model. Also, with the exception of Artuç et al. (2010) who study the United States, this literature focuses on the effects of trade liberalization in Latin American emerging economies, not developed countries.

[4] Consequently, we deliberately abstract from distinguishing between the intensive and extensive margin of international trade as for example in Chaney (2008) or Helpman et al. (2008b).

where n is the number of countries, σ is the elasticity of substitution in consumption, and β_i is a positive preference parameter measuring the product appeal for goods from country i.

International trade of goods from i to j imposes iceberg trade costs $t_{ij} > 1$. Profit maximization then implies that $p_{ij} = p_i t_{ij}$, where p_i denotes the factory gate price of the good in country i.

The representative consumer maximizes Equation (2.2) subject to the budget constraint $\tilde{y}_j = \sum_{i=1}^{n} p_i t_{ij} q_{ij}$, where $\tilde{y}_j = y_j(1+d_j)$, with y_j denoting nominal income in country j and d_j the share of the exogenously given trade deficit (if $d_j > 0$) or surplus (if $d_j < 0$) of country j in terms of GDP.[5] The value of aggregate sales of goods from country i to country j can then be expressed as

$$x_{ij} \;=\; p_i t_{ij} q_{ij} = \left(\frac{\beta_i p_i t_{ij}}{P_j}\right)^{1-\sigma} \tilde{y}_j, \tag{2.3}$$

and P_j is the standard CES price index given by $P_j = [\sum_{i=1}^{n} (\beta_i p_i t_{ij})^{1-\sigma}]^{1/(1-\sigma)}$. In general equilibrium, total sales correspond to nominal income, i.e., $y_i = \sum_{j=1}^{n} x_{ij}$. Assuming labor to be the only factor of production which produces one unit of output per worker, GDP in a world with imperfect labor markets is given by total production of the final output good multiplied with its price, i.e., $y_i = p_i(1-u_i)L_i$.[6]

This setup implies a gravity equation for bilateral trade flows. In general equilibrium, GDP is given by the sum of all sales, i.e.

$$y_i \;=\; \sum_{j=1}^{n} x_{ij} = \sum_{j=1}^{n} \left(\frac{\beta_i t_{ij} p_i}{P_j}\right)^{1-\sigma} \tilde{y}_j = (\beta_i p_i)^{1-\sigma} \sum_{j=1}^{n} \left(\frac{t_{ij}}{P_j}\right)^{1-\sigma} \tilde{y}_j. \tag{2.4}$$

Solving for scaled prices $\beta_i p_i$ and defining $y^W \equiv \sum_j y_j$, $\tilde{y}^W \equiv \sum_j \tilde{y}_j$ and income shares $\theta_j \equiv y_j/y^W$ and $\tilde{\theta}_j \equiv \tilde{y}_j/\tilde{y}^W$, we can write bilateral trade flows as given in Equation (2.3) as

$$x_{ij} = \frac{y_i \tilde{y}_j}{y^W} \left(\frac{t_{ij}}{\tilde{\Pi}_i \tilde{P}_j}\right)^{1-\sigma}, \quad \text{where} \tag{2.5}$$

[5] We allow for trade imbalances following Dekle et al. (2007). We also conducted all counterfactual scenarios assuming balanced trade, but our results changed very little. Detailed results can be found in Appendix B.3.

[6] For further reference, note that we measure (changes in) nominal variables like GDP in terms of the price index of the first country in our data set in our subsequent empirical analysis.

$$\tilde{\Pi}_i \equiv \left(\sum_{j=1}^{n} \left(\frac{t_{ij}}{\tilde{P}_j} \right)^{1-\sigma} \tilde{\theta}_j \right)^{1/(1-\sigma)}, \quad \tilde{P}_j \equiv \left(\sum_{i=1}^{n} \left(\frac{t_{ij}}{\tilde{\Pi}_i} \right)^{1-\sigma} \theta_i \right)^{1/(1-\sigma)}, \tag{2.6}$$

while we substituted equilibrium scaled prices into the definition of the price index to obtain the multilateral resistance terms $\tilde{P}_j$.

Note that this system of equations exactly corresponds to the system given in Equations (9)-(11) in Anderson and van Wincoop (2003) or Equations (5.32) and (5.35) in Feenstra (2004) assuming balanced trade, $d_i = 0$ for all i, even when labor markets are imperfect.[7]

The intuition for this result is that GDPs appear in Equation (2.5). Observed GDPs already include the actual number of employed people. Hence, it still holds that total spending equals total production. The only difference is that now total production is achieved by *employed workers*, not all workers, as is assumed with perfect labor markets. By adding a stochastic error term, Equation (2.5) can be written as

$$z_{ij} \equiv \frac{x_{ij}}{y_i \tilde{y}_j} = \exp\left(k - (1-\sigma)\ln t_{ij} - \ln \tilde{\Pi}_i^{1-\sigma} - \ln \tilde{P}_j^{1-\sigma} + \varepsilon_{ij}\right), \tag{2.7}$$

where ε_{ij} is a random disturbance term or measurement error of exports, assumed to be identically distributed and mean-independent of the remaining terms on the right-hand side of Equation (2.7), and k is a constant capturing the logarithm of world GDP. Country-specific importer and exporter fixed effects can be used to control for the outward and inward multilateral resistance terms $\tilde{\Pi}_i$ and $\tilde{P}_j$, respectively, as suggested by Anderson and van Wincoop (2003) and Feenstra (2004). Hence, even with labor market frictions, we can use established methods to estimate trade costs using the gravity equation, independently of the underlying labor market model. We summarize this result in Implication 1:

Implication 1 *The estimation of trade costs is unchanged when allowing for imperfect labor markets.*

To evaluate ex ante welfare effects of changes in trade policies, we need in addition to trade cost elasticity estimates the counterfactual changes in employment and GDP. To derive these, we have to take a stance on how to model the labor market, to which we turn in the next section.

[7] If trade is balanced, then $\tilde{\Pi}_i = \Pi_i$ and $\tilde{P}_i = P_i$. When, in addition, trade costs are symmetric, i.e., $t_{ij} = t_{ji}$, then $\tilde{\Pi}_i = \tilde{P}_i$ (see Anderson and van Wincoop, 2003).

2.2.2 Labor market

We model the labor market using a one-shot version of the search and matching framework (SMF, see Mortensen and Pissarides, 1994 and Pissarides, 2000).[8] Search-theoretic frameworks fit stylized facts of labor markets in developed economies as for example the simultaneous existence of unfilled vacancies and unemployed workers.[9]

The labor market is characterized by frictions. All potential workers in country j, L_j, have to search for a job, and firms post vacancies V_j in order to find workers. The number of successful matches between an employer and a worker, M_j, is given by $M_j = m_j L_j^{\mu} V_j^{1-\mu}$, where $\mu \in (0,1)$ is the elasticity of the matching function and m_j measures the overall efficiency of the labor market.[10] Only a fraction of open vacancies will be filled, $M_j/V_j = m_j (V_j/L_j)^{-\mu} = m_j \vartheta_j^{-\mu}$, and only a fraction of all workers will find a job, $M_j/L_j = m_j (V_j/L_j)^{1-\mu} = m_j \vartheta_j^{1-\mu}$, where $\vartheta_j \equiv V_j/L_j$ denotes the degree of labor market tightness in country j. This implies that the unemployment rate is given by[11]

$$u_j = 1 - m_j \vartheta_j^{1-\mu}. \tag{2.8}$$

As is standard in search models, we assume that every firm employs one worker. Similar to Helpman and Itskhoki (2010), this assumption does not lead to any loss of generality as long as the firm operates under perfect competition and constant returns to scale. In addition, we assume that all firms have the same productivity and produce a homogeneous good. In order to employ a worker (i.e. to enter the market), the firm has to post a vacancy at a cost of $c_j P_j$, i.e. in units of the final output good.[12] After paying these costs, a firm finds a worker with probability $m_j \vartheta^{-\mu}$. When a match between a worker and a firm has been established, we assume that they bargain over the total match surplus. Alternatively, we

[8] See Rogerson et al. (2005) for a survey of search and matching models, including an exposition of a simplified one-shot (directed) search model. For recent trade models using a similar static (non-directed search) approach, see for example Helpman and Itskhoki (2010). Felbermayr et al. (2013) use a similar labor market setup. However, they do not investigate its implications for the estimation of gravity equations nor do they use it for a structural quantitative analysis.

[9] They are less successful in explaining the cyclical behavior of unemployment and vacancies, see Shimer (2005). This deficiency is not crucial in our case as we purposely focus on the steady state.

[10] Note that we assume a constant returns to scale matching function in line with empirical studies, see Petrongolo and Pissarides (2001).

[11] Note that the matching efficiency has to be sufficiently low to ensure job finding rates and job filling rates between 0 and 1.

[12] This implies that not all of GDP is available for final consumption (and hence welfare) of workers.

consider minimum and efficiency wages in Appendices B.4 and B.5 as mechanisms for wage determination. All three approaches are observationally equivalent in our setting.

In the bargaining case, the match gain of the firm is given by its revenue from sales of one unit of the homogeneous product minus wage costs, $p_j - w_j$, as the firm's outside option is zero. The match surplus of a worker is given by $w_j - b_j$, where b_j is the outside option of the worker, i.e. the unemployment benefits (b_j) she receives when she is unemployed.[13]

We use a generalized Nash bargaining solution to determine the surplus splitting rule. Hence, wages w_j are chosen to maximize $(w_j - b_j)^{\xi_j}(p_j - w_j)^{1-\xi_j}$, where the bargaining power of the worker is given by $\xi_j \in (0,1)$. The unemployment benefits are expressed as a fraction γ_j of the market wage rate. Note that both the worker and the firm neglect the fact that in general equilibrium, higher wages lead to higher unemployment benefits, i.e., they both treat the replacement rate as exogenous (see Pissarides, 2000). The first order conditions of the bargaining problem yield $w_j - \gamma_j w_j = \xi_j/(1-\xi_j)\,(p_j - w_j)$. Solving for w_j results in the **wage curve** $w_j = \xi_j/(1 + \gamma_j\xi_j - \gamma_j)p_j$. Due to the one-shot matching, the wage curve does not depend on ϑ_j. The bargained wage increases in the value of output p_j, in the worker's bargaining power ξ_j, and in the replacement rate γ_j.

Given wages w_j, profits of a firm π_j are given by $\pi_j = p_j - w_j$. As we assume one worker firms and the probability of filling an open vacancy is $m_j\vartheta^{-\mu}$, expected profits are equal to $(p_j - w_j)m_j\vartheta^{-\mu}$. Firms enter the market until these expected profits cover the entry costs c_jP_j. Rewriting, one finds the **job creation curve** $w_j = p_j - P_jc_j/(m_j\vartheta_j^{-\mu})$. It is increasing in the value of output and decreasing in the expected recruiting costs $P_jc_j/(m_j\vartheta_j^{-\mu})$.

Combining the job creation and wage curves determines the equilibrium labor market tightness as

$$\vartheta_j = \left(\frac{p_j}{P_j}\right)^{1/\mu}\left(\frac{c_j}{m_j}\Omega_j\right)^{-1/\mu}, \tag{2.9}$$

where $\Omega_j \equiv \frac{1-\gamma_j+\gamma_j\xi_j}{1-\gamma_j+\gamma_j\xi_j-\xi_j} \geq 1$ summarizes the effective bargaining power of workers. Ω_j is increasing in the worker's bargaining power ξ_j and in the replacement rate γ_j. Labor market tightness decreases and the unemployment rate increases when m_j or c_j decrease or Ω_j increases.

[13] Unemployment benefits are financed via lump-sum transfers from employed workers to the unemployed. As we assume homothetic preferences and homogenous workers, this does not show up in the economy-wide budget constraint $\tilde{y}_j$, see equation (2.3).

The relative price p_j/P_j is determined by the demand and the supply of goods. It therefore provides the link between the labor and goods market.

2.2.3 Estimation of elasticities

We have now set the stage to derive expressions for our counterfactual welfare analysis—if we follow most of the gravity literature and merely assume plausible values for the elasticity of substitution, σ, and, in our case, the matching elasticity, μ. In the following, we demonstrate that in principle, both elasticities can be estimated within our quantitative framework, even though the main contribution of this paper is providing a structural gravity framework allowing for imperfect labor markets. Therefore, impatient (or unconvinced) readers may as well simply assume values for σ and μ and continue with Section 2.2.4. In addition for these readers, we present results of our counterfactual analysis for different assumed values of the elasticities in Table 2.4.

Estimating the elasticity of substitution

Bergstrand et al. (2013) show how to obtain estimates for σ within their proposed framework without relying on additional data besides the standard trade data. We show that a variant of their approach is also applicable when assuming imperfect labor markets. To estimate σ, in addition to the trade data we only need data on unemployment rates as well as civil labor force data.

First, note that we can rewrite trade flows as given in Equation (2.3) by observing that the variety price can be substituted by $p_i = y_i/[(1-u_i)L_i]$. This yields $x_{ij} = ((\beta_i y_i t_{ij})/((1-u_i)L_i P_j))^{1-\sigma}\,\tilde{y}_j$. Estimation of Equation (2.7) using observable determinants of bilateral trade costs generates estimates $\widehat{t_{ij}^{1-\sigma}}$. We next substitute $\widehat{t_{ij}^{1-\sigma}}$ in Equation (2.5) to generate $\hat{x}_{ij}$ and $\widehat{t_{mj}^{1-\sigma}}$ in its analogue to generate $\hat{x}_{mj}$. Using observed unemployment rates we end up with:

$$\frac{\hat{x}_{ij}}{\hat{x}_{mj}} = \frac{\widehat{t_{ij}^{1-\sigma}}}{\widehat{t_{mj}^{1-\sigma}}}\left(\frac{\beta_i y_i(1-u_m)L_m}{\beta_m y_m(1-u_i)L_i}\right)^{1-\sigma}. \tag{2.10}$$

We can solve Equation (2.10) for σ, where y_i, y_m, L_i, L_m, u_i, and u_m are observables. In addition, we assume that $\beta_i = \beta_m$. Then, we can calculate $n^2(n-1)$ values of σ by using all combinations i, j, and m ($m \neq i$). As a measure of central tendency, we use the average value of all estimates of $\sigma > 1$ as our summary estimate in order to ensure that trade costs do not counterfactually increase with rising distance. We use bootstrapped standard errors for σ.

Estimating the elasticity of the matching function

The other crucial parameter for our counterfactual analysis is the elasticity of the matching function, μ. As with the elasticity of substitution, there are a great many of plausible estimates of the matching elasticity available in the literature. Still, we demonstrate that it is also possible to obtain an estimate of μ within our structural gravity framework relying on the cross-country-pair variation in bilateral trade flows.

Using again Equations (2.8) and (2.9) and defining $\Xi_j \equiv m_j \left(\frac{c_j}{m_j}\Omega_j\right)^{\frac{\mu-1}{\mu}}$, we can write $1 - u_j = \Xi_j \left(p_j/\tilde{P}_j\right)^{(1-\mu)/\mu}$. As we observe u_j in the baseline, we may take ratios for two countries and the log of this ratio to obtain:

$$\ln\left(\frac{1-u_j}{1-u_m}\right) = \frac{1-\mu}{\mu}\left[\ln\left(\frac{p_j}{p_m}\frac{\tilde{P}_m}{\tilde{P}_j}\right) - \ln\left(\frac{c_j\Omega_j}{c_m\Omega_m}\right)\right] + \frac{1}{\mu}\ln\left(\frac{m_j}{m_m}\right). \tag{2.11}$$

We can solve Equation (2.11) for μ, where u_j, c_j and Ω_j are in principle observable. The unobservable variety prices p_j and the price indices P_j can be replaced by $(\beta_j p_j)^{1-\sigma} = (y^W/\tilde{y}^W)\theta_j\tilde{\Pi}_j^{\sigma-1} = (y^W/\tilde{y}^W)\mathfrak{t}_j$ and $\tilde{P}_j^{1-\sigma} = \sum_{i=1}^{n} t_{ij}^{1-\sigma}\mathfrak{t}_i$, respectively. $\mathfrak{t}_i$s can be recovered from solving the system of equations given in Equations (2.5) and (2.6) for observed trade flows using an estimate of $\widehat{t_{ij}^{1-\sigma}}$. In our application, we assume again that $\beta_j = \beta_m$. In addition, we assume identical recruiting costs, c_j, and matching efficiencies, m_j, across countries as empirical measures of recruiting costs and efficiencies which are comparable across countries are hard to come by. We also assume that the bargaining power of workers, ξ_j, is 0.5 in all countries. However, we use observed unemployment benefits across countries from OECD (2007).[14] Hence γ_j and thus Ω_j vary across countries and reflect the heterogeneity in this labor market institution across countries.

We can then calculate $n(n-1)$ such values of μ by using all combinations of j and m ($m \neq j$). As a summary estimate, we average over all estimated values of μ within the unit interval. We use bootstrapped standard errors for μ.[15]

[14] For further details on the data, see Section 2.3.
[15] We use analytical standard errors for the trade cost parameters.

2.2.4 Counterfactual analysis

While trade cost parameters can be recovered without assumptions concerning the labor market according to Implication 1, most researchers estimate gravity equations in order to evaluate counterfactual policy changes which take into account general equilibrium effects. This allows to analyze large policy changes which very likely violate the stable unit treatment assumption (SUTVA) and thus preclude interpreting gravity equation estimates as marginal effects. More importantly, a structural counterfactual analysis allows an ex ante evaluation of a potential policy change, whereas reduced form regressions are best suited for ex post evaluations of actually observed policies.

Having obtained consistent estimates of the trade cost parameters of t_{ij} as well as the elasticities μ and σ, our model structure allows us to conduct counterfactual analyses. Given these estimates, solving the system of equations given by Equation (2.6) for the multilateral resistance terms $\tilde{P}_j$ and $\tilde{\Pi}_i$ and using the actual observed GDPs to calculate world income shares θ_j gives us the solutions for the baseline scenario.[16] Resolving the system of equations after having changed e.g. the trade cost vector by abolishing all observed PTAs (i.e. setting the PTA dummy variable to 0) yields the multilateral resistance terms in the counterfactual scenario, $\tilde{P}_j^c$ and $\tilde{\Pi}_i^c$. When solving for the counterfactual, one has to take into account that world income shares change endogenously as implied by the model structure.

When calculating counterfactual GDP, all approaches to date neglect changes in the total number of employed workers. For example, in the framework of Anderson and van Wincoop (2003) with perfect labor markets, calculating GDP and corresponding shares in world GDP is easy as "*quantities produced are assumed fixed*" (p. 190). However, this assumption is also very restrictive, as it implies that GDP and welfare changes are solely due to changes in (real) prices. Hence, changes in a country's GDP only translate into price changes in the perfect labor market framework. Similarly, in Eaton and Kortum (2002) the number of employed workers remains constant.

In contrast, our model also leads to *employment* adjustments. When GDP falls, unemployment will rise, which in turn will impact wages. In essence, our model allows labor market variables to affect income. Hence, assuming perfect or imperfect labor markets matters for the proper counterfactual analysis.

[16] See Appendix B.6 for a detailed description of the solution of the system of multilateral resistance terms with asymmetric trade costs and trade deficits.

In the following, we derive and discuss in turn counterfactual welfare along the lines of Arkolakis et al. (2012), (un)employment, GDP, and trade flows as functions of the multilateral resistance terms in the baseline and counterfactual scenario.

Counterfactual welfare

We can now consider the welfare consequences of a counterfactual change in trade costs that leaves the ability to serve the own market, t_{jj}, unchanged as in Arkolakis et al. (2012). Additionally, we follow their normalization and set the wage in country j, w_j, equal to one. In our economy, (nominal) GDP is given by total production of the final output good multiplied with its price, i.e., $y_i = p_i(1-u_i)L_i$, whereas consumable income is given by $\breve{y}_j = (1+d_j)(1-u_j)w_jL_j$.[17] We then come up with the following sufficient statistics (see Appendix B.7 for the derivation):

Implication 2 *Welfare effects of trade liberalization in our model with imperfect labor markets can be expressed as*

$$\hat{W}_j = \hat{e}_j \hat{\lambda}_{jj}^{\frac{1}{1-\sigma}}.$$

Hence, welfare depends on the employment change, $\hat{e}_j$, the change in the share of domestic expenditures, $\hat{\lambda}_{jj}$, and the partial elasticity of imports with respect to variable trade costs, given in our case by $1/(1-\sigma)$. Note that in the case of perfect labor markets $\hat{e}_j = 1$ and $\hat{W}_j = \hat{\lambda}_{jj}^{1/(1-\sigma)}$, which is exactly Equation (6) in Arkolakis et al. (2012).

When $\hat{\lambda}_{jj}$ is observed, assuming imperfect or perfect labor markets would lead to different welfare predictions. The difference in the welfare change is given by $\hat{e}_j$. Hence, assuming perfect labor markets neglects the effects on employment and the corresponding welfare effects. Whether welfare increases or decreases in a particular country depends on the relative magnitude of trade creation and diversion.

While Implication 2 already describes how to calculate welfare within our framework, we can equivalently express the change in welfare as a function of the multilateral resistance terms by using the equivalent variation, i.e. the amount of income the representative consumer would need to make her as well off under current prices $\tilde{P}_j$ as in the counterfactual situation with price level $\tilde{P}_j^c$. The advantage of this formulation is that it allows for trade

[17] Total consumable income $\breve{y}_j$ consists of the income of employed workers $(1+d_j)(1-u_j)w_jL_j - B_j$, and the income of unemployed workers B_j where $B_j = u_jL_jb_j$, the total sum of unemployment benefits which is financed by a lump-sum transfer from employed workers to the unemployed.

imbalances and changes in labor market institutions. We can express the equivalent variation in percent as follows:

$$EV_j = \frac{\check{y}_j^c \frac{\tilde{P}_j}{\tilde{P}_j^c} - \check{y}_j}{\check{y}_j} = \frac{\check{y}_j^c}{\check{y}_j}\frac{\tilde{P}_j}{\tilde{P}_j^c} - 1 = \hat{\check{y}}_j \frac{\tilde{P}_j}{\tilde{P}_j^c} - 1. \tag{2.12}$$

Note that $\hat{\check{y}}_j = \hat{\upsilon}_j \hat{y}_j$ where $\upsilon_j \equiv \xi_j/(1+\gamma_j\xi_j - \gamma_j)$ and $\hat{\upsilon}_j \equiv \upsilon_j^c/\upsilon_j$. Hence welfare can be calculated by using the expressions for the price indices (which can be derived from the multilateral resistance terms) and the counterfactual change in GDP. To derive the counterfactual change in GDP, it turns out to be useful to first derive an expression for the counterfactual change in (un)employment.

Counterfactual (un)employment

Noting that variety prices p_j are not observed, we follow Anderson and van Wincoop (2003) and use Equation (2.4) to solve for scaled prices as follows:

$$(\beta_j p_j)^{1-\sigma} = \frac{y_j}{\sum_{i=1}^n \left(\frac{t_{ji}}{\tilde{P}_i}\right)^{1-\sigma} \tilde{y}_i} = \frac{y^W}{\tilde{y}^W}\theta_j \tilde{\Pi}_j^{\sigma-1} = \frac{y^W}{\tilde{y}^W}\mathfrak{k}_j, \tag{2.13}$$

where $\mathfrak{k}_j \equiv \theta_j \tilde{\Pi}_j^{\sigma-1}$. We then use the definition of u_j given in Equation (2.8), replacing ϑ_j by the expression given in Equation (2.9) and defining $\Xi_j \equiv m_j \left(\frac{c_j}{m_j}\Omega_j\right)^{\frac{\mu-1}{\mu}}$ and $\hat{\kappa}_j \equiv \Xi_j^c/\Xi_j$, where superscript c denotes counterfactual values:

$$\frac{e_j^c}{e_j} \equiv \frac{1-u_j^c}{1-u_j} = \hat{\kappa}_j \left(\frac{p_j^c}{p_j}\right)^{\frac{1-\mu}{\mu}} \left(\frac{\tilde{P}_j}{\tilde{P}_j^c}\right)^{\frac{1-\mu}{\mu}}, \tag{2.14}$$

where e_j denotes the employment rate. Noting the derivation of Equation (2.13) and remembering that $\tilde{P}_j^{1-\sigma} = \sum_i (y^W/\tilde{y}^W) t_{ij}^{1-\sigma}\mathfrak{k}_i$ (see the definition of the price index and (2.13)), we can express the ratios of the prices and price indices as functions of $\mathfrak{k}_i$ to end up with counterfactual (un)employment levels summarized in the following implication:

Implication 3 *Whereas in the setting with perfect labor markets (un)employment effects are zero by assumption, the (un)employment effects in our gravity* *system with imperfections on the labor market are given by:*

$$\begin{aligned}\hat{e}_j &\equiv \frac{e_j^c}{e_j} = \hat{\kappa}_j \left(\frac{\mathfrak{t}_j^c}{\mathfrak{t}_j}\right)^{\frac{1-\mu}{\mu(1-\sigma)}} \left(\frac{\sum_i t_{ij}^{1-\sigma}\mathfrak{t}_i}{\sum_i \left(t_{ij}^c\right)^{1-\sigma}\mathfrak{t}_i^c}\right)^{\frac{1-\mu}{\mu(1-\sigma)}}, \\ \Delta u_j &\equiv u_j^c - u_j = (1-u_j)(1-\hat{e}_j).\end{aligned}$$

Implication 3 reveals that a country can directly affect its (un)employment level by changes in its labor market institutions, as reflected by changes in $\hat{\kappa}_j$.[18] In addition, all trading partners are affected by such a labor market reform due to changes in prices as reflected by $\mathfrak{t}_i$. Direct effects are scaled by changes in relative prices $p_j/\tilde{P}_j$ which are proportional to $\left(\mathfrak{t}_j/\sum_i t_{ij}^{1-\sigma}\mathfrak{t}_i\right)^{1/(1-\sigma)}$, reflecting the spillovers of labor market reforms to other countries. Changes of relative prices due to trade liberalization therefore provide the link to the labor market.

Even with imperfect labor markets we just need one additional parameter alongside σ, namely μ, the elasticity of the matching function, in order to calculate counterfactual values once we have solved for the multilateral resistance terms. Note that μ plays a crucial role for the importance of the labor market frictions. To illustrate, assume that all labor market institutions remain the same and μ approaches one. Then, the (un)employment effects vanish.[19] A lower μ, i.e., higher labor market frictions, leads to larger changes in (un)employment for given relative price changes. Additionally, all (potential) changes in labor market policies are succinctly summarized in a reduced-form fashion in $\hat{\kappa}_j$.

Counterfactual GDP

We next derive counterfactual (nominal) GDPs. Using the definition of GDP, $y_j = p_j(1-u_j)L_j = p_j e_j L_j$, and taking the ratio of counterfactual GDP, y_j^c, and observed GDP, y_j, we can use Implication 3 and Equation (2.13) to come up with the following implication:

Implication 4 *Counterfactual GDPs are given by:*

$$\begin{aligned}&\textit{imperfect labor markets:} \quad \hat{y}_j = \left(\hat{D}^W\right)^{\frac{1}{1-\sigma}} \hat{\kappa}_j \left(\frac{\mathfrak{t}_j^c}{\mathfrak{t}_j}\right)^{\frac{1}{\mu(1-\sigma)}} \left(\frac{\sum_i t_{ij}^{1-\sigma}\mathfrak{t}_i}{\sum_i \left(t_{ij}^c\right)^{1-\sigma}\mathfrak{t}_i^c}\right)^{\frac{1-\mu}{\mu(1-\sigma)}}, \\ &\textit{perfect labor markets:} \quad \hat{y}_j = \left(\hat{D}^W\right)^{\frac{1}{1-\sigma}} \left(\frac{\mathfrak{t}_j^c}{\mathfrak{t}_j}\right)^{\frac{1}{1-\sigma}},\end{aligned}$$

[18] Note that employment changes are homogeneous of degree zero in prices, implying that a normalization does not matter for the employment effects.

[19] In this case the level of unemployment is given by $u_j = 1 - m_j$.

with $\hat{D}^W \equiv (y^{W,c}\tilde{y}^W)/(\tilde{y}^{W,c}y^W)$ indicating the endogenous change in the world trade deficit to keep trade deficit GDP shares d_js constant. It equals one in the case of balanced trade. In order to ensure a common numéraire, we normalize $\tilde{P}_1 = \tilde{P}_1^c = 1$, i.e., GDP changes are in terms of the price level of the first importer in the data set.[20] If we assume $\mu = 1$ and balanced trade, we end up with the case of perfect labor markets employed by Anderson and van Wincoop (2003).

It is illuminating to decompose the change in GDP as follows:

$$\hat{y}_j = \left(\hat{D}^W\right)^{\frac{1}{1-\sigma}} \underbrace{\left(\frac{\mathfrak{t}_j^c}{\mathfrak{t}_j}\right)^{\frac{\mu}{\mu(1-\sigma)}}}_{\text{price change}} \underbrace{\hat{\kappa}_j \left(\frac{\mathfrak{t}_j^c}{\mathfrak{t}_j}\right)^{\frac{1-\mu}{\mu(1-\sigma)}} \left(\frac{\sum_i t_{ij}^{1-\sigma}\mathfrak{t}_i}{\sum_i \left(t_{ij}^c\right)^{1-\sigma}\mathfrak{t}_i^c}\right)^{\frac{1-\mu}{\mu(1-\sigma)}}}_{\text{employment change}}, \qquad (2.15)$$

with the price change and the employment change as defined in Implication 3.

Let us focus on the numéraire country for a moment. As we use its price index as our numéraire, the last expression in brackets of Equation (2.15) is equal to one. Then, the equation simplifies to the change in the world deficit, and, when labor market institutions remain constant, i.e. $\hat{\kappa}_j = 0$, to two terms that are equal except for their exponents: the price change term rises to the power of μ and the employment change term to the power of $1 - \mu$. Hence, the relative importance of price and employment changes only depends on μ. If μ approaches one, the labor market rigidities vanish, and the total GDP change is due to the price change, as in models assuming perfect labor markets. With any value of μ between zero and one, the share of the GDP change attributable to the price change is μ and the share due to the employment change $1 - \mu$. To illustrate, let $\mu = 0.75$, then three-quarters of the change in GDP are due to the price change and one-quarter is due to the employment change. In all other countries, changes in price indices lead to a more complex relationship. A lower price index lowers recruiting costs and thus spurs employment. This effect is captured by the last bracket in Equation (2.15). On the other hand, lower variety prices render recruiting less attractive, which is reflected by the first term of the employment change. Hence, the overall effect is ambiguous.

[20] As mentioned in footnote 12 in Anderson and van Wincoop (2003), the solution of the multilateral resistance terms (MRTs) adopts a particular normalization. In general, this applied normalization may vary between the baseline MRTs and the counterfactual MRTs. In order to ensure the same normalization for the baseline and counterfactual scenario, we normalize $\tilde{P}_1 = \tilde{P}_1^c = 1$.

Taking logs, we can attribute the share of log change in GDP divided by $\left(\hat{D}^W\right)^{\frac{1}{1-\sigma}}$, $\hat{y}_j^\star$, due to changes in prices and employment as follows:

$$1 = \frac{\ln \hat{p}_j}{\ln \hat{y}_j^\star} + \frac{\ln \hat{e}_j}{\ln \hat{y}_j^\star}. \tag{2.16}$$

Alongside GDP changes, we will report this decomposition in all our counterfactual exercises.

Counterfactual trade flows

Finally, given estimates of $t_{ij}^{1-\sigma}$, data on y_i, and a value for σ, we can calculate (scaled) baseline trade flows as $x_{ij}y^W/(y_i\tilde{y}_j) = (t_{ij}/(\tilde{\Pi}_i\tilde{P}_j))^{1-\sigma}$, where $\tilde{\Pi}_i$ and $\tilde{P}_j$ are given by Equation (2.6). With counterfactual GDPs given by Implication 4, we can calculate counterfactual trade flows as $x_{ij}^c y^{W,c}/(y_i^c\tilde{y}_j^c) = (t_{ij}^c/(\tilde{\Pi}_i^c\tilde{P}_j^c))^{1-\sigma}$, where $\tilde{\Pi}_i^c$ and $\tilde{P}_j^c$ are defined analogously to their counterparts in the baseline scenario given in Equation (2.6).[21] Due to direct effects of changes in trade costs via t_{ij} and non-trivial changes in $\tilde{\Pi}_i$ and $\tilde{P}_j$, trade may change more or less when assuming imperfect labor markets in comparison with the baseline case of perfect labor markets.

2.3 Preferential trade agreements and labor market frictions

We now apply our framework to evaluate the trade effects of preferential trade agreements and labor market reforms in a sample of 28 OECD countries for the years 1950 to 2006. The trade data are from Head et al. (2010). We use internationally comparable harmonized unemployment rates as well as employment and civil labor force data from OECD (2011e). Internationally comparable gross average replacement rates are from OECD (2007).[22]

[21] Note that $\tilde{P}_j$ and $\tilde{P}_j^c$ are homogeneous of degree one in prices while $\tilde{\Pi}_i$ and $\tilde{\Pi}_i^c$ are homogeneous of degree minus one. Hence, scaled trade flows $x_{ij}y^W/(y_i\tilde{y}_j)$ and $x_{ij}^c y^{W,c}/(y_i^c\tilde{y}_j^c)$ are homogeneous of degree zero in prices. In other words, they do not depend on the normalization chosen.

[22] This OECD summary measure is defined as the average of the gross unemployment benefit replacement rates for two earnings levels, three family situations and three durations of unemployment (for details of its calculation see Martin, 1996). As Mexico does not have any unemployment insurance scheme but is characterized by a large informal employment share, its labor market institutions are markedly different to the other OECD countries in our sample. Consequently, no replacement rate data are available for Mexico. We therefore exclude it from our analysis. For all other countries, we use the simple average of replacement rates between 2005 and 2007 as data for 2006 are not available.

Table 2.1: Summary statistics

	Mean	Std. Dev.	Min.	Max.	N
x_{ij} (cur. mn U.S.$)	2,048.991	8,950.166	0	348,420.6	38,313
GDP (cur. mn U.S.$)	386,072.995	1,143,571.923	126.99	13,201,819	43,372
PTA	0.237	0.425	0	1	44,688
$\ln DIST$	7.863	1.213	4.201	9.880	44,688
$CONTIG$	0.077	0.266	0	1	44,688
$COMLANG$	0.074	0.262	0	1	44,688

Notes: Summary statistics for the OECD sample from 1950 to 2006. The 28 countries included are Australia, Austria, Belgium, Canada, the Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Japan, Korea, Netherlands, New Zealand, Norway, Poland, Portugal, the Slovak Republic, Spain, Sweden, Switzerland, Turkey, the United Kingdom, and the United States. Data are taken from Head et al. (2010).

To obtain an estimable gravity equation as given in Equation (2.7), we need to parameterize trade costs. We follow the literature and proxy t_{ij} by a vector of trade barrier variables as follows:

$$t_{ij\tau}^{1-\sigma} = \exp(\delta_1 PTA_{ij\tau} + \delta_2 \ln DIST_{ij} + \delta_3 CONTIG_{ij} + \delta_4 COMLANG_{ij}), \tag{2.17}$$

where $PTA_{ij\tau}$ is an indicator variable of preferential trade agreement membership between country pair ij in year τ, $DIST_{ij}$ is bilateral distance, $CONTIG_{ij}$ is a dummy variable indicating whether countries i and j are contiguous, and $COMLANG_{ij}$ indicates whether the two countries share a common official language.[23] The data for the PTA's are constructed from the notifications to the World Trade Organization (WTO) and augmented and corrected by using information from PTA secretariat webpages. Table 2.1 contains summary statistics of the data.

Obviously, countries do not randomly sign PTAs. This has long been recognized in the international trade literature, see for example Trefler (1993), Magee (2003), Baier and Bergstrand (2007), and references therein. Empirical evidence shows that the exogeneity assumption of PTAs is inappropriate when attempting to quantify the effects of regional trade agreements. To avoid potential endogeneity, we follow Baier and Bergstrand (2007) and Anderson and Yotov (2011) and use a two-step estimation approach to obtain consistent estimates of trade cost coefficients. In a first step, we estimate Equation (2.7) including (directional) bilateral fixed effects, i.e., we estimate

$$z_{ij\tau} = \exp\left(k + \delta_1 PTA_{ij\tau} + \varphi_{i\tau} + \phi_{j\tau} + \nu_{ij} + \varepsilon_{ij}\right), \tag{2.18}$$

[23] We do not use common colonizer indicators or similar variables regularly used in the literature as these have very little variation in our OECD sample.

where $\varphi_{i\tau}$ and $\phi_{j\tau}$ are exporter and importer time-varying fixed effects and ν_{ij} is a time-constant (directional) bilateral fixed effect.[24] Note that $\varphi_{i\tau}$ and $\phi_{j\tau}$ control for the multilateral resistance terms $\tilde{\Pi}_i$ and $\tilde{P}_j$, and the bilateral fixed effect also captures the time-invariant geography variables. In a second step, we re-estimate Equation (2.7) to obtain estimates for the coefficients of the time-invariant geography variables, δ_2 to δ_4. We therefore use only exporter- and importer-time-varying fixed effects and constrain the coefficient of PTA, δ_1, to the estimate of the first step, $\hat{\delta}_1$.

Finally, we use data from the last year in our sample, 2006, to estimate the elasticity of substitution and the elasticity of the matching function.

2.3.1 Estimation results

We present results estimating log-linearized trade flows by OLS as well as the Poisson pseudo-maximum-likelihood (PPML) estimator for the trade flows in levels following the recommendation by Santos Silva and Tenreyro (2006) in Table 2.2.

Columns (1)-(4) of Table 2.2 present results using bilateral fixed effects, i.e., assuming symmetric trade costs $t_{ij} = t_{ji}$ which is the same assumption made by Anderson and van Wincoop (2003). Columns (5)-(8) allow for asymmetric unobserved trade costs, i.e. $t_{ij} \neq t_{ji}$, by employing directional bilateral fixed effects. Each of these two blocks contains four specifications. Columns (1) and (5) report OLS estimates for scaled trade flows $z_{ij\tau}$ in logs. Column (2) and (6) present PPML estimates for the scaled trade flows in levels to control for heteroskedasticity and zero trade flows. Columns (3) and (7) reproduce Columns (1) and (5) for unscaled trade flows $x_{ij\tau}$. Finally, Columns (4) and (8) present PPML estimates for unscaled trade flows. The slightly larger number of observations for unscaled trade flows stems from the fact that GDP data are not available for all countries in all years where we have trade data and control variables.

Our estimates are in accordance with well-known results from the empirical trade literature. Distance is a large obstacle to trade, whereas contiguity, a common language and PTAs enhance trade. Comparing the results from Columns (1)-(4) with those of Columns (5)-(8) reveals that allowing for asymmetric trade costs does not substantially change our parameter estimates. Comparing with PPML estimates shows a clear pattern: distance coefficients are

[24] We report results for regressions including bilateral fixed effects, i.e.,$\nu_{ij} = \nu_{ji}$, and directional bilateral fixed effects, i.e., $\nu_{ij} \neq \nu_{ji}$.

Table 2.2: Estimation results for the OECD sample, 1950-2006

	(1) OLS $\ln z_{ij\tau}$	(2) PPML $z_{ij\tau}$	(3) OLS $\ln x_{ij\tau}$	(4) PPML $x_{ij\tau}$	(5) OLS $\ln z_{ij\tau}$	(6) PPML $z_{ij\tau}$	(7) OLS $\ln x_{ij\tau}$	(8) PPML $x_{ij\tau}$
Second stage								
$\ln DIST_{ij}$	-1.050***	-0.669***	-1.041***	-0.816***	-1.050***	-0.669***	-1.040***	-0.813***
	(0.009)	(0.027)	(0.010)	(0.010)	(0.009)	(0.027)	(0.010)	(0.010)
$CONTIG_{ij}$	0.097***	0.276***	0.116***	0.414***	0.097***	0.275***	0.115***	0.414***
	(0.019)	(0.030)	(0.019)	(0.018)	(0.019)	(0.030)	(0.019)	(0.018)
$COMLANG_{ij}$	0.386***	0.769***	0.387***	0.150***	0.386***	0.769***	0.387***	0.151***
	(0.019)	(0.049)	(0.019)	(0.017)	(0.019)	(0.049)	(0.019)	(0.017)
First stage								
$PTA_{ij\tau}$	0.274***	0.308***	0.267***	0.332***	0.274***	0.311***	0.276***	0.341***
	(0.016)	(0.019)	(0.017)	(0.019)	(0.014)	(0.016)	(0.015)	(0.013)
Estimated elasticities								
σ	2.349***	2.535***	2.349***	2.395***	2.349***	2.535***	2.350***	2.395***
	(0.303)	(0.051)	(0.024)	(0.728)	(0.352)	(0.195)	(0.255)	(0.476)
μ	0.946***	0.928***	0.947***	0.938***	0.946***	0.928***	0.947***	0.938***
	(0.003)	(0.007)	(0.001)	(0.009)	(0.005)	(0.007)	(0.003)	(0.008)
zero trade		X		X		X		X
symmetric $t_{ij\tau}$	X	X	X	X				
asymmetric $t_{ij\tau}$					X	X	X	X
N	36,945	37,741	37,493	38,313	36,945	37,741	37,493	38,313

Notes: Results for trade flows between 28 OECD countries between 1950 and 2006 estimated by ordinary least squares (OLS) and Poisson pseudo-maximum-likelihood (PPML). z_{ij} are trade flows standardized by importer and exporter GDPs. $\ln DIST$ is distance between exporting and importing country, $CONTIG$ is an indicator variable equal to 1 if the exporting and importing countries i and j share a common border, $COMLANG$ is an indicator variable equal to 1 if the exporting and importing country share a common official language, and PTA is an indicator variable equal to 1 if the exporting and importing country have signed a preferential trade agreement. All regressions control for multilateral resistance terms (MRTs) via exporter and importer fixed effects. (Robust) standard errors in parentheses, *** $p < 0.01$. Standard errors for σ and μ are bootstrapped using 200 replications.

smaller in absolute values, but all other coefficients are larger (except for the coefficients of $COMLANG$ in specifications (4) and (8)). The differences are larger for estimates using scaled trade rather than unscaled trade flows. Note that in the case of specifications using unscaled trade flows, GDP effects are captured by the time-varying importer- and exporter-fixed effects. Hence, those specifications implicitly allow for non-unitary GDP coefficients.

PTAs increase trade by 30.60 percent (Column (3)) to 40.64 percent (Column (8)) when neglecting general equilibrium effects.[25] The general equilibrium effects are accounted for in the counterfactual analysis, to which we turn in Section 2.3.2.

Turning to the elasticity of substitution, our significant estimates lie between 2.349 in Columns (1), (3), and (5) and 2.535 in Columns (2) and (6). These results are very much in line with recent evidence from Feenstra et al. (2012) who report estimates for the Armington elasticity between domestic and foreign goods of around 1 and between different foreign sources of 3.1. As our model forces these two elasticities to be equal, we would expect an estimate that lies in between these two estimates.[26]

Finally, our estimates of the matching elasticity vary between 0.928 and 0.947 and are significant at any standard level of significance. With our method, we find that the elasticity of labor markets in OECD countries indicates a very low level of labor market frictions and a very high matching elasticity compared to previous estimates. For example, Yashiv (2000) estimates μ between 0.2 and 0.6 for Israel for the years between 1975 and 1989. A literature review by Petrongolo and Pissarides (2001) reports estimates between 0.12 and 0.81 across studies focussing on several countries and time periods. Hall (2005) finds $\mu = 0.24$ for the United States for the years 2000 to 2002. Rogerson and Shimer (2011) estimate $\mu = 0.58$ for the same data for the years 2000 to 2009.[27] Even though our estimates are on the high side, note that our method infers the matching elasticity from (ratios) of bilateral trade flows using their cross-country-pair variation at one point in time. All other estimates of the matching elasticity in the literature use time series data on the number of matches, vacancies, and the unemployed from a single labor market. Hence, it is not too surprising that our estimates

[25] Effects are calculated as $(\exp(\hat{\delta}_{PTA}) - 1) \times 100$ percent.

[26] See Feenstra (2010) for a detailed discussion of estimates of the elasticity of substitution in international trade.

[27] Note that the literature reports both estimates of the matching elasticity with respect to the unemployed, as we do, or with respect to vacancies. In our discussion, we transformed the estimates when necessary assuming constant returns to scale in the matching process.

are somewhat different from the literature. In the counterfactual analysis, to which we turn next, we therefore provide results for alternative values of the matching elasticity.

2.3.2 Counterfactual analysis

We conduct two counterfactual experiments in our OECD sample. First, we evaluate the effects of PTAs. To this end, we compare a situation with PTAs as observed in 2006 with a counterfactual situation without any PTAs. Second, we evaluate improvements of labor market institutions in the United States and Germany.

Evaluating the effects of PTAs

Our first counterfactual experiment evaluates the effects of introducing PTAs as observed in 2006 compared to a counterfactual situation in which there are no PTAs. We base our counterfactual analysis on parameter estimates from Column (6) of Table 2.2 as they control for heteroskedasticity and impose unitary income elasticities for trade flows consistent with our framework.

The results are shown in Table 2.3.[28] It is organized as follows. Column (1), "PLM %GDP", gives the percentage change in nominal GDP in terms of the price index of Australia for the case of perfect labor markets. Column (2), "SMF %GDP", gives the same change within our search and matching framework. Columns (3) and (4) use Equation (2.16) and decompose the change in nominal GDP of Column (2) into price and employment changes. Column (5) reports the percentage change in the employment share for the case of imperfect labor markets, whereas Column (6) reports unemployment changes in percentage points. Finally, Columns (7) and (8) report the equivalent variation (EV) for the case of perfect and imperfect labor markets, respectively.

Table 2.3 reveals that all countries gain in terms of GDP when introducing PTAs as observed in 2006. This translates into an average gain in terms of GDP of 12.73 percent when assuming perfect labor markets. The average GDP gain increases by 4 percent to 13.28 percent when accounting for employment effects. Hidden behind these average effects is substantial heterogeneity. Some countries gain substantially more than the average, for example Canada with a gain of 20.70 percent, whereas other countries such as the United States experience a smaller increase of 9.92 percent. The decomposition of (log) GDP

[28] In Appendix B.8, we additionally provide results concerning the changes in trade flows across countries.

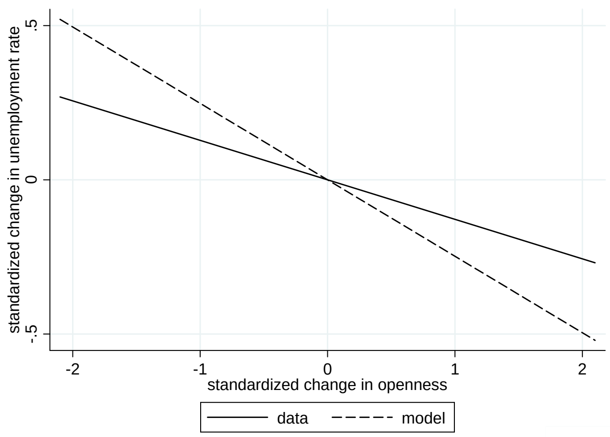


Figure 2.1: Implied regression lines of changes in openness and unemployment rates for both model and data.

change into (log) price and (log) employment changes highlights that for many of our sample countries, roughly 7 percent of the increase in GDP is driven by the increase in employment. Countries with only slight increases in GDP may even see negative employment effects, as can be seen in Column (5) of Table 2.3. Typically, welfare effects are magnified when taking into account employment effects. For example, the standard welfare estimate for Canada is about 5 percent larger when taking into account labor markets imperfections.

To assess the fit of our model, we first compare the implied changes in both openness (measured as imports plus exports over nominal GDP) and in unemployment rates predicted by our model with actually observed data for our sample. While it is straightforward to calculate these changes for our model, we cannot, of course, observe "real-world" counterfactual openness and unemployment rates. Thus, to compare model predictions with observed data, we take a simple and admittedly very crude approach: we calculate the observed change in openness and the unemployment rate as the change between the first year for which unemployment rate data are available and 2006.[29] Note that we standardized changes for comparison reasons. As can be seen from Figure 2.1, our model replicates the average negative correlation between openness and unemployment. The correlation between the fitted values of the two regression lines is 0.57.

[29] The first year is 1955 for the United States and Japan, 1956 for New Zealand, Ireland, France, and Canada, 1958 for Finland, 1959 for Italy, 1960 for Denmark and Turkey, 1961 for Greece, 1962 for Germany, 1964 for Australia and Austria, 1970 for Sweden, 1972 for Norway, Spain, and the United Kingdom, 1975 for Switzerland, 1983 for Belgium and the Netherlands, 1984 for Portugal, 1989 for Korea, 1990 for Poland, 1991 for Iceland, 1992 for Hungary, 1993 for the Czech Republic, and 1994 for the Slovak Republic. Note that all countries either had no or only a few PTAs in place for the first year in which we observe the unemployment rate, but all of them had experienced a tremendous increase in PTAs by 2006.

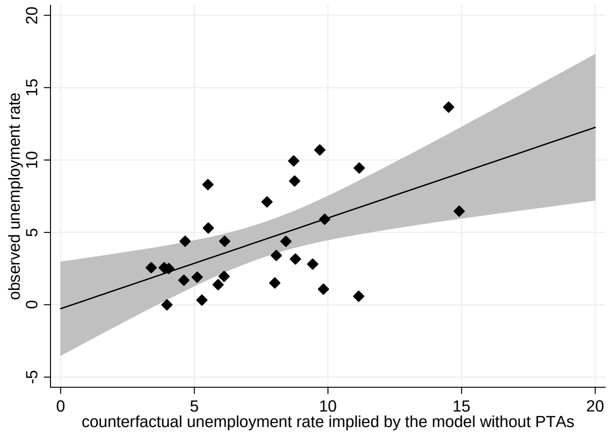


Figure 2.2: Regression of observed unemployment rate on the counterfactual unemployment rate implied by the model without PTAs.

As an additional validation of our results we compare observed unemployment rates in the first year available for our sample countries with the implied counterfactual unemployment rates without PTAs predicted by our model (see Figure 2.2). The correlation between the observed and predicted counterfactual unemployment rate is 0.54 which is tantamount to explaining 29 percent of the variation in the observed unemployment rate. Thus, although there is room for improving the model fit, we are the first to explain any of the observed variation in unemployment rates by changes in international trade policy changes.

As in every trade model, the resulting magnitudes of policy changes crucially depend on the exact values of the elasticities. We therefore test the sensitivity of our results to different values of the elasticity of substitution σ and the elasticity of the matching function μ. In the interest of brevity, we present only average effects in Table 2.4. The GDP, employment, and EV effects crucially depend on the values of σ and μ. When the elasticity of substitution increases, GDP, employment, and EV changes become smaller. This is because varieties are better substitutes, making trade less important. Hence, incepting PTAs leads to smaller predicted gains in terms of GDP, employment, and welfare. Changes in the elasticity of the matching function μ also show a clear pattern. Lower values of μ indicate higher GDP, employment, and welfare changes. A lower μ corresponds to larger labor market imperfections. When μ approaches 1 we end up in the case of perfect labor markets. The reason for this is that larger frictions on the labor market imply that firms have to post more vacancies in order to find a worker, effectively increasing recruiting costs. As trade liberalization decreases the overall price level, it also lessens a firm's recruiting costs.

This reduction of recruiting costs is more important in labor markets with higher frictions, making trade liberalization more attractive. Overall, Table 2.4 highlights that the extent of labor market frictions plays a crucial role in assessing the quantitative impact of free trade agreements.

Table 2.3: Comparative static effects of PTA inception controlling for trade imbalances in 2006

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	PLM	SMF	share %GDP SMF		SMF	SMF	PLM	SMF
	%GDP	%GDP	$\% \ln(\hat{p})$	$\% \ln(\hat{e})$	$\%\hat{e}$	Δu	%EV	%EV
Australia	16.45	17.40	92.75	7.25	1.17	-1.10	16.49	17.43
Austria	17.73	19.01	91.69	8.31	1.46	-1.37	20.59	22.12
Belgium	18.25	19.61	91.45	8.55	1.55	-1.40	21.92	23.57
Canada	20.70	22.16	90.60	9.40	1.90	-1.75	28.24	29.72
Czech Republic	17.29	18.50	91.95	8.05	1.38	-1.26	19.36	20.80
Denmark	16.71	17.84	92.28	7.72	1.28	-1.21	17.84	19.16
Finland	15.90	16.91	92.77	7.23	1.14	-1.04	15.72	16.90
France	15.70	16.71	92.88	7.12	1.11	-1.00	15.22	16.43
Germany	15.27	16.22	93.31	6.69	1.01	-0.90	13.77	14.91
Greece	15.62	16.60	92.92	7.08	1.10	-0.99	15.10	16.24
Hungary	16.79	17.92	92.24	7.76	1.29	-1.18	18.01	19.35
Iceland	15.36	16.26	93.17	6.83	1.04	-1.00	14.28	15.29
Ireland	16.19	17.20	92.66	7.34	1.17	-1.11	16.35	17.49
Italy	15.22	16.15	93.27	6.73	1.01	-0.94	13.83	14.94
Japan	9.25	9.28	101.03	-1.03	-0.09	0.09	-1.24	-1.26
Korea	9.39	9.44	100.71	-0.71	-0.06	0.06	-0.90	-0.89
Netherlands	16.86	18.01	92.32	7.68	1.28	-1.21	17.86	19.23
New Zealand	10.49	10.72	98.70	1.30	0.13	-0.13	1.61	1.85
Norway	16.38	17.45	92.55	7.45	1.21	-1.15	16.78	18.02
Poland	16.58	17.69	92.34	7.66	1.26	-1.07	17.53	18.83
Portugal	16.02	17.04	92.70	7.30	1.16	-1.06	16.03	17.21
Slovak Republic	17.05	18.22	92.08	7.92	1.34	-1.14	18.72	20.11
Spain	15.15	16.07	93.25	6.75	1.01	-0.92	13.86	14.93
Sweden	16.17	17.22	92.61	7.39	1.18	-1.09	16.39	17.62
Switzerland	18.50	19.89	91.31	8.69	1.59	-1.51	22.66	24.34
Turkey	15.58	16.54	93.00	7.00	1.08	-0.96	14.87	15.97
United Kingdom	13.61	14.31	94.49	5.51	0.74	-0.70	9.92	10.72
United States	9.92	10.08	99.63	0.37	0.04	-0.03	0.30	0.49
Average	12.73	13.28	96.59	3.41	0.55	-0.50	7.53	8.16

Notes: Counterfactual analysis is based on parameter estimates from column (6) of Table 2.2. PLM gives results assuming perfect labor markets. SMF gives results using a search and matching framework for the labor market. Averages are weighted averages using country GDP as weight.

Evaluating the effects of labor market reforms

In our second counterfactual experiment, we evaluate the effects of a hypothetical labor market reform which improves U.S. labor market institutions. We implement this by a 3

Table 2.4: Average comparative static effects of PTA inception controlling for trade imbalances for various parameter values

μ	σ	PLM %GDP	SMF %GDP	SMF $\%\hat{e}$	SMF $\%\Delta u$	PLM %EV	SMF %EV
0.2	5	4.81	16.68	11.91	-9.24	2.75	15.25
	10	2.13	7.11	5.00	-4.22	1.20	6.33
	15	1.37	4.51	3.16	-2.74	0.77	3.98
0.5	5	4.81	7.54	2.75	-2.41	2.75	5.67
	10	2.13	3.32	1.20	-1.08	1.20	2.44
	15	1.37	2.13	0.77	-0.70	0.77	1.55
0.75	5	4.81	5.69	0.90	-0.81	2.75	3.71
	10	2.13	2.52	0.40	-0.36	1.20	1.61
	15	1.37	1.62	0.25	-0.23	0.77	1.03
0.9	5	4.81	5.10	0.30	-0.27	2.75	3.07
	10	2.13	2.26	0.13	-0.12	1.20	1.34
	15	1.37	1.45	0.08	-0.08	0.77	0.85
0.99	5	4.81	4.83	0.03	-0.03	2.75	2.78
	10	2.13	2.14	0.01	-0.01	1.20	1.21
	15	1.37	1.37	0.01	-0.01	0.77	0.78

Notes: Table reports average changes in nominal GDP, employment, and the equivalent variation in percent assuming either a perfect labor market (PLM) or using a search and matching framework (SMF) for the labor market controlling for trade imbalances with varying elasticity of substitution σ and elasticity of the matching function μ. The remaining parameters are set to values from column (6) of Table 2.2.

percent increase in $\hat{\kappa}_j$ for the United States, i.e., we set $\hat{\kappa}_{U.S.}$ to 1.03. Given our estimate of the matching elasticity of $\mu = 0.928$, this change in $\hat{\kappa}_{U.S.}$ corresponds to either an increase of 2.8 percent in the overall matching efficiency m_j or a 32 percent reduction of recruiting costs in the United States. Note that within our framework we do not necessarily have to specify the explicit source of changes in labor market institutions. The results of this experiment are set out in Table 2.5.[30]

All countries gain in terms of GDP when U.S. labor market institutions improve. This highlights the positive spillover effects, recently theorized by Egger et al. (2012) and Felbermayr et al. (2013), and documented empirically in a reduced-form setting in Felbermayr et al. (2013). Of course, when perfect labor markets are assumed, it is not possible to evaluate any change in them. Therefore, Columns (1) and (7) are uninformative. The decomposition of (log) GDP into (log) price and (log) employment changes highlights that in the United States prices fall and all increases in GDP are due to increases in employment. For the trading partners of the United States, the positive GDP effects are composed of roughly

[30] Again, detailed results on the heterogeneous trade effects can be found in Appendix B.8.

Table 2.5: Comparative static effects of $\hat{\kappa}_{U.S.} = 1.03$ controlling for trade imbalances in 2006

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	PLM	SMF	share %GDP SMF		SMF	SMF	PLM	SMF
	%GDP	%GDP	$\% \ln(\hat{p})$	$\% \ln(\hat{e})$	$\%\hat{e}$	Δu	%EV	%EV
Australia	0.00	0.79	92.75	7.25	0.06	-0.05	0.00	0.77
Austria	0.00	0.50	98.72	1.28	0.01	-0.01	0.00	0.09
Belgium	0.00	0.48	99.41	0.59	0.00	-0.00	0.00	0.04
Canada	0.00	0.96	90.76	9.24	0.09	-0.08	0.00	1.21
Czech Republic	0.00	0.52	98.14	1.86	0.01	-0.01	0.00	0.13
Denmark	0.00	0.53	97.89	2.11	0.01	-0.01	0.00	0.15
Finland	0.00	0.56	97.15	2.85	0.02	-0.01	0.00	0.21
France	0.00	0.52	98.23	1.77	0.01	-0.01	0.00	0.12
Germany	0.00	0.52	98.28	1.72	0.01	-0.01	0.00	0.12
Greece	0.00	0.55	97.34	2.66	0.01	-0.01	0.00	0.20
Hungary	0.00	0.53	97.73	2.27	0.01	-0.01	0.00	0.16
Iceland	0.00	0.62	95.59	4.41	0.03	-0.03	0.00	0.37
Ireland	0.00	0.59	96.30	3.70	0.02	-0.02	0.00	0.29
Italy	0.00	0.53	97.81	2.19	0.01	-0.01	0.00	0.16
Japan	0.00	0.55	97.53	2.47	0.01	-0.01	0.00	0.18
Korea	0.00	0.55	97.34	2.66	0.01	-0.01	0.00	0.20
Netherlands	0.00	0.51	98.48	1.52	0.01	-0.01	0.00	0.10
New Zealand	0.00	0.73	93.58	6.42	0.05	-0.04	0.00	0.64
Norway	0.00	0.56	97.17	2.83	0.02	-0.01	0.00	0.21
Poland	0.00	0.53	97.78	2.22	0.01	-0.01	0.00	0.16
Portugal	0.00	0.56	96.88	3.12	0.02	-0.02	0.00	0.24
Slovak Republic	0.00	0.53	97.83	2.17	0.01	-0.01	0.00	0.16
Spain	0.00	0.55	97.23	2.77	0.01	-0.01	0.00	0.21
Sweden	0.00	0.55	97.44	2.56	0.01	-0.01	0.00	0.19
Switzerland	0.00	0.48	99.47	0.53	0.00	-0.00	0.00	0.03
Turkey	0.00	0.56	96.99	3.01	0.02	-0.01	0.00	0.23
United Kingdom	0.00	0.62	95.71	4.29	0.03	-0.02	0.00	0.36
United States	0.00	2.55	-16.54	116.54	2.97	-2.83	0.00	2.54
Average	0.00	1.30	55.11	44.89	1.11	-1.06	0.00	1.10

Notes: Counterfactual analysis is based on parameter estimates from column (6) of Table 2.2. PLM gives results assuming perfect labor markets. SMF gives results using a search and matching framework for the labor market. Averages are weighted averages using country GDP as weight.

97 percent of price changes and 3 percent changes in employment. This can also be seen when comparing the relative magnitudes of the employment changes reported in Column (5) of Table 2.5. Concerning welfare, obviously the United States profit the most from its improvements in labor market institutions, with an increase in welfare of 2.54 percent. However and importantly, all other countries also gain, with the highest gains for Canada at 1.21 percent.

We also analyzed the recent German labor market reforms implemented between 2003 and 2005.[31] These reforms reduced unemployment benefits to increase search incentives for unemployed workers and are thought to have increased the overall matching efficiency of German labor markets.[32] For our counterfactual scenario, we reduce the matching efficiency by 5 percent and increase the replacement rate to the level prevailing in 2003. We find that unemployment in Germany would be about 4 percentage points higher and GDP more than 4 percent lower were it to undo its recent labor market reforms.

2.4 Conclusion

State of the art frameworks for quantitative analyses of international trade policies to evaluate the trade and welfare implications of trade liberalization all assume perfect labor markets. However, net employment effects are at the heart of the political debate on trade integration. Accordingly, recent developments in international trade theory have highlighted the link between trade liberalization and labor market outcomes.

We build on these theoretical contributions to develop a quantitative framework of bilateral trade flows which takes into account labor market frictions within a search and matching framework. Our model allows counterfactual analysis of changes in trade costs and labor market reforms on trade flows, prices, employment, and welfare.

We apply our structural model to a sample of 28 OECD countries from 1950 to 2006 to evaluate the effects of preferential trade agreements (PTAs) and labor market reforms in the United States and Germany. We find that introducing PTAs as observed in 2006 leads to greater GDP increases when accounting for aggregate employment effects. Countries with only slight increases in GDP see negative employment effects. Our second counterfactual analysis assumes an improvement of labor market institutions in the United States. Average

[31] Results can be found in Appendix B.8.
[32] Fahr and Sunde (2009) estimate this increase to be about 5 percent.

welfare effects are substantially magnified when taking into account employment effects. U.S. GDP increases roughly five times more than GDP of the other countries. While the United States profits the most from improvements of its labor market institutions with an equivalent variation of 2.54 percent, all of its trading partners also experience an increase in welfare due to positive spillover effects.

As our approach does not require any information about the labor market except for the elasticity of the matching function, it can be easily applied to any other field in which the gravity equation is employed.

Chapter 3

Preferential Trade Agreements, Unemployment, and the Informal Sector*

3.1 Introduction

What are the welfare consequences of preferential trade agreements? And what are their employment effects? These questions are of major concern for policy makers in both developed and emerging economies. To answer the first question, trade economists have delevoped quantitative models of international trade which allow to analyze the effect of trade liberalization on aggregate trade flows and welfare, taking into account the interdependencies of trade flows between trading partners. Today, these structural gravity frameworks in the vein of Anderson and van Wincoop (2003) are the de facto industry standard to answer the first question. Interestingly, these frameworks have to remain silent on the second question, as they do not model employment, or assume full employment. Hence, in these type of models, trade liberalization cannot have any (net) employment effects, as all workers are assumed to be employed before and after a trade liberalization scenario.[1] An exception to this approach is Heid and Larch (2012a) who estimate employment effects of preferential

* A previous version of this paper has been circulated under the title "Trade Liberalization, Unemployment, and the Informal Sector".

[1] Whereas Anderson and van Wincoop (2003) present a model driven by love of variety considerations of consumers, Eaton and Kortum (2002) present a quantitative trade model with Ricardian technology differences across countries. Despite their differences in interpretation, Arkolakis et al. (2012) show that both models have the same quantitative welfare implications. Other quantitative trade models which can in principle be used for the evaluation of trade liberalization episodes which are not covered by the Arkolakis et al. (2012) equivalence are e.g. Waugh (2010) and Fieler (2011). All these frameworks assume full employment.

trade agreements for a sample of OECD countries by introducing a unified labor market, characterized by search and matching frictions, into a structural gravity framework.[2]

However, labor markets in emerging economies are remarkably distinct from labor markets in developed economies like the OECD countries. For example, irrespective of the variety of definitions used, informal employment comprises between 25 to more than 70 percent of the labor force in Latin American countries. Informal employment is not only restricted to Latin America, however: In general, the share of informal workers is higher in countries with lower GDP per capita (see Perry et al., 2007). The informal sector is characterized by low productivity, small scale establishments. Informal workers are often self-employed, or, when they work as employees, do not possess a written labor contract, or do not have access to social security or health insurance (see ILO, 2010). Therefore, informal sector employment has generally been seen as detrimental for the welfare of workers.[3]

In this paper, I extend the structural gravity framework of Heid and Larch (2012a) by introducing an informal sector to study the impact of trade liberalization on welfare, unemployment, as well as the size of the informal sector. To illustrate, I apply my quantitative framework to a set of 13 Latin American and Caribbean countries and use it to evaluate the welfare and employment effects of preferential trade agreements signed since 1950. I find that these preferential trade agreements have, on average, decreased welfare by 7.6 percent, decreased informal employment by 50.9 percent, and increased the official unemployment rate by 3.1 percentage points. These results are quantitatively and qualitatively different from standard frameworks assuming either full employment or a unified labor market with search and matching frictions.

The literature uses several definitions of informality or informal employment. Following Gasparini and Tornarolli (2009), informality can either be defined using a productive or legalistic definition. The productive definition declares a worker to be informal when she is an unskilled self-employed, is employed in a small scale establishment, or does not receive a monetary reward for her work but is paid in kind. According to the legalistic definition, a worker is declared informal if she does not possess a written labor contract, or does not have access to social security (mostly the pension system) or health insurance.[4] Both definitions can also focus on firms instead of individual workers, and both definitions have deficiencies.

[2] See Chapter 2.

[3] For example, Attanasio et al. (2004) find that informal employment is correlated with lower job satisfaction and generally worse job conditions in Colombia.

[4] For an in depth review of social security and its relation to informal employment see ILO (2010).

For example, small scale establishments need not necessarily be informal or employ informal workers. In addition, it may well be that larger firms partly employ informal workers, e.g. a firm may pay social security contributions for its manufacturing workers but employ a parking lot attendant informally. Therefore, depending on the specific definitions used, the share of informal workers as a percentage of the labor force varies; however, the measures correlate substantially (see Gasparini and Tornarolli, 2009). Irrespective of the respective definition used, informal employment is characterized by low productivity and hence low wages. Informal establishments are also characterized by no strict distinction between private and firm accounts, and often, workers are family members or close relatives (see de Laiglesia and Jütting, 2009 and de Mel et al., 2009).

Early attempts at modeling the informal sector theoretically treat it as a last resort for workers who did not manage to find a job in the formal part of the economy where regulatory restrictions like minimum wages prevent that workers can bid down wages (see Harris and Todaro, 1970). Maloney (2004) challenges this view by noting that informal employment is a multi-faceted phenomenon: Whereas informal employment is the last resort for some workers for want of better employment opportunities, others voluntarily leave the formal sector to start their own informal business. Accordingly, Albrecht et al. (2009) stress that worker differences in formal sector productivity can explain a voluntary sorting of high-skill workers into the formal sector. Empirical evidence about these two competing views is mixed. If informal employment collects workers which are queuing for formal sector jobs, then the share of informal workers should increase during recessions. Instead, if informal employment is a voluntary decision, it should not be related to the business cycle or could also be pro cyclical. Fiess et al. (2010) study the comovement of the informal sector with the overall business cycle in several Latin American countries and find that both views are supported by the data, depending on the country and time period studied.[5]

A different strand of the literature dealing with informal employment was started by Lewis (1954) who describes a model of an economy with two sectors: One modern "capitalist" sector of formal salaried workers, and a "subsistence" sector where workers engage in income sharing. Crucially, workers in the subsistence sector can leave the sector without reducing its output by much as the remaining workers can increase productivity by reorganizing jobs. While most of the subsequent literature has identified the latter with traditional agriculture, Lewis

[5] Günther and Launov (2012) also find that both views describe parts of the reality of informal employment in Côte d'Ivoire.

himself also envisaged petty workers in low productivity jobs which are nowadays associated with the informal sector.[6]

The empirical literature on the informality-trade nexus is rather small and has focused on case-studies for single countries, often using micro-level data sets of workers. Goldberg and Pavcnik (2003) find an increase in informality after trade liberalization episodes in the 1980s and 1990s in Colombia; they do not find such an effect in Brazil. Using time series data on the in- and outflows into and from informality, Bosch et al. (2012) also study the effect of trade liberalization during the same period in Brazil and find that it accounts for about an 1 to 2.5 percent increase in informal employment. Fiess et al. (2010) investigate the empirical implications of a small open economy macro model with a tradeable formal and a non-tradable informal sector for Argentina, Brazil, Colombia, and Mexico. In their model, trade liberalization can be interpreted as an increase in the productivity of the tradable sector which leads to a decline in informality along standard Stolper and Samuelson (1941) type arguments. Coşar et al. (2011) estimate a structural dynamic heterogeneous firm model to evaluate the impact of the trade liberalization episodes from the 1990s on informality in Colombia but find little to no effect. Arias et al. (2013) analyze the effects of a hypothetical tariff reduction on informal employment in Brazil and Mexico estimating dynamic discrete choice models for workers who chose in which sector to work. They find a slight increase in informal employment. Finally, Heid et al. (2013) use a calibrated heterogeneous firm model to study informality in Mexico during the 1990s and find that informality has slightly increased due to an increase in U.S. offshoring.[7]

All these studies stick to a small open economy assumption, i.e. they analyze the effect of trade liberalization for a single country. Hence they abstract from the interdependence of trade flows between trading countries as well as income effects, key features of the structural gravity models used for evaluating the welfare consequences of trade liberalization mentioned in the beginning. Importantly, as Egger et al. (2011) illustrate, these effects also matter quantitatively for the evaluation of preferential trade agreements.

The remainder of this paper is structured as follows: Section 3.2 presents a simple quantitative framework of international trade in the presence of search-generated unemployment and an informal sector. Section 3.3 illustrates how this framework can be used to counterfactually evaluate the effects of a change in trade costs brought about by e.g.

[6] The term “informal sector” only was used about 20 years later by Hart (1973). Harris and Todaro (1970) talk about “urban unemployment”, but do not use the terms “informal sector” or “informal employment”.

[7] See Chapter 4.

preferential trade agreements. Section 3.4 brings the model to the data, followed by the evaluation of the effects of preferential trade agreements signed between 13 Latin American Caribbean countries since 1950 in Section 3.5. Section 3.6 concludes.

3.2 The model

3.2.1 The decision of the worker

Every country j is populated by a representative household with labor endowment L_j. The household can decide how many members should work in the formal or informal sector, L_j^f and L_j^i, respectively.; hence $L_j = L_j^f + L_j^i$. Superscripts f and i will henceforth denote variables in the formal and informal sector, respectively. Once household members have chosen their sector, they cannot switch sectors.[8] Note that household members do not differ in terms of ability. As I am only interested in the impact of trade liberalization on the overall size of the informal sector, I abstract from the sorting of workers into different sectors.[9]

Workers who have chosen to work in the formal sector have to search for a job. Due to search frictions, a share $u_j^f L_j^f$ of formal sector workers is unemployed, where u_j^f denotes the probability that workers who chose to search in the formal sector will not find a formal job and hence will be unemployed. The unemployed receive a lump-sum transfer from the employed workers in the formal sector of $\gamma_j w_j^f$, where γ_j is the rate of unemployment benefits as a fraction of the formal sector wage w_j^f.

Workers who have chosen to work in the informal sector instantaneously find a job, as they can always become self-employed. Hence there is no informal unemployment. Several authors argue that informal employment is not subject to search frictions in the labor market: Zenou (2008) argues that formal employment is preceded by a more or less formal application process whereas informal workers can always set up shop in the informal sector and become

[8] While this is a strong assumption, allowing workers to switch between sectors is arguably important for modeling transitions of workers between formal and informal employment along the business cycle. This paper, however, focuses on the cross-country variation in experiences of the trade-informality nexus, following the international trade literature by deliberately abstracting from short-run fluctuations in economic activity. For a discussion of the cyclicality of informality, see e.g. Bosch and Maloney (2010), Fiess et al. (2010), and Bosch and Esteban-Pretel (2012).

[9] See Gasparini and Tornarolli (2009) for which types of workers sort into informality.

self-employed. Similar arguments are used by Wahba and Zenou (2005) and Heid et al. (2013).[10]

In equilibrium, a member of the risk-neutral household has to be indifferent between formal and informal employment, i.e.

$$(1-u_j^f)w_j^f + u_j^f \gamma_j w_j^f - f_j^f w_j^f = w_j^i, \tag{3.1}$$

which is similar to the setup in Helpman and Itskhoki (2010) which also essentially restate a variant of the equilibrium condition in Harris and Todaro (1970). Different to Harris and Todaro (1970), I abstract from employment in the agricultural sector. In both Helpman and Itskhoki (2010) and the present model, wages are not set exogenously but are determined in general equilibrium.[11]

In addition to the search effort, workers who have chosen to work in the formal sector have to incur a cost $f_j^f w_j^f$. These costs can be interpreted as moving costs, taxes and contributions to finance other social security provisions than unemployment benefits.[12] These taxes may even be wasteful, at least from the perspective of the worker. In many Latin American countries, formal sector social security and health care provisions often include free insurance for family members so that often only one family member works in the formal sector. For example, in Colombia, about 54 percent of informal self-employed workers do not contribute to health insurance as they have access through a relative, see Perry et al. (2007). Finally, it can also be the monetary equivalent of the cost of being a salaried worker instead of being one's own boss as a self-employed worker as stressed by Maloney (2004). The assumption of entry fixed costs of formal employment are also in line with empirical evidence provided by Arias et al. (2013) who find that entry costs into formal employment are substantially larger than for informal employment. In the empirical application, I will solve for $f_j^f w_j^f$ so that workers are indifferent between the two sectors using the observed data. Therefore, f_j^f

[10] Amaral and Quintin (2006) also reject the notion of search frictions or barriers to entry into the informal sector; instead, they argue that even formal labor markets are competitive.

[11] The household interpretation is needed in order to entice some workers to search for a job in the formal sector when there is no unemployment insurance, see Helpman and Itskhoki (2010). Unemployment insurance is scant at best or completely absent in most countries which are characterized by large rates of informal employment as e.g. Latin American countries. Therefore, self-employment acts as the de facto unemployment insurance at the household level in many developing and emerging countries.

[12] Note that I abstract from explicitly modeling the demand and supply of a public good like e.g. publicly provided health care or a public pension system. Hence, in the context of the model, the formal sector fixed costs are pure costs for formal sector workers.

captures in a catch-all way the several factors which prevent Equation (3.1) to hold without any entry costs.

3.2.2 Formal and informal firms

Firms in the formal sector have to pay a cost c_j to open their one worker firm.[13] They then have to search for a worker in order to start production. Hence this entry cost can be interpreted as vacancy posting costs for searching a worker as well as general fixed costs of production like complying with formal sector regulatory requirements like statistical duties etc. if we assume that firms are one-worker firms.[14] These costs are paid in terms of formal sector output whose aggregate price is P_j^f. Hence, they can also be interpreted as a form of capital requirement to set up a firm, as c_j is denoted not in terms of labor but in terms of the final output good. The formal labor market is characterized by search frictions according to a one-shot version of a Pissarides (2000) type model.[15] At the beginning of the period, all household members who have chosen the formal sector are unemployed. The number of successful matches M_j between unemployed workers L_j^f and formal sector vacancies V_j is characterized by the following constant returns to scale matching function:

$$M_j = m_j (L_j^f)^{\mu} V_j^{1-\mu}, \tag{3.2}$$

where μ is the elasticity of matches with respect to the number of the unemployed and m_j is a measure of the overall matching efficiency of the labor market. This implies that workers who search for a formal job will find formal employment with probability $M_j/L_j^f = m_j \vartheta_j^{1-\mu}$ where ϑ_j is a measure of the formal labor market tightness and is defined as $\vartheta_j \equiv V_j/L_j^f$. From this we can define the probability of not finding a job in the formal sector as $u_j^f = 1 - m_j \vartheta_j^{1-\mu}$. Note that this is not the overall or official unemployment rate in the economy which is reported by national statistical agencies. It is defined as the number of unemployed, U_j, divided by the labor force, hence $u_j^o = U_j/L_j$, where o is short for official. As informal sector

13 The following description of the behavior of formal firms draws heavily from Felbermayr et al. (2013) and Heid and Larch (2012a) as it borrows the labor market model used there.

14 This is without loss of generality if total setup costs of a firm are a linear function of the number of workers.

15 For a general discussion of one-shot models of search and matching frictions see Rogerson et al. (2005). One-shot labor market models are increasingly used in international trade if one is willing to abstract from the business cycle. Some examples are Keuschnigg and Ribi (2009), Helpman and Itskhoki (2010), and Felbermayr et al. (2013).

workers are not unemployed, this may explain low official unemployment rates in countries with a large informal sector.

The probability that a formal sector firm will fill its vacancy is given by $M_j/V_j = m_j\vartheta_j^{-\mu}$, and expected firm setup costs are $V_j/M_j c_j P_j^f$. After a successful match between a worker and a formal firm has been established, I assume that both parties bargain over the match surplus according to a generalized Nash bargaining solution. The surplus of the worker is the wage she gains minus her outside option. As the worker's decision for a sector is irreversible, her outside option in the formal sector is the unemployment benefit b_j, i.e. the worker's surplus is given by $w_j^f - b_j$. In equilibrium, $b_j = \gamma_j w_j^f$. Having sunk its setup costs, the surplus of the firm is the price for which it can sell the output minus the wage cost, i.e. $p_j - w_j^f$. Hence, the Nash bargaining solution wage maximizes $(w_j^f - b_j)^{\xi_j}(p_j^f - w_j^f)^{1-\xi_j}$, where ξ_j is the bargaining power of the worker and $\xi_j \in (0,1)$. The first order condition of the bargaining problem yields the formal wage curve $w_j^f = \xi_j/(1+\gamma_j\xi_j - \gamma_j)p_j^f$.[16] As the fraction on the right-hand side of the wage curve is always smaller than 1, workers get paid less than their marginal value product. Note that due to the one-shot nature of the model, the wage curve does not depend on the formal labor market tightness ϑ_j.

Firms enter the formal sector until expected setup costs equal firm profits, i.e. until

$$m_j^{-1}\vartheta_j^{\mu} c_j P_j^f = p_j^f - w_j^f, \tag{3.3}$$

which can be reformulated to get the job creation curve $w_j^f = p_j - c_j P_j^f m_j^{-1}\vartheta_j^{\mu}$.

Equilibrium formal labor market tightness is determined by the intersection of the wage and job creation curves and is given by

$$\vartheta_j = \left(\frac{p_j^f}{P_j^f}\right)^{1/\mu}\left(\frac{c_j}{m_j}\frac{1-\gamma_j+\gamma_j\xi_j}{1-\gamma_j+\gamma_j\xi_j-\xi_j}\right)^{-1/\mu}. \tag{3.4}$$

Equation (3.4) reveals that formal labor market tightness is determined by p_j^f/P_j^f, the real price of the formal sector output good. If country j consumes goods from abroad, any reduction in the prices of imports directly feeds into a reduction of the general price level in country j, which in turn affects the country's formal labor market tightness and hence

[16] Note that I follow Pissarides (2000) in assuming that both the firm and the worker do not take into account that their bargaining affects the level of unemployment benefits b_j. Felbermayr et al. (2013) and Heid and Larch (2012a) use the same model of the labor market but I extend their frameworks to include informality.

the probability of becoming unemployed in the formal sector.[17] Also note that formal labor market tightness does neither depend on the relative or absolute size of the formal sector. Hence the number of unemployed workers is determined only by institutional parameters of the formal labor market and the prevailing price level which is determined in general equilibrium.

Let us now turn to production in the informal sector. Workers who have chosen to become self-employed in the informal sector do not have to incur firm setup and worker search costs. They produce the same good as workers in the formal sector. Hence, the price of the good is the same, irrespective of whether it was produced in the formal or informal sector, i.e. $p_j^f = p_j^i = p_j$. This can be rationalized by the fact that consumers do not care about the working conditions under which a good has been produced, as I assume that consumers only derive utility from the consumption of a good. In principle, one could also assume that informal sector firms produce a different good, and that there exists some imperfect substitutability between the goods. However, I argue that this is not satisfactory on conceptual grounds. When one assumes that utility of consumers is given by a Cobb-Douglas or CES composite of formally and informally produced goods, the informal sector is assumed into existence by consumer preferences instead of institutional features of the labor market or the economy.[18]

The production mode in the informal sector is different to that in the formal sector. Whereas the formal sector is organized along capitalist lines where firms equate marginal benefit to marginal cost to determine how many workers to employ, informal sector firms engage in income sharing. Therefore, the informal sector wage, w_j^i, is equal to the average product of an informal sector firm. In addition, the informal sector is characterized by what Lewis (1954) described as 'surplus labor'. Informal sector establishments are often organized around families, do not distinguish between family and firm accounts and employ family members or workers who often do not get a monetary wage but are paid in kind. Crucially, Lewis argues that if an additional worker is employed in a informal establishment,

[17] The same mechanism is used in Heid and Larch (2012a) as well as Felbermayr et al. (2011a) and Felbermayr et al. (2013) but applies to the economy as a whole; Helpman and Itskhoki (2010) use a similar mechanism in a two sector setup with comparative advantage.

[18] Other authors who do not distinguish between consumption derived from formally and informally produced goods are e.g. Rauch (1991), Dessy and Pallage (2003), Goldberg and Pavcnik (2003), Amaral and Quintin (2006), Chong and Gradstein (2007), Marjit et al. (2007), Galiani and Weinschelbaum (2012), and Arias et al. (2013). Notable exceptions are Fiess et al. (2010) who assume that the non-tradable sector is identical to the informal sector and tradable and non-tradable goods are imperfect substitutes as well as Ulyssea (2010) who assumes that the final consumption good is a CES composite of formally and informally produced intermediate goods.

productivity is reduced as work is simply shared amongst the family members. Whereas the subsequent literature applied the concept of surplus labor or disguised unemployment to the traditional agricultural sector, Lewis himself emphasized that the same reasoning can be applied to petty workers which are today associated with the informal sector. Several authors provide micro foundations and develop the implications of this mode of production (see e.g. Sen, 1966 and Takagi, 1978 as well as the gentle introduction in Chapter 10 in Ray, 1998). I simply assume that productivity in the informal sector is a decreasing function of the relative size of the informal sector, i.e. informal sector output is given by

$$y_j^i = \left(\frac{L_j^f}{L_j}\right)^\alpha = \left(1 - \frac{L_j^i}{L_j}\right)^\alpha . \tag{3.5}$$

Hence, informal sector firms are less productive than formal sector firms which have a constant productivity of 1. α can be interpreted as the elasticity of an informal sector firm's productivity with respect to a change in the relative size of the formal sector. If the formal sector employment share increases by one percent, informal sector productivity increases by α percent.

This productivity disadvantage of informal firms is compensated by a lower informal sector wage. As mentioned before, informal sector firms engage in income sharing, i.e. the informal sector wage is the value of the average informal sector output:

$$w_j^i = \frac{p_j y_j^i L_j^i}{L_j^i} = p_j \left(\frac{L_j^f}{L_j}\right)^\alpha . \tag{3.6}$$

Equivalently, one can assume that informal products can only be sold at a discount due to their lower quality, or because consumers cannot enforce their contract in the sense that they cannot enforce producer liability in case the product does not meet its advertised standard. Both interpretations are consistent with the data which show that informal workers have, on average, lower wages.[19]

[19] I report the formality to informality wage ratio from the data set used in the empirical application in Table 3.2. For further evidence, see also Gasparini and Tornarolli (2009) and Pratap and Quintin (2006). In the data, informal workers tend to sort into the formal sector according to skill-levels in a Roy (1951) type fashion. Still, assuming a productivity penalty in the informal sector or assuming sorting of less productive workers into the informal sector is observationally equivalent if one is not interested in who selects into the informal sector but in the analysis of the aggregate effect of trade liberalization on informality.

Combining the wage curve and Equation (3.6) determines the equilibrium formal sector wage premium as

$$\frac{w_j^f}{w_j^i} = \left(\frac{L_j^f}{L_j}\right)^{-\alpha} \frac{1 + \gamma_j \xi_j - \gamma_j}{\xi_j}. \tag{3.7}$$

Inspection of Equation (3.7) shows that for $\alpha > 0$ the formal sector wage premium decreases with the size of the formal sector. The larger the formal sector, the smaller is the productivity disadvantage of the informal sector. This increases the informal sector wage and therefore reduces the formality premium.

3.2.3 Consumers and determination of trade flows

Having specified the labor market and the production structure, I describe preferences and consumer decisions which endogeneously determine international trade. I use the simplest model to generate trade between countries by following Armington (1969) who assumes that goods are differentiated across n countries.[20] The utility function of the representative household in country j is given by

$$U_j = \left[\sum_{i=1}^{n} \beta_i^{\frac{1-\sigma}{\sigma}} q_{ij}^{\frac{\sigma-1}{\sigma}}\right]^{\frac{\sigma}{\sigma-1}}, \tag{3.8}$$

where q_{ij} denotes the quantity of goods from country i consumed in country j, σ is the elasticity of substitution between varieties, and β_i is a preference parameter which reflects the relative attractiveness of goods from country i. Note that consumers do not differentiate between formally and informally produced goods. One could describe an observationally equivalent model where goods produced in the informal sector are of lower quality, when lower quality is interpreted as lower "effective" consumption of the good. Transporting goods from country i to j incurs (potentially asymmetric) iceberg-type transport costs t_{ij} such that the price of a good from country i in country j, p_{ij}, is given by $t_{ij}p_i$, where p_i is the price of the good at the factory gate.

[20] Arkolakis et al. (2012) show that the trade structure arising from this setting is observationally equivalent for a wider class of more complex trade models including Ricardian technology differences between countries as in Eaton and Kortum (2002) or heterogeneous firms as in Melitz (2003). Heid and Larch (2012a) demonstrate that this isomorphism is true even in models with aggregate employment effects similar to the model in this paper.

The representative household maximizes Equation (3.8) subject to its budget constraint $y_j = \sum_{i=1}^{n} p_i t_{ij} q_{ij}$, i.e. national income or GDP is given by the sum of sales. We can also generalize this budget constraint by allowing for exogenously given trade deficit shares. Then, the budget constraint becomes $\tilde{y}_j = \sum_{i=1}^{n} p_i t_{ij} q_{ij}$, where $\tilde{y}_j = y_j(1 + d_j)$, with y_j denoting nominal income in country j and d_j the share of the trade deficit (if $d_j > 0$) or surplus (if $d_j < 0$) of country j as a percentage of GDP.[21]

Note that sales include domestic and international sales by both formal and informal firms.[22] Utility maximization then yields the following expression for sales of goods from country i in country j:

$$x_{ij} = p_i t_{ij} q_{ij} = \left(\frac{\beta_i p_i t_{ij}}{P_j}\right)^{1-\sigma} \tilde{y}_j, \tag{3.9}$$

where P_j is the ideal price index given the CES utility function and is defined by $P_j = [\sum_{i=1}^{n} (\beta_i p_i t_{ij})^{1-\sigma}]^{1/(1-\sigma)}$. By using the general equilibrium adding-up constraint, $y_i = \sum_{i=1}^{n} x_{ij}$, in combination with Equation (3.9), Anderson and van Wincoop (2003) show that the utility-maximizing behavior of households implies a so-called gravity equation, one of the most robust empirical relations in economics.[23] We can write bilateral trade flows as

$$x_{ij} = \frac{y_i \tilde{y}_j}{y^W} \left(\frac{t_{ij}}{\tilde{\Pi}_i \tilde{P}_j}\right)^{1-\sigma}, \quad \text{where} \tag{3.10}$$

$$\tilde{\Pi}_i \equiv \left(\sum_{j=1}^{n} \left(\frac{t_{ij}}{\tilde{P}_j}\right)^{1-\sigma} \tilde{\theta}_j\right)^{1/(1-\sigma)}, \quad \tilde{P}_j \equiv \left(\sum_{i=1}^{n} \left(\frac{t_{ij}}{\tilde{\Pi}_i}\right)^{1-\sigma} \theta_i\right)^{1/(1-\sigma)}, \tag{3.11}$$

where we substituted equilibrium scaled prices into the definition of the price index to obtain the multilateral resistance terms $\tilde{P}_j$ $\tilde{\Pi}_i$ and defined $y^W \equiv \sum_j y_j$, $\tilde{y}^W \equiv \sum_j \tilde{y}_j$ and income shares $\theta_j \equiv y_j / y^W$ and $\tilde{\theta}_j \equiv \tilde{y}_j / \tilde{y}^W$.

The system of $2n$ equations given in (3.11) determines the $2n$ outward and inward multilateral resistance terms $\tilde{\Pi}_i$ and $\tilde{P}_j$. $\tilde{\Pi}_i$ and $\tilde{P}_j$ can be interpreted as weighted averages of

21 In the empirical analysis, I allow for trade imbalances similar to Dekle et al. (2007) as my sample only includes 13 countries, potentially exacerbating the importance of trade imbalances. Appendix C.3 reports results assuming balanced trade. Results are very similar.

22 Fiess et al. (2010) document that informal firms virtually never export. As in the present model international trade only implies iceberg trade costs, we can as well assume that only formal firms export without loss of generality.

23 For a recent in-depth survey of gravity equations, see Head and Mayer (2014).

export and import trade costs.[24] From these, we can derive the price levels in all n countries in general equilibrium.

3.3 Counterfactual analysis

We can now use the model to derive the general equilibrium effects of a reduction in bilateral tariffs and general trade costs brought about by preferential trade agreements. This reduction in trade costs impacts the price levels across all countries and, via the general equilibrium effects, also affects unemployment and informality levels. Specifically, I will evaluate the impact of preferential trade agreements on unemployment as well as informal employment across countries. As shown in Equation (3.4), the level of employment depends on the vector of price levels consistent with a given amount of trade costs. Given knowledge of the trade cost parameters as well as the labor market parameters like the formality premium, we can solve our model for the equilibrium price vectors, once for the trade costs observed in the data, i.e. with all PTAs which are currently signed between countries, and once in a counterfactual world where we abolish these trade agreements.[25] Given the price vectors in both the observed and counterfactual scenarios, we can calculate counterfactual changes in welfare, unemployment, and informal employment.

3.3.1 Counterfactual size of the formal sector

In equilibrium, the variety price charged by formal and informal firms is the same. Hence we can combine Equation (3.6) with the formal wage curve to receive the following expression for the counterfactual change in the number of formal sector workers when we assume that labor market parameters remain constant:

$$\hat{L}_j^f \equiv \frac{L_j^{f,c}}{L_j^f} = \left(\frac{\frac{w_j^{i,c}}{w_j^{f,c}} \frac{1+\gamma_j \xi_j - \gamma_j}{\xi_j}}{\frac{w_j^i}{w_j^f} \frac{1+\gamma_j \xi_j - \gamma_j}{\xi_j}} \right)^{\frac{1}{\alpha}} = \left(\frac{w_j^{f,c}/w_j^{i,c}}{w_j^f/w_j^i} \right)^{-\frac{1}{\alpha}}, \tag{3.12}$$

where the hat denotes a change and c denotes the counterfactual values. Note that as the labor force remains constant, this expression also gives the change in the formal employment

[24] For a discussion of the interpretation of multilateral resistance terms see Anderson and van Wincoop (2004).

[25] Details on the system of equations can be found in Appendix C.1.

share, L_j^f/L_j. The change in formal sector employment is inversely related to the change in the formal sector premium. When α decreases towards 0, implying a smaller reaction of informal sector productivity to changes in formal sector employment, the same percentage change in the formal sector wage premium is magnified.

Note that we can then calculate the change in the informal sector as

$$\hat{L}_j^i \equiv \frac{L_j^{i,c}}{L_j^i} = \frac{L_j - \hat{L}_j^f L_j^f}{L_j - L_j^f}. \tag{3.13}$$

3.3.2 Counterfactual formal employment probability

To derive the counterfactual change in formal employment, we express the change in the endogenous variables of interest in terms of the price vectors. Following Anderson and van Wincoop (2003), we can use the general equilibrium adding up constraint that total sales equal income, i.e. $y_i = \sum_{j=1}^n x_{ij}$, in combination with the definition of sales given in Equation (3.9) to express variety prices in a country as:

$$(\beta_j p_j)^{1-\sigma} = \frac{y_j}{\sum_{i=1}^n (\frac{t_{ji}}{\tilde{P}_i})^{1-\sigma} \tilde{y}_j} = \frac{y^W}{\tilde{y}^W} \theta_j \tilde{\Pi}_j^{\sigma-1} = \frac{y^W}{\tilde{y}^W} \mathfrak{k}_j, \tag{3.14}$$

where $\mathfrak{k}_j \equiv \theta_j \tilde{\Pi}_j^{\sigma-1}$ is determined by the system of equations given in Equation (3.11).[26]

Plugging Equation (3.4) into the definition of the probability of becoming unemployed, $u_j^f = 1 - m_j \vartheta_j^{1-\mu}$, and keeping labor market parameters constant, it can be shown that

$$\hat{e}_j^f \equiv \frac{e_j^{f,c}}{e_j^f} \equiv \frac{1 - u_j^{f,c}}{1 - u_j^f} = \left(\frac{p_j^c}{p_j}\right)^{\frac{1-\mu}{\mu}} \left(\frac{P_j}{P_j^c}\right)^{\frac{1-\mu}{\mu}} \tag{3.15}$$

$$= \left(\frac{\mathfrak{k}_j^c}{\mathfrak{k}_j}\right)^{\frac{1-\mu}{\mu(1-\sigma)}} \left(\frac{\sum_i t_{ij}^{1-\sigma} \mathfrak{k}_i}{\sum_i \left(t_{ij}^c\right)^{1-\sigma} \mathfrak{k}_i^c}\right)^{\frac{1-\mu}{\mu(1-\sigma)}}, \tag{3.16}$$

where e_j^f denotes the formal employment rate. Note that we can write the change in the probability of a formal sector worker becoming employed as

$$\hat{u}_j^f \equiv \frac{u_j^{f,c}}{u_j^f} = \frac{1 - e_j^{f,c}}{1 - e_j^f} = \frac{1 - e_j^f \hat{e}_j^f}{u_j^f}. \tag{3.17}$$

[26] Details on how to solve this system can be found in Appendix C.1.

The algebraic expression for $\hat{e}_j^f$ is identical to the expression of the counterfactual change of employment in Heid and Larch (2012a). The difference, however, lies in its interpretation: whereas in Heid and Larch (2012a) it gives the change for employment in the whole economy, in the present framework it only gives the change for the formal sector.

3.3.3 Counterfactual official unemployment rate

Note that $\hat{u}_j^f$ does not give the change in the official unemployment rate, $u_j^o = U_j/L_j$, as the latter depends on the absolute number of unemployed formal sector workers. It is given by

$$\hat{u}_j^o \equiv \frac{u_j^{o,c}}{u_j^o} = \frac{u_j^{f,c} L_j^{f,c}}{u_j^f L_j^f} = \hat{u}_j^f \hat{L}_j^f. \tag{3.18}$$

When trade is liberalized, and the price level in a country falls, then the probability of a formal sector worker finding a job increases, as the vacancy posting costs for formal firms are lower. The lower probability of becoming unemployed, however, makes the formal sector more attractive, as the expected formal sector wage is higher. Therefore, more workers leave the informal sector and seek formal employment. Whether the official unemployment rate decreases or increases depends on the interplay of the elasticities of the model: The elasticity of substitution, σ, the matching elasticity, μ, and the elasticity of informal sector productivity, α. Compared to Heid and Larch (2012a), who assume a single labor market in the whole economy, the reduction of the official unemployment rate is dampened by the rising attractiveness of the formal part of the economy. This may partly explain why empirical evidence on the observed correlation between official unemployment rates and changes in openness is mixed, and a relation between trade and unemployment is downplayed by some economists.[27]

[27] Felbermayr et al. (2011b) find that higher trade openness decreases unemployment. Similar conclusions can be drawn from Dutt et al. (2009) and Hasan et al. (2012). Heid and Larch (2012b), however, find no significant effect. Krugman (1993) argues that unemployment mainly is determined by macroeconomic factors like aggregate demand, whereas microeconomic factors like trade costs only play a minor role.

3.3.4 Counterfactual formal wage premium

Using the indifference condition of workers given in Equation (3.1), we can express the change in the formal wage premium as

$$\widehat{w_j^f/w_j^i} \equiv \frac{w_j^{f,c}/w_j^{i,c}}{w_j^f/w_j^i} = \frac{1-u_j^f+u_j^f\gamma_j-f_j^f}{1-u_j^f\hat{u}_j^f+u_j^f\hat{u}_j^f\gamma_j-f_j^f}. \tag{3.19}$$

3.3.5 Counterfactual (nominal) GDP

We now have everything in place to calculate the counterfactual change in (nominal) GDPs brought about by trade liberalization. GDP is given by

$$p_j(1-u_j^f)L_j^f + p_j L_j^i (L_j^f/L_j)^\alpha = p_j[(1-u_j^f)L_j^f + L_j^i(L^f/L_j)^\alpha]. \tag{3.20}$$

Hence we can write the counterfactual change in GDP in terms of changes in prices, formal employment as well as changes in the sectoral labor force composition:

$$\hat{y}_j = \frac{y_j^c}{y_j} = \frac{p_j^c}{p_j}\frac{[e_j^f\hat{e}_j^f(L_j-L_j^i\hat{L}_j^i)+L_j^i\hat{L}_j^i(1-(L_j^i\hat{L}_j^i)/L_j)^\alpha]}{[e_j^f(L_j-L_j^i)+L_j^i(1-L_j^i/L_j)^\alpha]}, \tag{3.21}$$

such that it can be expressed in terms of changes in prices using the derivations from above. Note that the change in the variety price can be deduced from Equation (3.14).

3.3.6 Counterfactual welfare

A model consistent welfare measure is the equivalent variation, i.e. the amount of income the representative consumer would need to make her as well off under current prices $\tilde{P}_j$ as in the counterfactual situation with price level $\tilde{P}_j^c$. We can express the equivalent variation in percent as follows:

$$EV_j = \frac{\breve{y}_j^c\frac{\tilde{P}_j}{\tilde{P}_j^c}-\breve{y}_j}{\breve{y}_j} = \frac{\breve{y}_j^c}{\breve{y}_j}\frac{\tilde{P}_j}{\tilde{P}_j^c}-1 = \hat{\breve{y}}_j\frac{\tilde{P}_j}{\tilde{P}_j^c}-1, \tag{3.22}$$

where $\hat{\breve{y}}_j$ is the change in consumable income $\breve{y}_j$ in country j. The change in the price indices can be recovered from the multilateral resistance terms.[28] As the vacancy posting costs of

[28] For computational details, see Appendix C.1.

formal sector firms consume part of the final output good, the change in consumamble income is not equal to the change in GDP. The former is given by the total wage sum augmented by the exogenous trade deficit share, $(1+d_j)[(1-u_j^f)w_j^f L_j^f + p_j(L_j^f/L_j)^\alpha L_j^i]$. Assuming that the trade deficit share is constant and exogenous, and using the formal sector wage curve, we can write the change in consumable income as:

$$\hat{\breve{y}}_j \equiv \frac{\breve{y}_j^c}{\breve{y}_j} = \frac{p_j^c}{p_j} \frac{\hat{e}_j^f e_j^f [\xi_j/(1+\gamma_j\xi_j-\gamma_j)]\hat{L}_j^f L_j^f + (\hat{L}_j^f L_j^f/L_j)^\alpha \hat{L}_j^i L_j^i}{e_j^f [\xi_j/(1+\gamma_j\xi_j-\gamma_j)] L_j^f + (L_j^f/L_j)^\alpha L_j^i}. \tag{3.23}$$

Hence welfare can be calculated by using the expressions derived previously as well as the changes in the variety price implied by Equation (3.14).

3.4 Bringing the model to the data

3.4.1 Estimation of trade agreement effects

To analyze the impact of signing a preferential trade agreement (PTA) on welfare, unemployment, and informal employment, we first need an estimate of the actual size of the reduction of trade costs brought about by a typical PTA. Whereas the previous literature has relied on direct measures of tariff reductions (see Goldberg and Pavcnik, 2003), it is well known that tariffs only make up a part of actual trade costs which also consist of non-tariff barriers like differences in languages, customs, culture etc. Similarly, trade agreements often include a considerable amount of harmonization of product standards and regulations as well as other measures which reduce non-tariff barriers and which are not measured by a change in tariff rates. Therefore, trade policy measures are only a very rough measure of actual trade cost reductions (see Anderson and van Wincoop, 2004). I therefore follow the standard approach in international trade and estimate the gravity equation of international trade implied by the theoretical model to get an estimate of the impact of a PTA on trade flows. In addition, gravity estimation allows to take into account the trade creation and diversion effects typical of PTAs.[29] As trade agreements are not signed randomly between countries, I follow the estimation approach outlined in Baier and Bergstrand (2007) and Anderson and Yotov (2011) to control for the potential endogeneity of the PTA measure.[30]

[29] For an overview of trade diversion and creation of PTAs, see Panagariya (2000).
[30] The same estimation approach is used in Heid and Larch (2012a).

Specifically, we can reformulate Equation (3.10), i.e. exports from country i to j, as

$$\frac{x_{ij\tau}}{y_{i\tau}y_{j\tau}} = \exp\left(y_\tau^W + (1-\sigma)\ln t_{ij\tau} - \ln \tilde{\Pi}_{i\tau}^{1-\sigma} - \ln \tilde{P}_{j\tau}^{1-\sigma} + \varepsilon_{ij\tau}\right), \tag{3.24}$$

where I have added a time superscript τ as well as a stochastic error term $\varepsilon_{ij\tau}$. I still have to specify the trade cost function $t_{ij\tau}$ which I assume is given by

$$t_{ij\tau} = \exp(\beta_1 PTA_{ij\tau} + \beta_2 \ln DIST_{ij} + \beta_3 CONTIG_{ij}),$$

where $PTA_{ij\tau}$ is an indicator variable of preferential trade agreement membership between country pair ij in year τ, $DIST_{ij}$ is bilateral distance, and $CONTIG_{ij}$ is a dummy variable indicating whether countries i and j are contiguous.[31]

I use data on trade flows between 13 Latin American and Carribbean countries for which also data on the informal sector are available.[32]

To account for the heteroscedasticity of trade flows, I follow the suggestion by Santos Silva and Tenreyro (2006) and use a Poisson Pseudo Maximum Likelihood (PPML) estimator to estimate the trade cost parameters. The approach by Anderson and Yotov (2011) proceeds in two steps: In a first estimation, Equation (3.24) is estimated including a set of exporter times year and importer times year dummies to control for the outward and inward multilateral resistance terms, $\ln \tilde{\Pi}_{i\tau}^{1-\sigma}$ and $\ln \tilde{P}_{j\tau}^{1-\sigma}$. In addition, a set of $n \times (n-1)/2$ dummies for each bilateral trade relation is included when one is willing to assume symmetric trade costs, and a set of $n \times (n-1)$ bidirectional dummies for each bilateral trade relation when one assumes that trade costs are asymmetric. Either way, the set of dummies controls for the special nature of a trade relation between two countries, effectively controlling for the endogeneity of the PTA variable caused by time-invariant unobserved factors influencing the probability that a specific country pair signs a preferential trade agreement. This first step regression drops regressors like bilateral distance and contiguity, and only β_1, the coefficient of the PTA variable, can be identified. Hence, in a second step, the coefficient β_1 is constrained

[31] Note that nearly all countries in the sample have Spanish as their official language; only Brazil has a different language, Portuguese. When including exporter and importer (times year) dummies, a common language dummy would be perfectly collinear. A similar argument applies to a common colonizer dummy. I hence omit these regressors which are normally used in the gravity literature.

[32] Trade and gravity variables except PTA are from CEPII and are described in Head et al. (2010). PTA is constructed from the notifications to the World Trade Organization (WTO) and augmented and corrected by using information from PTA secretariat webpages. The countries are: Argentina, Bolivia, Brazil, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Nicaragua, Paraguay, Peru, Uruguay, and Venezuela. Summary statistics of the gravity data set used can be found in Appendix C.2.

to its estimated value, the bilateral dummies are dropped and thus the influence of the time-invariant regressors $\ln DIST$ and $CONTIG$ can be identified. Results from the gravity estimations for the trade cost parameters can be found in Table 3.1. Columns (1) and (2) assume symmetric bilateral trade costs, whereas columns (3) and (4) assume symmetric trade costs. Columns (1) and (3) do not constrain the elasticity of trade flows with respect to exporter and importer GDP to unity by using simply trade flows as the dependent variable.[33] Columns (2) and (4) use scaled trade flows as a dependent variable, implicitly imposing unitary elasticities, consistent with the theoretical framework which assumes homothetic preferences. The coefficients in Table 3.1 can be interpreted as partial equilibrium average treatment effects. As the Poisson model is a log-linear model, coefficients can be interpreted directly as elasticities. Using this interpretation, all estimated coefficients have the correct sign and are in the expected ballpark: For example, an increase in the distance between two trading partners by one percent decreases bilateral trade flows by about 1.6 percent. Whether one assumes symmetric or asymmetric trade costs hardly affects the coefficient estimates. However, results for the other regressors are remarkably different, depending on whether one imposes the homotheticity assumption: Sharing a common border increases bilateral trade by about 6 percent assuming homothetic preferences, and by about 20 percent when not imposing the unitary income elasticities.[34] Interestingly, contiguity loses its significance assuming homothetic preferences. When two countries have signed a preferential trade agreement, bilateral trade flows increase between 47 (column (3)) and 179 (column (4)) percent on average.

For the counterfactual general equilibrium analysis, I also need a value of σ. Bergstrand et al. (2013) use a structural gravity model with full employment to derive an estimator for σ. I use their estimate and set $\sigma = 7.1$. This is also broadly in line with the estimate of $\sigma = 9.3$ from Eaton and Kortum (2002).[35]

[33] Note that I cannot report coefficients for importer and exporter GDP as these are controlled for by the exporter and importer times year dummies.

[34] I calculate partial equilibrium average treatment effects of discrete regressors as $[\exp(\hat{\beta}_k) - 1] \times 100$.

[35] Eaton and Kortum (2002) use a Ricardian model of trade to derive a gravity equation for trade flows which depend on the comparative advantage parameter θ. Their model is observationally equivalent to a model with Armington (1969) preferences where $\sigma = 1 + \theta$, see Arkolakis et al. (2012). A considerably lower estimate of $\sigma = 3.8$ can be found in Bernard et al. (2003) who use plant-level export data.

Table 3.1: Estimation results for a sample of 13 Latin American and Caribbean countries, 1950-2006

	(1) PPML $x_{ij\tau}$	(2) PPML $z_{ij\tau}$	(3) PPML $x_{ij\tau}$	(4) PPML $z_{ij\tau}$
First stage				
$PTA_{ij\tau}$	0.396***	0.951***	0.382***	1.025***
	(0.080)	(0.147)	(0.068)	(0.109)
Second stage				
$\ln DIST_{ij}$	−1.578***	−1.645***	−1.579***	−1.637***
	(0.041)	(0.052)	(0.041)	(0.052)
$CONTIG_{ij}$	0.185***	0.063	0.186***	0.064
	(0.059)	(0.074)	(0.059)	(0.074)
symmetric $t_{ij\tau}$	X	X		
asymmetric $t_{ij\tau}$			X	X
N	8,743	8,743	8,743	8,743

Notes: Results for trade flows between 13 Latin American and Caribbean countries between 1950 and 2006 estimated by Poisson pseudo-maximum-likelihood (PPML). z_{ij} are trade flows standardized by importer and exporter GDPs. $\ln DIST$ is distance between exporting and importing country, $CONTIG$ is an indicator variable equal to 1 if the exporting and importing countries i and j share a common border, and PTA is an indicator variable equal to 1 if the exporting and importing country have signed a preferential trade agreement. All regressions control for multilateral resistance terms (MRTs) via exporter-time and importer-time fixed effects. Robust standard errors in parentheses, *** $p < 0.01$.

3.4.2 Labor market data

For the counterfactual analysis, I need data on the following characteristics of countries' labor markets: The unemployment rate, the rate of unemployment benefits, the size of the total labor force, the rate of employment in the (urban) informal sector as well as information about the (urban) formality premium, i.e. the wage of formal sector workers relative to informal sector workers. I use the year 2006 for all data or the year closest to 2006 available in the data.[36] If there are different measures from surveys at the national and sub-national level available for a country, I always use the survey on the national level.

The main data source on informality is the Socio-Economic Database for Latin America and the Caribbean (SEDLAC) from CEDLAS and The World Bank (2013).[37] It contains data on the unemployment rate, the share of adults in informal jobs, as well as formal and informal hourly wages. I use the data based on a legalistic definition of informality. Hence individuals are considered to work in the informal sector when they do not have the right to a pension when they retire.

To transform the share data into data in levels, I use data on total population and labor force participation rates from the World Development Indicators (WDI) from The World Bank (2013).[38] As the model abstracts from the agricultural sector, I use urban informality shares and assume that the number of informal workers in the economy is given by the share of urban informal workers times the labor force.

Data on the rate of unemployment benefits are hard to come by for Latin American countries. In addition, many Latin American countries rely on severance payments instead of a system of unemployment insurance with mandatory or voluntary contributions. Finally, some countries have individual insurance accounts.[39] Therefore, focusing on a single instrument of unemployment insurance may hinder the comparability across countries. Instead, I use data from ILO (2010) on the effective share of unemployed workers who are covered by some form of income support system.[40]

[36] I use data on the share of adults in the labor force for 2007 for Bolivia and for 2005 for Nicaragua. Wage rates are for 2008 for Colombia. Data for Argentina are the simple average of the two waves of the same survey available for 2006.

[37] The database can be accessed via `http://sedlac.econo.unlp.edu.ar/`. I use the data as of 08/16/2013.

[38] The database can be accessed via `http://data.worldbank.org/data-catalog/world-development-indicators`. I use the data as of 08/16/2013.

[39] For a detailed overview, see OECD (2011f).

[40] The share is for the latest available year at the time of publication of ILO (2010), no further details are provided. The use of this data can be rationalized in terms of the model if we assume for simplicity that workers who receive some form of support when they are unemployed receive the full going wage; however,

I also have to set the bargaining power of formal sector workers. I follow Heid and Larch (2012a) and set it equal to 0.5 in all countries.

Finally, I need an estimate of the elasticity of the matching function with respect to the unemployed, μ. Papers which structurally estimate matching functions exclusively focus on labor markets in developed countries and estimate μ in a range between 0.12 and 0.81 (see the survey by Petrongolo and Pissarides, 2001). In addition, most studies use OLS which suffers from several biases. Also, the literature discusses data measurement issues which may also bias the estimates (see Petrongolo and Pissarides, 2001; Yashiv, 2007, and Borowczyk-Martins et al., 2013 as well as the references cited therein for a discussion). Most recent estimates use data on U.S. job vacancies from the Job Openings and Labor Turnover Survey (JOLTS) and lie in the range between 0.32 and 0.72 (see Rogerson and Shimer, 2011 and Borowczyk-Martins et al., 2013). I set $\mu = 0.52$, which is the midpoint of the two most recent estimates for the United States, in the empirical application.

3.4.3 Solving for the entry fixed costs into the formal sector

To bring the model to the data, I first solve for the level of the entry fixed costs into the formal sector, f_j^f, by using Equation (3.1). For this, I calculate the formality wage premium, w_j^f/w_j^i, as well as the probability of becoming unemployed in the formal sector, u_j. Following the model, the latter is given by the ratio of the number of unemployed workers to the number of workers in the formal sector, as all informal sector workers cannot become unemployed. I report these in Table 3.2.

3.4.4 Estimating the informal sector productivity elasticity

To get an estimate of the informal sector productivity elasticity, we can log-linearize Equation (3.7) and shuffle terms to receive an estimable equation for α:

$$\ln\left(\frac{w_j^f}{w_j^i}\right) - \ln\left(\frac{1+\gamma_j\xi_j - \gamma_j}{\xi_j}\right) = \alpha_0 - \alpha \ln\left(\frac{L_j^f}{L_j}\right) + \eta_j, \tag{3.25}$$

where I have added a constant term α_0 as well as a stochastic error term η_j. I report OLS estimates of this regression using data from 2006 in Table 3.3. Estimates are not significant,

only with probability γ_j. If the probability of becoming unemployed is independent of the probability of receiving the unemployment benefit, γ_j is exactly the share of unemployed workers covered.

Table 3.2: Formal and informal sector statistics

Country	u_j^f	L_j^i/L_j	w_j^f/w_j^i	$f_j^f w_j^f/w_j^i$
Argentina	0.11	0.34	1.77	0.59
Bolivia	0.11	0.57	2.22	0.97
Brazil	0.09	0.25	2.13	0.96
Colombia	0.15	0.38	2.52	1.15
Costa Rica	0.04	0.21	1.74	0.67
Dominican Rep.	0.05	0.36	1.59	0.50
Ecuador	0.11	0.46	1.58	0.41
El Salvador	0.06	0.32	2.11	0.99
Nicaragua	0.14	0.51	1.41	0.22
Paraguay	0.16	0.63	2.11	0.78
Peru	0.08	0.47	2.42	1.22
Uruguay	0.09	0.17	1.85	0.70
Venezuela	0.11	0.34	1.36	0.21

Notes: Formal and informal sector statistics for 13 Latin American and Caribbean countries in (roughly) 2006. u_j^f is the probability of becoming unemployed in the formal sector. L_j^i/L_j is the share of informal workers. w_j^f/w_j^i is the formal to informal sector wage ratio. $f_j^f w_j^f/w_j^i$ is the monetary formal sector entry cost in multiples of the informal sector wage. For details about the data sources used and the calculation see Section 3.4.2.

which is not too surprising given the low number of observations. In principle, one could expand the data set to a panel for efficiency gains. More importantly, Equation (3.25) suffers from a potential endogeneity bias as the relative formal sector size is determined by the formality wage premium. In principle, one could instrument the formal sector employment share; however, given the data availability, one is hard pressed to come up with an instrument. Still, the estimate $\hat{\alpha}$ is still a good estimate in the sense that it is the best linear predictor of α in the data set and therefore fits the data best.

3.5 Evaluation of Latin American preferential trade agreements

In the following, I will evaluate the welfare and employment effects of the preferential trade agreements which have been signed between the 13 Latin American and Caribbean countries since 1950. Figure 1 shows the proliferation of preferential trade agreements between these countries by depicting the share of country pairs with an agreement. The first trade agreement was signed in 1961, after which the number of agreements slowly increased. In 1981, the share of country pairs with an agreement jumped from little more than 10 percent to more than 50 percent. Since then, there was a steady increase to reach more than 60 percent

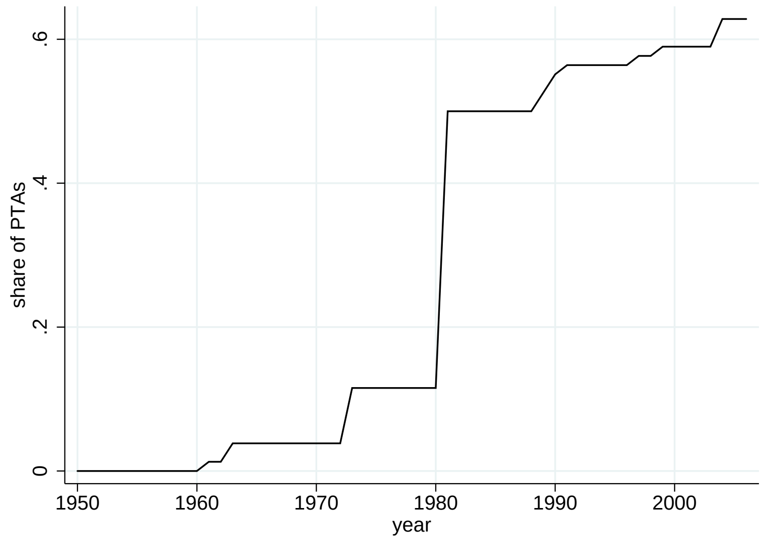
Table 3.3: Estimation results for the informal sector productivity elasticity for a sample of 13 Latin American and Caribbean countries in 2006

	(1) OLS
α_0	-0.097
	(17.294)
α	0.084
	(9.719)
N	13

Notes: Results for the regression given in Equation (3.25) for 13 Latin American and Caribbean countries in 2006 estimated by OLS. Standard errors in parentheses.

of all country pairs at the turn of the century. The counterfactual situation I will consider is a world without any preferential trade agreement. I will compare this situation with the observed agreements in place in 2006. I report results from this counterfactual exercise in Table 3.4. It shows the effect of trade liberalization, i.e. changes are calculated as moving from the counterfactual scenario to the observed data. The table is organized as follows: The column labeled $\Delta\% e_j^f$ reports the percentage change in the probability of finding a job in the formal sector. $\Delta\% pts\, u_j^f$ gives the according change in the probability of becoming unemployed in the formal sector in percentage points. $w_j^f/w_j^i\Delta\%$ gives the percentage change in the formality premium, and $L_j^i\Delta\%$ the accompanying percentage change in the size of the informal sector. $\Delta\% pts\, u_j^o$ gives the change in the official unemployment rate in percentage points. $\Delta\% EV$ gives the percentage change in the equivalent variation. For comparison, I report the equivalent variation implied by the framework with a perfect labor market (PLM) by Anderson and van Wincoop (2003) as well as for the framework which assumes a unified labor market with search and matching frictions (SMF) from Heid and Larch (2012a). I use the same elasticity of substitution, $\sigma = 7.1$, for the calculation of all three equivalent variation measures, and set the same elasticity of the matching function, $\mu = 0.52$, for both the model with informality and the framework from Heid and Larch (2012a). Besides values for individual countries, I report weighted average effects which use a country's labor force as weight.

On average, I find that switching on preferential trade agreements increases employment in the formal sector by 5.8 percent, and the according probability of becoming unemployed in



Notes: Share of country pairs covered by a preferential trade agreement (PTA) in a sample of 13 Latin American and Caribbean countries. For a description of data sources, see Section 3.4.

Figure 3.1: Share of country pairs covered by preferential trade agreements

the formal sector decreases by 4.7 percentage points. As trade liberalization brought about by the preferential trade agreements makes the formal sector more attractive, the indifference condition given in Equation (3.1) implies that the formality wage premium has to decrease in order to restore the equilibrium. On average, I find that the formality premium decreases by 9.3 percent due to the shrinking productivity gap between the formal and informal sector. The change in the formality premium in turn implies a change in the share of formal workers. On average, the informal sector is reduced by 50.9 percent. This is a large effect of trade liberalization, compared with results from other studies which find a much more modest effect of trade liberalization on informality, if at all. For example, Attanasio et al. (2004) find that informal employment in Colombia in 1998 is 4.4 percentage points larger due to tariff reductions compared to 1984. Bosch et al. (2012) find that trade liberalization accounts for 1-2.5 percent of the increase in informality in Brazil during the 1990s. Contrary to that, Goldberg and Pavcnik (2003) find no effect of trade liberalization on informality in Brazil during the same period. Note, however, that these results are not directly comparable as they use tariff reductions which might be unilateral and not necessarily linked to preferential trade agreements. Also, the mentioned papers only study trade liberalization episodes in the mid 1980s to 1990s, whereas I evaluate the effect of preferential trade agreements signed since 1950. As the largest increase in the number of preferential trade agreements happened in 1981, it may well be that the effects of trade liberalization on informal employment are different from those from later periods of trade liberalization.

As the probability of becoming unemployed is reduced by the preferential trade agreements, workers move into the formal sector. The combined effect on the official unemployment rate is given in column $\%pts\, u_j^o$ in percentage points. On average, the official unemployment rate is 3.1 percentage points higher, implying that the absolute number of unemployed workers has increased even though the probability of becoming unemployed in the formal sector has decreased. This may explain the fact that politicians fear a net increase in unemployment due to trade liberalization, especially in countries with a large informal sector.

Finally, we can turn to the changes in the equivalent variation, our welfare measure. I find that on average, welfare decreases by 7.6 percent. Why can trade liberalization decrease welfare? In the model, the decrease in the size of the informal sector increases the latter's productivity, which has a positive effect on welfare via the increase in the informal sector wage. However, now more workers have to negotiate their wage in the formal sector. As they cannot leave the sector, their outside option are the unemployment benefits. As these are zero for many countries in the sample, negotiated wages in the formal sector are rather low. In addition, formal sector firms have to pay a larger amount of vacancy posting costs as the relative size of the formal sector has increased. The net effect is such that the positive productivity gain in the informal sector as well as the lower price indices for consumers is more than outweighed by the higher share of firms which have to pay vacancy posting costs and the bad bargaining position of formal sector workers. Still, Bolivia, the Dominican Republic, Nicaragua as well as Peru benefit from the trade liberalization brought about by the preferential trade agreements they have signed. These findings stand in contrast to the frameworks which either assume full employment or a unified labor market with search frictions. Both frameworks find that welfare on average is increased by 6.3 percent with full employment and 12.6 percent with search frictions, in line with the relative magnitudes of effects by Heid and Larch (2012a).

The average effects hide substantial heterogeneity in the effects of preferential trade agreements. Uruguay sees its informal sector reduced by 82 percent, whereas the Dominican Republic actually experiences an increase in the informal sector by 13.8 percent. As a robustness check, I redid the counterfactual analysis assuming balanced trade between the 13 countries. Results hardly change. I report these results in Appendix C.2.

Summing up, I find that preferential trade agreements have reduced the informal sector, increased the official unemployment rate, and decreased welfare in most countries in the sample. Obviously, the presented, highly stylized framework should no be taken as a literal description of the reality of experiences in Latin American and the Caribbean brought about by trade liberalization. However, the large quantitative and qualitative difference in the welfare effects highlights the importance of assumptions about the structure of labor markets for the evaluation of preferential trade agreements and trade liberalization in general.

3.6 Conclusion

The standard tools to evaluate the welfare effects of trade liberalization episodes and preferential trade agreements are structural gravity models. State of the art quantitative frameworks assume perfect labor markets. Recently, Heid and Larch (2012a) introduced search and matching frictions into a structural gravity model and evaluate preferential trade agreements between developed OECD countries. I extend their framework to include an informal sector, a decisive feature of labor markets in emerging economies. I apply this framework to a set of 13 Latin American and Caribbean countries to evaluate the employment and welfare effects of preferential trade agreements. I find that the preferential trade agreements which have been signed since 1950 have, on average, decreased welfare by 7.6 percent, decreased informal employment by 50.9 percent, and increased the official unemployment rate by 3.1 percentage points. These results are quantitatively and qualitatively different from standard frameworks assuming either full employment or a unified labor market with search and matching frictions.

Similar to single country studies by Goldberg and Pavcnik (2003) and Attanasio et al. (2004), my results highlight the importance of labor market institutions for evaluating the consequences of trade liberalization for welfare in general and informal employment in particular.

A potential avenue for future research is to consider the agricultural sector to quantify the classic Harris and Todaro (1970) view of informality. In such a framework, workers would choose between secure employment in the agricultural sector or in the urban manufacturing sector where there is a probability of becoming unemployed. The urban unemployed work in the informal sector. In combination with a multi-sector framework for trade flows, this setup would allow to evaluate the effect of preferential trade agreements in developing and

Table 3.4: Comparative static effects of switching on all observed PTAs in 2006

Country	$\Delta\% e_j^f$	$\Delta\% pts\, u_j^f$	$w_j^f / w_j^i \Delta\%$	$\Delta\% L_j^i$	$\Delta\% pts\, u_j^o$	$\Delta\% EV$	$\Delta\% EV(PLM)$	$\Delta\% EV(SMF)$
Argentina	9.9	-8.0	-12.9	-60.7	4.9	-10.4	10.8	21.5
Bolivia	22.6	-16.4	-36.3	-42.9	4.7	60.9	25.0	49.4
Brazil	2.0	-1.8	-3.6	-52.0	1.4	-14.3	2.2	4.4
Colombia	8.7	-6.8	-17.3	-59.3	7.8	-9.2	9.5	18.8
Costa Rica	0.4	-0.3	-0.6	-20.5	-0.0	-4.1	0.4	0.9
Dominican Rep.	-0.4	0.4	0.6	13.8	0.0	2.0	-0.4	-0.6
Ecuador	9.0	-7.4	-11.7	-47.7	3.6	-5.1	9.9	19.8
El Salvador	0.5	-0.5	-1.0	-19.4	0.2	-4.8	0.5	1.1
Nicaragua	13.4	-10.2	-14.4	-45.0	4.9	2.9	14.9	29.6
Paraguay	21.1	-14.7	-30.9	-36.3	5.7	49.3	23.4	46.4
Peru	10.0	-8.3	-20.1	-51.1	3.7	5.3	10.8	21.5
Uruguay	19.1	-14.5	-23.5	-82.0	7.0	-5.8	21.5	42.8
Venezuela	6.1	-5.1	-6.9	-53.3	2.9	-14.0	6.7	13.2
Average (L_j-weighted)	5.8	-4.7	-9.3	-50.9	3.1	-7.6	6.3	12.6

Notes: Counterfactual analysis based on parameter estimates from column (4) of Table 3.1. PLM gives results assuming perfect labor markets. SMF gives results using a unified labor market with search and matching frictions. Averages are weighted averages using a country's labor force as weight.

emerging economies. In such a framework, if trade liberalization decreases the probability of becoming unemployed, the informal sector may increase or decrease, depending on the net effect of rural to urban migration, similar to the effect on the official unemployment rate in the present manuscript.

Chapter 4

The Rise of the Maquiladoras: A Mixed Blessing*

4.1 Introduction

Over the past three decades Mexico has undergone a dramatic transformation that has made it one of the most open developing countries in the world today. One of the key drivers behind Mexico's impressive export growth has been the *maquila* sector.

Maquila plants, or *maquiladoras* for short, are export assembly plants which are mostly located along a 20km strip along the US-Mexico border (Cañas and Gilmer, 2009). The defining characteristic of *maquiladoras* is their exclusive focus on assembling imported intermediate inputs which are then re-exported either for further assembly or as finished goods, mostly to the US. Although the *maquiladora* program formally started in 1965, it was not until the end of the 1980s, after Mexico's first round of trade and investment liberalization reforms, that the sector started booming (see Bergin et al., 2009 and Waldkirch, 2010; for an in-depth account of Mexican trade and investment liberalization see Kehoe, 1995). With the sector's value-added growing at an average of 10% per year during the 1990s (in comparison to a 3% per year growth rate of real GDP, see Hanson, 2002), *maquiladoras* have come to

* This chapter is based on joint work with Mario Larch and Alejandro Riaño. It is based on the article "The Rise of the Maquiladoras: A Mixed Blessing" in: *Review of Development Economics*, 2013, 17(2), 252-267. It is a revised version of CESifo Working Paper No. 3689, 2011 which circulated under the title "Maquiladoras and Informality: A Mixed Blessing". A previous version of this paper has been circulated under the title "The Rise of the Maquiladoras: Labor Market Consequences of Offshoring in Developing Countries".

account for 8.3% of manufacturing value-added, 47.1% of manufacturing employment and 52.9% of aggregate exports by 2004.[1]

One of the main goals of the *maquiladora* program was to increase employment of unskilled workers (see Martin, 2000). Although Mexico's unemployment rate has always been particularly low,[2] around 30 to 50% of the labor force is employed in the informal sector, an array of small-scale, low-productivity establishments, where workers earn wages substantially lower than in formal firms. The fact that such a large share of the labor force participates in this sector is regarded as undesirable, since it is widely assumed that workers only turn to informality as a last measure when they cannot find a formal sector job.[3]

We develop a quantitative model that allows us to explore the implications of an expansion in the *maquila* sector for Mexico's industrial structure and labor market outcomes, such as the skill premium, the share of the labor force employed in the informal sector and overall welfare. We calibrate our small open economy, two-sector, two-factor model of trade with firm heterogeneity and the possibility of informal employment for unskilled workers to match key cross-sectional moments of the Mexican economy.

Our model takes into account the fact that *maquiladoras* differ substantially from non-*maquiladora* manufacturing plants across several dimensions. Namely, *maquiladoras* (i) are less skill-intensive (their share of production workers in total employment tends to be higher than that of non-*maquila* manufacturing plants)[4], (ii) use a high share of imported intermediate inputs, (iii) are more likely to be foreign-owned, and (iv) are on average larger in terms of total employment than non-*maquila* manufacturing plants.

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1 Data from CNIME (*Consejo Nacional de la Industria Maquiladora y Manufacturera de Exportación*, National Council of Maquiladora Industries), Mexico's central bank, *Banco de México*, and *Sistema de Cuentas Nacionales de México* (Mexican National Income and Production accounts) from Mexico's national statistical agency, *Instituto Nacional de Estadística y Geografía* (INEGI).

2 At the height of the Tequila crisis in 1995, the unemployment rate reached a peak of 7%.

3 For a different view see Maloney (2004) who stresses the positive entrepreneurial aspects of the informal sector.

4 Robertson (2007) using data from Mexico's Monthly Industrial Survey for 1994 and 2004, shows that the non-production/production (N/P) employment ratio for *maquiladoras* is lower than for non-*maquila* plants in almost all industries where *maquiladoras* operate. This fact seems at odds with Feenstra and Hanson (1997) who find that during the 1980's the relative demand for non-production workers was higher in regions where *maquiladoras* expanded most rapidly. However, Bernard et al. (2010b) find that controlling for industry, *maquiladora* plants do employ a higher N/P ratio than non-*maquila* manufacturing plants. The reason behind these seemingly contradictory facts is that *maquiladoras* are concentrated in low-skill intensive industries. Since in our model we treat *maquila* as a completely separate industry from non-*maquila* manufacturing, we assume that the *maquila* sector is relatively low-skill intensive.

Concerning informality, our model seeks to incorporate three main stylized facts about the Mexican labor market: (i) a large share of the labor force is employed in the informal sector, (ii) the vast majority of informal workers has low educational attainment, and (iii) there is a formality premium: on average, informal workers earn lower wages than comparable individuals employed in the formal sector.[5]

We use our model to simulate an exogenous increase in the foreign demand for *maquila* output that replicates the observed increase in the sector's share of GDP during the 1990s. Our results suggest that the rise of the *maquiladoras* has been more of a mixed blessing than a panacea for Mexico. We find that despite *maquila* production being relatively intensive in unskilled labor, the expansion of the sector is accompanied by a much larger contraction in non-*maquila* manufacturing. This ultimately results in a smaller number of open vacancies and higher informality. The response of factor rewards resembles a Stolper-Samuelson effect: the increase in demand for the low-skill intensive *maquila* output induces a reduction in the skill premium. Although the reduction in the skilled wage follows directly from the contraction of the skill-intensive manufacturing sector, the increase in the unskilled wage is due to an increase in the recruitment costs of unskilled workers. This result is in turn a consequence of lower average productivity and a higher price index in Mexican manufacturing caused by the expansion of the *maquila* sector. Given the magnitude of the changes in skilled and unskilled wages as well as the increase in informality and the price index faced by Mexican consumers, our model predicts a reduction in real income, our welfare measure.

Our study of the expansion of the *maquiladoras* in an economy with an informal sector contributes to three separate strands of the literature seeking to understand how globalization shapes labor market outcomes. Despite their considerable importance to aggregate exports in several developing countries, the behavior of export processing firms like *maquiladoras* has not been explored in models of international trade that combine firm heterogeneity and labor market frictions such as those by Felbermayr et al. (2011a) and Helpman and Itskhoki (2010). Similarly, models that study the causes, consequences and implications of informality in developing countries using a search and matching framework (Zenou, 2008; Satchi and Temple, 2009; Albrecht et al., 2009) have also overlooked export-processing plants. Moreover, since these models assume a very stylized view of the production side of the economy, usually considering only one-worker firms, they are unable to take into account

[5] For a more detailed description of the stylized facts about *maquiladoras* and the informal sector in Mexico, please refer to the working paper version of this article, Heid et al. (2011).

the significant differences between *maquiladoras* and other manufacturing plants highlighted above. Finally, incorporating the informal sector and its importance in Mexico allows us to shed new light on the aggregate implications of the *maquila* phenomenon, an area of inquiry that has been studied by Feenstra and Hanson (1997) and Bergin et al. (2009).

While this paper focuses on the case of Mexico, we believe our model can also be applied to other developing countries where export processing zones (EPZs) similar to the *maquiladora* program have been instrumental in attracting large FDI inflows. By 2006, 130 countries had established more than 3,500 EPZs accounting for 66 million employees world-wide.[6] Crucially, many of these countries are also characterized by large informal sectors as described in depth by Gasparini and Tornarolli (2009) and de Laiglesia and Jütting (2009).

The remainder of the paper is structured as follows: Section 4.2 presents our model. Section 4.3 provides details on chosen parameter values, presents the empirical moments matched in the calibration and evaluates the model's fit to the data. We present our counterfactual experiment evaluating the rise of the *maquiladoras* during the 1990s in Section 4.4. Section 4.5 concludes.

4.2 The model

In this section we present a model that combines the setup of Bernard et al. (2007) and Felbermayr et al. (2011a) and extends these models to incorporate an informal sector arising from search frictions as well as export processing firms which can differ substantially from regular manufacturing firms along several dimensions such as size, ownership status and skill-intensity. Our heterogeneous-firm framework features resource reallocation between and within industries in response to exogenous changes in foreign demand, which in turn result in labor market adjustments which are important determinants for evaluating the implications of the rise of the *maquiladoras* on labor market outcomes in Mexico.

We assume that Mexico is a small open economy and treat the US as the rest of world, abstracting from all other trade partners. This is not unduly restrictive, since 80% of all Mexican exports are shipped to the US.[7] Thus, we only model Mexico explicitly and take the foreign price indices, expenditure shares and prices of the imported goods as given.

[6] China alone accounts for 40 million employees, Latin America for 5.5 million employees, the transition economies in Eastern Europe for 1.4 million employees; for further details, see Singa Boyenge (2007).

[7] In 1991, 79.4% of all exports were shipped to the US; in 2009, 80.5%.

We assume that production in Mexico takes place in two sectors, *maquila*, $j = 1$, and non-*maquila* manufacturing, $j = 2$, both populated by firms that are heterogeneous with respect to their productivity.[8] There are two types of labor, skilled and unskilled, and we assume that Mexico is abundant in unskilled labor.

Due to the existence of search and matching frictions, not all low-skill individuals can gain employment in *maquiladoras* or manufacturing firms, which means that that a share of them has to resort to informality. We assume that the matching process between unskilled individuals and formal firms is governed by only one matching function, that is, we assume that the labor market for unskilled workers is unified. This in turn means that what determines the probability of an unskilled worker finding a formal job is the total number of vacancies open in the formal sector (i.e. in the *maquila* and manufacturing sector altogether), and that a matched unskilled worker earns the same wage working in a *maquiladora* or in a manufacturing firm. The labor market for skilled workers, on the other hand, is assumed to be perfectly competitive, which is in line with the low share of skilled informal workers observed in the data.

4.2.1 Consumption

Mexican households only consume goods produced in the manufacturing sector, which means that *maquila* output is exported in its entirety. Consumers maximize

$$C_2 = M_2^{\frac{1}{1-\sigma}} \left[\int_{\omega \in \Omega_{2d}} [q_{2d}(\omega)]^{\frac{\sigma-1}{\sigma}} d\omega + \int_{\omega' \in \Omega_{2f}} [q_{2f}(\omega')]^{\frac{\sigma-1}{\sigma}} d\omega' \right]^{\frac{\sigma}{\sigma-1}}, \tag{4.1}$$

where Ω_{2d} is the set of varieties produced in the manufacturing sector in Mexico, and Ω_{2f} the set of varieties imported from the US, $\sigma > 1$ is the elasticity of substitution and M_2 denotes the total mass of manufacturing varieties available in Mexico.[9] We follow Blanchard and Giavazzi (2003) and normalize utility by $M_2^{\frac{1}{1-\sigma}}$ in order to ensure that an increase in the size of an economy does not mechanically translate into a smaller informal sector.

[8] Hereafter we will refer to the non-*maquila* manufacturing as manufacturing sector for short.

[9] The total number of manufacturing varieties available for consumption in Mexico is $M_2 = M_{2d} + M_{2x}^f$ where M_{2x}^f denotes the mass of imported varieties.

Taking into account the existence of iceberg transportation costs $\tau_2 \geq 1$ for imported varieties, the price index corresponding to the composite C_2 is given by:

$$P_2 = M_2^{\frac{1}{\sigma-1}} \left[\int_{\omega \in \Omega_{2d}} [p_{2d}(\omega)]^{1-\sigma} d\omega + \int_{\omega' \in \Omega_{2f}} [\tau_2 p_{2f}(\omega')]^{1-\sigma} d\omega' \right]^{\frac{1}{1-\sigma}}. \tag{4.2}$$

Inverse demand for domestic and imported foreign varieties from sector 2 is then given by:

$$p_{2d}(\omega) = \left(\frac{Y}{M_2}\right)^{\frac{1}{\sigma}} P_2^{\frac{\sigma-1}{\sigma}} q_{2d}(\omega)^{-\frac{1}{\sigma}}, \quad p_{2f}(\omega) = \left(\frac{\tau_2 Y}{M_2}\right)^{\frac{1}{\sigma}} P_2^{\frac{\sigma-1}{\sigma}} q_{2f}(\omega)^{-\frac{1}{\sigma}}, \tag{4.3}$$

where Y denotes total expenditure in Mexico. Note that we define $p_{2f}(\omega)$ as the cif price in the US and $q_{2f}(\omega)$ is the total quantity produced, including the quantity lost in transit due to the iceberg transportation costs.

4.2.2 Production

Firms in both sectors are heterogeneous with respect to their idiosyncratic productivity φ as in Melitz (2003). Since each firm produces a unique variety, we index firm-level variables by φ.

Manufacturing firms. There is an unbounded mass of potential entrants in the domestic manufacturing sector. To enter, producers pay a sunk cost f_{e2}. All costs in the model are denominated in terms of units of the manufacturing good.[10] After incurring this cost, firms draw their productivity from a Pareto distribution with density $g(\varphi) = ak^a\varphi^{-(a+1)}$ for $\varphi \geq k$. Firms that choose to operate need to pay a fixed cost f_2 per period. Having set up a plant, manufacturing firms produce their output by combining skilled labor s and unskilled labor l in a Cobb-Douglas form,

$$q_2(\varphi) = \varphi(s_2)^{\beta_{2s}}(l_2)^{1-\beta_{2s}}, \tag{4.4}$$

where β_{2s} is the labor cost share of skilled workers.

Firms sell their output domestically but can also incur an additional fixed cost f_{x2} to serve the foreign market through exports. We borrow the notion of a small open economy under monopolistic competition from Flam and Helpman (1987), and the extension to a heterogeneous-firm environment proposed by Demidova and Rodríguez-Clare (2009). This assumption implies that, despite the fact that firms located in Mexico face a downward-

[10] Note that this implies that not all output produced can be used for consumption.

sloping demand schedule for their exports, their pricing decisions do not affect the price index, expenditure nor the mass of firms operating abroad. Demidova and Rodríguez-Clare (2013) show that this small country setup is the limit case of a large two-country model in which the labor endowment share of the small country tends to zero. However, the subset of firms exporting to Mexico, M^f_{2x}, is endogenous.[11] Thus, inverse demand for Mexican manufacturing exports abroad is given by

$$p_{2x}(\varphi) = A_{2x}^{1/\sigma} \left(\frac{q_{2x}(\varphi)}{\tau_2} \right)^{-\frac{1}{\sigma}}, \tag{4.5}$$

where A_{2x} is a demand-shifter parameter that is taken as given by Mexican manufacturing firms. Hence, we define total revenue for a Mexican manufacturing firm with productivity φ as:

$$\begin{aligned} r_2(\varphi) &= r_{2d}(\varphi) + \mathbb{I}_x(\varphi) r_{2x}(\varphi) \\ &= \left(\frac{Y}{M_2} \right)^{\frac{1}{\sigma}} P_2^{\frac{\sigma-1}{\sigma}} q_{2d}(\varphi)^{\frac{\sigma-1}{\sigma}} + \mathbb{I}_x(\varphi) A_{2x}^{1/\sigma} \left(\frac{q_{2x}(\varphi)}{\tau_2} \right)^{\frac{\sigma-1}{\sigma}}, \end{aligned} \tag{4.6}$$

where $\mathbb{I}_x(\varphi)$ is an indicator function that takes the value one if a manufacturing firm with productivity φ exports and zero otherwise.

***Maquiladora* firms.** We model *maquiladoras* in a similar fashion to manufacturing firms, therefore in this section we just highlight the differences between the two sectors, namely that (i) *maquila* plants are foreign-owned, (ii) export all their output and (iii) use foreign manufacturing goods as intermediate inputs for production.

A foreign investor pays a sunk entry cost in Mexico to set up a *maquiladora* plant.[12] *Maquiladoras* draw their productivity from the same Pareto distribution as Mexican manufacturing firms. Since *maquiladoras* export all their output, there is no meaningful distinction between domestic and exporting fixed costs. We assume that *maquiladoras* use foreign manufacturing goods as intermediate inputs for production, denoted by i, as well as skilled and unskilled labor. Thus, production of *maquiladora* with productivity φ takes the form

$$q_1(\varphi) = \varphi (s_1)^{\beta_{1s}} (l_1)^{\beta_{1l}} (i_1)^{1-\beta_{1s}-\beta_{1l}}, \tag{4.7}$$

[11] Demidova and Rodríguez-Clare (2009)'s framework needs an endogenous variable to clear the trade balance. There, the price index and expenditure abroad are unaffected by Mexican firms but the share of US firms exporting to Mexico is endogenous.

[12] The fixed costs of entry, operation and vacancy posting for unskilled workers are incurred in Mexico and are denominated in units of the Mexican manufacturing good.

where β_{1s} and β_{1l} are the skilled and unskilled labor cost shares for *maquiladoras*, respectively. Inverse demand for *maquila* variety φ abroad is given by

$$p_{1x}(\varphi) = A_{1x}^{1/\sigma} \left(\frac{q_{1x}(\varphi)}{\tau_1} \right)^{-\frac{1}{\sigma}}, \tag{4.8}$$

where A_{1x} is a foreign demand shifter that *maquiladora* plants take as given and which has a similar interpretation to A_{2x} defined above. $\tau_1 > 1$ are the iceberg transportation costs to ship a *maquila* variety to the US. Total revenues for a *maquiladora* with productivity φ are given by

$$r_1(\varphi) = r_{1x}(\varphi) = A_{1x}^{1/\sigma} \left(\frac{q_{1x}(\varphi)}{\tau_1} \right)^{\frac{\sigma-1}{\sigma}}. \tag{4.9}$$

Unlike Mexican-owned firms in the manufacturing sector, profits derived from the operation of *maquiladoras* are repatriated abroad.

4.2.3 Labor market

Since most individuals employed in the informal sector are unskilled, we assume that search and matching frictions only affect these workers, whereas skilled workers face a perfectly competitive labor market. Thus in our model only unskilled workers are employed in the informal sector. Although we recognize that there are several ways in which informality can be incorporated into a search and matching framework,[13] there is empirical evidence that suggests that informational frictions play a prominent role in the labor market for low-skill and informal occupations.[14]

Following Satchi and Temple (2009), unskilled individuals that are unable to get matched with neither a firm in the formal manufacturing sector nor in the formal *maquiladora* sector become informal workers. These individuals earn income bw_l, with $b \in (0, 1)$, financed by lump-sum transfers from employed individuals, so we can interpret $1 - b$ as the formality wage premium for unskilled workers.[15]

[13] For instance, Zenou (2008) assumes that search and matching frictions only affect the formal labor market, while the informal labor market is assumed to be fully competitive and accessible for everybody. Satchi and Temple (2009) assume that unmatched urban workers become informal as in our model, but they assume the existence of an outside agricultural sector along the lines of the traditional Harris-Todaro model.

[14] Assaad (1993) provides evidence of the importance of kinship and social networking in regulating informal employment in Egypt. Similarly, Wahba and Zenou (2005) find that information sharing through friends and relatives relative to other methods of finding a job is more important for uneducated individuals.

[15] See Appendix D and `http://alejandroriano.weebly.com/research.html` for a variant of the model where workers in the informal sector produce non-traded manufacturing varieties to earn their wage.

In order to hire unskilled workers, firms need to post vacancies v at a cost c per vacancy. As is common in the search and matching literature, we assume that the matching technology is a constant returns to scale Cobb-Douglas function, $m(\theta) = \overline{m}\theta^{-\gamma}$, with $\gamma \in (0,1)$ and where $\theta \equiv v/u$ is the vacancy-informality ratio, and $\overline{m}$ determines the overall efficiency of the matching process in the economy. The probability that a vacancy is filled is given by $m(\theta)$, which is decreasing in θ, and the probability that an unskilled individual in the informal sector finds a job in a formal firm is $\theta m(\theta)$ which is increasing in θ. We follow Keuschnigg and Ribi (2009) and consider a one-shot, static version of the search and matching framework in which the entire population of unskilled workers has just one opportunity to get matched with firms.

The optimal labor demand decision for a manufacturing firm solves the following program:

$$\pi_2(\varphi) = \max_{l_2, s_2} \left\{ r_2(\varphi) - w_l l_2 - w_s s_2 - cP_2\left(\frac{l_2}{m(\theta)}\right) - f_2 P_2 - f_{x2} P_2 \mathbb{I}_x(\varphi) \right\}, \tag{4.10}$$

where we have also made use of the fact that a manufacturing firm wishing to hire l_2 unskilled workers needs to post $l_2/m(\theta)$ vacancies.[16]

The solution to program (4.10) yields two policy rules, one for skilled labor demand, which is the usual condition that the marginal revenue product of skilled labor has to be equal to the skilled wage, w_s, and a second one for unskilled employment, which shows that firms have monopsony power and take into account that their vacancy posting has an impact on the wage rate for unskilled workers:

$$\frac{\partial r_2(\varphi)}{\partial l_2} = w_l + \frac{\partial w_l}{\partial l_2} l_2 + \frac{cP_2}{m(\theta)}. \tag{4.11}$$

As in Stole and Zwiebel (1996) we assume that unskilled workers bargain individually with their employers about their wage and are all treated as the marginal worker. Total surplus of a worker-employer match is split according to a generalized Nash bargaining solution in each sector j, i.e. $(1-\mu)[E(\varphi) - U] = \mu \partial \pi_j(\varphi)/\partial l_j$ where $E(\varphi)$ denotes the income of an unskilled worker being employed at a firm with productivity φ, U is the income of a worker in the informal sector, and $\mu \in (0,1)$ measures the bargaining power of a worker.

Following the same procedure as in Felbermayr et al. (2011a) and Larch and Lechthaler (2011) (i.e. combining the first-order conditions for unskilled employment by plants in both

[16] The labor demand program for *maquila* plants is almost identical to equation (4.10), the only difference being that *maquiladoras* also need to choose how much foreign intermediate inputs to use for production.

sectors together with the surplus-splitting rule), yields a set of two job-creation conditions (one for each sector):

$$w_l + \frac{cP_2}{m(\theta)} = \left[\frac{\beta_{1l}(\sigma-1)}{\sigma - \beta_{1l}\mu + \beta_{1l}\sigma\mu - \sigma\mu}\right]\varphi p_{1x}(\varphi)s_1(\varphi)^{\beta_{1s}}l_1(\varphi)^{\beta_{1l}-1}i_1(\varphi)^{1-\beta_{1s}-\beta_{1l}}, \tag{4.12}$$

$$w_l + \frac{cP_2}{m(\theta)} = \left[\frac{(1-\beta_{2s})(\sigma-1)}{\sigma + \beta_{2s}\mu - \mu - \beta_{2s}\sigma\mu}\right]\varphi p_{2d}(\varphi)\left(\frac{s_2(\varphi)}{l_2(\varphi)}\right)^{\beta_{2s}}, \tag{4.13}$$

and the wage curve is given by:

$$w_l = \frac{\mu c P_2}{(1-\mu)(1-b)}\left[\theta + \frac{1}{m(\theta)}\right]. \tag{4.14}$$

Note that since we assume that the labor market for unskilled workers is unified, this implies that wages for unskilled formal workers are the same in both manufacturing and *maquiladora* firms. The same holds for skilled workers.

4.2.4 Productivity cutoffs and entry

As described in Section 4.2.2, the production side in our model closely follows Melitz (2003). Because $\pi_j(\varphi)$ is a strictly increasing function of φ, only firms with high enough productivity to earn non-negative profits will start production. Thus the usual productivity cutoff for production in sector j is defined implicitly by $\pi_j(\varphi_j^*) = 0$. In the manufacturing sector, where firms need to incur a fixed cost to serve the foreign market, an export cutoff is similarly defined as $\pi_{2x}(\varphi_{2x}^*) = 0$. We follow Melitz (2003) and define average productivity in sector j as:

$$\tilde{\varphi}_j \equiv \left[\frac{1}{1-G(\varphi_j^*)}\int_{\varphi_j^*}^{\infty} \varphi^{\sigma-1}g(\varphi)d\varphi\right]^{\frac{1}{\sigma-1}}, \quad j = 1, 2. \tag{4.15}$$

Using the cutoff productivity of the least productive exporting manufacturing firm φ_{2x}^*, we can define the average productivity for manufacturing exporters analogously. Finally, let $\chi_2 \equiv [1-G(\varphi_{2x}^*)]/[1-G(\varphi_2^*)]$ denote the ex-ante probability that a manufacturing firm exports, conditional on successful entry. Using these definitions we can write the free-entry condition for firms in sector j as $[1-G(\varphi_j^*)]\overline{\pi}_j = f_{ej}P_2$.[17]

[17] For *maquiladoras* $\overline{\pi}_1 = \pi_1(\tilde{\varphi}_1)$ and for manufacturing firms $\overline{\pi}_2 = \pi_{2d}(\tilde{\varphi}_2) + \chi_2\pi_{2x}(\tilde{\varphi}_{2x})$.

4.2.5 Aggregate variables

The equilibrium share of informal workers in the labor force follows from the one-period equivalent of the Beveridge curve and is given by $u = 1/[1 + \theta m(\theta)]$. The mass of firms operating in sector j in Mexico, M_{jd}, is pinned down by the labor market clearing condition for unskilled workers:

$$M_{1d} = \frac{L_1}{l_1(\tilde{\varphi}_1)}; \qquad M_{2d} = \frac{L_2}{l_{2d}(\tilde{\varphi}_2) + \chi_2 l_{2x}(\tilde{\varphi}_{2x})}, \tag{4.16}$$

with $L_1 + L_2 = (1-u)\overline{L}$, where L_j denotes total unskilled employment in sector j and $\overline{L}$ is the total endowment of unskilled labor in the economy. Market clearing for skilled labor is given by $M_{1d}s_1(\tilde{\varphi}_1) + M_{2d}\big[s_{2d}(\tilde{\varphi}_2) + \chi_2 s_{2x}(\tilde{\varphi}_{2x})\big] = \overline{S}$. Finally, the trade balance condition reads:

$$\underbrace{M_{1d}r_1(\tilde{\varphi}_1)}_{\text{value of } maquila \text{ exports}} + \underbrace{\chi_2 M_{2d} r_{2x}(\tilde{\varphi}_{2x})}_{\text{value of manufacturing exports}} =$$

$$\underbrace{\tau_2^{1-\sigma}\left(\frac{Y}{M_2}\right)\left(\frac{P_2}{P_2^f}\right)^{\sigma-1}}_{\text{value of manufacturing imports}} + \underbrace{\tau_2 P_2^f M_{1d} i_1(\tilde{\varphi}_1)}_{\text{value of intermediate imports}} + \underbrace{M_{1d}\pi_1(\tilde{\varphi}_1)}_{\text{aggregate } maquila \text{ profits}} . \tag{4.17}$$

We define the foreign price index for manufacturing goods, P_2^f, as the *numéraire.* Note that aggregate profits in the manufacturing sector remain in Mexico, since firms in this sector are domestically owned.

4.3 Bringing the model to the data

We calibrate parameters in order to match observations both at the aggregate and at the cross-sectional level for the Mexican economy.[18] Table 4.1 presents the parameters used in the benchmark solution of the model.

We normalize the endowment of unskilled labor $\overline{L}$ to 1,500, and choose the endowment of skilled labor to match an employment share of production workers in Mexican manufacturing of 0.825. Factor shares in each sector $\{\beta_{jk}\}_{j=1,2}^{k=s,l}$ are calibrated using national accounts data. In order to be consistent with our model, we take the gross value of production in the *maquila*

[18] Unless otherwise noted, all figures correspond to the year 2000.

Table 4.1: Parameters for the baseline economy

Parameter	Description	Value
σ	Elasticity of substitution	3.800
Foreign market		
P_2^f	Price index manufacturing abroad (*numéraire*)	1.000
$\mathbf{p_{1x}}(\tilde{\varphi}_1)$	**Variety price of the average *maquila* exporter**	2.858
$\mathbf{p_{2f}}(\varphi_2^{*f})$	**Variety price of the marginal US mfg. exporter**	16.680
$\mathbf{A_{1x}}$	**Foreign demand shifter *maquila***	33,527.635
$\mathbf{A_{2x}}$	**Foreign demand shifter manufacturing**	1,691.753
Labor market		
$\overline{L}$	Unskilled labor endowment	1,500.000
$\overline{S}$	Skilled labor endowment	318.864
μ	Bargaining power unskilled workers	0.500
γ	Matching function elasticity	0.500
$1-b$	Formality premium	0.290
$\mathbf{c}$	**Vacancy posting fixed cost**	0.001
$\overline{\mathbf{m}}$	**Efficiency of matching function**	0.603
Factor shares		
β_{1l}	Unskilled labor share *maquila*	0.089
β_{1s}	Skilled labor share *maquila*	0.028
β_{1i}	Foreign intermediates share *maquila*	0.884
β_{2l}	Unskilled labor share manufacturing	0.571
β_{2s}	Skilled labor share manufacturing	0.429
Productivity distribution		
a	Pareto distribution shape parameter	3.400
k	Pareto distribution lower bound	0.200
Transport costs		
$\{\tau_j\}_{j=1,2}$	Iceberg transportation costs in sector j	1.000
Fixed costs		
f_{e2}	Fixed entry cost manufacturing	1.000
$\mathbf{f_{e1}}$	**Fixed entry cost *maquila***	42.266
$\mathbf{f_1}$	**Fixed cost of production *maquila***	64.264
$\mathbf{f_2}$	**Fixed cost of production manufacturing**	0.311
$\mathbf{f_{x2}}$	**Fixed cost of exporting manufacturing**	0.135

Note: Parameters in bold are chosen to match calibration targets defined in Table 4.2.

sector to be composed of wage payments and consumption of foreign intermediate goods, which yields $\beta_{1l} = 0.089$, $\beta_{1s} = 0.028$ and $\beta_{1i} = 1 - \beta_{1l} - \beta_{1s} = 0.884$. In the manufacturing sector, the gross value of production is entirely accounted for by wage payments, resulting in $\beta_{2l} = 1 - \beta_{2s} = 0.571$ and $\beta_{2s} = 0.429$. Thus, $\beta_{2s}/\beta_{2l} > \beta_{1s}/\beta_{1l}$, implying that the manufacturing sector's production is more skill-intensive than that of *maquiladoras*.

Since, as Satchi and Temple (2009) note, there are no studies that estimate search and matching models for Mexico, we choose to set both the elasticity of the matching function, γ, and the bargaining power of unskilled workers, μ, to 0.5, a common parametrization used in the calibration of search and matching models as exemplified by Petrongolo and Pissarides (2001), Albrecht et al. (2009) or Felbermayr et al. (2011a). The parameter b that determines the income that unskilled workers earn in the informal sector is pinned down by the estimate of Binelli and Attanasio (2010) of a 29% formality premium for male employees in Mexico.[19]

The parameters characterizing the distribution from which both *maquiladoras* and manufacturing firms draw their productivity, the shape parameter a and the lower bound of the support k, as well as the elasticity of substitution σ, are chosen following Bernard et al. (2007). Thus, $a = 3.4$, $k = 0.2$ and $\sigma = 3.8$, satisfying the condition that $a > \sigma - 1$, which insures that the variance of the sales distribution is finite. Note that we normalize the fixed entry costs of manufacturing plants f_{e2} to 1. This allows us to interpret the matched magnitudes of the remaining fixed costs as multiples of f_{e2}.

We set the iceberg transportation costs in both sectors $\{\tau_j\}_{j=1,2}$ to 1, reflecting the fact that by 2001, after several rounds of unilateral trade liberalization and NAFTA provisions coming into place, both the average tariff faced by Mexican exporters selling in the US and the average import tariff for manufacturing imports coming from the US into Mexico were below 1.3% as documented by Kose et al. (2004). Due to the proximity of Mexico and the US, we abstract from additional transportation costs. Table 4.2 presents the set of moments that we use to calibrate the remaining parameters of the model which appear in boldface in Table 4.1.

To provide a better sense of how our model fits the data, we present equilibrium variables produced by our model that have not been used as targets in the calibration. Since our model

[19] Binelli and Attanasio (2010) calculate the formality premium as the ratio of mean formal to informal wages for male employees aged between 25 and 60. A worker is considered informal if she does not pay any social security contribution in either the private or public sector. Based on their productive definition of informality, Gasparini and Tornarolli (2009) report a formality premium of 21.9% in Mexico for males with primary education, controlling for age and region, and a 30% premium based on their legalistic definition.

Table 4.2: Calibration targets

#	Statistic to match	Target
1	Share of exporters, manufacturing	0.389
2	Mean plant size, *maquila*	371
3	Mean plant size, manufacturing	214
4	Aggregate trade openness	0.600
5	Share of *maquila* exports in total exports	0.549
6	Yearly transition rate informal $\rightarrow$ formal	0.210
7	Share of informal workers	0.366
8	*Maquila* value added to GDP ratio	0.093
9	Intermediate imports to GDP ratio	0.106
10	Mexican to US GDP ratio	0.091

Note: The share of exporting plants (1) comes from Iacovone and Javorcik (2010). Mean size of *maquila* plants (2) comes from CNIME (*Consejo Nacional de la Industria Maquiladora y Manufacturera de Exportación*, National Council of Maquiladora Industries). Mean plant size for manufacturing (3) is from INEGI, EIA (*Encuesta Industrial Anual*, Annual Manufacturing Survey). Aggregate trade openness (4) is calculated from the World Bank's World Development Indicators. The share of *maquila* exports in total exports (5) comes from CNIME. Both the yearly transition rate from informal to formal employment (6) and the share of informal workers (7) come from Gong et al. (2004). The *maquila* value added to GDP ratio (8) is from INEGI, *Sistema de Cuentas Nacionales de México* (Mexican National Income and Production accounts). The share of intermediate imports for *maquiladoras* in Mexican GDP (9) is from *Banco de México* Balance of Payments statistics. The ratio of Mexican to US GDP (10) is measured in PPP in current US dollars from the World Bank's World Development Indicators.

features a direct relationship between size (measured in terms of employment) and productivity, this implies that *maquiladoras* are the most productive firms in Mexico, being 15% more productive than local manufacturing exporters and 52 % more productive than domestic producers. Unfortunately, since INEGI records plant-level variables for *maquiladoras* and non-*maquiladora* manufacturing plants in different surveys, to the best of our knowledge no study has yet compared the performance of these two types of firms in terms of productivity. Focusing on the manufacturing sector, our model predicts an exporter size premium of 43.5%, which is very close to the 47.4% average reported by Verhoogen (2008) for Mexican manufacturing plants for the period 1993-2001.

To compare the fixed costs of setting up and operating a plant in each sector, we scale them by average sales, thus facilitating the comparison with other studies. Using this metric, our results indicate that the fixed cost of opening a *maquiladora* and the fixed costs of operation account for 21.7% and 33.0% of average sales respectively. The fixed costs paid by Mexican manufacturing firms are substantially smaller. This result is in line with theoretical models in which firms choose whether to serve foreign markets by exporting or through a subsidiary as in Helpman et al. (2004), which assume that the fixed costs associated with FDI are larger than those of exporting. Entry and operation costs for firms operating only in the domestic market amount to 6.8% of total sales. Fixed costs of serving the foreign market by exporting amount to 1.6% of average export sales. The low estimates for the fixed cost of exporting are in line with structural estimates for Colombia reported by Das et al. (2007). Using a structural estimation technique, Riaño (2009) finds the fixed costs of production and exporting for Mexican manufacturing firms to be around 33% of average labor costs and 5% of export sales revenues respectively.[20] Finally, recruitment costs for the average Mexican manufacturing firm are 1.4% of its wage-bill (or 1.2% of its sales), a very close figure to that used by Satchi and Temple (2009) who report vacancy costs of 1.2% of formal sector output in their calibrated model with homogeneous one-worker firms.

Our model is less successful at matching aggregate labor outcomes. The skill premium implied by our model, which is the wage of skilled workers relative to unskilled workers employed in the formal sector, is 1.7, whereas in the data, Robertson (2007) finds the average wage of non-production workers relative to production workers in the Mexican manufacturing sector to be close to 2.7 in 2000. Our model also underestimates the *maquila* sector's share of manufacturing employment (3.5% in our model versus 20% in the data), although this result could easily be overcome if we allowed the manufacturing sector to use intermediate inputs as well. Finally, the informal sector accounts for 22% of GDP in our model, an estimate that falls between INEGI's own conservative estimate of 13% and estimates from Buehn and Schneider (2012) of 30%.

4.4 The rise of the *maquiladoras* during the nineties

We use our quantitative model to evaluate the impact that the extraordinary expansion of *maquiladoras* had on the size of the informal sector, the skill premium and welfare. To do so, we present an experiment in which we increase the exogenous foreign demand shifter for *maquila* goods so as to reproduce the observed increase in the *maquila* sector's share of GDP from 4.2% to 9.9% during the 1990s. This entails increasing A_{1x} from 0.6 to 1.4 times the value used in our benchmark calibration. Table 4.3 summarizes the response of the main endogenous variables to the increase in demand for *maquila* output.

[20] In our model, fixed costs of domestic production correspond to 8% of the total wage-bill for the average domestic manufacturing firm.

Table 4.3: Change in endogenous variables due to an increase in *maquila* goods demand

Variable	**% Change**
***Maquila* sector**	
Average productivity	0
Mass of firms	133.3
Exports	133.3
Unskilled employment	131.7
Skilled employment	138.5
Manufacturing sector	
Average productivity	-0.1
Mass of firms	-5.8
Exports	-5.0
Unskilled employment	-3.9
Skilled employment	-1.1
Share of Mexican exporters	-2.1
Share of US exporters	6.9
Consumer price index in Mexico	3.1
Labor market	
Vacancy-informality ratio	-2.9
Unskilled wage	0.6
Skilled wage	-2.1
Skill premium	-2.7
Share of labor force in informality	0.9
Welfare	-3.7

Note: Table depicts percentage changes in endogenous variables due to an exogenous increase in the foreign demand parameter for *maquila* goods, A_{1x}, by 130%, i.e. from 0.6 to 1.4 times the value used for the benchmark calibration. This increase resembles the rapid expansion of the *maquila* sector during the 1990s, roughly an increase in the *maquila* sector's share of GDP from 4.2 to 9.9%. All other parameters remain at the values from the benchmark calibration.

To evaluate the welfare implications of the expansion of the *maquila* sector for Mexico, we use real wage income as our welfare measure. Because we allow for free entry of firms in both sectors, there are no aggregate profits in equilibrium, as in Melitz (2003). In the *maquila* sector, variable profits are transferred abroad and cover the fixed entry costs of setting up *maquila* plants paid by US investors. Variable profits in the domestically-owned manufacturing sector do not leave Mexico but are also used to pay for entry costs. Informal sector wages are completely financed by the wage income of formal sector workers via lump sum transfers. Due to our assumption of homothetic preferences, consumption patterns of informal sector workers do not differ from those of formal workers. Hence, welfare, stated in terms of the indirect utility function, is simply real wage income:

$$W = \frac{(1-u)w_l\bar{L} + w_s\bar{S}}{P_2}. \tag{4.18}$$

Because by definition *maquiladoras* export all their output, the decision whether to operate or not is characterized by just one productivity cutoff, above which it is profitable for a firm to produce and export, instead of the usual two (one for domestic production, another for exporting) featured in trade models with firm heterogeneity. Moreover, because of our assumption that firms' productivity is drawn from a Pareto distribution, it is easily shown that both the production cutoff and average productivity for *maquiladoras* are independent of A_{1x}. Thus, the increase in demand for *maquila* output leads to an adjustment on the extensive margin (the mass of firms), but not on the intensive margin (firm size) in the *maquila* sector.[21] Thus, our model produces a one-to-one increase in both the mass of *maquiladora* firms and the value of *maquila* exports, both increasing by a factor of 2.3.

How does the expansion of the *maquila* sector affect non-*maquila* manufacturing and labor market outcomes? Since the *maquila* sector always presents a trade surplus, it follows that its expansion needs to be balanced by an increase in the manufacturing sector's trade deficit in order to maintain equilibrium in the balance of payments. This adjustment occurs on two fronts: the share of US-based manufacturing firms exporting to Mexico increases by 6.9%, while at the same time the share of Mexican manufacturing exporters falls by 2.1%. In contrast to the *maquila* sector, there is a within-sector reallocation of market shares in manufacturing. Lower expected profits in the foreign market for Mexican manufacturing

[21] This contrasts with the usual result in heterogeneous-firm models, in which increasing the profitability of exporting, by reducing iceberg transportation costs, for instance, produces a within-industry reallocation of resources from low to high-productivity firms.

firms are compensated by higher domestic profits, which are reflected in a lower cutoff of production for the domestic market, inducing entry of firms in the lower end of the productivity distribution.

As foreign demand for Mexican manufacturing goods weakens following the expansion of the *maquila* sector, the mass of manufacturing firms and average productivity in this sector fall by 5.8% and 0.1% respectively, resulting in an increase in the manufacturing price index of 3.1%. From the labor market perspective, because the manufacturing sector is relatively skill-intensive, we observe that it sheds 1.1% of its skilled employment, while reducing its unskilled employment by 3.9%. Some of the unskilled workers that leave manufacturing will find a job in the *maquila* sector, whereas the unlucky ones that are unable get matched will join the informal sector.

As we mention in the previous section, because of the high cost share of foreign intermediates in the production of *maquila* output, this sector only accounts for 4.1% of total unskilled employment in our model. This means that in aggregate, the contraction of the manufacturing sector dominates the increase in demand for unskilled workers in the *maquila* sector, resulting in a reduction in the number of vacancies opened for unskilled workers and an increase in informality of 0.9%. This effect is reinforced by the fact that the higher manufacturing price index increases the cost of recruiting unskilled workers.

In terms of wages, the reduction in the demand for skilled labor caused by the contraction in manufacturing leads to a reduction in the skilled workers' wage of 2.1%. For the wage of unskilled workers, there are two effects at work that operate in opposite directions. On the one hand, the reduction in the total number of vacancies decreases the vacancy/informality ratio, θ, curtailing the bargaining power of unskilled workers. A lower θ means unskilled workers find it more difficult to get matched with firms in the formal sector, which reduces the share of the match's surplus that they can retain when negotiating their wage. On the other hand, a higher recruitment cost cP_2 means that matched workers are rewarded for reducing firms' recruitment costs as noted by Pissarides (2000). In our quantitative model, the second effect dominates, and wages of unskilled workers increase by 0.6%. These predictions are in line with Waldkirch (2010), who finds that a 10% increase in *maquila* FDI reduces wages of skilled workers by 0.19% without having any significant effect on the wages of unskilled workers. In our model, a 10% increase in the foreign demand for *maquila* output decreases the wage of skilled workers by 0.27%, increasing the wage of unskilled workers by just 0.08%.

The movements in absolute wages imply a 2.7% reduction in the skill premium. This is consistent with the observed pattern of the average relative wage of non-production workers in Mexican manufacturing documented by Robertson (2007).[22] The skill premium started to fall gradually after 1994, following the tremendously rapid increase of more than 30% that characterized the second half of the 1980s and early 1990s, when most of Mexico's unilateral trade and investment liberalization reforms took place. Robertson suggests that the steady rise in the price of *maquila* output relative to that of non-*maquila* manufacturing observed after 1995 could explain the fall in the skill premium via a Stolper-Samuelson mechanism. Our quantitative model suggests that although the expansion of the *maquila* sector might not have been large enough to reduce informality, it could have contributed to the fall in the skill premium.

Finally, since the rise of the *maquiladoras* increases both the price index faced by consumers and the share of unskilled workers in informality, while at the same time reducing the wage of skilled workers and, to a lesser extent, increasing the wage of unskilled workers, we find that real income, our welfare measure for the Mexican economy, falls by 3.7%.

4.5 Conclusion

This paper investigates how the rise of the *maquila* sector during the 1990s affected informality, the skill premium, and welfare in Mexico. Using a quantitative model with heterogeneous firms and imperfect labor markets calibrated to match key stylized facts of the Mexican economy, we find that the expansion of the *maquila* sector during the 1990s increased the size of the informal sector and reduced overall welfare in Mexico by 0.9% and 3.7% respectively, while at the same time reducing the skill premium by 2.7%. Thus, our quantitative model suggests that the expansion of the *maquila* sector may have been a mixed blessing for Mexico.

[22] Similarly, Airola (2008) finds only weak evidence that growth in *maquila* employment has increased the skill premium using Mexican household survey data.

Chapter 5

Migration, Trade and Unemployment*

5.1 Introduction

Does immigration lead to higher unemployment rates in the destination country? As immigration and trade exposure of a country are highly correlated, it is the aim of this paper to study the effect of immigration on unemployment in OECD countries explicitly taking into account trade volumes of receiving countries.

This question is of eminent political importance as its answer, or at least what policy makers perceive as its correct answer, has direct consequences for millions of potential migrants across the globe. For example, as a reaction to rising unemployment rates in the wake of the financial crisis, several countries implemented voluntary return programs (VRPs) for migrants with entitlements to domestic unemployment benefit schemes. These programs offered financial incentives like a free one way return ticket as well as lump sum payments if immigrants left the host country and did not return for at least three years. Even though few of the migrants eligible for the programs did actually participate, according to Manzano and Vaccaro (2009), the Spanish government spent 21 € million in 2009 on this kind of program.[1]

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[1] Besides Spain also the Czech Republic and Japan have introduced VRPs. For further details see Fix et al. (2009).

At the European level, in the vein of the last two enlargements of the European Union, both treaties of accession[2] contained clauses about a transition period before workers from the new member states could be employed on equal, non-discriminatory terms in the old member states as policy makers feared negative effects on labor markets in the EU-15 countries. The old member states had the possibility to impose restrictions for worker immigration for a transitional period of two years. Afterwards, they could decide to extend it for another three years. After five years, if the country informed the European Commission of serious disruptions on its labor market the period could be extended for the last time for two more years.[3]Austria and Germany were the only member states which used up the whole seven year period for shutting off their labor markets from inflows from eight of the ten accession countries from 2004 (from all but Malta and Cyprus). This seven year period ended on May 1st, 2011.

How did Austria and Germany actually argue for the serious disruption on their labor markets? Basically, two arguments where brought forward defending transitional immigration restrictions. First, Germany's State Secretary for Employment Gerd Andres defended Germany's decision to maintain restrictions by pointing out that the disruptions brought about by adjustment effects would be too high without the transitional restrictions. Second, he argued that "the geographical position is very different for Germany and Austria than it is for France or the UK".[4] EU-Employment Commissioner Vladimir Špidla accepted the application for prolongation of the restrictions from both Austria and Germany by arguing that both countries "are undergoing serious disturbance of their labour markets as a consequence of the general economic downturn."[5] In essence, the reports to the European Commission only argued for the supposedly existing disruptive consequences of what was perceived as a premature opening of labor markets. However, to the best of our knowledge, no evidence was provided which would back up the causal link between higher immigration and unemployment or any other detrimental labor market effects. The causality, it seems, was taken for granted.[6]

[2] The "Treaty of Accession 2003" was the agreement between the European Union and ten countries (Czech Republic, Estonia, Cyprus, Latvia, Lithuania, Hungary, Malta, Poland, Slovenia, Slovak Republic) concerning these countries' accession into the EU that took place 2004. The "Treaty of Accession 2005" is an agreement between the European Union and Bulgaria and Romania concerning accession into the EU of the latter two countries that took place 2007.
[3] For more details, see European Commission (2012).
[4] See EurActiv.com (2009).
[5] See Slegers (2009).
[6] See European Commission (2006).

This example illustrates the widely held belief that on average, immigration has detrimental effects on the labor market in the destination country.[7] This is in contrast with much of the current empirical evaluations of the effects of migration on wages of domestic workers in the destination country. These studies can be grouped into three types. The first uses the elementary model of labor demand and carries out simulations in order to quantify the effects (see for example Borjas, 1999). The second approach uses natural experiments, i.e. supposedly exogenous inflows of migrants, like a short episode of easier Cuban immigration to Miami (Mariel boat lift study by Card, 1990) or the immigration to France in the wake of the Algerian independence (Hunt, 1992).[8] The third approach estimates parameters of a regression of (changes) in wages or employment on the number of migrants and a set of control variables to identify the causal effect of immigration (Borjas et al., 1997; Borjas, 1999, and Friedberg and Hunt, 1995). All three approaches usually find very modest effects of immigration on workers in the destination country.[9] Not surprisingly, the Czech government opposed the prolongation of immigration restrictions in Germany and Austria as these were against "available evidence".[10]

All these empirical studies of the effects of immigration were done by labor economists. To the contrary, analysis of the process of European integration, or more broadly globalization in general, typically falls in the domain of trade economists. While trade economists paid only scant attention to labor market frictions for a long time, the effects of globalization on unemployment featured more prominently in recent trade models. This recent literature focuses on models with heterogeneous firms and increasing returns to scale (see Egger and Kreickemeier 2009, 2012; Felbermayr et al., 2011a; Helpman and Itskhoki, 2010; Helpman et al. 2008a, 2010a, 2010b). One of the main findings in this literature is that trade liberalization is likely to reduce unemployment rates.

This literature also spurred new empirical investigations into the trade and unemployment nexus. Dutt et al. (2009) as well as Felbermayr et al. (2011b) investigate empirically the

[7] Using European Social Survey data, Dustmann and Preston (2004) show that EU citizens believe that average wages are brought down by immigrants. In addition, even though Europeans do not think that immigrants take away jobs from domestic workers, they do not think that immigration can relieve labor shortages.

[8] Recent studies have also used the mass inflow of German expellees into West Germany after World War II and of ethnic Germans from former socialist countries after the fall of the Iron Curtain as quasi-natural experiments to identify the causal labor market effects of immigration (see Braun and Mahmoud, 2011 and Glitz, 2012). Also internal migration caused by the Great Depression in the US during the 1930s has been identified as a quasi-natural experiment to study the labor market consequences of immigration, see Boustan et al. (2010).

[9] For a very recent survey on the economic impacts of immigration, see Kerr and Kerr (2011).

[10] See EurActiv.com (2009).

trade and unemployment nexus using high-quality OECD cross-section and panel data. They both find support for a negative relationship between openness and unemployment levels.

While both papers use a battery of labor-market related control variables in their regressions, none considers the effects of (im)migration. This is astonishing as it is well known since Mundell (1957) that "[c]ommodity movements are at least to some extent a substitute for factor movement". In a standard two goods, two factors trade model without trade costs, factor prices will equalize through goods trade. Hence, goods trade has the same effect as if factors could wander freely between countries. In other words, immigration has the same impact on factor prices as trade. When factor prices cannot fully adjust, there will be additional effects on the quantity of labor used, i.e. the unemployment rate. Hence, (factor) price differences between countries will trigger both, trade and immigration flows, implying that trade and immigration are not statistically independent and therefore correlated. While standard neoclassical trade theory predicts that price differentials can be mitigated by either migration or trade which leads to a negative correlation between trade and migration, recent evidence has suggested that immigration may actually spur trade (e.g. Gould, 1994; Felbermayr et al., 2010b). Theoretical predictions concerning the effects of immigration on unemployment are ambiguous and depend inter alia on factor endowments, production and market structure and differences in institutions. In the labor demand model with one sector and rigid wages, immigration leads to an increase of unemployment (see Boeri and van Ours, 2008, pp. 178ff.). In general equilibrium trade models with capital and labor as production factors, constant returns to scale and perfect competition, immigration has an ambiguous effect on aggregate unemployment (see for example Brecher and Chen, 2010). To the contrary, with increasing returns to scale and monopolistic competition, immigration leads to a fall of unemployment (see Epifani and Gancia, 2005 and Südekum, 2005). There are many good surveys about international migration and trade. Gaston and Nelson (2011) is a particular useful one in the context of this paper as it surveys current theoretical and empirical research on international migration with a particular emphasis on the links between trade theory and labor empirics.

In the light of this discussion the question arises why goods trade should have a statistically significant effect on unemployment whilst (im)migration has not. And when trade decreases unemployment, should not (im)migration, too? If the answer to these questions is in the affirmative, one has to conclude that previous studies may suffer from a potential omitted variable bias. The direction of this bias is not clear a priori, as it depends on whether trade and migration are substitutes or complements.[11]

We want to contribute to both the trade and immigration literature and address this omitted variable bias by considering not only the effects of goods trade flows on unemployment, but also of migration flows. In order to do so, we have to deal with the problem that migrants do not select their destination countries randomly. Rather, it is likely that they migrate into countries with better economic conditions, including countries with lower unemployment rates. This creates an endogeneity problem. We deal with it by using dynamic panel regressions as well as a Frankel and Romer (1999) type instrument. It uses the fact that immigration flows are to a large part determined by geographic variables like the distance between sending and receiving country, i.e. factors which are arguably exogenous to the determination of the unemployment rate.[12]

Finally, note that we do not distinguish between the impact of immigrants of different skill groups on unemployment as panel data for different immigrant skill classes for a large set of countries and a sufficient time span are not available. We therefore focus on aggregate migration flows to address the concern of policy makers and the public at large which presupposes a positive impact of immigration on the level of the unemployment rate *on average*. Accordingly, the transition periods of the EU accession treaties also presuppose on average a positive impact and do not distinguish between workers of different skill levels. By this we offer an alternative empirical strategy which complements the more micro-level based empirical studies typically undertaken by labor economists.

The remainder of the paper is structured as follows. Section 5.2 describes the database and gives suggestive evidence. Section 5.3 describes the empirical specification. Section 5.4 provides the empirical results. The last section concludes.

[11] Relatedly, Ortega and Peri (2011) also argue that previous studies of the effects of both trade and migration suffer from an omitted variable bias as both trade and migration are highly correlated. They use data on OECD countries from 1980 to 2007 to study the effects of trade and immigration on GDP per capita. However, they do not study effects on unemployment rates.

[12] Ottaviano et al. (2013) study the impact of both migration and offshoring in the US on employment of US workers using a theoretical trade framework as basis for their empirical analysis across manufacturing industries. However, they do not study overall unemployment.

5.2 Data and descriptive evidence

To examine the relationship between migration, trade and unemployment we collected a panel dataset from 1997 to 2007 for 24 OECD countries.[13] The selection of countries as well as the time period is driven by concerns of data availability. In addition, we try to follow Felbermayr et al. (2011b) and use the same control variables in order to replicate their results on the trade and openness link for our dataset. The dataset has the advantage that it allows to control for time-invariant country-specific effects and the dynamics (persistence) of unemployment rates. The variables used are summarized in Table 5.1 and 5.2. We describe each variable in turn in the following.

5.2.1 Unemployment rates, immigration and trade openness

The dependent variable is the yearly average harmonized unemployment rate (as percentage of the civilian labor force) from the OECD (2011d) Key Short-Term Economic Indicators, the same data as used in Felbermayr et al. (2011b). These data have the advantage that they are available for the whole time period under consideration and for all OECD member countries. In addition, the OECD has ensured that unemployment rates are comparable across countries.

The migration data are from the OECD (2011b) International Migration Database. It contains bilateral data both on flows and stocks of immigrants. Note that the data do not contain information on illegal migration. Even though data for some countries are available before 1997, broad coverage only starts then and we therefore opt to start our analysis with this year. Specifically, it contains data on the inflows and outflows of immigrants from country i to j defining a migrant as someone with a different nationality than the receiving country. From these data we construct total inflows of immigrants by collapsing the bilateral data. Also note that outflows do only include foreigners, i.e. return migrants. It does not include nationals leaving their home country. Hence net inflows are inflows of foreign nationals. Note that our regressions only include the receiving countries of immigrants. However, to construct the inflow data we use information about the immigrants from all

[13] The countries included are Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Hungary, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Spain, Sweden, Switzerland, United Kingdom, and United States.

Table 5.1: Summary statistics for migration, trade and unemployment dataset

Variable	Mean	Std. Dev.	Min.	Max.	*N*
Total unemployment rate	6.735	2.896	2.245	19.025	207
Migration data					
Net immigrant inflows (ln)	10.834	1.381	7.651	14.041	207
Total immigrant inflows (ln)	11.308	1.292	8.598	14.041	207
Stock of immigrants (foreign nationals) (ln)	13.386	1.319	9.222	15.711	150
Stock of immigrants (foreign born) (ln)	14.041	1.541	11.365	17.441	111
Net inflows (ln) (prediction)	11.566	1.269	8.949	14.044	207
Openness measures					
Total trade openness	78.883	41.244	22.884	217.786	207
Total current price openness	80.491	38.867	18.188	184.308	207
Merchandise curr. price open.	31.218	17.046	8.236	91.566	207
Merchandise openness	30.325	16.847	8.535	106.512	207
Labor market data					
Wage distortion (index)	57.170	18.418	25.187	92.17	207
EPL (index)	2.008	0.818	0.170	4.330	207
Union density (index)	32.755	20.362	7.617	81.285	207
High corporatism (index)	2.546	1.364	0	6	207
PMR (index)	2.348	0.728	0.900	4.700	207
Other control variables					
Population (ln)	16.749	1.228	15.127	19.525	207
Output gap (%)	0.487	1.562	-2.901	4.752	207
Civil liberties (index)	1.159	0.367	1	2	207
Years since perm. trade lib./1945	42.304	12.423	12	62	207

Table 5.2: Summary statistics for gravity dataset

Variable	Mean	Std. Dev.	Min.	Max.	*N*
Bilateral immigrant inflows	1,286	6,316	0	218,822	41,545
Bilateral geographical distance (ln)	8.570	0.885	5.081	9.880	41,545
Contiguity	0.025	0.157	0	1	41,545
Common official language	0.120	0.325	0	1	41,545
Colonial relationship after 1945	0.019	0.138	0	1	41,545

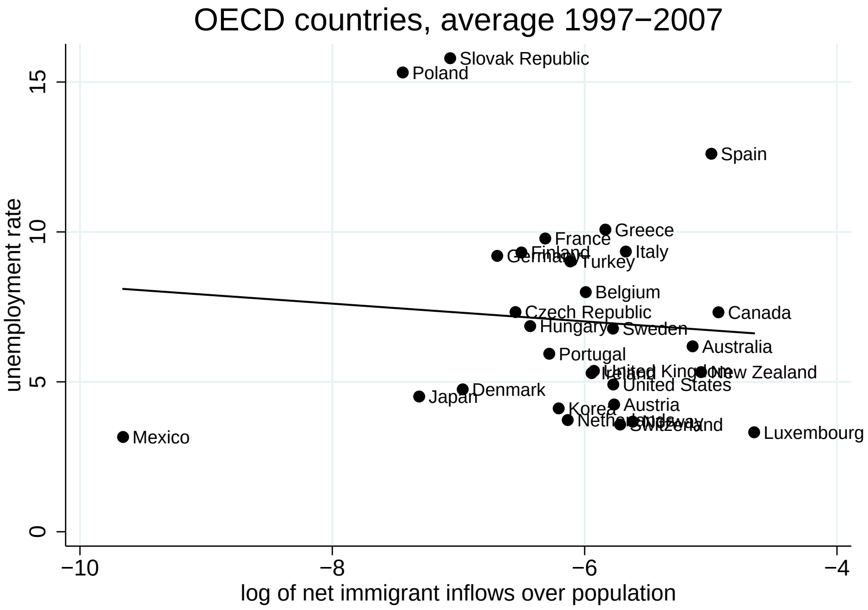

Figure 5.1: Average unemployment and log of net immigrant inflows over population of the receiving country

198 sending countries of immigrants.[14] The huge discrepancy between the high number of sending countries but low number of receiving countries stems from the fact that few countries provide accurate data on immigration. However, those countries that do report these data also have data on the nationalities of all the persons immigrating into the country. Therefore, our data set includes immigrants from all major immigrant sending countries like China, India, North-African and Latin American countries.[15] In addition, the data contain information on the total stock of immigrants, using either an immigrant definition based on the nationality of the person or its country of birth. Note that stock data are only available for a different set of countries as national governments differ in their used definitions of migrants and hence do not necessarily collect data using both definitions.[16]

[14] The complete list of sending countries can be found in Table 5.3.

[15] Countries with flow data used in the regressions are Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Hungary, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Spain, Sweden, Switzerland, United Kingdom, and United States. Outflows are not available (for a subset of years) for Canada, France, Ireland, Italy, Korea, Poland, Slovak Republic, Spain, and Turkey. For these cases, we treat total inflows as net inflows, in effect overstating the number of migrants entering the country. Our main results are robust to this treatment.

[16] Stock data based on nationality are available (for at least a subset of years) for Austria, Belgium, Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Japan, Netherlands, Norway, Poland, Portugal, Slovak Republic, Spain, Sweden, Switzerland, United Kingdom; stock data based on country of birth are available (for at least a subset of years) for Australia, Austria, Belgium, Canada, Denmark, Finland, France, Hungary, Ireland, Netherlands, New Zealand, Norway, Slovak Republic, Spain, Sweden, Switzerland, United Kingdom, and United States.

Table 5.3: List of sending countries of immigrants to construct the inflow data

List of sending countries
Afghanistan, Albania, Algeria, Andorra, Angola, Antigua and Barbuda, Argentina, Armenia, Australia, Austria, Azerbaijan, Bahamas, Bahrain, Bangladesh, Barbados, Belarus, Belgium, Belize, Benin, Bermuda, Bhutan, Bolivia, Bosnia and Herzegovina, Botswana, Brazil, Brunei Darussalam, Bulgaria, Burkina Faso, Burundi, Cambodia, Cameroon, Canada, Cape Verde, Central African Republic, Chad, Chile, China, Chinese Taipei, Colombia, Comoros, Congo, Cook Islands, Costa Rica, Croatia, Cuba, Cyprus, Czech Republic, Côte d'Ivoire, Democratic People's Republic of Korea, Democratic Republic of the Congo, Denmark, Djibouti, Dominica, Dominican Republic, Ecuador, Egypt, El Salvador, Equatorial Guinea, Eritrea, Estonia, Ethiopia, Fiji, Finland, Former Yug. Rep. of Macedonia, France, Gabon, Gambia, Georgia, Germany, Ghana, Greece, Grenada, Guatemala, Guinea, Guinea-Bissau, Guyana, Haiti, Honduras, Hong Kong (China), Hungary, Iceland, India, Indonesia, Iran, Iraq, Ireland, Israel, Italy, Jamaica, Japan, Jordan, Kazakhstan, Kenya, Kiribati, Korea, Kuwait, Kyrgyzstan, Laos, Latvia, Lebanon, Lesotho, Liberia, Libya, Lithuania, Luxembourg, Macao, Madagascar, Malawi, Malaysia, Maldives, Mali, Malta, Marshall Islands, Mauritania, Mauritius, Mexico, Micronesia, Moldova, Mongolia, Morocco, Mozambique, Myanmar, Namibia, Nauru, Nepal, Netherlands, New Zealand, Nicaragua, Niger, Nigeria, Niue, Norway, Oman, Pakistan, Palau, Palestinian administrative areas, Panama, Papua New Guinea, Paraguay, Peru, Philippines, Poland, Portugal, Puerto Rico, Qatar, Romania, Russian Federation, Rwanda, Saint Kitts and Nevis, Saint Lucia, Saint Vincent and the Grenadines, Samoa, San Marino, Sao Tome and Principe, Saudi Arabia, Senegal, Serbia and Montenegro, Seychelles, Sierra Leone, Singapore, Slovak Republic, Slovenia, Solomon Islands, Somalia, South Africa, Spain, Sri Lanka, Sudan, Suriname, Swaziland, Sweden, Switzerland, Syria, Tajikistan, Tanzania, Thailand, Timor-Leste, Togo, Tokelau, Tonga, Trinidad and Tobago, Tunisia, Turkey, Turkmenistan, Tuvalu, Uganda, Ukraine, United Arab Emirates, United Kingdom, United States, Uruguay, Uzbekistan, Vanuatu, Venezuela, Viet Nam, Yemen, Zambia, Zimbabwe.

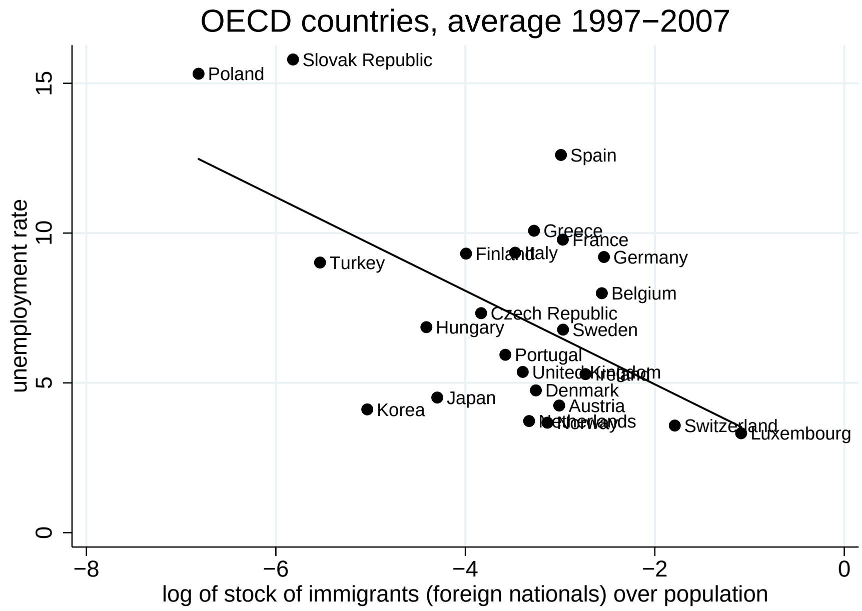

Figure 5.2: Average unemployment and log of stock of immigrants (foreign nationals) over population of the receiving country

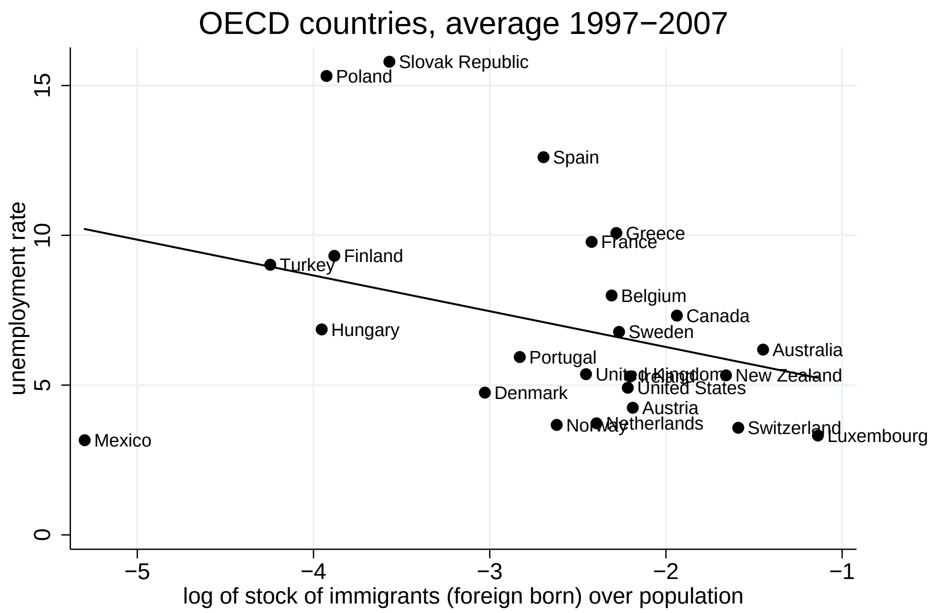

Figure 5.3: Average unemployment and log of stock of immigrants (foreign born) over population of the receiving country

Figure 5.1 provides a first look on the unemployment migration nexus. It plots the average unemployment rate over the period of 1997 to 2007 against the average logged immigration net inflows over the population of the receiving country for the period of 1997 to 2007. As we see, this figure suggests a negative relationship between immigration and unemployment. Figures 5.2 and 5.3 plot average unemployment rates against the average of the logged stock of foreign nationals over the population of the receiving country and the logged stock of foreign born immigrants over population, respectively. Again, we find a negative relationship between immigration and unemployment.

This correlation between unemployment and migration may be misleading due to two main effects: i) It is an unconditional correlation, ignoring potential heterogeneity of countries and other driving factors, and ii) the endogeneity of migration flows and unemployment.

Concerning migration from the perspective of an individual, two questions arise: The first question is whether to migrate at all, and the second question, given that one decided to migrate, where to migrate. The labor literature typically models those two decisions sequentially, where the second step depends on expected wage differences between the origin and destination country, accounting for unemployment differences (see Cahuc and Zylberberg, 2004 and Boeri and van Ours, 2008 for an overview). In other words, immigration will be larger all else equal into countries with lower unemployment rates. This is consistent with Figures 5.1, 5.2, and 5.3.

However, we are interested in the causal effect of immigration on unemployment. Hence, we have to control for the reversed causality. In order to get rid of the endogeneity problem due to reversed causality, we are pursuing two strategies. First, we control for time-invariant and country-specific effects. This wipes out all level effects between countries. Hence, these regressions only use the change in unemployment levels and immigration inflows to identify the coefficients.

Figure 5.4 plots the change of unemployment against the change of immigration inflows over the population of the receiving country. This transformation removes the unobserved time-invariant country-specific heterogeneity. Again, the figure suggests a negative relationship between immigration and unemployment. Our second approach to control for the endogeneity due to reversed causality is to instrument the migration flows. Besides using external instruments, we will follow the established methodology used in Dutt, Mitra, and Ranjan (2009) and Felbermayr, Prat, and Schmerer (2011b) and rely on dynamic panel

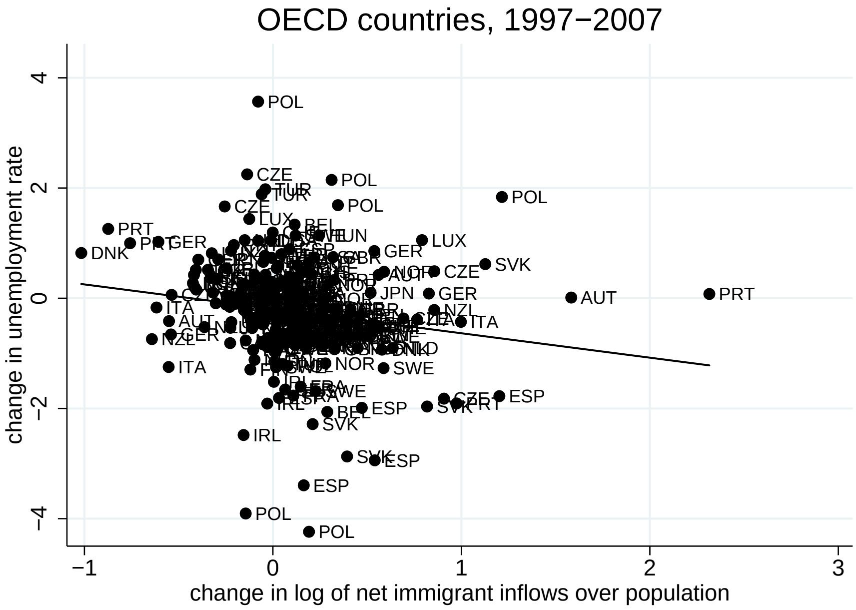

Figure 5.4: Change in unemployment and change in log of net immigrant inflows over population of the receiving country

estimators which use internal instruments, i.e. suitable lags of regressors in both differences and levels, in order to control for both unobserved time-invariant heterogeneity as well as endogeneity of the migration variable. In addition, these estimators allow to control for the possible endogeneity of other control variables like e.g. the openness measure capturing trade linkages of the migration receiving country.

Let us next describe our second main explanatory variable, trade openness. We follow Alcalá and Ciccone (2004) and Felbermayr et al. (2011b) and construct a real openness measure, labeled *total trade openness*. It is defined as the sum of total imports and exports in exchange rate US-$ over GDP in purchasing power parity US-$. We construct it by multiplying the current price openness measure (*total current price openness*) times the GDP price level from the Penn World Tables, edition 7.0 from Heston et al. (2011).[17]

In addition, we construct openness measures for merchandise trade only using data from the OECD (2011c) International Trade by Commodity Statistics database. Again, we calculate a real and current price openness measure (*merchandise openness* and *merchandise current price openness*, respectively). We will describe the used control variables in the following section.

[17] Felbermayr et al. (2011b) use the Penn World Tables, edition 6.2.

5.2.2 Controls

We closely follow Felbermayr et al. (2011b) in our choice of control variables.

Wage distortion is the sum of the tax wedge and the average replacement rate. The tax wedge is the average tax wedge on labor as a percentage of total labor compensation and is computed for a couple with two children and averages across different situations regarding the wage of the second earner. Tax wedge data are from the OECD. Specifically, we use the tax wedge data from Bassanini and Duval (2009) until 2003 and the publicly available data for 2004 to 2007 from the OECD (2011g) Taxing Wages database. Note that for the overlapping years, data from both sources do not perfectly match for some countries. In general, however, data are nearly or even exactly the same. We therefore merge the data to fill up the variable for the whole sample period. The average replacement rates are from the Benefits and Wages study from OECD (2007). They are defined as the average of the gross unemployment benefit replacement rates for two earnings levels, three family situations and three durations of unemployment and is comparable across countries. As data are available only for odd years, we follow Bassanini and Duval (2009) and interpolate the data for even years. For a detailed description of the OECD replacement rate measures, see Martin (1996).

EPL is an employment protection legislation index which is comparable across countries and is from the Going for Growth database from OECD (2010). It measures protection for regular employment and ranges from 0 to 6 from weakest to strongest protection.

Union density corresponds to the ratio of wage and salary earners that are trade union members to the total number of wage and salary earners and is from the OECD (2011e) Labour Force Statistics.

High corporatism is an index variable from the Database on Institutional Characteristics of Trade Unions, Wage Setting, State Intervention and Social Pacts which is compiled at the Amsterdam Institute for Advanced Labour Studies (AIAS) at the University of Amsterdam by Visser (2011). It measures the degree of coordination of wage bargaining in the respective country where 1 indicates firm-level wage bargaining and 5 equals economy-wide bargaining.[18]

[18] Note that Felbermayr et al. (2011b) only use a dummy variable from Bassanini and Duval (2009) to indicate high wage coordination. These data, however, are only available until 2003 and do not vary across our sample period and hence would be dropped from the regression. We therefore use the index measure which contains more information.

PMR is a measure of product market regulation on a scale from 0 to 6 indicating increasing regulatory restrictions to competition from Conway et al. (2006). We again follow Felbermayr et al. (2011b) and use the OECD data on regulation in seven sectors—telecoms, electricity, gas, post, rail, air passenger transport, and road freight—to measure overall product market regulation. As manufacturing sectors are less regulated and open to foreign competition, and most anti-competitive legislation is concentrated in the considered sectors, the measures do reflect an important part of the overall degree of product market regulation in a country, see Conway et al. (2006). The measures are based on regulation-related policies in the respective countries and are specifically constructed to allow cross-country comparisons. Further details on these measures can be found in Nicolette and Conway (2006).[19]

Population is the population of the receiving country from the OECD (2011e) Labour Force Statistics.

Output gap is the output gap in percent as reported in the OECD (2011a) Economic Outlook No. 89 data.

In additional regressions, we include control variables from Dutt et al. (2009). *Civil liberties* is an index from Freedom House (2011) which gives the amount of civil liberties in a country. It runs from 1 to 7 where 1 indicates a maximum of liberties. Dutt et al. (2009) include a dummy which is 1 in the years after a country has permanently liberalized trade. In our sample, all countries have free trade according to this index, hence we cannot include this dummy as it does not have variation. Therefore, we construct the variable *years since liberalization* which measures the years since a country has permanently liberalized its trade. It is based on data collected by Wacziarg and Welch (2008).[20]

To generate the instrumental variable, the predicted bilateral migration flows from a gravity-type migration regression, we use indicators for contiguity, common official language, and common colonial relationship after 1945 as well as the weighted bilateral distance between economic centers of the receiving and sending countries. All variables are from CEPII, see Head et al. (2010). Summary statistics for the gravity dataset can be found in Table 5.2.

[19] Note that the OECD also compiles data on economy-wide measures of product market regulation. These, however, are only collected irregularly, prohibiting their use in a panel study. The used measure is highly correlated with the economy-wide measure for the years where it is available.

[20] We assume the year 1945 for all countries where Wacziarg and Welch (2008) report "always" instead of a specific year as the permanent liberalization year. In our sample, these countries are Norway, Portugal, Switzerland, United Kingdom, and United States.

5.3 Empirical specification

We follow Nickell et al. (2005) and Felbermayr et al. (2011b) and estimate variants of the following dynamic model:

$$\begin{aligned} u_{it} &= \rho u_{i,t-1} + \alpha NETINFLOW_{it} + \gamma OPENNESS_{it} \\ &\quad + \delta \boldsymbol{CONTROLS}_{it} + \nu_i + \nu_t + \varepsilon_{it}, \end{aligned} \tag{5.1}$$

where u_{it} is the unemployment rate in country i at time t, $NETINFLOW_{it}$ is the net inflow of immigrants into country i at time t, $OPENNESS_{it}$ is a standard openness measure (the sum of imports and exports over GDP), $\boldsymbol{CONTROLS}_{it}$ is a vector of control variables, and ν_i, ν_t, ε_{it} are country and period effects and an error term, respectively. In contrast to Felbermayr et al. (2011b) we do not use five-year averages for our regressions as we would lose a lot of observations given the short time-series of the migration data. Additionally, we also want to capture the short-term transitional effects of migration on unemployment in our dynamic specifications which precludes us from taking averages over years.

The standard estimator for dynamic panel models with unobserved time-invariant heterogeneity is the difference GMM estimator as proposed by Arellano and Bond (1991). However, this estimator suffers from potentially huge small sample bias when the number of time periods is small and the dependent variable shows a high degree of persistence, see Alonso-Borrego and Arellano (1999). As unemployment numbers are very persistent, we follow Arellano and Bover (1995) and Blundell and Bond (1998) and also present estimates of the model using system GMM which circumvents the finite sample bias if one is willing to assume a mild stationarity assumption on the initial conditions of the underlying data generating process.[21] This estimator uses moment conditions for the model both in differences and in levels to reap significant efficiency gains. However, efficiency gains do not come without a cost: The number of instruments tends to increase exponentially with the number of time periods. This proliferation of instruments leads to an overfitting of endogenous variables and increases the likelihood of false positive results and suspiciously high pass rates of specification tests like Hansen's J-test, a routinely used statistic to check the validity of the dynamic panel model, see Roodman (2009a). We therefore follow his

[21] Specifically, the deviations from the long-run mean of the dependent variable have to be uncorrelated with the stationary individual-specific long-run mean itself, see Blundell and Bond (1998).

advice and present results with a collapsed instrument matrix for both the difference and system GMM estimators.[22] We also use the Windmeijer (2005) finite sample correction for standard errors.

As described above, $NETINFLOW_{it}$ is likely to be endogenous. Hence, we instrument this variable by suitable lags. In addition, we use an external instrument. To find an external valid instrument, we have to look for other determinants of migration besides destination country unemployment. A natural candidate are predicted migrant inflows, a method inspired by Romer and Frankel (1999) who use predicted trade flows as an instrument for trade flows.[23] The predictions of migrant flows are obtained by estimating a gravity equation. The gravity equation has a long history in the literatures on bilateral aggregate trade and migration flows. In fact, the earliest uses of the gravity equation were to model migration flows, see (Ravenstein, 1885, 1889). Since then, the gravity equation has been used extensively to model migration flows, see Zipf (1946), Stewart (1948), Isard (1975), Sen and Smith (1995). The gravity model was first adopted for studying international trade flows in Tinbergen (1962) and Linnemann (1966), and is well established in the trade literature.

More precisely, bilateral international migration $INFLOW_{ijt}$ is specified as a function of geographic variables, GDPs and so called "multilateral resistance" terms (see Anderson and van Wincoop, 2003):

$$INFLOW_{ijt} = \frac{Y_{it}Y_{jt}}{Y_{wt}} \frac{DIST_{ij}}{P_{it}P_{jt}}, \tag{5.2}$$

where Y_{it} and Y_{jt} are the GDPs of the origin and destination, Y_{wt} is world income, $DIST_{ij}$ is a (potentially multidimensional) time-invariant distance measure between country i and j, and P_{it} and P_{jt} are the measures for origin and destination market potential, or "multilateral resistance" terms.

Typically, Y_{it}/P_{it} and Y_{jt}/P_{jt} are replaced by origin×year and destination×year fixed effects (which also take account of Y_{wt}) and one takes logs of Equation (5.2) in order to get an empirical specification linear in the parameters, allowing to estimate the parameters via ordinary least squares. However, as migration data are likely to be heteroskedastic and contain zero migration flows, taking logs is no longer feasible.[24] Fortunately, there are a couple of recent contributions concerning gravity equation estimation taking into account

[22] All GMM estimations are carried out using the `xtabond2` package in Stata, see Roodman (2009b).

[23] See Felbermayr et al. (2010a) who also use a Romer and Frankel (1999) instrument for immigration to investigate the effect of immigration on per capita income.

[24] Some authors replace zero values by a unit value for the migration flow. In general, this leads to inconsistent estimates.

heteroskedasticity and zero trade flows. Helpman et al. (2008b) propose a sample selection model to account for zero trade flows and show that omitting zero trade flows leads to biased estimates.

Santos Silva and Tenreyro (2006) suggest to estimate the gravity model in multiplicative form employing a Poisson pseudo maximum likelihood estimator in order to account for the "log of gravity". The "log of gravity" says that taking logs of the right and left hand side of the gravity equation may lead to inconsistent and biased estimates because of Jensen's inequality, i.e., $E(\ln INFLOW_{ijt}) \neq \ln E(INFLOW_{ijt})$. This is for example the case in the presence of heteroskedasticity, which is very likely the case with migration and trade data.

In order to account for the heterogeneity and zeros in the bilateral migration flow data, we follow the approach of Santos Silva and Tenreyro (2006). Our empirical specification for the first step gravity model of international bilateral migration flows is therefore:

$$INFLOW_{ijt} = \exp\left(DIST_{ij} + \nu_{it} + \nu_{jt}\right)\varepsilon_{ijt}, \tag{5.3}$$

where ν_{it} and ν_{jt} are origin×year and destination×year fixed effects, and ε_{ijt} is a multiplicative remainder error term. Note that the fixed effects also control for origin and destination variables commonly used in Romer and Frankel (1999) type regressions like the land area covered by the respective country as well as its population.[25]

We specify $DIST_{ij}$ to consist of bilateral geographical distance $(GDIST_{ij})$[26], a contiguity dummy between countries $(CONTIG_{ij})$, a dummy for a common official primary language $(COMLANG\_OFF_{ij})$, and a dummy indicating whether the two countries had a colonial relationship after 1945 $(COL45_{ij})$, i.e.

$$\begin{aligned} DIST_{ij} &= \varrho_1 \ln\left(\mathrm{GDIST}_{ij}\right) + \varrho_2 CONTIG_{ij} + \varrho_3 COMLANG\_OFF_{ij} \\ &\quad + \varrho_4 COL45_{ij}. \end{aligned} \tag{5.4}$$

As our migration data are bilateral but our second stage regression for explaining the unemployment rate has only country-time but no bilateral variation, we sum up our predictions of migration flows $\widehat{INFLOW}_{ijt}$ over all origin countries, i.e.,

[25] The gravity equation explains bilateral total flows of migrants. Hence, we use bilateral total inflows as dependent variable in specification (5.3).

[26] We use the simple weighted bilateral distance measure as proposed by Head and Mayer (2000) which is provided by CEPII and which is defined as distance between the regions in the respective countries weighted by the economic size of the regions.

$\widehat{INFLOW}_{jt} = \sum_{i=1}^{N} \widehat{INFLOW}_{ijt}$ where N is 198, the number of sending countries of immigrants.[27]

5.4 Regression results

In this section, we present our results. In the first subsection we present regression results from our benchmark specification using different estimators. The second subsection discusses several robustness checks concerning different measures of migration, trade openness as well as using additional control variables and sample definitions.

5.4.1 Benchmark results

Table 5.4 presents eight different specifications which all use as dependent variable the unemployment rate and some or all of the following explanatory variables: the net inflows of migrants into the country (in logs), a measure of total trade openness, an index of wage distortion, a measure of employment protection legislation, a measure of union density, an index of the centralization of the wage bargaining process, a measure of product market regulation, a country's size as measured by its population (in log), as well as a measure of the output gap to control for business cycle effects. For the dynamic panel estimators, this list of regressors is augmented by the lagged dependent variable.

Column (1) reproduces column (1) in Table 1 of Felbermayr et al. (2011b) for our sample using a fixed effects estimator (FE). Qualitatively, results are exactly the same as in Felbermayr et al. (2011b). However, in our case only *population* and the *output gap* are significant.

Column (2) adds real total openness as defined in Alcalá and Ciccone (2004). Contrary to Dutt et al. (2009) and Felbermayr et al. (2011b), we find a significant positive effect of international trade on unemployment. However, this does not imply that our results are necessarily at odds with empirical findings in the literature. Both Dutt et al. (2009) and Felbermayr et al. (2011b) use data for a different time period (1985–2004 and 1980–2003, respectively) and also for a larger set of countries with vastly differing levels of development. In addition we do not use five-year averages of the data. Our sample only focuses on a subset

[27] Note that we even do not need to estimate the parameters of the migration equation consistently to use $\widehat{INFLOW}_{jt}$ as a valid instrument. The only assumption we need is that $\widehat{INFLOW}_{jt}$ is a constructed exogenous measure of migration stocks or flows. For a similar argument, see Felbermayr et al. (2011b).

Table 5.4: Benchmark regressions

Dependent variable: unemployment rate	(1) FE	(2) FE	(3) FE	(4) FE-IV	(5) Diff-GMM	(6) Diff-GMM	(7) Sys-GMM	(8) Sys-GMM
Lag dep. var.					2.067	1.424	0.953***	0.951***
					(8.917)	(2.867)	(0.083)	(0.082)
Net inflow (ln)			-0.099	-0.562		-2.832		-0.888**
			(0.287)	(0.393)		(23.82)		(0.390)
Total trade openness		0.028*	0.027*	0.025***	0.376	0.183	0.005	0.003
		(0.013)	(0.014)	(0.010)	(3.444)	(1.415)	(0.008)	(0.007)
Wage distortion (index)	0.019	0.028	0.029	0.035**	0.991	1.260	-0.001	-0.011
	(0.024)	(0.021)	(0.020)	(0.017)	(9.842)	(10.319)	(0.040)	(0.029)
EPL (index)	-1.259	-1.069	-1.125	-1.387*	50.209	26.510	1.205	0.948
	(1.023)	(0.951)	(0.968)	(0.753)	(486.914)	(226.431)	(0.996)	(0.641)
Union density (index)	0.144	0.170	0.173	0.191***	1.012	-0.017	-0.002	-0.003
	(0.105)	(0.105)	(0.106)	(0.060)	(9.368)	(1.388)	(0.027)	(0.021)
High corporatism (index)	-0.023	-0.013	-0.011	-0.002	5.859	-0.415	-0.152	0.345
	(0.079)	(0.083)	(0.081)	(0.115)	(58.163)	(7.215)	(0.473)	(0.473)
PMR (index)	0.487	0.741	0.719	0.614**	0.068	1.852	-0.100	-0.250
	(0.543)	(0.592)	(0.594)	(0.272)	(4.919)	(13.485)	(0.438)	(0.299)
Population (ln)	-20.241*	-18.786*	-18.880*	-19.319***	66.471	48.194	0.240	1.136**
	(11.354)	(10.110)	(9.938)	(4.935)	(459.307)	(222.346)	(0.282)	(0.443)
Output gap	-0.461***	-0.468***	-0.464***	-0.446***	0.190	-0.334	-0.319***	-0.205**
	(0.103)	(0.102)	(0.106)	(0.070)	(4.483)	(1.498)	(0.121)	(0.085)
Observations	224	224	224	224	181	181	207	207
Countries	24	24	24	24	24	24	24	24
Instruments				1	18	19	25	27
R^2 (within)	0.529	0.548	0.549					
R^2 (between)	0.017	0.018	0.018					
R^2 (overall)	0.016	0.015	0.015					
Hansen test (OID)							0.702	0.971
AR(1)					0.924	0.923	0.815	0.587
AR(2)					0.997	0.883	0.337	0.934

Notes: Robust standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1; all models control for unobserved country and period effects. H_0 for AR(1) and AR(2) is no autocorrelation. Openness, output gap, wage distortion, and net inflow treated as endogenous in GMM regressions. Maximum number of lags used is 1. Instrument matrix was collapsed as proposed by Roodman (2009a). Constant estimated but not reported.

of OECD countries for a recent 10-year period due to data availability of the migration data. Differences in the results may therefore simply be due to the specifics of the sample under study. Also remember that we treat all variables as exogenous in specifications (1) and (2), so our results could simply be a result of the endogeneity of our regressors.

In column (3), we add the net inflow of migrants to the specification given in column (2), again using fixed effects. It turns out that the sign of the coefficient of the immigration flow is negative but statistically insignificant. The sign is in line with predictions from new trade theory models with international migration but seems to be in contradiction with predictions based on the labor demand model with wage rigidities. Hence, immigration seems at least not to increase unemployment in the destination country. However, our specification given in column (3) may suffer from an endogeneity bias. As stated in the introduction, migrants might select into countries with lower unemployment rates.

Hence, in column (4) we take as instrument the predicted migration flows based on the Romer and Frankel (1999) instrument described in Section 5.3. We use an instrumental variables panel estimator with fixed effects (FE-IV). Instrumenting migration flows preserves the negative sign but still does not lead to a precise estimate. The coefficient implies that a 1 percent increase in migration inflows leads to a decrease in the unemployment rate of 0.006 percentage points. Openness still has a significant positive effect on unemployment.

Specification (4) still ignores both the persistence of unemployment rates as well as the potential endogeneity of other control variables like trade openness and wage distortion. In addition, the exogeneity of the instrument could be debated as it is inter alia a proxy for the remoteness of a country. It is well known that general remoteness to foreign markets is a determinant of many aggregate variables and therefore could influence unemployment directly. We therefore investigate the effect of migration on unemployment presenting difference and system GMM estimates in columns (5) to (8).

Column (5) presents the specification in column (2) augmented by the lagged dependent variable where we treat openness, wage distortion, *EPL*, as well as the high corporatism measure as endogenous variables using the Arellano and Bond (1991) difference GMM estimator (Diff-GMM). In this specification we do not find a significant effect of the lagged unemployment rate. Additionally, the estimated coefficient on the lag implies a non-stationary behavior and openness is again not significant.

Column (6) adds migration inflows to specification (5) which we also treat as endogenous. It turns out to be non-significant again but still negative. However, one concern in this specification is the high coefficient of the lagged dependent variable. As soon as the dependent variable is highly persistent (our estimates would even imply an explosive behavior of the unemployment rate), the difference GMM estimator has poor small sample properties, see Alonso-Borrego and Arellano (1999). This is reflected in the high standard errors of the estimates.

A suggestion for highly persistent dependent variables is the system GMM estimator due to Arellano and Bover (1995) and Blundell and Bond (1998) which exploits more information conveyed by additional moment conditions. Column (7) repeats specification (5) estimated with system GMM (Sys-GMM). Here, the output gap is significantly negative. In addition, the lagged dependent variable becomes highly significant. It also implies a very high degree of persistence in unemployment rates as expected.

In column (8) we add migration flows to specification (7). Now, migration flows are again negative and also significant on the 5% level. Openness still has a positive impact on unemployment rates but not significantly so. Additionally, the coefficient of the lagged dependent variable has the same magnitude as in previous studies. This is our preferred specification as it allows for the endogeneity of various regressors and can handle the persistence of our dependent variable. It implies that a one percent increase in migration inflows leads to a 0.009 percentage point decrease in the unemployment rate in the short-run. In the long-run, a one percent increase of the total inflow of migrants would amount to a 0.18 percentage point decrease in the unemployment rate.[28]

Note that we report p-values of Hansen's overidentifying restrictions test as well as tests on autocorrelation in the first and second differences of the residuals for both the difference and system GMM estimates. The null hypothesis of valid overidentifying restrictions is not rejected, indicating a well specified model. We do not find autocorrelation in neither the first nor the second differences. Even though one would expect to detect autocorrelation in the first differences when specifying a dynamic panel model, this is not necessary to apply dynamic panel estimators. Autocorrelation in the second differences would be more problematic as it would render some instruments invalid. In any case, it is well known that both the Hansen test as well as the autocorrelation tests suffer from potentially large losses

[28] The long-run effects are found by dividing the coefficient by one minus the coefficient on the lagged dependent variable.

in power for small sample sizes, see Roodman (2009b). He explicitly states that for sample sizes as the ones used in our study with only few time periods, reliance on asymptotic distributions of the test statistics is “worrisome”. As there exists ample evidence on the persistence of unemployment rates, we nevertheless are confident that the system GMM estimator for the dynamic panel model is appropriate.

To sum up, we find no robust statistically significant effect of migration inflows on unemployment rates. Hence, empirical evidence based on a cross-section of aggregate migration flows does not support the widely held belief that immigration is detrimental to employment prospects of workers in the destination country on average.

5.4.2 Robustness checks

In this section we describe two tables with robustness checks. While Table 5.5 presents regressions using different migration measures than used in Table 5.4, Table 5.6 gives results for different trade openness measures, additional control variables and varying sample definitions.

Migration measures

In Table 5.4 we used net inflows of migrants as migration measure where a migrant was defined as a person which does not have the citizenship of the receiving country. Column (1) in Table 5.5 reproduces our preferred specification (8) from Table 5.4 for convenience of comparison. By subtracting return migrants from total immigrants, we assume that it is only the net number of migrants which influences the unemployment rate. From a theoretical point of view, it is not entirely clear whether net or total migration flows should be used. If labor markets are characterized by search frictions, total inflows may be the appropriate measure especially for quantifying the short-run impact as every new migrant has to search for a job. However, in the medium- to long-run or when labor markets are very flexible, net inflows may be more appropriate.

Hence, in column (2) we use total inflow of migrants instead of net inflows. Now, immigration flows are no longer significant but still negative. Interestingly, openness now has a negative but still non-significant impact.

Table 5.5: Robustness checks: Different migration measures

Dependent variable: unemployment rate					
	(1) Sys-GMM	(2) Sys-GMM	(3) Sys-GMM	(4) Sys-GMM	(5) Sys-GMM
Lag dep. var.	0.951***	0.948***	1.039***	1.013***	0.927***
	(0.082)	(0.139)	(0.108)	(0.262)	(0.127)
Net inflow (ln)	-0.888**				0.083
	(0.390)				(0.389)
Total inflows (ln)		-0.059			
		(0.654)			
Total stock (nationality) (ln)			0.007		
			(0.008)		
Total stock (c. of birth) (ln)				3.085	
				(1.956)	
Total trade openness	0.003	-0.006	0.019**	0.051	0.020
	(0.007)	(0.012)	(0.008)	(0.041)	(0.017)
Wage distortion (index)	-0.011	0.059	0.074	0.067	-0.017
	(0.029)	(0.161)	(0.111)	(0.051)	(0.048)
EPL (index)	0.948	3.361	1.520	-1.040	1.308
	(0.641)	(1.936)	(1.232)	(0.966)	(0.752)
Union density (index)	-0.003	0.027	0.001	-0.011	0.014
	(0.021)	(0.033)	(0.066)	(0.027)	(0.039)
High corporatism (index)	0.345	0.315	-0.521	-1.069	-0.210
	(0.473)	(0.751)	(1.238)	(1.340)	(0.446)
PMR (index)	-0.250	-0.364	0.872	1.164	-0.141
	(0.299)	(0.371)	(0.543)	(0.855)	(0.405)
Population (ln)	1.136**	0.442	0.389	-2.975	0.641
	(0.443)	(0.501)	(1.524)	(1.914)	(0.726)
Output gap	-0.205**	-0.358***	-0.126	-0.869*	-0.466**
	(0.085)	(0.128)	(0.098)	(0.481)	(0.226)
Observations	207	207	155	111	207
Countries	24	24	21	18	24
Instruments	27	27	27	27	28
Hansen test (OID)	0.971	0.597	0.996	1.000	0.618
AR(1)	0.587	0.937	0.077		0.753
AR(2)	0.934	0.977	0.009	0.678	0.286

Notes: Standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1; all models control for unobserved country and period effects. H_0 for AR(1) and AR(2) is no autocorrelation. Openness, output gap, wage distortion, and net inflow treated as endogenous in GMM regressions. Maximum number of lags used is 1. Instrument matrix was collapsed as proposed by Roodman (2009a). Constant estimated but not reported. Total stock (nationality) (ln) is multiplied by 10 for numerical stability.

So far we used migration flows, implying that we identify our parameters by exploiting the variation in the change of migration flows over time in the difference GMM and system GMM specifications. As a robustness check we also investigate how the stock of migrants affects the unemployment rate, exploiting the variation in the change of migrant stocks, that is, migration flows. In columns (3) and (4) of Table 5.5 we therefore replace the migration flows by the stock of foreign citizens and the stock of foreign-born persons, respectively. It turns out that migrant stocks have a positive effect on the unemployment rate, but not significantly so. However, in these regressions the coefficient on the lagged unemployment rate is larger than one and highly significant, implying an explosive behavior of the unemployment rate. In addition the test statistic on autocorrelation in the second difference of residuals implies rejection of no autocorrelation in column (3), hinting at a violation of one of the system GMM assumptions. This may well be due to the limited availability of stock data which reduces our sample considerably. Overall, the regressions with migration flows seem to fit our dynamic specification better as they do not imply a counter factual explosive behavior for the unemployment rate.

Our employed dynamic GMM estimator does account for the endogeneity of migration flows by relying on internal instruments based on suitable lags of the respective variables. However, it is also possible to additionally include external instruments such as our predicted migration flows. Specification (5) in Table 5.5 shows the estimates from the specification given in column (1) augmented by the additional exogenous variable. The results change as now net inflows are no longer significant and have a positive impact on the unemployment rate.

Trade openness measures and additional controls

All regressions until now employed a real openness measure as proposed by Alcalá and Ciccone (2004). It is defined as the sum of imports and exports in exchange rate US-$ over GDP in purchasing power parity US-$. Traditionally, openness measures are constructed by dividing by GDP in current US-$. In order to provide comparable results, we therefore use the latter openness measure in column (1) in Table 5.6. Interestingly, we now can corroborate the findings of Felbermayr et al. (2011b) that openness reduces the unemployment rate. Note though that these authors argue against using these openness measures and use total trade openness instead as we do in our benchmark regressions. Still, immigration remains to have a reducing effect on unemployment. Both variables are significant at the 5% level.

As services are very hard to measure and therefore not very well comparable across countries, see e.g. Francois and Hoekman (2010), using total trade flows including services may render openness a noisy measure for actual trade openness. Therefore we re-run our preferred specification using an openness measure based on merchandise trade only. In column (2) we present the Alcalá and Ciccone (2004) real openness measure using only merchandise trade. While trade openness again turns out to have a negative influence on unemployment, it is no longer significant. Immigration has a negative impact on unemployment but is not significant. Immigration becomes negatively significant again in column (3), where we use the standard trade openness based on merchandise trade measured in current US-$ GDP. Here, openness remains negative and not significant.

In columns (4) to (6) in Table 5.6 we introduce additional control variables following Dutt et al. (2009). As openness measure, we return to the total trade openness measure from our preferred specification. We add an index of civil liberties and an additional measure of trade liberalization. Specifically, we add the years since permanent trade liberalization of the country as a control. To allow for a non-linear impact of trade liberalization on unemployment we include the variable both in levels and squared. The inclusion of the civil liberty index renders net immigration non-significant. So does the inclusion of the liberalization variable. Both variables are not significant, though. Openness again turns to have a positive impact but is again not significant. If we include both variables simultaneously, the effect of immigration becomes positive again and we estimate an autoregressive parameter which again implies an explosive behavior of the unemployment rate.

In unreported regressions, we use a different output gap measure. The output gap can also be calculated as the difference between log GDP and log trend GDP. We calculate GDP by multiplying real GDP per capita (chain) by the population from the Penn World Table, edition 7.0. The trend series is calculated by Hodrick-Prescott filtering. We use 6.25 as the smoothing factor for annual data as recommended by Ravn and Uhlig (2002).[29]

Furthermore (again not reported), we re-run our preferred specification from column (8) in Table 5.4 only for EU receiving countries. The coefficient for net inflows (ln) is 1.425 with a standard error of 1.328. Splitting the sample in the years before and after the eastern EU enlargement of 2004 leads to a net inflow immigration coefficient of 0.160 (standard error 0.228) and -0.361 (standard error 0.933) for the pre- and post-accession period, respectively. Finally, we augment our preferred specification by an interaction term between net inflows

[29] Felbermayr et al. (2011b) use it for some regressions as well. However, they use 400 as smoothing factor.

Table 5.6: Robustness checks: Different control variables

Dependent variable: unemployment rate						
	(1) Sys-GMM	(2) Sys-GMM	(3) Sys-GMM	(4) Sys-GMM	(5) Sys-GMM	(6) Sys-GMM
Lag dep. var.	0.918***	0.954***	0.939***	0.946***	0.959***	1.094**
	(0.105)	(0.081)	(0.093)	(0.159)	(0.145)	(0.547)
Net inflow (ln)	-0.388*	-0.628	-0.429*	-1.312	-0.722	0.537
	(0.217)	(0.384)	(0.259)	(0.805)	(0.681)	(7.046)
Total curr. price open.	-0.013*					
	(0.007)					
Merchandise open.		-0.012				
		(0.028)				
Merch. curr. price open.			-0.013			
			(0.019)			
Total trade openness				0.004	0.009	-0.008
				(0.018)	(0.019)	(0.025)
Wage distortion (index)	0.022	0.016	0.015	0.026	-0.020	-0.013
	(0.034)	(0.028)	(0.024)	(0.068)	(0.077)	(0.239)
EPL (index)	0.249	1.306	0.784	0.105	0.900	3.787
	(0.878)	(1.243)	(0.858)	(1.610)	(1.620)	(15.165)
Union density (index)	-0.007	-0.010	-0.005	-0.027	0.015	0.017
	(0.015)	(0.027)	(0.027)	(0.033)	(0.056)	(0.189)
High corporatism (index)	-0.330	-0.192	-0.295	0.270	0.194	-1.415
	(0.226)	(0.391)	(0.476)	(1.118)	(0.707)	(5.686)
PMR (index)	0.248	-0.267	0.096	-0.248	0.017	-0.130
	(0.384)	(0.771)	(0.694)	(0.508)	(0.618)	(0.779)
Population (ln)	0.305	0.547	0.424	1.475**	1.275*	-0.651
	(0.365)	(0.504)	(0.636)	(0.633)	(0.758)	(6.850)
Output gap	-0.275***	-0.276**	-0.242	-0.222	-0.235***	-0.992
	(0.095)	(0.119)	(0.169)	(0.188)	(0.088)	(2.951)
Civil liberties				-1.420		-0.251
				(1.188)		(2.803)
Yrs. since lib.					-0.090	0.127
					(0.172)	(0.800)
(Yrs. since lib.)^2					0.001	-0.001
					(0.002)	(0.007)
Observations	207	207	207	207	207	207
Countries	24	24	24	24	24	24
Instruments	27	27	27	28	29	30
Hansen test (OID)	0.919	0.930	0.944	0.929	0.986	0.970
AR(1)	0.456	0.818	0.576	0.716	0.773	0.814
AR(2)	0.182	0.246	0.154	0.553	0.853	0.818

Notes: Standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1; all models control for unobserved country and period effects. H_0 for AR(1) and AR(2) is no autocorrelation. Openness, output gap, wage distortion, and net inflow treated as endogenous in GMM regressions. Maximum number of lags used is 1. Instrument matrix was collapsed as proposed by Roodman (2009a). Constant estimated but not reported.

(ln) and total trade openness. The value for the interaction term is 0.003 (standard error 0.022), while the net inflow coefficient is -1.281 (standard error 2.304). Hence, we do not find a significant interaction between trade openness and immigration.

To again summarize our results, we find no statistically significant impact of immigration on the unemployment rate across a range of specifications and using different definitions of the control variables.

5.5 Conclusion

How do international trade and immigration affect unemployment in the destination country? While there is ample evidence that trade openness reduces unemployment, to the best of our knowledge the literature has so far not investigated the effect of immigration on unemployment explicitly taking into account a country's exposure to trade. This is astonishing as it is well known since at least Mundell (1957) that goods trade implies implicit factor movements. Hence, when one is interested in the effect of trade on unemployment it seems important to control for additional movement of workers.

In this paper we present the first evidence of the effects of trade and migrant inflows on unemployment in the destination country taking into account that immigration and trade exposure of a country are highly correlated and therefore not statistically independent. In our sample, we find no significant aggregate effect of immigrant inflows on unemployment rates in destination countries on average.

This finding seems to be at odds with the widely held belief of a detrimental effect of immigration on unemployment amongst politicians and the public at large. More importantly, our findings leave us puzzled about how easy European decision makers willingly accepted to erect barriers to the freedom of movement: One of the corner stones of the European Common Market Policy is that workers be employed on equal, non-discriminatory terms in all member states of the European Union. Even though restrictions to this right could only be sustained for a seven year transitional period if the country informed the European Commission about serious disruptions on its labor market, two countries (Austria and Germany) actually achieved shielding their labor markets from inflows for the full seven year period. Given our results, the feared detrimental effect of immigration on domestic labor markets seems dubious at best, at least on average. In the worst case it may have hindered welfare gains for the respective countries due to more efficient allocation of labor across countries. Taking our results even a step further, on average it may have even forced additional workers in Austria and Germany into unemployment, contrary to the well-meant original intention.

Chapter 6

Income and Democracy: Evidence from System GMM Estimates*

6.1 Introduction

Higher levels of income cause the establishment of democratic regimes. This cornerstone of "modernization theory" (see Lipset, 1959) is increasingly accepted by economists and political scientists alike. Reviewing the existing literature reveals that the empirical evidence overwhelmingly supports modernization theory.[1] However, a recent paper by Acemoglu et al. (2008) argues that the empirically observed correlation is spurious. They show that the relationship between democracy and income breaks down when controlling for country and time-fixed effects using a postwar period (1960–2000) sample of countries. Instead, both democracy and higher income are caused by underlying changes in institutional arrangements and are contingent on specific historic events. This alternative view is dubbed the "critical junctures hypothesis" (for a short review see Acemoglu et al., 2009).

Empirical evidence supporting modernization theory relies on SUR regressions, fixed effects and non-linear panel specifications whereas Acemoglu et al. (2008) employ the dynamic

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[1] For example, Barro (1999) uses a SUR regression framework, Gundlach and Paldam (2009) use repeated cross-sectional analysis, Corvalan (2010) uses a panel probit estimator, Boix (2011) and Treisman (2011) use a fixed effects panel estimator, Benhabib et al. (2011) use non-linear panel estimators and Moral-Benito and Bartolucci (2012) use the Arellano and Bond (1991) estimator as well as a limited information maximum likelihood approach (LIML).

panel estimator by Arellano and Bond (1991). All these studies do not take into account the high persistence of income and democracy.

We therefore follow Arellano and Bover (1995) as well as Blundell and Bond (1998) and present empirical evidence using system GMM which performs well with highly persistent data under mild assumptions. We show that even in the smaller postwar period sample with up to 150 countries used by Acemoglu et al. (2008), we find a statistically significant positive relation between income and democracy.[2]

6.2 Econometric methods and data

Acemoglu et al. (2008) estimate the following dynamic panel model:

$$d_{it} = \alpha d_{it-1} + \gamma y_{it-1} + \mathbf{x}'_{it-1}\beta + \delta_i + \mu_t + u_{it}, \tag{6.1}$$

where d_{it} is the democracy level of country i, y_{it-1} is the lagged log GDP per capita, $\mathbf{x}_{it-1}$ is a vector of lagged control variables, δ_i and μ_t denote sets of country dummies and time effects and u_{it} is an error term with $E(u_{it}) = 0$ for all i and t.

Acemoglu et al. (2008) use the difference GMM estimator as proposed by Arellano and Bond (1991) to estimate Equation (6.1).[3] However, this estimator suffers from potentially huge small sample bias when the number of time periods is small and the dependent variable is highly persistent (see Alonso-Borrego and Arellano, 1999). The literature tries to mitigate this persistence by using five year intervals or averages. This reduces the number of observations considerably, while income and democracy are still substantially persistent. We follow Arellano and Bover (1995) and Blundell and Bond (1998) and present system GMM estimates which circumvent the finite sample bias if one accepts a mild stationarity assumption.[4]

The asymptotic efficiency gains of the additional orthogonality conditions of the system GMM estimator do not come without a cost: The number of instruments increases expo-

[2] In a similar fashion, Bobba and Coviello (2007) show that the estimated effect of education on democracy changes its sign when using system GMM.

[3] For a good textbook treatment of (dynamic) panel estimators see Baltagi (2008).

[4] Specifically, the deviations from the long-run mean of the dependent variable have to be uncorrelated with the stationary individual-specific long-run mean itself (see Blundell and Bond, 1998). As there are no a priori reasons to believe that the speed of change in a country's political system is related to its current level of democracy this stationarity condition does not seem unduly restrictive.

nentially with the number of time periods which leads to finite sample bias and increases the likelihood of false positive results as well as suspiciously high pass rates of specification tests like the Hansen (1982) J-test (see Roodman, 2009a). We follow Roodman (2009a) and also present results with a collapsed instrument matrix and use only two lags for both the difference and system GMM estimators.[5] We use Windmeijer (2005) finite sample corrected standard errors.

We employ an unbalanced panel with five-year interval data from 1960 to 2000 taken from Acemoglu et al. (2008). We use two different measures for democracy: the Freedom House index and the composite Polity IV index. The Freedom House index is normalized between zero and one, with one corresponding to the most democratic institutions. It uses data from the non-governmental organization Freedom House and is augmented by data taken from Bollen (2001) for the years 1950, 1955, 1960 and 1965. It is constructed from a checklist of questions concerning both political and civil rights, such as free and fair elections and the prevalence of the rule of law.[6] The main advantage of this index is its broad coverage of countries. For reasons of comparison, we follow Acemoglu et al. (2008) and use the Freedom House index as our main measure of democracy.

The Freedom House index is not without problems. One issue is that it includes too many components, such as socio-economic rights, freedom from war and freedom from gross socioeconomic inequalities, thus leading to a maximalist definition of democracy potentially harming its discriminatory power. Another problem is that the exact coding rules for the indicator are not made publicly available. We therefore contrast our results with an alternative minimalist measure of democracy, the Polity IV index from the Polity IV research project.[7] The composite Polity IV index is also normalized between zero and one, with one corresponding again to the most democratic institutions. It combines the scores of democracy and autocracy indices to a single regime indicator including information on competitiveness of political participation and constraints on the chief executive.[8]

[5] All GMM estimations are carried out using the `xtabond2` package in Stata (see Roodman, 2009b).
[6] For more information see `http://freedomhouse.org/report/freedom-world/freedom-world-2012.`
[7] For more information see `http://www.systemicpeace.org/polity/polity4.htm.`
[8] For a discussion of existing democracy indices and measurement problems of democracy see Munck and Verkuilen (2002).

6.3 Results

Table 6.1 reports the baseline results of estimation of Equation (6.1) using the Freedom House index as dependent variable. Column (1) and (2) show the results of the pooled OLS and fixed effects (within) OLS estimator. Both regressions use robust standard errors clustered by country. These estimates provide the lower and upper bound for the autoregressive coefficient (for details see Bond, 2002). The lower bound is equal to 0.379 whereas the upper bound is 0.706. Both are positive and highly statistically significant. Concerning lagged log GDP per capita we find a positive and significant effect for pooled OLS and no systematic influence using fixed effects.

Columns (3) to (5) employ difference GMM estimators. In column (3) the results from the one-step difference GMM estimator are reported, whereas in columns (4) and (5) we report the results from the two-step difference GMM estimator. All GMM regressions use robust standard errors and treat the lagged democracy measure as predetermined. In the two-step GMM estimates, the Windmeijer (2005) finite sample correction for standard errors is employed. In column (5) also log GDP per capita is treated as endogenous. Note that column (3) reproduces column (2) in Table 2 of Acemoglu et al. (2008). While in all difference GMM estimates the autoregressive coefficient lies within the bounds given by columns (1) and (2), the sign of the coefficient for lagged log GDP per capita becomes negative and weakly significant. However, as motivated in the introduction and when discussing our identification strategy, the one- and two-step differenced GMM estimators do not take into account the high persistence of income and democracy.

We therefore present system GMM estimates in columns (6) to (8). Whereas column (6) reproduces column (5) using the system GMM estimator, column (7) follows the advice given in Roodman (2009a) and collapses the instrument matrix and only uses two lags as instruments. Column (8) includes lagged log population, lagged education and lagged age structure as additional controls. All specifications estimate an autoregressive coefficient that lies between the two bounds given in columns (1) and (2). However, lagged log GDP per capita has now a positive and significant effect on democracy. The point estimate of lagged log GDP in the specification given in column (6) is 0.118, implying that a one percent increase of lagged GDP increases the steady-state value of democracy by 0.26 percentage points.[9]

[9] The long-run effect is calculated as $\gamma/(1-\alpha)$.

The row for the Hansen J-test reports the p-values for the null hypothesis of the validity of the overidentifying restrictions. In all specifications we do not reject the null hypothesis. The values reported for the Diff-in-Hansen test are the p-values for the validity of the additional moment restrictions necessary for system GMM. Again, we do not reject the null that the additional moment conditions are valid. The values reported for AR(1) and AR(2) are the p-values for first and second order autocorrelated disturbances in the first-differenced equation. As expected, there is high first order autocorrelation, and no evidence for significant second order autocorrelation. To sum up, our test statistics hint at a proper specification.

In Table 6.2 we check the robustness of our results against using the second democracy measure and including additional external instruments as used by Acemoglu et al. (2008). The first three columns reestimate specifications (6) to (8) from Table 6.1 using the Polity IV index. GDP per capita still turns out to be positive and significant in columns (1) and (2) albeit a bit smaller in magnitude. In specification (3), GDP per capita is still positive but no longer significant. This is similar as in specification (8) of Table 6.1, where significance was also lower than in specifications (6) and (7). This may well be due to the lower number of observations. The specification tests indicate well-specified models. Hence, the choice of the democracy measure does not influence our qualitative result.

Columns (4) and (5) in Table 6.2 use the trade-weighted world income of the respective country as an additional external instrument. We report the system GMM estimates as in columns (6) and (7) in Table 6.1. Again, as in Table 6.1 the coefficient of GDP per capita changes its sign going from the difference GMM (not reported) to the system GMM estimates. With system GMM, it turns out to be positive and significant again. Again, all the specification tests indicate a well-specified model. In columns (6) and (7) we use the second lag of the savings rate of the countries as an additional external instrument instead. Here, we again find a change in the sign from negative to positive on the GDP per capita variable when moving from difference (not reported) to system GMM estimates. The model specification tests also indicate a well-specified model across the different specifications. Only the Diff-in-Hansen test for the system GMM estimates using the collapsed instrument matrix in column (7) rejects the null of the validity of the additional overidentifying restrictions. However, the autocorrelation tests indicate that the model is well specified. This could well be due to the use of the collapsed instruments as the asymptotic behavior of this ad hoc method is not well understood (see Roodman, 2009a). As the Hansen tests are known to have weak power and all results are in line with our previous ones, we still believe that we have properly identified the influence of GDP on democracy.

6.4 Conclusion

When studying the relationship between income and democracy, one has to account for the dynamic nature and the high persistence of the data. Employing system GMM, we find a significant positive relation between income and democracy for a postwar period sample of up to 150 countries. Our results are robust to different measures of democracy and instrument sets.

Table 6.1: Baseline results

	Pooled OLS (1)	FE OLS (2)	Diff-1 GMM (AJRY) (3)	Diff-2 GMM (4)	Diff-2 GMM END (5)	Sys-2 GMM END (6)	Sys-2 GMM END CL (7)	Sys-2 GMM END CL (8)
	Dependent variable is $Democracy_t$ (Freedom House index)							
$Democracy_{t-1}$	0.706***	0.379***	0.489***	0.528***	0.432***	0.548***	0.568***	0.546***
	(0.035)	(0.051)	(0.085)	(0.105)	(0.085)	(0.053)	(0.063)	(0.076)
$Log\ GDP\ per\ capita_{t-1}$	0.072***	0.010	-0.129*	-0.012	-0.097*	0.118***	0.136***	0.110*
	(0.010)	(0.035)	(0.076)	(0.065)	(0.053)	(0.020)	(0.023)	(0.060)
Controls	No	No	No	No	No	No	No	Yes
Instruments			55	55	90	108	16	21
Hansen J-test			[0.260]	[0.260]	[0.273]	[0.131]	[0.778]	[0.614]
Diff-in-Hansen test						[0.298]	[0.791]	[0.268]
AR(1)			[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
AR(2)			[0.448]	[0.421]	[0.540]	[0.332]	[0.297]	[0.875]
Observations	945	945	838	838	838	945	945	676
Countries	150	150	127	127	127	150	150	95

Notes: Base sample – taken from Acemoglu et al. (2008) – is an unbalanced panel spanning from 1960–2000 with data at five-year intervals, where the start date of the panel refers to the dependent variable. The dependent variable is the Augmented Freedom House Political Rights index. Standard errors are in parentheses, p-values in brackets. Pooled and FE OLS regressions use robust standard errors clustered by country. All GMM regressions use robust standard errors and treat the lagged democracy measure as predetermined. In addition to that, regressions with suffix "END" treat lagged log GDP per capita as endogenous and regressions with suffix "CL" follow Roodman (2009a) and collapse the instrument matrix and use only two lags. In the case of two-step GMM, the Windmeijer (2005) finite sample correction for standard errors is employed. In the last column, lagged log population, lagged education (average years of total schooling) and lagged age structure are added as controls. Age structure is specified as median age of the population at $t-1$ and four covariates corresponding to the percent of the population at $t-1$ in the following age groups: 0–15, 15–30, 30–45, and 45–60. *, ** and *** denote significance at the 10%-, 5%- and 1%-level, respectively. The row for the Hansen J-test reports the p-values for the null hypothesis of instrument validity. The values reported for the Diff-in-Hansen test are the p-values for the validity of the additional moment restriction necessary for system GMM. The values reported for AR(1) and AR(2) are the p-values for first and second order autocorrelated disturbances in the first differences equations.

Table 6.2: Robustness checks

	Sys-2 GMM END (1)	Sys-2 GMM END CL (2)	Sys-2 GMM END CL (3)	Sys-2 GMM END (4)	Sys-2 GMM END CL (5)	Sys-2 GMM END (6)	Sys-2 GMM END CL (7)
	Dependent variable is $Democracy_t$						
	Polity IV index			Freedom House index			
				World income instrument		Savings rate instrument	
$Democracy_{t-1}$	0.616***	0.694***	0.735***	0.547***	0.578***	0.584***	0.575***
	(0.071)	(0.080)	(0.082)	(0.053)	(0.066)	(0.054)	(0.072)
$Log\ GDP\ per\ capita_{t-1}$	0.079***	0.067***	0.104	0.110***	0.128***	0.110***	0.114***
	(0.019)	(0.023)	(0.067)	(0.023)	(0.024)	(0.018)	(0.023)
Controls	No	No	Yes	No	No	No	No
Instruments	108	16	21	109	17	107	16
Hansen J-test	[0.139]	[0.379]	[0.207]	[0.158]	[0.597]	[0.213]	[0.058]
Diff-in-Hansen test	[0.859]	[0.167]	[0.068]	[0.185]	[0.331]	[0.630]	[0.037]
AR(1)	[0.000]	[0.000]	[0.001]	[0.000]	[0.000]	[0.000]	[0.000]
AR(2)	[0.361]	[0.329]	[0.322]	[0.367]	[0.320]	[0.441]	[0.436]
Observations	854	854	640	895	895	891	891
Countries	136	136	92	124	124	134	134

Notes: Base sample – taken from Acemoglu et al. (2008) – is an unbalanced panel spanning from 1960–2000 with data at five-year intervals, where the start date of the panel refers to the dependent variable. The dependent variable in columns (1)–(3) is the composite Polity IV index, the dependent variable in columns (4)–(7) is the Augmented Freedom House Political Rights index. Standard errors are in parentheses, p-values in brackets. All GMM regressions use robust standard errors and treat the lagged democracy measure as predetermined and either the second lag of the savings rate or trade-weighted world income as additional external instrument. In addition to that, regressions with suffix "END" treat lagged log GDP per capita as endogenous and regressions with suffix "CL" follow Roodman (2009a) and collapse the instrument matrix and use only two lags. In the case of two-step GMM, the Windmeijer (2005) finite sample correction for standard errors is employed. *, ** and *** denote significance at the 10%-, 5%- and 1%-level, respectively. The row for the Hansen J-test reports the p-values for the null hypothesis of instrument validity. The values reported for the Diff-in-Hansen test are the p-values for the validity of the additional moment restriction necessary for system GMM. The values reported for AR(1) and AR(2) are the p-values for first and second order autocorrelated disturbances in the first differences equations.

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Appendix A

Appendix to Chapter 1

A.1 Empirical probability of exporting

Table A.1: Empirical probability of exporting—firm level sample

Rank	Country	Probability	Rank	Country	Probability
1	Japan	0.75695	16	New Zealand	0.01853
2	South Korea	0.25367	17	Republic of South Africa	0.01718
3	Singapore	0.07008	18	Switzerland	0.01602
4	Australia	0.05367	19	Sri Lanka	0.01467
5	Vietnam	0.04691	20	Chile	0.01293
6	Thailand	0.04305	21	Panama	0.01236
7	Malaysia	0.03552	22	Egypt	0.01120
8	United Arab Emirates	0.03185	23	Cambodia	0.01062
9	Indonesia	0.03127	24	Mexico	0.00965
10	Philippines	0.02529	25	Pakistan	0.00907
11	Saudi Arabia	0.02201	26	Israel	0.00888
12	Russia	0.02162	27	Kuwait	0.00753
13	Bangladesh	0.02143	28	Brazil	0.00714
14	Myanmar	0.02124	29	Norway	0.00676
15	India	0.01873	30	Ukraine	0.00579

Turkey, Guatemala, Morocco, Madagascar, Jordan, Kenya, Algeria, Honduras, Venezuela, Romania, Ghana, El Salvador, Sudan, Mongolia, Togo, Peru, Nigeria, Mozambique, Lebanon, Nepal, Djibouti, Yemen, Tanzania, Benin, Nicaragua, Jamaica, Croatia, Zimbabwe, Congo (Republic of), Sierra Leone, Argentina, Iran, Syria, Mauritius, Mauritania, Papua New Guinea, Colombia, Kazakstan, Bermuda, Bahrain, Tunisia, Iceland, Angola, Fiji, Senegal, Mali, Uganda, Liberia, Ecuador, Serbia, Oman, Costa Rica, Azerbaijan, Guinea Bissau, Guinea, Gabon, Afghanistan, Gambia, Trinadad and Tabago, Ethiopia, Iraq, Laos, Congo (Democratic Republic), Swaziland, Cameroon, Côte d'Ivoire, Cuba, Paraguay, Lesotho, Dominican Republic, Brunei, Puerto Rico, Niger, Rwanda, Bulgaria, Samoa, Guyana, Suriname, Uruguay, Central African Republic, Botswana, Barbados, Bolivia, Zambia, Tajikistan, Comoros Islands, Libya, Micronesia (Federated States of), Antigua and Barbuda, Malawi, Albania, Eritrea, Chad, New Caledonia, Macedonia, Maldive Islands, Belize, Kiribati and Tuvalu, Moldova, São Tomé and Principe, Grenada, Haiti, Palau, Bahamas, Vanuatu and New Hebrides, Burundi, Solomon Islands, Bhutan, Tonga, Burkina, Turkmenistan, Cape Verde Islands, Namibia, Marshall Islands, Georgia, Uzbekistan, Bosnia Herzegovina, Seychelles, Dominica, Armenia.

Notes: Table gives the observed frequencies of exporting firms in the firm-level regression sample for the top 30 export destinations outside the MFA countries in descending order. The rest of the 150 export destinations considered in our sample are given, again in descending order. A detailed description of our sample is provided in Section 1.2.

A.2 Construction of explanatory variables

We construct different contiguity indicators $\mathbb{I}(N_{ij,t-1} > 0)_{ijt}$ using common border, common language, common colonizer, common income group, and common continent contiguity indicators from data provided by CEPII, see Mayer and Zignago (2011). For the different contiguity measures $\mathbb{I}(N_{ij,t-1} > 0)_{ijt} = 1$ is defined as follows:

Common border: $\mathbb{I}(N_{ij,t-1} > 0)_{ijt} = 1$ for firm i if country j shares a land border with at least one export destination of firm i in $t-1$ and 0 otherwise.

Common language: $\mathbb{I}(N_{ij,t-1} > 0)_{ijt} = 1$ for firm i if country j shares a language with at least one export destination of firm i in $t-1$ and 0 otherwise which is spoken by at least 9 percent of the population in both countries.

Common colonizer: $\mathbb{I}(N_{ij,t-1} > 0)_{ijt} = 1$ for firm i if country j shares a common colonizer after 1945 with at least one export destination of firm i in $t-1$ and 0 otherwise.

Common income group: $\mathbb{I}(N_{ij,t-1} > 0)_{ijt} = 1$ for firm i if country j is in the same income group with at least one export destination of firm i in $t-1$ and 0 otherwise. The four different categories (very low income, low income, medium income, and high income) follow the World Bank's 2006 World Development Indicators (WDI) classification.

Common continent: $\mathbb{I}(N_{ij,t-1} > 0)_{ijt} = 1$ for firm i if country j is located on the same continent as at least one export destination of firm i in $t-1$ and 0 otherwise.

C_j is defined accordingly.

A.3 Descriptive statistics

Table A.2: Descriptive statistics—firm level sample

Variable	Mean	Std. Dev.	Min.	Max.
y_{ijt}	0.01189	0.10837	0	1
C_j defined according to...				
common border	0.08000	0.27129	0	1
common language	0.64000	0.48000	0	1
common colonizer	0.36667	0.48189	0	1
common income group	0.27333	0.44567	0	1
common continent	0.34000	0.47371	0	1
N_j defined according to...				
common border	0.15333	0.64017	0	5
common language	2.48667	2.30575	0	7
common colonizer	0.78667	1.05569	0	3
common income group	3.38667	6.45785	0	20
common continent	3.74667	8.37392	0	25
$\mathbb{I}(N_{ij,t-1} > 0)$ defined according to...				
common border	0.01176	0.10781	0	1
common language	0.15567	0.36254	0	1
common colonizer	0.06478	0.24614	0	1
common income group	0.20637	0.40470	0	1
common continent	0.26975	0.44383	0	1
$N_{ij,t-1}$ defined according to...				
common border	0.01344	0.13153	0	5
common language	0.23624	0.72363	0	22
common colonizer	0.10450	0.52529	0	20
common income group	0.32240	0.84828	0	17
common continent	0.41640	0.93765	0	20
# of firms				1,295
# of observations				770,000

Notes: Table gives descriptive statistics of the dependent and the explanatory variables used in our empirical analysis at the firm level. A detailed description of our sample is provided in Section 1.2.

Table A.3: Descriptive statistics—firm-product couple level sample

Variable	Mean	Std. Dev.	Min.	Max.
y_{ijt}	0.00987	0.09885	0	1
$\mathbb{I}(N_{ij,t-1}^{sameproduct} > 0)$ defined according to...				
common border	0.00763	0.08699	0	1
common language	0.11944	0.32431	0	1
common colonizer	0.04459	0.20640	0	1
common incomegroup	0.17092	0.37644	0	1
common continent	0.26266	0.44008	0	1
$\mathbb{I}(N_{ij,t-1}^{otherproducts} > 0)$ defined according to...				
common border	0.01420	0.11831	0	1
common language	0.07914	0.26995	0	1
common colonizer	0.03716	0.18916	0	1
common incomegroup	0.06440	0.24547	0	1
common continent	0.05955	0.23666	0	1
$N_{ij,t-1}^{sameproduct}$ defined according to...				
common border	0.00839	0.10025	0	4
common language	0.17192	0.57271	0	19
common colonizer	0.06833	0.38434	0	15
common incomegroup	0.24435	0.68145	0	17
common continent	0.35050	0.74127	0	19
$N_{ij,t-1}^{otherproducts}$ defined according to...				
common border	0.05293	0.73395	0	45
common language	1.02048	4.73813	0	98
common colonizer	0.49207	4.37562	0	131
common incomegroup	1.91062	8.16292	0	135
common continent	2.73516	9.17680	0	154
# of firms				6,573
# of firms-product-couples				1,965
# of observations				3,943,800

Notes: Table gives descriptive statistics of the dependent and the explanatory variables used in our empirical analysis at the firm-product couple level. A detailed description of our sample is provided in Section 1.4.

Appendix B

Appendix to Chapter 2

B.1 Introduction to the Appendix

In this Appendix, we present further results and robustness checks.

In Section B.2, we present an alternative model setup in the vein of the Ricardian model of international trade by Eaton and Kortum (2002) and show that our results from the main text hold when reinterpreting the elasticity of substitution as the technology dispersion parameter used in Eaton and Kortum (2002).

Section B.3 presents results for the counterfactual analyses in Section 3 from the main text under the assumption of balanced trade.

In Section B.4, we present a variant of our model where wages are determined by a binding minimum wage instead of bargaining once the match between a worker and firm is established. We derive counterfactual changes in employment and show that for constant labor market institutions, calculated employment changes are identical to the ones assuming wage bargaining as in the main text.

In Section B.5, we assume that the wage setting process is determined within an efficiency wage framwork. Again, when labor market institutions remain unchanged, calculated changes in employment and GDP are identical to the model presented in the main text.

In Section B.6, we derive the solution of the system of asymmetric multilateral resistance equations.

In Section B.7, we derive sufficient statistics for welfare with imperfect labor markets and show that in the case of imperfect labor markets, the welfare statistics presented in Arkolakis et al. (2012) are augmented by the net employment change.

Finally, Section B.8 presents further results on trade flow and employment changes for the evaluation of PTAs and labor market reforms in the United States as well as detailed results for labor market reforms in Germany as presented in Section 3 from the main text.

B.2 A Ricardian trade model with imperfect labor markets following Eaton and Kortum (2002)

In the following, we introduce search and matching frictions in the Ricardian model of international trade by Eaton and Kortum (2002) and show that this leads to expressions for counterfactual changes in GDP, employment, trade flows, and welfare which are isomorphic to those in the main text. Note that in the following we assume balanced trade.

The representative consumer in country j is again characterized by the utility function U_j. As in Eaton and Kortum (2002), we assume a continuum of goods $k \in [0,1]$. Consumption of individual goods is denoted by $q(k)$, leading to the following utility function

$$U_j = \left[\int_0^1 q(k)^{\frac{\sigma-1}{\sigma}} dk\right]^{\frac{\sigma}{\sigma-1}}, \tag{B.1}$$

where σ is the elasticity of substitution in consumption. Again, international trade of goods from i to j imposes iceberg trade costs $t_{ij} > 1$.

Countries differ in the efficiency with which they can produce goods. We denote country i's efficiency in producing good $k \in [0,1]$ as $\mathfrak{z}_i(k)$. Denoting input costs in country i as $\mathfrak{c}_i$, the cost of producing a unit of good k in country i is then $\mathfrak{c}_i/\mathfrak{z}_i(k)$.

Taking trade barriers into account, delivering a unit of good k produced in country i to country j costs

$$p_{ij}(k) = \left(\frac{\mathfrak{c}_i}{\mathfrak{z}_i(k)}\right) t_{ij}. \tag{B.2}$$

Assuming perfect competition, $p_{ij}(k)$ is the price which consumers in country j would pay if they bought good k from country i. With international trade, consumers can choose from which country to buy a good. Hence, the price they actually pay for good k is $p_j(k)$, the lowest price across all sources i:

$$\underline{p}_j(k) = \min\left\{p_{ij}(k); i = 1, \cdots, n\right\}, \tag{B.3}$$

where n denotes the number of countries.

Let country i's efficiency in producing good k be the realization of an independently drawn Fréchet random variable with distribution $F_i(\mathfrak{z}) = e^{-T_i \mathfrak{z}^{-\theta}}$, where T_i is the location parameter (also called "state of technology" by Eaton and Kortum, 2002) and θ governs the variation within the distribution and thereby also the comparative advantage within the continuum of goods.

Plugging Equation (B.2) in $F_i(\mathfrak{z})$ leads to $G_{ij}(p) = Pr[P_{ij} \leq p] = 1 - e^{-[T_i(\mathfrak{c}_i t_{ij})^{-\theta}]p^\theta}$. Noting that the distribution of prices for which a country j buys is given by $G_j(p) = Pr[P_j \leq p] = 1 - \prod_{i=1}^n [1 - G_{ij}(p)]$ leads to:

$$G_j(p) = 1 - e^{-\Phi_j p^\theta}, \tag{B.4}$$

where $\Phi_j = \sum_{i=1}^n T_i \left(\mathfrak{c}_i t_{ij}\right)^{-\theta}$.

The probability that country i provides good k at the lowest price to country j is given by (see Eaton and Kortum, 2002, page 1748):

$$\pi_{ij} = \frac{T_i \left(\mathfrak{c}_i t_{ij}\right)^{-\theta}}{\Phi_j}. \tag{B.5}$$

With a continuum of goods between zero and one this is also the fraction of goods that country j buys from country i. Eaton and Kortum (2002) show that the price of a good that country j actually buys from any country i is also distributed $G_j(p)$, and that the exact price index is given by $P_j = \tilde{\Gamma}\Phi_j^{-1/\theta}$ with $\tilde{\Gamma} = \left[\Gamma\left(\frac{\theta+1-\sigma}{\theta}\right)\right]^{\frac{1}{1-\sigma}}$ where Γ is the Gamma function.

The fraction of goods that country j buys from country i, π_{ij}, is also the fraction of its expenditures on goods from country i, x_{ij}, due to the fact that the average expenditures per good do not vary by source. Hence,

$$x_{ij} = \frac{T_i(\mathfrak{c}_i t_{ij})^{-\theta}}{\Phi_j} y_j = \frac{T_i(\mathfrak{c}_i t_{ij})^{-\theta}}{\sum_{k=1}^n T_k(\mathfrak{c}_k t_{kj})^{-\theta}} y_j, \tag{B.6}$$

where y_j is country j's total spending.

Assuming balanced trade, exporters' total sales (including home sales) are equal to total expenditure and are simply given by:

$$y_i = \sum_{j=1}^n x_{ij} = T_i \mathfrak{c}_i^{-\theta} \sum_{j=1}^n \frac{t_{ij}^{-\theta}}{\Phi_j} y_j. \tag{B.7}$$

Solving for $T_i\mathfrak{c}_i^{-\theta}$ leads to:

$$T_i \mathfrak{c}_i^{-\theta} = \frac{y_i}{\sum_{j=1}^n \frac{t_{ij}^{-\theta}}{\Phi_j} y_j}. \tag{B.8}$$

Replacing $T_i\mathfrak{c}_i^{-\theta}$ in Equation (B.6) with this expression leads to:

$$x_{ij} = \frac{t_{ij}^{-\theta}}{\Phi_j \left(\sum_{j=1}^n \frac{t_{ij}^{-\theta}}{\Phi_j} y_j\right)} y_i y_j.$$

Using $P_j = \tilde{\Gamma}\Phi_j^{-\frac{1}{\theta}}$ to replace Φ_j in both terms of the denominator leads to:

$$x_{ij} = \frac{t_{ij}^{-\theta}}{\tilde{\Gamma}^{\theta} P_j^{-\theta} \left(\sum_{j=1}^n \frac{t_{ij}^{-\theta}}{\tilde{\Gamma}^{\theta} P_j^{-\theta}} y_j\right)} y_i y_j.$$

Define

$$\Pi_i = \left(\sum_{j=1}^n \left(\frac{t_{ij}}{P_j}\right)^{-\theta} \theta_j\right)^{-\frac{1}{\theta}},$$

and note that we can express P_j also as follows:

$$\begin{aligned} P_j &= \left(\tilde{\Gamma}^{-\theta}\Phi_j\right)^{-\frac{1}{\theta}} = \left(\tilde{\Gamma}^{-\theta} \sum_{i=1}^n T_i(\mathfrak{c}_i t_{ij})^{-\theta}\right)^{-\frac{1}{\theta}} = \left(\tilde{\Gamma}^{-\theta} \sum_{i=1}^n \frac{t_{ij}^{-\theta} y_i}{\sum_{l=1}^n \frac{t_{il}^{-\theta}}{\Phi_l} y_l}\right)^{-\frac{1}{\theta}}, \\ &= \left(\sum_{i=1}^n \left(\frac{t_{ij}}{\Pi_i}\right)^{-\theta} \theta_i\right)^{-\frac{1}{\theta}}, \end{aligned}$$

where $\theta_j = y_j/y^W$ with $y^W = \sum_j y_j$. Then we can write:

$$x_{ij} = \frac{y_i y_j}{y^W} \left(\frac{t_{ij}}{\Pi_i P_j}\right)^{-\theta}.$$

Replacing $-\theta$ by $1-\sigma$ we end up with exactly the same system as in the model by Anderson and van Wincoop (2003).

Hence, our approach can be applied to both worlds with the only difference that the interpretation differs and the roles of θ and σ have to be exchanged.

B.2.1 Counterfactual GDP in the Eaton and Kortum (2002) framework with perfect labor markets

We assume that there are no intermediates and one unit of the final good is produced with one unit of labor, hence $\mathfrak{c}_i = w_i$. Equation (B.8) can be written as

$$T_i w_i^{-\theta} = \frac{y_i}{\sum_{j=1}^n \frac{t_{ij}^{-\theta}}{\Phi_j} y_j} = \frac{\theta_i}{\sum_{j=1}^n \tilde{\Gamma}^{-\theta} \left(\frac{t_{ij}}{P_j}\right)^{-\theta} \theta_j} = \tilde{\Gamma}^{\theta} \theta_i \Pi_i^{\theta}.$$

Solving for w_i leads to:

$$w_i = \tilde{\Gamma}^{-1} T_i^{\frac{1}{\theta}} \theta_i^{-\frac{1}{\theta}} \Pi_i^{-1}.$$

As $y_i = w_i L_i$, the change in GDP is given by $y_i^c / y_i = w_i^c / w_i$. Hence,

$$\frac{y_i^c}{y_i} = \frac{\tilde{\Gamma} T_i^{\frac{1}{\theta}} \left(\theta_i^c\right)^{-\frac{1}{\theta}} \left(\Pi_i^c\right)^{-1}}{\tilde{\Gamma} T_i^{\frac{1}{\theta}} \theta_i^{-\frac{1}{\theta}} \Pi_i^{-1}} = \frac{\left(\theta_i^c\right)^{-\frac{1}{\theta}} \left(\Pi_i^c\right)^{-1}}{\theta_i^{-\frac{1}{\theta}} \Pi_i^{-1}} = \left(\frac{\mathfrak{k}_i^c}{\mathfrak{k}_i}\right)^{-\frac{1}{\theta}},$$

where $\mathfrak{k}_i = \theta_i \Pi_i^{\theta}$.

B.2.2 Counterfactuals in the Eaton and Kortum (2002) framework with imperfect labor markets

We assume that there are no intermediates and $\mathfrak{z}_i$ units of the final good k are produced using one unit of labor. For simplicity, we omit the product index k in the following. Denoting the net price earned by the producer by $p_i = p_{ij}/t_{ij}$, the total surplus of a successful match is given by $\mathfrak{z}_i p_i - b_i$, while the firm's rent is given by $\mathfrak{z}_i p_i - w_i$ and the worker's by $w_i - b_i$. Nash bargaining leads to $w_i - b_i = \xi_i/(1-\xi_i)(\mathfrak{z}_i p_i - w_i)$. Using $b_i = \gamma_i w_i$ and combining leads to

$$w_i = \frac{\xi_i}{1-\gamma_i+\xi_i\gamma_i} \mathfrak{z}_i p_i = \frac{\xi_i}{1-\gamma_i+\xi_i\gamma_i} \mathfrak{c}_i. \tag{B.9}$$

Firms create vacancies until all rents are dissipated. The free entry (zero profit) condition is given by $M_i/V_i(\mathfrak{z}_i p_i - w_i) = P_i c_i$. Rewriting leads to the job creation curve

$$w_i = \mathfrak{z}_i p_i - \frac{P_i c_i}{m_i \vartheta_i^{-\mu}} = \mathfrak{c}_i - \frac{P_i c_i}{m_i \vartheta_i^{-\mu}}. \tag{B.10}$$

We can combine Equations (B.9) and (B.10) to write the wage paid by a firm as

$$w_i = \frac{\xi_i}{1-\gamma_i+\gamma_i\xi_i-\xi_i} \frac{P_i c_i}{m_i \vartheta^{-\mu}}. \tag{B.11}$$

The wage paid by a firm producing variety k is solely determined by parameters and aggregate variables and does neither depend on its variety-specific price nor on productivity. Hence, as wages are equalized across firms, Equation (B.10) then implies that also $\mathfrak{c}_i$ is the same across firms, irrespective of the variety they produce. Hence the job creation and wage curve are the same for all firms and we can thus determine aggregate labor market tightness ϑ_i as the locus of intersection of both curves:

$$\vartheta_i = \left(\frac{\mathfrak{c}_i}{P_i}\right)^{1/\mu} \left(\frac{c_i}{m_i}\Omega_i\right)^{-1/\mu}. \tag{B.12}$$

Equation (B.8) can be written as

$$T_i \mathfrak{c}_i^{-\theta} = \frac{y_i}{\sum_{j=1}^{n} \frac{t_{ij}^{-\theta}}{\Phi_j} y_j} = \frac{\theta_i}{\sum_{j=1}^{n} \tilde{\Gamma}^{-\theta} \left(\frac{t_{ij}}{P_j}\right)^{-\theta} \theta_j} = \tilde{\Gamma}^{\theta} \theta_i \Pi_i^{\theta}.$$

Solving for $\mathfrak{c}_i$ leads to:

$$\mathfrak{c}_i = \tilde{\Gamma}^{-1} T_i^{\frac{1}{\theta}} \theta_i^{-\frac{1}{\theta}} \Pi_i^{-1}. \tag{B.13}$$

As $y_i = \mathfrak{c}_i(1-u_i)L_i$, assuming a constant labor force the change in GDP is given by $y_i^c/y_i = (1-u_i^c)\mathfrak{c}_i^c/[(1-u_i)\mathfrak{c}_i]$ leading to

$$\begin{aligned} \frac{y_i^c}{y_i} &= \frac{(1-u_i^c)\tilde{\Gamma} T_i^{\frac{1}{\theta}} (\theta_i^c)^{-\frac{1}{\theta}} (\Pi_i^c)^{-1}}{(1-u_i)\tilde{\Gamma} T_i^{\frac{1}{\theta}} \theta_i^{-\frac{1}{\theta}} \Pi_i^{-1}} \\ &= \frac{(1-u_i^c)\,(\theta_i^c)^{-\frac{1}{\theta}} (\Pi_i^c)^{-1}}{(1-u_i)\theta_i^{-\frac{1}{\theta}} \Pi_i^{-1}} \\ &= \frac{(1-u_i^c)}{(1-u_i)} \left(\frac{\mathfrak{k}_i^c}{\mathfrak{k}_i}\right)^{-\frac{1}{\theta}}, \end{aligned} \tag{B.14}$$

where $\mathfrak{k}_i = \theta_i \Pi_i^{\theta}$.

For the change in employment (the first fraction on the right-hand side of Equation (B.14)) the same relationship holds as is given in the main text in Equation (2.14) when we remember once more that $-\theta = 1 - \sigma$. Hence, we end up with

$$\frac{y_i^c}{y_i} = \hat{\kappa}_i \left(\frac{\mathfrak{k}_i^c}{\mathfrak{k}_i}\right)^{-\frac{1}{\mu\theta}} \left(\frac{\sum_i t_{ij}^{-\theta}\mathfrak{k}_i}{\sum_i \left(t_{ij}^c\right)^{-\theta} \mathfrak{k}_i^c}\right)^{-\frac{1-\mu}{\mu\theta}}, \tag{B.15}$$

which is the same relationship as given in Implication 3 in the main text when we remember that we assumed balanced trade and again replace $1-\sigma$ by $-\theta$.

Besides counterfactual employment, also counterfactual trade flows and welfare can be calculated as in the main text.

B.3 Results with balanced trade

The following Tables present the results for the same counterfactual experiments as presented in Section 3.2 in the main text but we assume balanced trade throughout, i.e. $\tilde{y}_j = y_j$ and $\tilde{\theta}_j = \theta_j$. Results basically remain the same, both qualitatively and quantitatively. Note that imposing balanced trade also affects the estimates for σ and μ, whereas the estimated trade cost coefficients do not change by construction (see Table 2.2).

Table B.1: Estimation results for OECD sample, 1950-2006, assuming balanced trade

	(1) OLS $\ln z_{ij\tau}$	(2) PPML $z_{ij\tau}$	(3) OLS $\ln x_{ij\tau}$	(4) PPML $x_{ij\tau}$	(5) OLS $\ln z_{ij\tau}$	(6) PPML $z_{ij\tau}$	(7) OLS $\ln x_{ij\tau}$	(8) PPML $x_{ij\tau}$
Second stage								
$\ln DIST_{ij}$	-1.050***	-0.669***	-1.041***	-0.816***	-1.050***	-0.669***	-1.040***	-0.813***
	(0.009)	(0.027)	(0.010)	(0.010)	(0.009)	(0.027)	(0.010)	(0.010)
$CONTIG_{ij}$	0.097***	0.276***	0.116***	0.414***	0.097***	0.275***	0.115***	0.414***
	(0.019)	(0.030)	(0.019)	(0.018)	(0.019)	(0.030)	(0.019)	(0.018)
$COMLANG_{ij}$	0.386***	0.769***	0.387***	0.150***	0.386***	0.769***	0.387***	0.151***
	(0.019)	(0.049)	(0.019)	(0.017)	(0.019)	(0.049)	(0.019)	(0.017)
First stage								
$PTA_{ij\tau}$	0.274***	0.308***	0.267***	0.332***	0.274***	0.311***	0.276***	0.341***
	(0.016)	(0.019)	(0.017)	(0.019)	(0.014)	(0.016)	(0.015)	(0.013)
Estimated elasticities								
σ	2.361***	2.506***	2.371***	2.383***	2.362***	2.507***	2.373***	2.383***
	(0.174)	(0.045)	(0.606)	(0.560)	(0.019)	(0.048)	(0.491)	(0.449)
μ	0.947***	0.928***	0.946***	0.939***	0.947***	0.928***	0.945***	0.939***
	(0.003)	(0.007)	(0.009)	(0.006)	(0.001)	(0.007)	(0.006)	(0.005)
zero trade		X		X		X		X
symmetric $t_{ij\tau}$	X	X	X	X				
asymmetric $t_{ij\tau}$					X	X	X	X
N	36,945	37,741	37,493	38,313	36,945	37,741	37,493	38,313

Notes: Results for trade flows between 28 OECD countries between 1950 and 2006 estimated by ordinary least squares (OLS) and Poisson pseudo-maximum-likelihood (PPML). z_{ij} are trade flows standardized by importer and exporter GDPs. $\ln DIST$ is distance between exporting and importing country, $CONTIG$ is an indicator variable equal to 1 if the exporting and importing countries i and j share a common border, $COMLANG$ is an indicator variable equal to 1 if the exporting and importing country share a common official language, and PTA is an indicator variable equal to 1 if the exporting and importing country have signed a preferential trade agreement. All regressions control for multilateral resistance terms (MRTs) via exporter and importer fixed effects. (Robust) standard errors in parentheses, *** p <0.01. Standard errors for σ and μ are bootstrapped using 200 replications.

B.3.1 Introducing PTAs as observed in 2006

Table B.2 presents the results from incepting PTAs as observed in 2006 starting from a counterfactual situation without any PTAs assuming balanced trade. Tables B.3 and B.4 present the changes in trade flows for both perfect and imperfect labor markets, similar to Tables B.10 and B.11.

B.3.2 Different parameter values for elasticities

Table B.5 presents the robustness checks for different parameter values for the elasticity of substitution and the matching elasticity assuming balanced trade.

B.3.3 Evaluating the effects of a labor market reform in the U.S.

Tables B.6 and B.7 present the results from the counterfactual labor market reform in the U.S. assuming balanced trade.

B.3.4 Evaluating the effects of counterfactually undoing the recent German labor market reforms

Tables B.8 and B.9 present the results of counterfactually undoing the recent labor market reforms in Germany assuming balanced trade.

B.4 Minimum wages within the search and matching framework

In this Section, we introduce minimum wages in our search and matching framework. The binding minimum wage replaces the bargaining of workers and firms that are matched. We then show that this leads to expressions for counterfactual changes in GDP, employment, trade flows, and welfare which are isomorphic to those in the main text.

We assume balanced trade for the following derivations. Let us first consider the bounds for a binding minimum wage. If the minimum wage is above the wage that a firm and a worker agree upon, it is not binding and hence not relevant. The lower bound for a binding minimum wage, denoted by $\underline{w}_j$, is therefore given by the **wage curve** from the main text

$$\underline{w}_j = w_j = \frac{\xi_j}{1 + \gamma_j \xi_j - \gamma_j} p_j. \tag{B.16}$$

The upper bound for a minimum wage, denoted by $\overline{w}_j$, is given by the job's output, as firms would not be able to recover recruiting costs. Hence, $\overline{w}_j = p_j$.

A well defined equilibrium with a binding minimum wage $\tilde{w}_j$ exists if $\underline{w}_j < \tilde{w}_j < \overline{w}_j$. With a given binding minimum wage, the wage curve is no longer relevant. ϑ_j can be solved

Table B.2: Comparative static effects of PTA inception assuming balanced trade in 2006

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	PLM	SMF	share %GDP SMF		SMF	SMF	PLM	SMF
	%GDP	%GDP	$\% \ln(\hat{p})$	$\% \ln(\hat{e})$	$\%\hat{e}$	Δu	%EV	%EV
Australia	16.69	17.64	92.78	7.22	1.18	-1.11	16.69	17.64
Austria	18.31	19.61	91.78	8.22	1.48	-1.39	21.05	22.60
Belgium	18.79	20.17	91.53	8.47	1.57	-1.42	22.37	24.04
Canada	21.05	22.53	90.63	9.37	1.92	-1.77	28.68	30.16
Czech Republic	17.82	19.06	92.04	7.96	1.40	-1.28	19.74	21.19
Denmark	17.25	18.40	92.37	7.63	1.30	-1.23	18.19	19.54
Finland	16.44	17.48	92.87	7.13	1.16	-1.05	16.04	17.24
France	16.26	17.30	92.97	7.03	1.13	-1.02	15.56	16.79
Germany	15.65	16.61	93.39	6.61	1.02	-0.91	13.94	15.10
Greece	16.22	17.23	93.01	6.99	1.12	-1.01	15.45	16.62
Hungary	17.30	18.46	92.33	7.67	1.31	-1.19	18.34	19.70
Iceland	15.88	16.80	93.27	6.73	1.05	-1.01	14.54	15.56
Ireland	16.63	17.65	92.77	7.23	1.18	-1.12	16.52	17.67
Italy	15.69	16.64	93.37	6.63	1.03	-0.95	14.05	15.17
Japan	9.59	9.62	101.02	-1.02	-0.09	0.09	-1.27	-1.29
Korea	9.74	9.79	100.70	-0.70	-0.07	0.06	-0.92	-0.91
Netherlands	17.20	18.36	92.39	7.61	1.29	-1.22	18.06	19.44
New Zealand	10.79	11.02	98.71	1.29	0.13	-0.13	1.63	1.88
Norway	16.80	17.88	92.64	7.36	1.22	-1.16	16.99	18.23
Poland	17.14	18.28	92.43	7.57	1.28	-1.09	17.90	19.23
Portugal	16.55	17.59	92.80	7.20	1.17	-1.07	16.33	17.52
Slovak Republic	17.57	18.77	92.18	7.82	1.35	-1.16	19.06	20.47
Spain	15.73	16.67	93.35	6.65	1.03	-0.93	14.15	15.24
Sweden	16.70	17.77	92.70	7.30	1.20	-1.10	16.71	17.96
Switzerland	19.06	20.48	91.40	8.60	1.61	-1.53	23.11	24.81
Turkey	16.11	17.09	93.10	6.90	1.10	-0.97	15.15	16.27
United Kingdom	14.21	14.93	94.57	5.43	0.76	-0.71	10.20	11.02
United States	10.26	10.43	99.60	0.40	0.04	-0.04	0.35	0.54
Average	13.14	13.70	96.61	3.39	0.56	-0.51	7.68	8.32

Notes: Counterfactual analysis based on parameter estimates from column (6) of Table 2.2. PLM gives results assuming perfect labor markets. SMF gives results using a search and matching framework for the labor market. Averages are weighted averages using country GDP as weight.

Table B.3: Heterogeneity of comparative static trade effects of PTA inception with perfect labor markets assuming balanced trade in 2006

	Changes in exports in percent by importer quantiles						
Exporting country	Min.	0.025	0.25	0.5	0.75	0.975	Max.
Australia	-29.91	-29.39	-24.14	-22.93	-21.50	20.63	20.88
Austria	-32.29	-31.15	-3.53	0.33	2.32	6.34	6.92
Belgium	-32.98	-31.85	-4.26	-0.68	1.28	5.27	5.84
Canada	-33.36	-33.29	-30.74	-29.71	-28.40	5.31	10.25
Czech Republic	-31.59	-30.44	-2.53	1.37	3.37	7.44	8.03
Denmark	-30.75	-29.59	-1.33	2.61	4.64	8.76	9.35
Finland	-29.55	-28.36	0.39	4.27	6.47	10.66	11.26
France	-29.27	-28.08	0.78	4.68	6.78	11.10	11.70
Germany	-28.33	-27.12	2.13	6.08	8.13	12.56	13.19
Greece	-29.21	-28.01	0.87	4.77	6.80	11.19	11.80
Hungary	-30.83	-29.67	-1.45	2.50	4.52	8.64	9.23
Iceland	-28.68	-27.47	2.32	5.69	7.78	22.41	24.49
Ireland	-29.82	-28.64	-0.00	3.87	6.05	10.23	10.83
Italy	-28.39	-27.18	2.04	5.99	8.04	12.48	13.09
Japan	-17.98	-17.37	-11.23	-9.95	-8.27	4.67	4.89
Korea	-18.25	-17.56	-11.42	-9.90	0.20	24.38	24.49
Netherlands	-30.68	-29.51	-1.23	2.72	4.75	8.88	9.47
New Zealand	-20.18	-19.59	-13.61	-12.24	-10.61	16.46	19.44
Norway	-30.08	-28.90	0.31	3.61	5.87	20.00	22.03
Poland	-30.59	-29.42	-1.11	2.72	4.88	9.01	9.61
Portugal	-29.71	-28.52	0.16	4.03	6.22	10.40	11.01
Slovak Republic	-31.23	-30.07	-2.01	1.91	3.92	8.02	8.61
Spain	-28.45	-27.24	1.95	5.89	7.95	12.38	12.99
Sweden	-29.93	-28.75	-0.16	3.71	5.89	10.06	10.66
Switzerland	-33.36	-32.24	-3.42	-0.52	0.91	14.38	16.31
Turkey	-29.03	-27.84	1.81	5.16	7.24	21.80	23.87
United Kingdom	-26.03	-24.79	5.39	9.47	11.59	13.17	13.19
United States	-15.79	-15.71	-12.48	-10.97	-9.45	19.02	20.88
Average	-28.44	-27.42	-3.80	-0.55	1.69	12.35	13.37

Notes: Counterfactual analysis based on parameter estimates from column (6) of Table 2.2. Table depicts changes in normalized exports, i.e. exports divided by source and origin GDPs.

Table B.4: Heterogeneity of comparative static effects of PTA inception with imperfect labor markets and assuming balanced trade in 2006

	Changes in exports in percent by importer quantiles						
Exporting country	Min.	0.025	0.25	0.5	0.75	0.975	Max.
Australia	-29.68	-29.21	-24.15	-22.91	-21.51	20.83	21.08
Austria	-32.25	-31.12	-3.84	0.06	2.04	6.10	6.70
Belgium	-32.96	-31.84	-4.59	-0.98	0.98	5.00	5.59
Canada	-33.33	-33.27	-30.68	-29.62	-28.34	5.61	10.54
Czech Republic	-31.54	-30.40	-2.84	1.11	3.11	7.22	7.82
Denmark	-30.69	-29.54	-1.63	2.36	4.39	8.55	9.16
Finland	-29.47	-28.29	0.11	4.01	6.24	10.47	11.09
France	-29.22	-28.04	0.46	4.37	6.52	10.85	11.47
Germany	-28.29	-27.09	1.78	5.74	7.81	12.30	12.94
Greece	-29.13	-27.95	0.59	4.51	6.55	11.00	11.62
Hungary	-30.77	-29.62	-1.74	2.25	4.27	8.42	9.03
Iceland	-28.55	-27.36	2.11	5.53	7.62	22.53	24.65
Ireland	-29.70	-28.53	-0.22	3.66	5.89	10.10	10.72
Italy	-28.33	-27.13	1.73	5.69	7.75	12.25	12.88
Japan	-17.65	-17.10	-11.17	-9.80	-8.18	5.02	5.26
Korea	-17.94	-17.31	-11.39	-9.84	0.40	24.53	24.65
Netherlands	-30.64	-29.49	-1.56	2.44	4.47	8.63	9.24
New Zealand	-19.97	-19.43	-13.67	-12.26	-10.66	16.68	19.65
Norway	-30.00	-28.83	0.03	3.38	5.68	20.04	22.12
Poland	-30.53	-29.37	-1.40	2.44	4.64	8.80	9.41
Portugal	-29.62	-28.45	-0.11	3.78	6.01	10.23	10.85
Slovak Republic	-31.17	-30.03	-2.31	1.65	3.67	7.79	8.40
Spain	-28.37	-27.18	1.67	5.62	7.69	12.18	12.82
Sweden	-29.86	-28.69	-0.44	3.43	5.65	9.86	10.48
Switzerland	-33.33	-32.22	-3.73	-0.79	0.65	14.33	16.30
Turkey	-28.94	-27.75	1.55	4.95	7.03	21.86	23.97
United Kingdom	-25.92	-24.68	5.15	9.24	11.38	12.93	12.94
United States	-15.89	-15.81	-12.55	-11.02	-9.54	19.23	21.08
Average	-28.35	-27.35	-4.03	-0.75	1.51	12.26	13.30

Notes: Counterfactual analysis based on parameter estimates from column (6) of Table 2.2. Table depicts changes in normalized exports, i.e. exports divided by source and origin GDPs.

Table B.5: Average comparative static effects of PTA inception assuming balanced trade for various parameter values

μ	σ	PLM %GDP	SMF %GDP	SMF $\%\hat{e}$	SMF $\%\Delta u$	PLM %EV	SMF %EV
	5	4.86	16.76	11.90	-9.23	2.75	15.23
0.2	10	2.15	7.13	5.00	-4.22	1.20	6.33
	15	1.38	4.53	3.16	-2.74	0.77	3.98
	5	4.86	7.60	2.75	-2.41	2.75	5.66
0.5	10	2.15	3.35	1.20	-1.08	1.20	2.44
	15	1.38	2.15	0.77	-0.70	0.77	1.55
	5	4.86	5.75	0.90	-0.81	2.75	3.71
0.75	10	2.15	2.55	0.40	-0.36	1.20	1.61
	15	1.38	1.64	0.25	-0.23	0.77	1.03
	5	4.86	5.15	0.30	-0.27	2.75	3.07
0.9	10	2.15	2.28	0.13	-0.12	1.20	1.34
	15	1.38	1.47	0.08	-0.08	0.77	0.85
	5	4.86	4.89	0.03	-0.03	2.75	2.78
0.99	10	2.15	2.16	0.01	-0.01	1.20	1.21
	15	1.38	1.39	0.01	-0.01	0.77	0.78

Notes: Table reports average changes in nominal GDP, employment, and the equivalent variation in percent assuming either a perfect labor market (PLM) or using a search and matching framework (SMF) for the labor market with varying elasticity of substitution σ and elasticity of the matching function μ.

Table B.6: Comparative static effects of $\hat{\kappa}_{U.S.} = 1.03$ assuming balanced trade in 2006

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	PLM	SMF	share %GDP SMF		SMF	SMF	PLM	SMF
	%GDP	%GDP	$\% \ln(\hat{p})$	$\% \ln(\hat{e})$	$\%\hat{e}$	Δu	%EV	%EV
Australia	0.00	0.79	92.78	7.22	0.06	-0.05	0.00	0.79
Austria	0.00	0.51	98.69	1.31	0.01	-0.01	0.00	0.09
Belgium	0.00	0.49	99.36	0.64	0.00	-0.00	0.00	0.04
Canada	0.00	0.96	90.80	9.20	0.09	-0.08	0.00	1.23
Czech Republic	0.00	0.52	98.12	1.88	0.01	-0.01	0.00	0.14
Denmark	0.00	0.53	97.88	2.12	0.01	-0.01	0.00	0.16
Finland	0.00	0.56	97.15	2.85	0.02	-0.01	0.00	0.22
France	0.00	0.52	98.20	1.80	0.01	-0.01	0.00	0.13
Germany	0.00	0.52	98.25	1.75	0.01	-0.01	0.00	0.13
Greece	0.00	0.55	97.32	2.68	0.01	-0.01	0.00	0.20
Hungary	0.00	0.54	97.71	2.29	0.01	-0.01	0.00	0.17
Iceland	0.00	0.62	95.60	4.40	0.03	-0.03	0.00	0.38
Ireland	0.00	0.59	96.29	3.71	0.02	-0.02	0.00	0.30
Italy	0.00	0.53	97.79	2.21	0.01	-0.01	0.00	0.16
Japan	0.00	0.54	97.50	2.50	0.01	-0.01	0.00	0.19
Korea	0.00	0.55	97.32	2.68	0.01	-0.01	0.00	0.20
Netherlands	0.00	0.51	98.45	1.55	0.01	-0.01	0.00	0.11
New Zealand	0.00	0.73	93.60	6.40	0.05	-0.04	0.00	0.65
Norway	0.00	0.56	97.17	2.83	0.02	-0.02	0.00	0.22
Poland	0.00	0.54	97.76	2.24	0.01	-0.01	0.00	0.17
Portugal	0.00	0.57	96.87	3.13	0.02	-0.02	0.00	0.25
Slovak Republic	0.00	0.53	97.81	2.19	0.01	-0.01	0.00	0.16
Spain	0.00	0.55	97.22	2.78	0.02	-0.01	0.00	0.21
Sweden	0.00	0.55	97.43	2.57	0.01	-0.01	0.00	0.19
Switzerland	0.00	0.48	99.41	0.59	0.00	-0.00	0.00	0.04
Turkey	0.00	0.56	96.98	3.02	0.02	-0.02	0.00	0.24
United Kingdom	0.00	0.62	95.73	4.27	0.03	-0.02	0.00	0.36
United States	0.00	2.54	-16.66	116.66	2.97	-2.83	0.00	2.54
Average	0.00	1.30	55.06	44.94	1.11	-1.06	0.00	1.10

Notes: Counterfactual analysis based on parameter estimates from column (6) of Table 2.2. PLM gives results assuming perfect labor markets. SMF gives results using a search and matching framework for the labor market. Averages are weighted averages using country GDP as weight.

Table B.7: Heterogeneity of comparative static trade effects of $\hat{\kappa}_{U.S.} = 1.03$ with imperfect labor markets and assuming balanced trade in 2006

	Changes in exports in percent by importer quantiles						
Exporting country	Min.	0.025	0.25	0.5	0.75	0.975	Max.
Australia	-0.98	-0.91	-0.10	-0.05	-0.01	0.08	0.08
Austria	-0.36	-0.29	0.48	0.56	0.60	0.71	0.71
Belgium	-0.31	-0.24	0.53	0.60	0.65	0.74	0.75
Canada	-0.98	-0.98	-0.49	-0.45	-0.40	-0.31	-0.31
Czech Republic	-0.40	-0.33	0.44	0.52	0.56	0.67	0.67
Denmark	-0.41	-0.35	0.43	0.50	0.56	0.65	0.65
Finland	-0.47	-0.40	0.37	0.45	0.50	0.59	0.59
France	-0.39	-0.32	0.45	0.53	0.57	0.67	0.67
Germany	-0.39	-0.32	0.45	0.53	0.57	0.68	0.68
Greece	-0.46	-0.39	0.38	0.47	0.52	0.61	0.61
Hungary	-0.43	-0.36	0.41	0.49	0.55	0.64	0.64
Iceland	-0.61	-0.54	0.27	0.31	0.36	0.45	0.45
Ireland	-0.55	-0.48	0.33	0.38	0.43	0.52	0.52
Italy	-0.42	-0.35	0.42	0.50	0.55	0.64	0.64
Japan	-0.44	-0.37	0.40	0.47	0.53	0.62	0.62
Korea	-0.46	-0.39	0.38	0.47	0.51	0.61	0.61
Netherlands	-0.37	-0.30	0.47	0.54	0.59	0.69	0.69
New Zealand	-0.85	-0.78	0.03	0.07	0.12	0.21	0.21
Norway	-0.47	-0.40	0.37	0.45	0.50	0.59	0.59
Poland	-0.42	-0.35	0.42	0.49	0.55	0.64	0.64
Portugal	-0.50	-0.43	0.35	0.43	0.48	0.57	0.57
Slovak Republic	-0.42	-0.35	0.42	0.50	0.55	0.64	0.65
Spain	-0.47	-0.40	0.37	0.46	0.51	0.60	0.60
Sweden	-0.45	-0.38	0.39	0.47	0.52	0.62	0.62
Switzerland	-0.31	-0.24	0.53	0.61	0.65	0.75	0.75
Turkey	-0.49	-0.42	0.35	0.44	0.49	0.58	0.58
United Kingdom	-0.60	-0.53	0.28	0.32	0.37	0.46	0.46
United States	-0.98	-0.91	-0.10	-0.06	-0.01	0.08	0.08
Average	-0.51	-0.45	0.32	0.39	0.44	0.54	0.54

Notes: Counterfactual analysis based on parameter estimates from column (6) of Table 2.2. Table depicts changes in normalized exports, i.e. exports divided by source and origin GDPs.

by using the **job creation curve** given in the main text

$$\tilde{w}_j = p_j - \frac{P_j c_j}{m_j \vartheta_j^{-\mu}} \Rightarrow$$
$$\vartheta_j = \left(\frac{p_j - \tilde{w}_j}{P_j}\right)^{1/\mu} \left(\frac{c_j}{m_j}\right)^{-1/\mu}, \tag{B.17}$$

which corresponds to Equation (2.9) in the main text. By replacing u_j by Equation (2.8) from the main text and using Equation (B.17), GDP in country j can be written as:

$$y_j = p_j(1-u_j)L_j = p_j m_j \left(\frac{p_j - \tilde{w}_j}{P_j}\right)^{\frac{1-\mu}{\mu}} \left(\frac{c_j}{m_j}\right)^{\frac{\mu-1}{\mu}} L_j. \tag{B.18}$$

Assuming that the nominal minimum wage is indexed to prices, we can express it as a share of prices, i.e. $\tilde{w}_j = \xi_j p_j$. This allows us to express GDP solely as a function of prices and parameters. Similarly, (counterfactual) employment can be rewritten using Equation (2.8) in the main text and Equation (B.17). Then, defining $\tilde{\Xi}_j = m_j \left(\frac{c_j}{m_j}\right)^{\frac{\mu-1}{\mu}}$ and $\hat{\tilde{\kappa}}_j = \tilde{\Xi}_j^c / \tilde{\Xi}_j$, we get

$$\frac{1-u_j^c}{1-u_j} = \hat{\tilde{\kappa}}_j \left(\frac{p_j^c - \tilde{w}_j}{p_j - \tilde{w}_j}\right)^{\frac{1-\mu}{\mu}} \left(\frac{P_j}{P_j^c}\right)^{\frac{1-\mu}{\mu}}. \tag{B.19}$$

Using again that $\tilde{w}_j = \xi_j p_j$, the last expression simplifies to

$$\frac{1-u_j^c}{1-u_j} = \hat{\tilde{\kappa}}_j^* \left(\frac{p_j^c}{p_j}\right)^{\frac{1-\mu}{\mu}} \left(\frac{P_j}{P_j^c}\right)^{\frac{1-\mu}{\mu}}, \tag{B.20}$$

where $\hat{\tilde{\kappa}}_j^* = \hat{\tilde{\kappa}}_j((1-\xi_j^c)/(1-\xi_j))^{(1-\mu)/\mu}$. Equation (B.20) exactly corresponds to Equation (2.14) in the main text except for the replacement of $\hat{\kappa}_j$ by $\hat{\tilde{\kappa}}_j^*$. Hence, when assuming that labor market institutions (here: minimum wage levels) do not change, we can proceed as with bargained wages to calculate employment effects.

Note that in the case of binding minimum wages, all GDP changes are due to employment changes. Hence, counterfactual GDP changes correspond to employment changes.

Counterfactual trade flows and welfare can be calculated as in the case of bargained wages.

B.5 Efficiency wages within the search and matching framework

In this Section, we show how efficiency wages in the spirit of Stiglitz and Shapiro (1984) can be introduced into our search and matching framework by replacing the bargaining of workers and firms with the no-shirking condition. Note that we assume balanced trade and risk neutral workers in the following.

We first derive the utility for a shirker, s, and a non-shirker, ns. The non-shirker ns earns wage w_j while exerting effort e_j. Hence, her utility in our one-shot framework is given by

$$E_j^{ns} = w_j - e_j. \tag{B.21}$$

A shirker s also earns wage w_j but does not exert any effort e_j. However, a share α_j of shirkers is detected by firms and gets fired, which leads to unemployment. When the worker is unemployed she earns $\gamma_j w_j$, and hence the expected utility for a shirker can be written as

$$E_j^s = (1 - \alpha_j) w_j + \alpha_j \gamma_j w_j. \tag{B.22}$$

The no-shirking condition $E^{ns} \geq E^s$ leads to $E^{ns} = E^s$ in equilibrium. Hence, using Equations (B.21) and (B.22), the wage can be written as:

$$w_j = \frac{1}{\alpha_j(1 - \gamma_j)} e_j. \tag{B.23}$$

As in the case of bargaining, wages can be solved without knowledge of ϑ_j. ϑ_j can be solved by using the **job creation curve** given in the main text:

$$\begin{aligned} \frac{1}{\alpha_j(1-\gamma_j)} e_j &= p_j - \frac{P_j c_j}{m_j \vartheta_j^{-\mu}} \Rightarrow \\ \vartheta_j^{\mu} &= \left(\frac{m_j}{P_j c_j}\right)\left(p_j - \frac{1}{\alpha_j(1-\gamma_j)} e_j\right). \end{aligned} \tag{B.24}$$

Now assume that effort e_j can be expressed in terms of prices p_j as $e_j = \xi_j p_j$. Then we can simplify Equation (B.24) to:

$$\vartheta_j = \left(\frac{p_j}{P_j}\right)^{1/\mu} \left(\frac{c_j}{m_j} \check{\Omega}_j\right)^{-1/\mu}, \tag{B.25}$$

with $\check{\Omega}_j = \frac{\alpha_j(1-\gamma_j)}{\alpha_j(1-\gamma_j)-\xi_j}$, which corresponds to Equation (2.9).

Counterfactual employment can be calculated using the definition of u_j given in Equation (2.8) in the main text, replacing ϑ_j by the expression given in Equation (B.25) and defining $\check{\Xi}_j = m_j \left(\frac{c_j}{m_j}\check{\Omega}_j\right)^{\frac{\mu-1}{\mu}}$ and $\hat{\check{\kappa}}_j = \check{\Xi}_j^c / \check{\Xi}_j$:

$$\frac{1 - u_i^c}{1 - u_i} = \hat{\check{\kappa}}_j \left(\frac{p_i^c}{p_i}\right)^{\frac{1-\mu}{\mu}} \left(\frac{P_i}{P_i^c}\right)^{\frac{1-\mu}{\mu}}, \tag{B.26}$$

which exactly corresponds to Equation (2.14) in the main text except for the replacement of $\hat{\kappa}_j$ by $\hat{\check{\kappa}}_j$. Hence, when assuming that labor market institutions do not change, we can proceed as with bargained wages to calculate employment effects.

Using the definition of $\breve{\Xi}_j$, GDP can be expressed as:

$$y_j = p_j e_j L_j = p_j m_j \left(\frac{p_j}{P_j}\right)^{\frac{1-\mu}{\mu}} \left(\frac{c_j}{m_j}\breve{\Omega}_j\right)^{\frac{\mu-1}{\mu}} L_j = p_j \left(\frac{p_j}{P_j}\right)^{\frac{1-\mu}{\mu}} \breve{\Xi}_j L_j. \tag{B.27}$$

Now take the ratio of counterfactual GDP, y_j^c, and observed GDP, y_j, and note that the labor force, L_j, stays constant:

$$y_j^c = \hat{\breve{\kappa}}_j \frac{p_j^c \left(\frac{p_j^c}{P_j^c}\right)^{\frac{1-\mu}{\mu}}}{p_j \left(\frac{p_j}{P_j}\right)^{\frac{1-\mu}{\mu}}} = \hat{\breve{\kappa}}_j \left(\frac{p_j^c}{p_j}\right)^{\frac{1}{\mu}} \left(\frac{P_j}{P_j^c}\right)^{\frac{1-\mu}{\mu}} y_j, \tag{B.28}$$

where $\hat{\breve{\kappa}}_j = \breve{\Xi}_j^c/\breve{\Xi}_j$. Then, using Equation (2.13) from the main text and the fact that $\tilde{P}_j^{1-\sigma} = \sum_i (y^W/\tilde{y}^W) t_{ij}^{1-\sigma}\mathfrak{t}_i$, we end up with exactly the same expression as given in the result in Implication 3 in the main text except for the replacement of $\hat{\kappa}_j$ by $\hat{\breve{\kappa}}_j$. Hence, we can calculate counterfactual GDP as in the case of bargained wages. Similarly, counterfactual trade flows and welfare can be calculated as in the case with bargained wages.

B.6 Solution of asymmetric multilateral resistance equations

Using Equation (2.6), we can write $\tilde{\Pi}_i^{1-\sigma} = \sum_{j=1}^n t_{ij}^{1-\sigma}\tilde{P}_j^{\sigma-1}\tilde{\theta}_j$. Defining $\mathfrak{P}_j = \tilde{\theta}_j\tilde{P}_j^{\sigma-1}$ leads to $\tilde{\Pi}_i^{1-\sigma} = \sum_{j=1}^n t_{ij}^{1-\sigma}\mathfrak{P}_j$. Similarly, $\tilde{P}_j$ can be written as $\tilde{P}_j^{1-\sigma} = \sum_{i=1}^n t_{ij}^{1-\sigma}\tilde{\Pi}_i^{\sigma-1}\theta_i$. Defining $\mathfrak{t}_i = \theta_i\tilde{\Pi}_i^{\sigma-1}$ leads to $\tilde{P}_j^{1-\sigma} = \sum_{i=1}^n t_{ij}^{1-\sigma}\mathfrak{t}_i$. Now dividing $\tilde{\Pi}_i^{1-\sigma} = \sum_{j=1}^n t_{ij}^{1-\sigma}\mathfrak{P}_j$ by $\tilde{\Pi}_i^{1-\sigma}$ and using again $\mathfrak{t}_i = \theta_i\tilde{\Pi}_i^{\sigma-1}$ leads to $\theta_i = \mathfrak{t}_i\sum_{j=1}^n t_{ij}^{1-\sigma}\mathfrak{P}_j$ which can be rearranged to $\theta_i = \mathfrak{t}_i\sum_{j=1}^n t_{ij}^{1-\sigma}\mathfrak{P}_j$. Similarly, dividing $\tilde{P}_j^{1-\sigma} = \sum_{i=1}^n t_{ij}^{1-\sigma}\mathfrak{t}_i$ by $\tilde{P}_j^{1-\sigma}$ and using again $\mathfrak{P}_j = \tilde{\theta}_j\tilde{P}_j^{\sigma-1}$ leads to $\tilde{\theta}_j = \mathfrak{P}_j\sum_{i=1}^n t_{ij}^{1-\sigma}\mathfrak{t}_i$ which can be rearranged to $\tilde{\theta}_j = \mathfrak{P}_j\sum_{i=1}^n t_{ij}^{1-\sigma}\mathfrak{t}_i$. $\theta_i = \mathfrak{t}_i\sum_{j=1}^n t_{ij}^{1-\sigma}\mathfrak{P}_j$ and $\tilde{\theta}_j = \mathfrak{P}_j\sum_{i=1}^n t_{ij}^{1-\sigma}\mathfrak{t}_i$ define a system of $2n$ equations that can be solved for the $2n$ unknowns $\mathfrak{t}_i$ and $\mathfrak{P}_j$.

B.7 Sufficient statistics for welfare with imperfect labor markets

Defining real income as $W_j \equiv \breve{y}_j/P_j$ and taking logs, the total differential is given by $d\ln W_j = d\ln\breve{y}_j - d\ln P_j$. As $y_j = p_j(1-u_j)L_j$, we can write analogously $d\ln y_j = d\ln p_j - u_j/(1-u_j)d\ln u_j = -u_j/(1-u_j)d\ln u_j$ assuming that the labor force remains constant. The second expression on the right-hand side uses the wage curve $w_j = \xi_j/(1+\gamma_j\xi_j-\gamma_j)p_j$, implying $d\ln w_j = d\ln p_j$ holding all labor market parameters constant and choice of numéraire w_j. Assuming that $d_j = 0$, i.e. that there are no trade imbalances, it holds that $d\ln\breve{y}_j = d\ln y_j$.

The total differential of $\ln P_j = \ln\left\{\left[\sum_{i=1}^n(\beta_i p_i t_{ij})^{1-\sigma}\right]^{\frac{1}{1-\sigma}}\right\}$ is given by

$$d\ln P_j = \sum_{i=1}^n \left(\left(\frac{\beta_i p_i t_{ij}}{P_j}\right)^{1-\sigma} d\ln p_i + \left(\frac{\beta_i p_i t_{ij}}{P_j}\right)^{1-\sigma} d\ln t_{ij}\right).$$

Using $x_{ij} = ((\beta_i p_i t_{ij})/P_j)^{1-\sigma} y_j$ and defining $\lambda_{ij} = x_{ij}/y_j = ((\beta_i p_i t_{ij})/P_j)^{1-\sigma}$, yields

$$d \ln P_j = \sum_{i=1}^{n} \lambda_{ij} \left(d \ln p_i + d \ln t_{ij} \right). \tag{B.29}$$

Noting again that $d \ln p_i = d \ln w_i$ holds, we can also write: $d \ln P_j = \sum_{i=1}^{n} \lambda_{ij} (d \ln w_i + d \ln t_{ij})$. Combining terms leads to $d \ln W_j = d \ln y_j - d \ln P_j = -\frac{u_j}{1-u_j} d \ln u_j - \sum_{i=1}^{n} \lambda_{ij} (d \ln w_i + d \ln t_{ij})$. Taking the ratio of λ_{ij} and λ_{jj} we can write $\lambda_{ij}/\lambda_{jj} = [(\beta_i p_i t_{ij})/(\beta_j p_j t_{jj})]^{1-\sigma}$. Noting that $dt_{jj} = 0$ by assumption and that w_j is the numéraire, so that $dw_j = dp_j = 0$, the log-change of this ratio is given by $d \ln \lambda_{ij} - d \ln \lambda_{jj} = (1-\sigma)(d \ln t_{ij} + d \ln p_i)$. Combining this with Equation (B.29) leads to:

$$d \ln P_j = \frac{1}{1-\sigma} \left(\sum_{i=1}^{n} \lambda_{ij} d \ln \lambda_{ij} - d \ln \lambda_{jj} \sum_{i=1}^{n} \lambda_{ij} \right).$$

Noting that $y_j = \sum_{i=1}^{n} x_{ij}$, it follows that $\sum_{i=1}^{n} \lambda_{ij} = 1$ and $d \sum_{i=1}^{n} \lambda_{ij} = \sum_{i=1}^{n} d\lambda_{ij} = 0$. Hence, $\sum_{i=1}^{n} \lambda_{ij} d \ln \lambda_{ij} = \sum_{i=1}^{n} d\lambda_{ij} = 0$. Using these facts, the above expression simplifies to $d \ln P_j = -\frac{1}{1-\sigma} d \ln \lambda_{jj}$. The welfare change can than be expressed as $d \ln W_j = -\frac{u_j}{1-u_j} d \ln u_j + \frac{1}{1-\sigma} d \ln \lambda_{jj}$. Integrating between the initial and the counterfactual situation we get $\ln \hat{W}_j = \ln \hat{e}_j + \frac{1}{1-\sigma} \ln \hat{\lambda}_{jj}$, where $e_j = 1 - u_j$ is the share of employed workers. Taking exponents leads to $\hat{W}_j = \hat{e}_j \hat{\lambda}_{jj}^{\frac{1}{1-\sigma}}$. Moving from any observed level of trade to autarky, i.e., $\lambda_{jj}^c = 1$, yields $\hat{W}_j = \hat{e}_j (\lambda_{jj})^{-\frac{1}{1-\sigma}}$. Note, however, that in contrast to the case with perfect labor markets considered in Arkolakis et al. (2012), even this expression needs information about employment changes.

B.8 Further results for counterfactual analyses

B.8.1 Further results for introducing PTAs as observed in 2006

This section reports additional results for the counterfactual analysis presented in Section 3.2.1 in the main text.

Tables B.10 and B.11 report goods trade changes for perfect and imperfect labor markets, respectively. Trade changes are heterogeneous across importers and exporters. To summarize this heterogeneity, we present quantiles of calculated trade flow changes across all destination countries for all exporters. Both tables report the minimum and maximum changes, along with the 0.025, 0.25, 0.5, 0.75, and 0.975 quantiles. Comparing numbers across columns for each row reveals the heterogeneity across importers, while comparing numbers across rows for each column highlights the heterogeneity across exporters.

In general, every country experiences both positive and negative bilateral trade flow changes. For example, the introduction of PTAs as observed in 2006 implies that the change in trade flows for the United Kingdom is larger than 11.94% for 25% of all countries importing goods from the United Kingdom. Turning to the trade flow results of our model with imperfect labor markets (Table B.11), we find a similar pattern for trade flow changes. Again, changes are heterogeneous across importers and exporters and, again, small and remote countries experience larger changes. The implied trade flow changes differ from the case with perfect labor markets but are of similar magnitude.

The employment effects of incepting PTAs from column (5) of Table 3 in the main text are illustrated graphically in Figure B.1.

B.8.2 Further results for a labor market reform in the U.S.

Table B.12 summarizes the trade effects of the hypothetical labor market reform in the U.S. presented in Section 3.2.2 in the main text. A labor market reform in the United States spurs trade changes across the whole sample. The effects of exports by the United States range between -0.98% and 0.08%. Effects across other exporters range from -0.98% for Australia to 0.77% for Belgium and Switzerland. On average, 50% of trade flow changes are larger than 0.41%. The size pattern of the spillover effects of labor market reforms in the United States clearly depend on the distance from and trade volume of the corresponding country and the United States.

The employment effects of the counterfactual U.S. labor market reform from column (5) of Table 5 are graphically illustrated in Figure B.2.

B.8.3 Evaluating the effects of counterfactually undoing the recent German labor market reforms

In the following, we present the results of counterfactually undoing the recent labor market reforms in Germany as alluded to in the last paragraph of Section 3.2.2 in the main text.

Table B.13 presents the main results, and Table B.14 the corresponding trade effects. As can be seen, undoing the German labor market reforms would increase unemployment in Germany by about 4 percentage points, and welfare would be more than 3 percent lower. Most importantly, we see that abolishing German labor market reforms would have negative spillover effects in all trading partners of Germany. Whereas the net effect on unemployment rates in the trading partners is negligible given our parameter estimates, welfare effects are not: Austria's welfare would be about 0.9 percent lower without German labor market reforms. This is also reflected in the trade effects reported in Table B.14. Austria's exports would change between 0.5 and 1.2 percent across its importing partners. Again, trade effects are heterogeneous across countries.

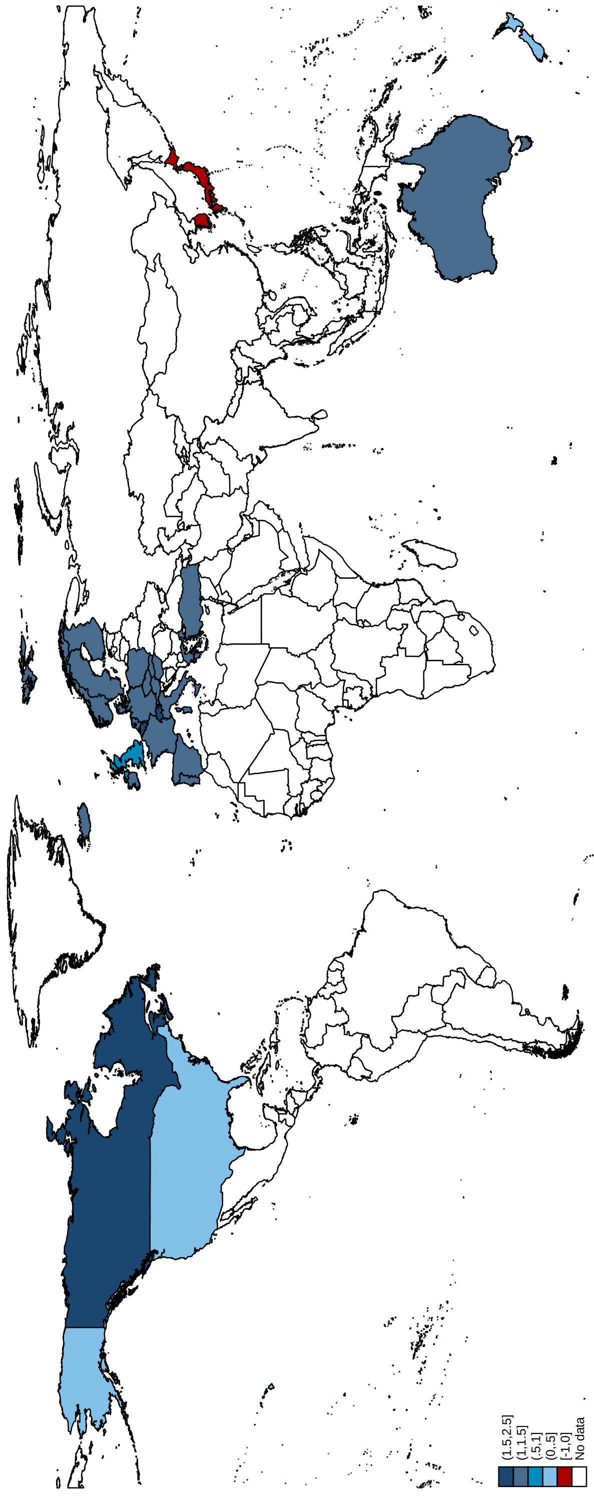

Figure B.1: Employment effects of incepting PTAs as observed in 2006.

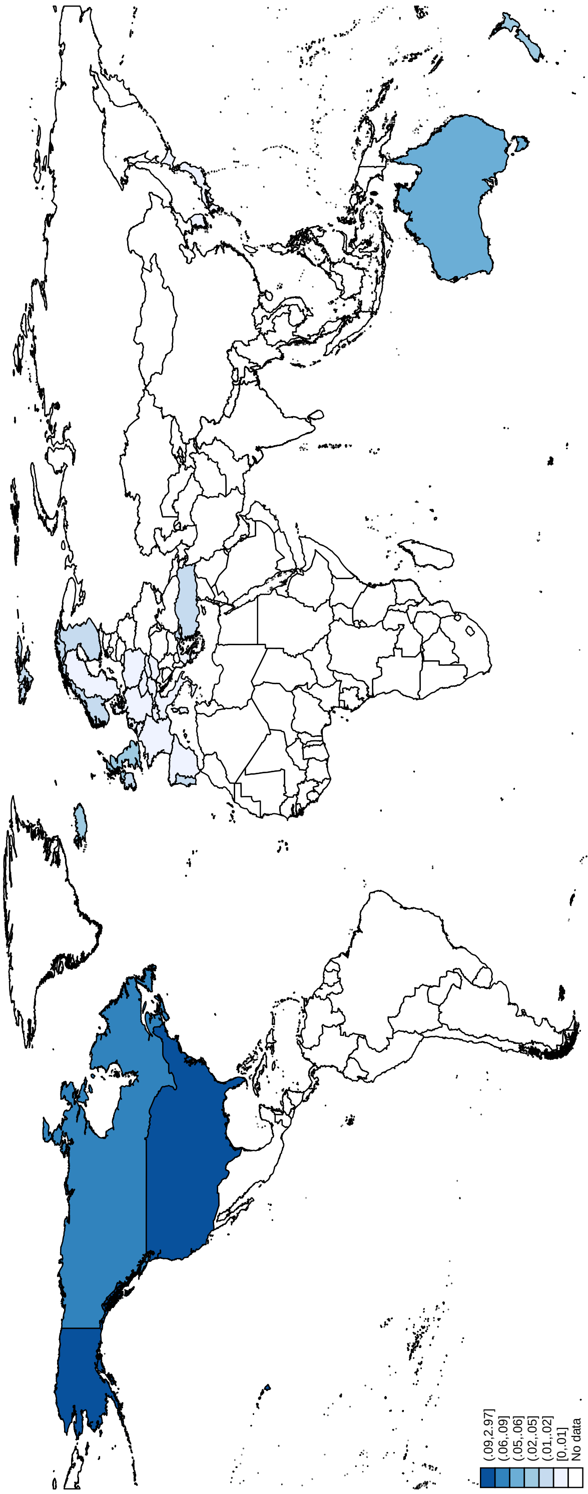

Figure B.2: Employment effects of a hypothetical labor market reform in the United States ($\hat{\kappa}_{U.S.} = 1.03$).

Table B.8: Comparative static effects of undoing recent German labor market reforms assuming balanced trade in 2006

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	PLM	SMF	share %GDP SMF		SMF	SMF	PLM	SMF
	%GDP	%GDP	$\% \ln(\hat{p})$	$\% \ln(\hat{e})$	$\%\hat{e}$	Δu	%EV	%EV
Australia	0.00	-0.03	92.78	7.22	-0.00	0.00	0.00	-0.03
Austria	0.00	-0.39	82.93	17.07	-0.07	0.06	0.00	-0.91
Belgium	0.00	-0.31	83.13	16.87	-0.05	0.05	0.00	-0.73
Canada	0.00	-0.02	99.31	0.69	-0.00	0.00	0.00	-0.00
Czech Republic	0.00	-0.25	83.41	16.59	-0.04	0.04	0.00	-0.57
Denmark	0.00	-0.24	83.44	16.56	-0.04	0.04	0.00	-0.56
Finland	0.00	-0.11	85.05	14.95	-0.02	0.02	0.00	-0.23
France	0.00	-0.16	84.19	15.81	-0.02	0.02	0.00	-0.34
Germany	0.00	-4.58	-37.40	101.08	-4.63	4.15	0.00	-3.11
Greece	0.00	-0.10	85.33	14.67	-0.02	0.01	0.00	-0.21
Hungary	0.00	-0.14	84.42	15.58	-0.02	0.02	0.00	-0.31
Iceland	0.00	-0.10	85.40	14.60	-0.01	0.01	0.00	-0.20
Ireland	0.00	-0.07	87.07	12.93	-0.01	0.01	0.00	-0.12
Italy	0.00	-0.12	84.98	15.02	-0.02	0.02	0.00	-0.24
Japan	0.00	-0.03	91.78	8.22	-0.00	0.00	0.00	-0.04
Korea	0.00	-0.04	91.06	8.94	-0.00	0.00	0.00	-0.05
Netherlands	0.00	-0.23	83.49	16.51	-0.04	0.04	0.00	-0.54
New Zealand	0.00	-0.03	92.76	7.24	-0.00	0.00	0.00	-0.03
Norway	0.00	-0.14	84.54	15.46	-0.02	0.02	0.00	-0.29
Poland	0.00	-0.22	83.58	16.42	-0.04	0.03	0.00	-0.50
Portugal	0.00	-0.09	85.83	14.17	-0.01	0.01	0.00	-0.18
Slovak Republic	0.00	-0.15	84.35	15.65	-0.02	0.02	0.00	-0.32
Spain	0.00	-0.10	85.59	14.41	-0.01	0.01	0.00	-0.19
Sweden	0.00	-0.14	84.46	15.54	-0.02	0.02	0.00	-0.30
Switzerland	0.00	-0.27	83.31	16.69	-0.04	0.04	0.00	-0.62
Turkey	0.00	-0.11	85.25	14.75	-0.02	0.01	0.00	-0.22
United Kingdom	0.00	-0.11	85.21	14.79	-0.02	0.01	0.00	-0.22
United States	0.00	-0.04	91.25	8.75	-0.00	0.00	0.00	-0.04
Average	0.00	-0.44	78.90	18.15	-0.39	0.35	0.00	-0.39

Notes: Counterfactual analysis based on parameter estimates from column (6) of Table 2.2. PLM gives results assuming perfect labor markets. SMF gives results using a search and matching framework for the labor market. Averages are weighted averages using country GDP as weight.

Table B.9: Heterogeneity of comparative static trade effects of undoing recent German labor market reforms assuming balanced trade with imperfect labor markets in 2006

	Changes in exports in percent by importer quantiles						
Exporting country	Min.	0.025	0.25	0.5	0.75	0.975	Max.
Australia	-0.26	-0.26	-0.10	-0.05	0.15	0.53	0.56
Austria	0.53	0.54	0.65	0.74	0.84	1.18	1.20
Belgium	0.37	0.37	0.49	0.58	0.68	1.15	1.20
Canada	-0.26	-0.26	-0.13	-0.08	0.13	0.50	0.53
Czech Republic	0.22	0.23	0.34	0.43	0.53	1.02	1.05
Denmark	0.21	0.22	0.33	0.42	0.52	1.01	1.04
Finland	-0.08	-0.08	0.04	0.13	0.33	0.71	0.74
France	0.02	0.02	0.14	0.23	0.43	0.81	0.84
Germany	0.01	0.01	0.12	0.21	0.42	0.80	0.83
Greece	-0.10	-0.10	0.02	0.11	0.31	0.69	0.72
Hungary	-0.01	-0.01	0.10	0.19	0.40	0.78	0.81
Iceland	-0.11	-0.10	0.01	0.11	0.31	0.69	0.71
Ireland	-0.18	-0.18	-0.02	0.03	0.23	0.61	0.64
Italy	-0.08	-0.07	0.04	0.13	0.34	0.72	0.75
Japan	-0.26	-0.25	-0.10	-0.04	0.16	0.54	0.57
Korea	-0.25	-0.24	-0.09	-0.04	0.17	0.54	0.57
Netherlands	0.19	0.20	0.31	0.40	0.50	0.99	1.02
New Zealand	-0.26	-0.26	-0.10	-0.05	0.15	0.53	0.56
Norway	-0.03	-0.02	0.09	0.18	0.39	0.77	0.79
Poland	0.16	0.17	0.28	0.37	0.47	0.96	0.99
Portugal	-0.13	-0.13	-0.01	0.08	0.28	0.66	0.69
Slovak Republic	-0.00	0.00	0.11	0.20	0.41	0.79	0.82
Spain	-0.12	-0.12	-0.00	0.09	0.30	0.67	0.70
Sweden	-0.02	-0.02	0.10	0.19	0.40	0.77	0.80
Switzerland	0.27	0.27	0.39	0.47	0.57	1.06	1.09
Turkey	-0.10	-0.09	0.02	0.12	0.32	0.70	0.73
United Kingdom	-0.09	-0.09	0.02	0.12	0.32	0.70	0.73
United States	-0.25	-0.25	-0.09	-0.04	0.17	0.54	0.57
Average	-0.02	-0.02	0.11	0.19	0.37	0.77	0.79

Notes: Counterfactual analysis based on parameter estimates from column (6) of Table 2.2 in the main text. Table depicts changes in normalized exports, i.e. exports divided by source and origin GDPs.

Table B.10: Heterogeneity of comparative static trade effects of PTA inception with perfect labor markets and controlling for trade imbalances in 2006

	Changes in exports in percent by importer quantiles						
Exporting country	Min.	0.025	0.25	0.5	0.75	0.975	Max.
Australia	-30.19	-29.69	-24.57	-23.39	-21.97	20.12	20.37
Austria	-32.09	-30.93	-3.37	0.47	2.42	6.46	7.02
Belgium	-32.84	-31.70	-4.21	-0.65	1.29	5.27	5.83
Canada	-33.64	-33.57	-31.02	-30.04	-28.74	4.98	9.92
Czech Republic	-31.44	-30.27	-2.45	1.44	3.41	7.48	8.05
Denmark	-30.58	-29.40	-1.23	2.70	4.70	8.82	9.40
Finland	-29.34	-28.14	0.53	4.46	6.57	10.76	11.35
France	-29.03	-27.82	0.98	4.93	6.93	11.25	11.84
Germany	-28.36	-27.14	1.94	5.92	7.89	12.27	12.90
Greece	-28.91	-27.70	1.15	5.11	7.06	11.44	12.03
Hungary	-30.69	-29.51	-1.38	2.54	4.54	8.65	9.23
Iceland	-28.49	-27.28	2.46	5.79	7.85	22.56	24.66
Ireland	-29.78	-28.58	-0.08	3.82	5.91	10.08	10.66
Italy	-28.27	-27.05	2.06	6.05	8.02	12.44	13.04
Japan	-17.92	-17.34	-11.32	-9.96	-8.41	4.63	4.83
Korea	-18.20	-17.52	-11.49	-10.00	0.20	24.21	24.32
Netherlands	-30.80	-29.63	-1.54	2.37	4.36	8.47	9.05
New Zealand	-20.24	-19.67	-13.83	-12.48	-10.85	16.41	19.42
Norway	-30.08	-28.89	0.18	3.44	5.67	19.84	21.89
Poland	-30.37	-29.19	-0.93	2.94	5.01	9.14	9.72
Portugal	-29.53	-28.33	0.27	4.19	6.29	10.47	11.06
Slovak Republic	-31.08	-29.91	-1.94	1.97	3.95	8.04	8.61
Spain	-28.17	-26.95	2.21	6.20	8.18	12.60	13.20
Sweden	-29.75	-28.56	-0.05	3.86	5.95	10.12	10.70
Switzerland	-33.20	-32.07	-3.32	-0.46	0.95	14.50	16.45
Turkey	-28.84	-27.63	1.97	5.28	7.33	21.98	24.06
United Kingdom	-25.67	-24.41	5.76	9.90	11.94	13.58	13.61
United States	-15.89	-15.80	-12.57	-11.10	-9.54	19.13	21.00
Average	-28.33	-27.31	-3.78	-0.52	1.68	12.35	13.36

Notes: Counterfactual analysis based on parameter estimates from column (6) of Table 2.2 in the main text. Table depicts changes in normalized exports, i.e. exports divided by source and origin GDPs.

Table B.11: Heterogeneity of comparative static trade effects of PTA inception with imperfect labor markets and controlling for trade imbalances in 2006

	Changes in exports in percent by importer quantiles						
Exporting country	Min.	0.025	0.25	0.5	0.75	0.975	Max.
Australia	-29.96	-29.51	-24.58	-23.37	-21.98	20.31	20.56
Austria	-32.05	-30.91	-3.69	0.19	2.15	6.21	6.79
Belgium	-32.82	-31.69	-4.54	-0.95	0.99	5.00	5.58
Canada	-33.61	-33.54	-30.97	-29.83	-28.67	5.28	10.21
Czech Republic	-31.39	-30.23	-2.75	1.17	3.15	7.25	7.84
Denmark	-30.52	-29.35	-1.52	2.45	4.45	8.61	9.20
Finland	-29.26	-28.08	0.26	4.20	6.34	10.57	11.17
France	-28.98	-27.79	0.66	4.61	6.67	11.00	11.61
Germany	-28.32	-27.12	1.59	5.59	7.57	12.02	12.65
Greece	-28.83	-27.64	0.87	4.84	6.81	11.24	11.85
Hungary	-30.63	-29.47	-1.68	2.29	4.29	8.43	9.02
Iceland	-28.37	-27.16	2.25	5.63	7.69	22.69	24.83
Ireland	-29.66	-28.47	-0.30	3.62	5.75	9.95	10.56
Italy	-28.21	-27.01	1.75	5.75	7.73	12.21	12.82
Japan	-17.61	-17.08	-11.28	-9.86	-8.33	4.96	5.19
Korea	-17.90	-17.28	-11.47	-9.96	0.38	24.35	24.47
Netherlands	-30.76	-29.60	-1.86	2.10	4.09	8.23	8.82
New Zealand	-20.03	-19.51	-13.88	-12.50	-10.91	16.62	19.63
Norway	-30.00	-28.82	-0.09	3.22	5.48	19.89	21.98
Poland	-30.31	-29.14	-1.23	2.65	4.76	8.93	9.52
Portugal	-29.44	-28.26	0.01	3.94	6.08	10.29	10.89
Slovak Republic	-31.03	-29.87	-2.24	1.71	3.69	7.81	8.40
Spain	-28.09	-26.88	1.92	5.93	7.92	12.40	13.02
Sweden	-29.68	-28.50	-0.33	3.58	5.71	9.92	10.52
Switzerland	-33.18	-32.06	-3.64	-0.73	0.69	14.45	16.44
Turkey	-28.74	-27.55	1.71	5.07	7.12	22.04	24.17
United Kingdom	-25.55	-24.30	5.52	9.67	11.73	13.34	13.37
United States	-15.99	-15.90	-12.64	-11.14	-9.63	19.34	21.20
Average	-28.25	-27.24	-4.01	-0.72	1.49	12.26	13.30

Notes: Counterfactual analysis based on parameter estimates from column (6) of Table 2.2 in the main text. Table depicts changes in normalized exports, i.e. exports divided by source and origin GDPs.

Table B.12: Heterogeneity of comparative static trade effects of $\hat{\kappa}_{U.S.} = 1.03$ controlling for trade imbalances with imperfect labor markets in 2006

	Changes in exports in percent by importer quantiles						
Exporting country	Min.	0.025	0.25	0.5	0.75	0.975	Max.
Australia	-0.98	-0.91	-0.10	-0.06	-0.01	0.08	0.08
Austria	-0.34	-0.27	0.50	0.58	0.62	0.72	0.72
Belgium	-0.30	-0.23	0.55	0.63	0.66	0.76	0.77
Canada	-0.97	-0.97	-0.49	-0.45	-0.40	-0.31	-0.31
Czech Republic	-0.39	-0.32	0.46	0.54	0.58	0.68	0.68
Denmark	-0.41	-0.34	0.44	0.52	0.57	0.66	0.66
Finland	-0.47	-0.40	0.38	0.46	0.51	0.60	0.60
France	-0.38	-0.31	0.47	0.55	0.58	0.69	0.69
Germany	-0.37	-0.30	0.47	0.55	0.59	0.69	0.69
Greece	-0.44	-0.37	0.40	0.48	0.53	0.62	0.62
Hungary	-0.42	-0.35	0.43	0.51	0.56	0.65	0.65
Iceland	-0.61	-0.54	0.27	0.32	0.36	0.45	0.45
Ireland	-0.54	-0.47	0.34	0.39	0.44	0.52	0.52
Italy	-0.41	-0.34	0.44	0.52	0.57	0.66	0.66
Japan	-0.44	-0.38	0.40	0.48	0.53	0.62	0.62
Korea	-0.46	-0.39	0.39	0.47	0.52	0.61	0.61
Netherlands	-0.36	-0.29	0.48	0.56	0.60	0.70	0.70
New Zealand	-0.85	-0.78	0.03	0.07	0.12	0.20	0.21
Norway	-0.46	-0.39	0.38	0.46	0.51	0.60	0.60
Poland	-0.41	-0.34	0.43	0.51	0.56	0.65	0.65
Portugal	-0.48	-0.41	0.36	0.44	0.49	0.58	0.58
Slovak Republic	-0.41	-0.34	0.44	0.52	0.57	0.65	0.66
Spain	-0.45	-0.38	0.39	0.47	0.52	0.61	0.61
Sweden	-0.44	-0.37	0.40	0.48	0.53	0.62	0.62
Switzerland	-0.29	-0.22	0.55	0.63	0.67	0.76	0.77
Turkey	-0.47	-0.41	0.37	0.45	0.50	0.59	0.59
United Kingdom	-0.60	-0.53	0.28	0.33	0.37	0.46	0.46
United States	-0.98	-0.91	-0.10	-0.06	-0.01	0.07	0.08
Average	-0.50	-0.44	0.33	0.41	0.45	0.54	0.54

Notes: Counterfactual analysis based on parameter estimates from column (6) of Table 2.2 in the main text. Table depicts changes in normalized exports, i.e. exports divided by source and origin GDPs.

Table B.13: Comparative static effects of undoing recent German labor market reforms controlling for trade imbalances in 2006

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	PLM	SMF	share %GDP SMF		SMF	SMF	PLM	SMF
	%GDP	%GDP	$\% \ln(\hat{p})$	$\% \ln(\hat{e})$	$\%\hat{e}$	Δu	%EV	%EV
Australia	0.00	-0.02	92.75	7.25	-0.00	0.00	0.00	-0.03
Austria	0.00	-0.35	82.14	17.86	-0.06	0.06	0.00	-0.89
Belgium	0.00	-0.29	82.44	17.56	-0.05	0.05	0.00	-0.72
Canada	0.00	-0.01	98.28	1.72	-0.00	0.00	0.00	-0.01
Czech Republic	0.00	-0.22	82.44	17.56	-0.04	0.04	0.00	-0.56
Denmark	0.00	-0.22	82.59	17.41	-0.04	0.04	0.00	-0.55
Finland	0.00	-0.09	84.05	15.95	-0.02	0.02	0.00	-0.23
France	0.00	-0.13	83.00	17.00	-0.02	0.02	0.00	-0.34
Germany	0.00	-4.58	-37.14	100.89	-4.63	4.16	0.00	-3.13
Greece	0.00	-0.08	83.99	16.01	-0.01	0.01	0.00	-0.21
Hungary	0.00	-0.12	83.20	16.80	-0.02	0.02	0.00	-0.30
Iceland	0.00	-0.08	84.47	15.53	-0.01	0.01	0.00	-0.20
Ireland	0.00	-0.05	85.86	14.14	-0.01	0.01	0.00	-0.12
Italy	0.00	-0.09	83.56	16.44	-0.02	0.02	0.00	-0.24
Japan	0.00	-0.03	92.24	7.76	-0.00	0.00	0.00	-0.04
Korea	0.00	-0.03	91.42	8.58	-0.00	0.00	0.00	-0.05
Netherlands	0.00	-0.21	82.72	17.28	-0.04	0.04	0.00	-0.53
New Zealand	0.00	-0.02	92.76	7.24	-0.00	0.00	0.00	-0.03
Norway	0.00	-0.12	83.57	16.43	-0.02	0.02	0.00	-0.29
Poland	0.00	-0.20	82.61	17.39	-0.04	0.03	0.00	-0.49
Portugal	0.00	-0.07	84.47	15.53	-0.01	0.01	0.00	-0.17
Slovak Republic	0.00	-0.12	83.13	16.87	-0.02	0.02	0.00	-0.31
Spain	0.00	-0.08	84.19	15.81	-0.01	0.01	0.00	-0.19
Sweden	0.00	-0.12	83.49	16.51	-0.02	0.02	0.00	-0.30
Switzerland	0.00	-0.24	82.24	17.76	-0.04	0.04	0.00	-0.60
Turkey	0.00	-0.09	84.07	15.93	-0.02	0.01	0.00	-0.21
United Kingdom	0.00	-0.09	84.21	15.79	-0.02	0.01	0.00	-0.22
United States	0.00	-0.03	91.18	8.82	-0.00	0.00	0.00	-0.05
Average	0.00	-0.43	78.53	18.52	-0.39	0.35	0.00	-0.39

Notes: Counterfactual analysis based on parameter estimates from column (6) of Table 2.2. PLM gives results assuming perfect labor markets. SMF gives results using a search and matching framework for the labor market. Averages are weighted averages using country GDP as weight.

Table B.14: Heterogeneity of comparative static trade effects of undoing recent German labor market reforms controlling for trade imbalances with imperfect labor markets in 2006

	Changes in exports in percent by importer quantiles						
Exporting country	Min.	0.025	0.25	0.5	0.75	0.975	Max.
Australia	-0.25	-0.25	-0.08	-0.02	0.18	0.56	0.59
Austria	0.49	0.49	0.62	0.71	0.81	1.15	1.17
Belgium	0.34	0.34	0.47	0.56	0.66	1.14	1.19
Canada	-0.25	-0.25	-0.10	-0.05	0.16	0.54	0.57
Czech Republic	0.18	0.19	0.32	0.41	0.51	1.01	1.04
Denmark	0.18	0.19	0.31	0.40	0.50	1.00	1.03
Finland	-0.10	-0.09	0.03	0.13	0.34	0.72	0.75
France	-0.01	-0.00	0.12	0.21	0.42	0.81	0.84
Germany	-0.02	-0.02	0.11	0.20	0.41	0.80	0.83
Greece	-0.12	-0.12	0.01	0.11	0.31	0.70	0.72
Hungary	-0.04	-0.04	0.09	0.18	0.40	0.78	0.81
Iceland	-0.12	-0.12	0.01	0.11	0.31	0.70	0.73
Ireland	-0.20	-0.19	-0.02	0.03	0.24	0.62	0.65
Italy	-0.10	-0.10	0.03	0.12	0.34	0.72	0.75
Japan	-0.24	-0.24	-0.07	-0.01	0.19	0.58	0.61
Korea	-0.24	-0.23	-0.07	-0.01	0.20	0.58	0.61
Netherlands	0.17	0.17	0.30	0.39	0.49	0.99	1.02
New Zealand	-0.25	-0.25	-0.08	-0.02	0.18	0.56	0.59
Norway	-0.05	-0.04	0.08	0.17	0.39	0.77	0.80
Poland	0.13	0.14	0.26	0.35	0.45	0.95	0.98
Portugal	-0.15	-0.15	-0.02	0.08	0.28	0.67	0.70
Slovak Republic	-0.03	-0.03	0.10	0.19	0.40	0.79	0.82
Spain	-0.14	-0.14	-0.01	0.09	0.30	0.68	0.71
Sweden	-0.04	-0.03	0.09	0.18	0.40	0.78	0.81
Switzerland	0.22	0.23	0.35	0.44	0.55	1.04	1.07
Turkey	-0.12	-0.11	0.01	0.11	0.32	0.70	0.73
United Kingdom	-0.11	-0.10	0.02	0.12	0.33	0.71	0.74
United States	-0.24	-0.24	-0.07	-0.01	0.19	0.57	0.60
Average	-0.04	-0.04	0.10	0.18	0.37	0.77	0.80

Notes: Counterfactual analysis based on parameter estimates from column (6) of Table 2.2 in the main text. Table depicts changes in normalized exports, i.e. exports divided by source and origin GDPs.

Appendix C

Appendix to Chapter 3

C.1 Computational details about the system of equations of the multilateral resistance terms and the counterfactuals

In the following, I describe the algorithm for computing the counterfactual changes. It is essentially identical to the one given in Anderson and van Wincoop (2003).

For convenience, I repeat the system of equations given in (3.11):

$$\tilde{\Pi}_i \equiv \left(\sum_{j=1}^{n} \left(\frac{t_{ij}}{\tilde{P}_j} \right)^{1-\sigma} \tilde{\theta}_j \right)^{1/(1-\sigma)}, \quad \tilde{P}_j \equiv \left(\sum_{i=1}^{n} \left(\frac{t_{ij}}{\tilde{\Pi}_i} \right)^{1-\sigma} \theta_i \right)^{1/(1-\sigma)}.$$

For computational reasons, it is convenient to rewrite this system of equations as

$$\theta_i = \mathfrak{t}_i \sum_{j=1}^{n} t_{ij}^{1-\sigma} \mathfrak{P}_j, \quad \tilde{\theta}_j = \mathfrak{P}_j \sum_{i=1}^{n} t_{ij}^{1-\sigma} \mathfrak{t}_i, \tag{C.1}$$

where I have defined $\mathfrak{P}_j \equiv \tilde{P}_j^{\sigma-1} \tilde{\theta}_j$ and $\mathfrak{t}_i \equiv \tilde{\Pi}_i^{\sigma-1} \theta_i$.[1] The equations given in (C.1) constitute a system of $2n$ equations in the n unknowns $\mathfrak{P}_j$ and the n unknowns $\mathfrak{t}_i$ which can be solved by standard nonlinear equation solvers.

The steps to compute the counterfactual values are as follows:

1. Having estimated the gravity equation given in Equation (3.24), one can obtain an estimate of the trade cost matrix (risen to the power of $1-\sigma$), $t_{ij}^{1-\sigma}$, in the observed baseline scenario. Given this estimate as well as the observed income shares, θ_js and $\tilde{\theta}_j$s, in the data, one can solve the system of equations given in Equation (C.1) for the vector of unknown $\mathfrak{t}_i$s and $\mathfrak{P}_j$s in the baseline scenario.

2. After changing the trade cost matrix (or any other model parameter) to the values of the unobserved counterfactual, one has to resolve the system of equations given in Equation (C.1) for the now counterfactual values of the $\mathfrak{t}_i$s and $\mathfrak{P}_j$s. However, in

[1] For a derivation see Heid and Larch (2012a).

this solution for the counterfactual situation, one has to take into account general equilibrium effects, i.e. the changes in GDPs, and the associated income shares in the counterfactual, θ_j^c. To calculate the change in the income shares, one has to take into account the counterfactual change in GDPs by multiplying the observed GDPs in the data by the counterfactual change in GDP implied by the model as given in Equation (3.21).

3. Having obtained the solution, one can calculate counterfactual changes according to the formulae given in the main text, normalizing nominal variables by a numéraire, as the system of equations given in Equation (C.1) determines the solutions only up to a scalar due to Walras' law.

C.2 Summary statistics gravity data set

Table C.1: Summary statistics

Variable	Mean	Std. Dev.	Min.	Max.	N
$x_{ij\tau}$	76.500	428.768	0	12,885.180	8,743
$z_{ij\tau}$	4.61×10^{-7}	2.08×10^{-6}	0	4.34×10^{-5}	8,743
$PTA_{ij\tau}$	0.282	0.450	0	1	8,743
$\ln DIST_{ij\tau}$	7.864	0.676	5.854	8.759	8,743
$CONTIG_{ij}$	0.220	0.414	0	1	8,743

Notes: Summary statistics for the sample of 13 Latin American and Caribbean countries from 1950 to 2006. The 13 countries included are: Argentina, Boliva, Brazil, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Nicaragua, Paraguay, Peru, Uruguay, Venezuela. Data are from Head et al. (2010).

C.3 Comparative static results with balanced trade

Table C.2: Comparative static effects of switching on all observed PTAs in 2006, balanced trade

Country	$\Delta\%e_j^f$	$\Delta\%pts\,u_j^f$	$w_j^f/w_j^i\Delta\%$	$\Delta\%L_j^i$	$\Delta\%pts\,u_j^o$	$\Delta\%EV$	$\Delta\%EV(PLM)$	$\Delta\%EV(SMF)$
Argentina	9.9	-8.0	-12.9	-60.7	4.9	-10.4	10.8	21.5
Bolivia	22.5	-16.3	-36.2	-42.9	4.7	60.4	24.9	49.2
Brazil	2.0	-1.8	-3.6	-51.9	1.4	-14.2	2.2	4.4
Colombia	8.7	-6.8	-17.2	-59.2	7.8	-9.3	9.4	18.7
Costa Rica	0.3	-0.3	-0.6	-19.5	-0.0	-3.9	0.3	0.9
Dominican Rep.	-0.4	0.4	0.6	14.1	0.0	2.1	-0.5	-0.7
Ecuador	9.1	-7.4	-11.7	-47.8	3.7	-5.0	10.0	19.9
El Salvador	0.5	-0.5	-1.0	-19.1	0.1	-4.8	0.5	1.1
Nicaragua	13.7	-10.4	-14.7	-45.1	4.9	3.4	15.3	30.3
Paraguay	21.3	-14.8	-31.2	-36.3	5.7	50.2	23.7	46.8
Peru	10.0	-8.4	-20.2	-51.2	3.7	5.6	10.9	21.6
Uruguay	19.3	-14.6	-23.7	-82.0	7.0	-5.4	21.7	43.3
Venezuela	6.1	-5.1	-6.9	-53.4	2.9	-14.0	6.7	13.3
Average (L_j-weighted)	5.8	-4.7	-9.3	-50.9	3.1	-7.5	6.3	12.6

Notes: Counterfactual analysis based on parameter estimates from column (4) of Table 3.1. PLM gives results assuming perfect labor markets. SMF gives results using a unified labor market with search and matching frictions. Averages are weighted averages using a country's labor force as weight.

Appendix D

Appendix to Chapter 4

D.1 Model variant with production in the informal sector

D.1.1 General remarks

In the following we describe a model of a small open economy, Mexico, with a foreign-owned *maquila* and a domestically-owned standard manufacturing sector in the presence of a unified informal labor market for unskilled workers. The key difference to the model presented in Chapter 4 is that we allow unskilled workers in the informal sector to produce varieties of the standard manufacturing good, i.e. the good which is consumed in Mexico. The informal sector varieties are assumed to be non-tradeable. This is in line with evidence from recent representative surveys of small scale enterprises typically associated with the informal sector, which indicate that more than 99% of these enterprises do not engage in any exporting activities in Mexico.[1] Formal sector standard manufacturing varieties remain tradable.

Specifically, we model the informal sector as an endogenously determined mass of homogeneous firms à la Krugman (1980), which employ all the informal unskilled workers who did not get a job at a formal sector firm. The mass of firms is pinned down by a free-entry condition, which also implies that there are no profits in the informal sector.

D.1.2 Consumption

Mexican households only consume goods produced in the manufacturing sector, which means that *maquila* output is exported in its entirety. Consumers maximize

$$C_2 = M_2^{\frac{1}{1-\sigma}} \left[\int_{\omega \in \Omega_{2d}} [q_{2d}(\omega)]^{\frac{\sigma-1}{\sigma}} d\omega + \int_{\omega' \in \Omega_{2f}} [q_{2f}(\omega')]^{\frac{\sigma-1}{\sigma}} d\omega' + \int_{\omega'' \in \Omega_2^{inf}} [q_2^{inf}(\omega'')]^{\frac{\sigma-1}{\sigma}} d\omega'' \right]^{\frac{\sigma}{\sigma-1}}, \tag{D.1}$$

where Ω_{2d} is the set of varieties produced in the formal manufacturing sector in Mexico, Ω_{2f} the set of varieties imported from the US, and Ω_2^{inf} the set of manufacturing varieties

[1] Encuesta NAcional de MIcroNegocios, ENAMIN. This survey is comprised of a representative sample of Mexican enterprises with less than seven employees (including the owner).

produced in the informal sector. $\sigma > 1$ is the elasticity of substitution and M_2 denotes the total mass of manufacturing varieties available in Mexico.[2] We follow Blanchard and Giavazzi (2003) and normalize utility by $M_2^{\frac{1}{1-\sigma}}$ in order to ensure that an increase in the size of the economy does not mechanically translate into a smaller informal sector.

Taking into account the existence of iceberg transportation costs $\tau_2 \geq 1$ for imported varieties, the price index corresponding to the composite C_2 is given by:

$$P_2 = M_2^{\frac{1}{\sigma-1}} \left[\int_{\omega \in \Omega_{2d}} [p_{2d}(\omega)]^{1-\sigma} d\omega + \int_{\omega' \in \Omega_{2f}} [\tau_2 p_{2f}(\omega')]^{1-\sigma} d\omega' + \int_{\omega'' \in \Omega_2^{inf}} [p_2^{inf}(\omega'')]^{1-\sigma} d\omega'' \right]^{\frac{1}{1-\sigma}}. \tag{D.2}$$

Inverse demand for formally produced domestic and imported foreign varieties from sector 2 is then given by:

$$p_{2d}(\omega) = \left(\frac{Y}{M_2}\right)^{\frac{1}{\sigma}} P_2^{\frac{\sigma-1}{\sigma}} q_{2d}(\omega)^{-\frac{1}{\sigma}}, \quad p_{2f}(\omega) = \left(\frac{\tau_2 Y}{M_2}\right)^{\frac{1}{\sigma}} P_2^{\frac{\sigma-1}{\sigma}} q_{2f}(\omega)^{-\frac{1}{\sigma}}, \tag{D.3}$$

where Y denotes total expenditure in Mexico. Note that we define $p_{2f}(\omega)$ as the cif price in the US and $q_{2f}(\omega)$ is the total quantity produced, including the quantity lost in transit due to the iceberg transportation costs.

Inverse demand for manufacturing varieties produced in the informal sector is given by:

$$p_2^{inf}(\omega) = \left(\frac{Y}{M_2}\right)^{\frac{1}{\sigma}} P_2^{\frac{\sigma-1}{\sigma}} q_2^{inf}(\omega)^{-\frac{1}{\sigma}}. \tag{D.4}$$

D.1.3 Production

Formal firms in both sectors are heterogeneous with respect to their idiosyncratic productivity φ as in Melitz (2003). Since each firm produces a unique variety, we index firm-level variables by φ.

Manufacturing firms

There is an unbounded mass of potential entrants in the domestic formal manufacturing sector. To enter, producers pay a sunk cost f_{e2}. All costs in the model are denominated in terms of the manufacturing good.[3] After incurring this cost, formal firms draw their productivity from a Pareto distribution with density $g(\varphi) = ak^a \varphi^{-(a+1)}$ for $\varphi \geq k$.[4] Formal firms that choose to operate need to pay a fixed cost f_2 per period. Having set up the plant, formal manufacturing firms produce their output by combining skilled labor s and unskilled labor l in a Cobb-Douglas form,

$$q_2(\varphi) = \varphi (s_2)^{\beta_{2s}} (l_2)^{1-\beta_{2s}}, \tag{D.5}$$

where β_{2s} is the labor cost share of skilled workers.

[2] The total number of manufacturing varieties available for consumption in Mexico is $M_2 = M_{2d} + M_{2x}^f + M_2^{inf}$ where M_{2x}^f denotes the mass of imported varieties, and M_2^{inf} the mass of varieties produced in the informal sector.

[3] Note that this implies that not all output produced can be used for consumption.

[4] We also restrict $a > \sigma - 1$ to ensure that the variance of the sales distribution is finite.

Formal firms sell their output domestically but can also incur an additional fixed cost f_{x2} to serve the foreign market through exports. We borrow the notion of a small open economy under monopolistic competition from Flam and Helpman (1987), and the extension to a heterogeneous-firm environment proposed by Demidova and Rodríguez-Clare (2009). This assumption implies that, despite the fact that formal firms located in Mexico face a downward-sloping demand schedule for their exports, their pricing decisions do not affect the price index, expenditure nor the mass of firms operating abroad, however, the subset of formal firms exporting to Mexico, M_{2x}^{f}, is endogenous.[5] Thus, foreign inverse demand for Mexican manufacturing exports by formal manufacturing firms is given by

$$p_{2x}(\varphi) = A_{2x}^{1/\sigma}\left(\frac{q_{2x}(\varphi)}{\tau_2}\right)^{-\frac{1}{\sigma}}, \tag{D.6}$$

where A_{2x} is a demand-shifter parameter that is taken as given by Mexican formal manufacturing firms. Hence, we define total revenue for a Mexican formal manufacturing firm with productivity φ as:

$$\begin{aligned} r_2(\varphi) &= r_{2d}(\varphi) + \mathbb{I}_x(\varphi) r_{2x}(\varphi) \\ &= \left(\frac{Y}{M_2}\right)^{\frac{1}{\sigma}} P_2^{\frac{\sigma-1}{\sigma}} q_{2d}(\varphi)^{\frac{\sigma-1}{\sigma}} + \mathbb{I}_x(\varphi) A_{2x}^{1/\sigma}\left(\frac{q_{2x}(\varphi)}{\tau_2}\right)^{\frac{\sigma-1}{\sigma}}, \end{aligned} \tag{D.7}$$

where $\mathbb{I}_x(\varphi)$ is an indicator function that takes the value one if a formal manufacturing firm with productivity φ exports and zero otherwise.

Maquiladora firms

We model *maquiladoras* in a similar fashion to formal manufacturing firms, therefore in this section we just highlight the differences between the two formal sectors, namely that (i) *maquila* plants are foreign-owned, (ii) export all their output and (iii) use foreign manufacturing goods as intermediate inputs for production.

A foreign investor pays a sunk entry cost in Mexico to set up a *maquiladora* plant.[6] *Maquiladoras* draw their productivity from the same Pareto distribution as Mexican manufacturing firms. Since *maquiladoras* export all their output, there is no meaningful distinction between domestic and exporting fixed costs. We assume that *maquiladoras* use foreign manufacturing goods as intermediate inputs, denoted by i, for production along with skilled and unskilled labor. Thus, production of *maquila* output for a plant with productivity φ takes the form

$$q_1(\varphi) = \varphi (s_1)^{\beta_{1s}} (l_1)^{\beta_{1l}} (i_1)^{1-\beta_{1s}-\beta_{1l}}, \tag{D.8}$$

where β_{1s} and β_{1l} are the skilled and unskilled labor cost shares for *maquila* plants, respectively.

Inverse demand for *maquila* variety φ abroad is given by

$$p_{1x}(\varphi) = A_{1x}^{1/\sigma}\left(\frac{q_{1x}(\varphi)}{\tau_1}\right)^{-\frac{1}{\sigma}}, \tag{D.9}$$

[5] Demidova and Rodríguez-Clare (2009)'s framework needs an endogenous variable that clears the trade balance. In Demidova and Rodríguez-Clare (2009) the price index and expenditure abroad are unaffected by Mexican firms but the share of US firms exporting to Mexico is endogenous.

[6] The fixed costs of entry, operation and vacancy posting for unskilled workers are incurred in Mexico and are denominated in units of the Mexican manufacturing good.

where A_{1x} is a foreign demand shifter that *maquiladora* plants take as given and has a similar interpretation to A_{2x} defined above. $\tau_1 > 1$ are the iceberg transportation costs to ship a *maquila* variety to the US. Total revenues for a *maquiladora* plant with productivity φ are given by

$$r_1(\varphi) = r_{1x}(\varphi) = A_{1x}^{1/\sigma}\left(\frac{q_{1x}(\varphi)}{\tau_1}\right)^{\frac{\sigma-1}{\sigma}}. \tag{D.10}$$

Unlike Mexican-owned plants in the formal manufacturing sector, profits derived from the operation of *maquila* plants are repatriated abroad.

Informal sector manufacturing firms

In contrast to formal sector manufacturing firms, informal sector firms are not heterogeneous in their productivity. Instead, we model firms as in Krugman (1980). This reflects the fact that informal sector establishments tend to be rather homogeneous in the sense that they are mostly small and unproductive. If an informal sector firm were very productive and hence very large, it would very likely be detected by government authorities. As we do not explicitly model any tax evasion incentives for firms in order to keep the informal sector production as simple as possible, our way of modeling informal sector firms as homogeneous should be seen as a reduced form way of capturing the stylized facts on informal sector establishments.

We assume that informal sector firms produce manufacturing good varieties which are only consumed in Mexico and which cannot be exported to the US market. In order to set up production, an informal sector firm has to pay a fixed cost f_2^{inf}. Once this cost is incurred, the production function is given by

$$q_2^{inf}(\omega) = \frac{1}{\varphi^{inf}} l_2^{inf}(\omega) \tag{D.11}$$

were l_2^{inf} is the labor demand of an informal sector firm, and φ^{inf} is a productivity parameter. This production function assumes that informal sector firms only use unskilled workers, reflecting the stylized fact that skilled workers are predominantly employed in the formal parts of the Mexican economy.

Profit maximization then implies that all informal firms charge the same price

$$p_2^{inf} = \frac{\sigma}{\sigma-1}\varphi^{inf} b w_l, \tag{D.12}$$

where $1-b$ is the now endogenous formal sector wage premium and w_l is the wage of unskilled workers in the formal economy. In the model version without informal sector production, $1-b$ is exogenous.

We assume that there is free entry in the informal sector for additional establishments so that operating profits equal fixed costs in equilibrium, i.e.

$$\frac{p_2^{inf} l_2^{inf}/\varphi^{inf}}{\sigma} = f_2^{inf} P_2. \tag{D.13}$$

D.1.4 Labor market

Since most individuals employed in the informal sector are unskilled, we assume that search and matching frictions only affect these workers, whereas skilled workers face a perfectly competitive labor market. Thus in our model only unskilled workers are employed in the informal sector. Following Satchi and Temple (2009), unskilled individuals that are unable to get matched with neither a plant in the formal manufacturing sector nor in the formal *maquiladora* sector become informal workers. These individuals earn income bw_l, with $b \in (0, 1)$, by working in informal sector manufacturing firms as described above, so we can interpret $1 - b$ as the formality wage premium for unskilled workers.

In order to hire unskilled workers, firms in the formal sectors need to post vacancies v at a cost c per vacancy. As is common in the search and matching literature, we assume that the matching technology is a constant returns to scale Cobb-Douglas function, $m(\theta) = \overline{m}\theta^{-\gamma}$, with $\gamma \in (0, 1)$ and where $\theta \equiv v/u$ is the vacancy-informality ratio, and $\overline{m}$ determines the overall efficiency of the matching process in the economy. The probability that a vacancy is filled is given by $m(\theta)$, which is decreasing in θ, and the probability that an unskilled individual in the informal sector finds a job in a formal plant is $\theta m(\theta)$ which is increasing in θ. We follow Keuschnigg and Ribi (2009) and consider a one-shot, static version of the search and matching framework in which the entire population of unskilled workers has just one opportunity to get matched with formal sector firms.

The optimal labor demand decision for a formal manufacturing firm solves the following program:

$$\pi_2(\varphi) = \max_{l_2, s_2} \left\{ r_2(\varphi) - w_l l_2 - w_s s_2 - cP_2 \left(\frac{l_2}{m(\theta)} \right) - f_2 P_2 - f_{x2} P_2 \mathbb{I}_x(\varphi) \right\}, \tag{D.14}$$

where we have also made use of the fact that a formal manufacturing plant wishing to hire l_2 unskilled workers needs to post $l_2/m(\theta)$ vacancies.[7]

The solution to program (D.14) yields two policy rules, one for skilled labor demand, which is the usual condition that the marginal revenue product of skilled labor has to be equal to the skilled wage, w_s, and a second one for formal unskilled employment that shows that firms have monopsony power and take into account that their vacancy posting has an impact on the wage rate for formal unskilled workers:

$$\frac{\partial r_2(\varphi)}{\partial l_2} = w_l + \frac{\partial w_l}{\partial l_2} l_2 + \frac{cP_2}{m(\theta)}. \tag{D.15}$$

As in Stole and Zwiebel (1996) we assume that unskilled workers bargain individually with their formal employers (in both formal sectors) about their wage and are all treated as the marginal worker. Total surplus of a worker-employer match is split according to a generalized Nash bargaining solution in each sector j, i.e. $(1 - \mu)[E(\varphi) - U] = \mu \partial \pi_j(\varphi)/\partial l_j$ where $E(\varphi)$ denotes the income of an unskilled worker being employed at a plant with productivity φ, U is the income of a worker in the informal sector, and $\mu \in (0, 1)$ measures the bargaining power of a worker.

Following the same procedure as Felbermayr et al. (2011a) and Larch and Lechthaler (2011), i.e. combining the first-order conditions for unskilled employment by plants in both sectors together with the surplus-splitting rule yields a set of two job-creation conditions

[7] The labor demand program for *maquila* plants is almost identical to equation (D.14), the only difference being that *maquiladoras* also need to choose how much foreign intermediate inputs to use for production.

(one for each sector):

$$w_l + \frac{cP_2}{m(\theta)} = \left[\frac{\beta_{1l}(\sigma-1)}{\sigma - \beta_{1l}\mu + \beta_{1l}\sigma\mu - \sigma\mu}\right]\varphi p_{1x}(\varphi)s_1(\varphi)^{\beta_{1s}}l_1(\varphi)^{\beta_{1l}-1}i_1(\varphi)^{1-\beta_{1s}-\beta_{1l}}, \tag{D.16}$$

$$w_l + \frac{cP_2}{m(\theta)} = \left[\frac{(1-\beta_{2s})(\sigma-1)}{\sigma + \beta_{2s}\mu - \mu - \beta_{2s}\sigma\mu}\right]\varphi p_{2d}(\varphi)\left(\frac{s_2(\varphi)}{l_2(\varphi)}\right)^{\beta_{2s}}, \tag{D.17}$$

and the wage curve is given by:

$$w_l = \frac{\mu c P_2}{(1-\mu)(1-b)}\left[\theta + \frac{1}{m(\theta)}\right]. \tag{D.18}$$

Note that since we assume that wages for unskilled formal workers are the same in manufacturing and *maquiladora* firms, we assume that the labor market for unskilled workers is unified. The same holds for skilled workers.

D.1.5 Productivity cutoffs and entry

As described in Section D.1.3, the production side of formal sector firms in our model closely follows Melitz (2003). Because $\pi_j(\varphi)$ is a strictly increasing function of φ, only plants with high enough productivity to earn non-negative profits will start production. Thus the usual productivity cutoff for production in sector j is defined implicitly by $\pi_j(\varphi_j^*) = 0$. In the formal manufacturing sector, where plants need to incur a fixed cost to serve the foreign market, an export cutoff is similarly defined as $\pi_{2x}(\varphi_{2x}^*) = 0$. We follow Melitz (2003) and define average productivity in formal sector j as:

$$\tilde{\varphi}_j \equiv \left[\frac{1}{1-G(\varphi_j^*)}\int_{\varphi_j^*}^{\infty} \varphi^{\sigma-1}g(\varphi)d\varphi\right]^{\frac{1}{\sigma-1}}, \quad j = 1, 2. \tag{D.19}$$

Using the cutoff productivity of the least productive exporting manufacturing firm φ_{2x}^*, we can define the average productivity for formal manufacturing exporters analogously. Finally, let $\chi_2 \equiv [1-G(\varphi_{2x}^*)]/[1-G(\varphi_2^*)]$ denote the ex-ante probability that a manufacturing plant exports, conditional on successful entry. Using these definitions we can write the free-entry condition for plants in sector j as $[1-G(\varphi_j^*)]\overline{\pi}_j = f_{ej}P_2$.[8]

D.1.6 Aggregate variables

The equilibrium share of informal workers in the labor force follows from the one-period equivalent of the Beveridge curve and is given by $u = 1/[1+\theta m(\theta)]$. The mass of formal firms operating in sector j in Mexico, M_{jd}, is pinned down by the labor market clearing condition for unskilled workers:

$$M_{1d} = \frac{L_1}{l_1(\tilde{\varphi}_1)}; \qquad M_{2d} = \frac{L_2}{l_{2d}(\tilde{\varphi}_2) + \chi_2 l_{2x}(\tilde{\varphi}_{2x})}, \tag{D.20}$$

with $L_1 + L_2 = (1-u)\overline{L}$, where L_j denotes total unskilled formal employment in sector j and $\overline{L}$ is the total endowment of unskilled labor in the economy. Market clearing for skilled labor is given by $M_{1d}s_1(\tilde{\varphi}_1) + M_{2d}\big[s_{2d}(\tilde{\varphi}_2) + \chi_2 s_{2x}(\tilde{\varphi}_{2x})\big] = \overline{S}$.

[8] For *maquiladoras* $\overline{\pi}_1 = \pi_1(\tilde{\varphi}_1)$ and for manufacturing plants $\overline{\pi}_2 = \pi_{2d}(\tilde{\varphi}_2) + \chi_2\pi_{2x}(\tilde{\varphi}_{2x})$.

The overall mass of informal sector firms is then given by

$$l_2^{inf} M_2^{inf} = u\bar{L}. \tag{D.21}$$

Finally, the trade balance condition reads:

$$\underbrace{M_{1d} r_1(\tilde{\varphi}_1)}_{\text{value of } maquila \text{ exports}} + \underbrace{\chi_2 M_{2d} r_{2x}(\tilde{\varphi}_{2x})}_{\text{value of manufacturing exports}} =$$

$$\underbrace{\tau_2^{1-\sigma}\left(\frac{Y}{M_2}\right)\left(\frac{P_2}{P_2^f}\right)^{\sigma-1}}_{\text{value of manufacturing imports}} + \underbrace{\tau_2 P_2^f M_{1d} i_1(\tilde{\varphi}_1)}_{\text{value of intermediate imports}} + \underbrace{M_{1d}\pi_1(\tilde{\varphi}_1)}_{\text{aggregate } maquila \text{ profits}} . \tag{D.22}$$

We define the foreign price index for manufacturing goods, P_2^f, as the *numéraire*. Note that aggregate profits in the manufacturing sector remain in Mexico, since this sector is domestically owned.

Curriculum Vitae

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Current positions

04/2010–today	Researcher, Chair of Economics VI: Empirical Economics, Prof. Dr. Mario Larch, University of Bayreuth
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