

ABSTRACT

Title of dissertation: **TRADE SHOCKS AND LOCAL
LABOR MARKETS' LINKAGES:
THEORY AND EVIDENCE**

Marisol Rodríguez Chatruc, Doctor of Philosophy, 2016

Dissertation directed by: **Professor Nuno Limão**
Department of Economics

In the past few years, there has been a concern among economists and policy makers that increased openness to international trade affects some regions in a country more than others. Recent research has found that local labor markets more exposed to import competition through their initial employment composition experience worse outcomes in several dimensions such as, employment, wages, and poverty. Although there is evidence that regions within a country exhibit variation in the intensity with which they trade with each other and with other countries, trade linkages have been ignored in empirical analyses of the regional effects of trade, which focus on differences in employment composition.

In this dissertation, I investigate how local labor markets' trade linkages shape the response of wages to international trade shocks. In the second chapter, I lay out a standard multi-sector general equilibrium model of trade, where domestic regions trade with each other and with the rest of the world. Using this benchmark, I decompose a region's wage change resulting from a national import cost shock into

a direct effect on prices, holding other endogenous variables constant, and a series of general equilibrium effects. I argue the direct effect provides a natural measure of exposure to import competition within the model since it summarizes the effect of the shock on a region's wage as a function of initial conditions given by its trade linkages. I call my proposed measure linkage exposure while I refer to the measures used in previous studies as employment exposure. My theoretical analysis also shows that the assumptions previous studies make on trade linkages are not consistent with the standard trade model.

In the third chapter, I calibrate the model to the Brazilian economy in 1991—at the beginning of a period of trade liberalization—to perform a series of experiments. In each of them, I reduce the Brazilian import cost by 1 percent in a single sector and I calculate how much of the cross-regional variation in counterfactual wage changes is explained by exposure measures. Over this set of experiments, employment exposure explains, for the median sector, 2 percent of the variation in counterfactual wage changes while linkage exposure explains 44 percent.

In addition, I propose an estimation strategy that incorporates trade linkages in the analysis of the effects of trade on observed wages. In the model, changes in wages are completely determined by changes in market access, an endogenous variable that summarizes the real demand faced by a region. I show that a linkage measure of exposure is a valid instrument for changes in market access within Brazil. By using observed wage changes in Brazil between 1991-2000, my estimates imply that a region at the 25th percentile of the change in domestic market access induced by trade liberalization, experiences a 0.6 log points larger wage decline (or smaller

wage increase) than a region at the 75th percentile. The estimates from a regression of wages changes on exposure imply that a region at the 25th percentile of exposure experiences a 3 log points larger wage decline (or smaller wage increase) than a region at the 75th percentile. I conclude that estimates based on exposure overstate the negative impact of trade liberalization on wages in Brazil.

In the fourth chapter, I extend the standard model to allow for two types of workers according to their education levels: skilled and unskilled. I show that there is substantial variation across Brazilian regions in the skill premium. I use the exogenous variation provided by tariff changes to estimate the impact of market access on the skill premium. I find that decreased domestic market access resulting from trade liberalization resulted in a higher skill premium. I propose a mechanism to explain this result: that the manufacturing sector is relatively more intensive in unskilled labor and I show empirical evidence that supports this hypothesis.

TRADE SHOCKS AND LOCAL LABOR MARKETS'
LINKAGES : THEORY AND EVIDENCE

by

Marisol Rodríguez Chatruc

Dissertation submitted to the Faculty of the Graduate School of the
University of Maryland, College Park in partial fulfillment
of the requirements for the degree of
Doctor of Philosophy
2016

Advisory Committee:

Professor Nuno Limão, Chair/Advisor

Professor Sebastián Galiani, Advisor

Professor Judith K. Hellerstein, Advisor

Professor Luca D. Opromolla

Professor Ernesto F. Calvo

© Copyright by
Marisol Rodríguez Chatruc
2016

Dedication

A mi mamá y mis hermanas

Acknowledgements

I am deeply grateful to my advisor, Professor Nuno Limão, for his constant guidance and support and for his enormous patience throughout this process. I am greatly indebted to the rest of my dissertation committee, Professor Sebastián Galiani and Professor Judith Hellerstein, who were always willing to discuss ideas with me and gave me valuable suggestions. I thank Luca Opromolla and Lourenço Paz for their insightful comments. I thank Ernesto Calvo for kindly agreeing to be in my committee.

I cannot imagine having gone through this process without my friends: Jeronimo Carballo, Jingting Fan, Camilo Morales-Jimenez, and Stephanie Rennane. They provided direct help with my dissertation and gave me support, motivation, and encouragement.

Finally, none of this would have been possible without the constant emotional support of my mother and my sisters, who despite living thousands of miles away, were aware of every deadline and presentation. They always kept me down to earth and reminded me about what is truly important in life.

Contents

<i>List of Abbreviations</i>	ix
<i>1. Introduction</i>	1
<i>2. Theory</i>	13
2.1 Introduction	13
2.2 Environment, behavior, and equilibrium	13
2.3 Decomposing regional wage changes	21
2.4 The wage elasticity and employment composition	25
2.4.1 The partial-equilibrium wage elasticity	26
2.4.2 From partial-equilibrium elasticities to employment-weighted averages of shocks	29
2.5 An illustration using simulations	35
2.6 Conclusion	39
2.7 Figures	40
<i>3. Trade shocks, linkages, and wages in Brazil</i>	42
3.1 Introduction	42
3.2 Data sources and measurement	43
3.3 Counterfactuals	46
3.3.1 Approach	47
3.3.2 A 1% import cost reduction in each sector	54
3.3.2.1 Nominal wage decomposition	54
3.3.2.2 Wages, exposure, and linkages	56
3.3.3 Trade liberalization	60
3.3.4 Real wages	65
3.4 Estimating the wage elasticity with respect to market access	67
3.4.1 Exposure as an instrument for changes in market access	68
3.4.2 Estimation results	75
3.4.3 Economic significance across different approaches	79
3.4.4 Robustness	81
3.5 Conclusion	84
3.6 Tables	85
3.7 Figures	104

4. <i>Market access and the skill premium in Brazil</i>	114
4.1 Introduction	114
4.2 Model	115
4.3 Descriptive evidence	119
4.4 Estimation	122
4.4.1 Empirical approach	122
4.4.2 Results	122
4.4.3 Mechanism	126
4.5 Conclusions	129
4.6 Tables	130
4.7 Figures	135
5. <i>Conclusion</i>	140
 <i>Appendix</i>	 142
A. <i>Theory</i>	143
A.1 Derivation of the wage elasticity	143
A.2 Shock affecting the two sectors	144
A.3 Parameter values for simulations	145
A.4 Omitted variable bias	146
A.5 Solution algorithm	147
A.6 Price index changes	149
A.7 Assumptions on residual sectors and implications for relative wages .	150
A.7.1 Constant returns to scale non-traded sector. Balanced trade. .	150
A.7.2 Constant returns to scale non-traded sector. Unbalanced trade.	151
A.7.3 Freely traded residual sector	152
B. <i>Data</i>	154
B.1 Data sources and definitions	154
B.1.1 Wages and employment	154
B.1.2 Tariffs	155
B.1.3 Elasticities of substitution	155
B.1.4 Interstate and international trade	156
B.1.5 Distances	157
B.1.6 Rest of the world wages and employment	157
B.1.7 Industry crosswalks	159
B.2 Supplementary tables	161
 <i>Bibliography</i>	 166

List of Tables

3.1	Descriptive statistics of microregion level variables.	85
3.2	Descriptive statistics of ad-valorem tariffs.	85
3.3	Gravity equation: 1999. PPML estimates.	86
3.4	Tariff levels and tariff changes by sector: 1991-1995 and 1991-1998 . .	87
3.5	Counterfactual wage change decomposition.	88
3.6	R-squared decomposition from synthetic OLS regressions. 1% shock in each sector.	89
3.7	Distribution of counterfactual wage changes and price index changes. Trade liberalization.	90
3.8	Correlation coefficient between exposure measures	90
3.9	Standard exposure measures and change in counterfactual wages in Brazilian microregions. Trade liberalization. OLS estimates.	91
3.10	Linkages-based exposure measures and change in counterfactual wages in Brazilian microregions. Trade liberalization. OLS estimates. . . .	92
3.11	Exposure measures and change in counterfactual wages in Brazilian microregions. Trade liberalization. OLS estimates.	93
3.12	R-squared decomposition from synthetic OLS regressions. Trade lib- eralization.	94
3.13	Changes in market access and changes in wages in Brazilian microre- gions. OLS and 2SLS estimates. 1991-2000	95
3.14	Changes in market access and changes in wages in Brazilian microre- gions. Other IVs. 2SLS estimates. 1991-2000	96
3.15	Changes in market access and changes in wages in Brazilian microre- gions. First stage estimates. 1991-2000	97
3.16	Exposure measures and changes in wages in Brazilian microregions. OLS (reduced form) estimates. 1991-2000	98
3.17	Changes in market access and changes in wages in Brazilian microre- gions. Robustness checks. 2SLS estimates. 1991-2000	99
3.18	Changes in market access and changes in wages in Brazilian microre- gions. Robustness checks. OLS estimates. 1991-2000	100
3.19	Changes in market access and in-migration rates in Brazilian mi- croregions. Robustness checks. 2SLS estimates.	101
3.20	Estimated wage change implied by an interquartile range change in market access and exposure	102
3.21	Descriptive statistics of variables in wage regressions	103

4.1	Changes in domestic market access and changes in the skill premium in Brazilian microregions. OLS and 2SLS estimates. 1991-2000	130
4.2	Changes in domestic market access and changes in wages in Brazilian microregions. OLS and 2SLS estimates. 1991-2000	130
4.3	Changes in domestic market access and changes in the skill premium in Brazilian microregions. 2SLS estimates. 1991-2000	131
4.4	Changes in domestic market access and changes in wages in Brazilian microregions. 2SLS estimates. 1991-2000	131
4.5	Exposure and changes in the skill premium and wages in Brazilian microregions. Reduced form estimates. 1991-2000	132
4.6	Descriptive statistics of variables in regressions	133
4.7	Share of hours and share of wage bill by sector, 1991 and 2000	134
B.1	Elasticities of substitution and Cobb-Douglas shares	161
B.2	Imports and exports to value added ratio, 1991 and 2000.	162
B.3	Calibrated border effects: 1991 and 2000.	163
B.4	Correspondence between industrial classifications.	164
B.5	Data Sources and Classifications.	165

List of Figures

2.1	Simulated wage response to a national shock to sector 1	40
2.2	Simulated wage response to region-specific shocks to sector 1	40
2.3	Simulated wage response to a shock to sector 1 under complete specialization	41
3.1	Regional share of hours in manufacturing sectors (in %): 25th and 75th pctile	104
3.2	Evolution of applied Brazilian tariffs: 1987-1998	105
3.3	ETC exposure measure and change in counterfactual wages in Brazilian microregions. 1% shock in each sector.	106
3.4	LTC exposure measure and change in counterfactual wages in Brazilian microregions. 1% shock in each sector.	107
3.5	ETC exposure measure and change in counterfactual real wages in Brazilian microregions. 1% shock in each sector.	108
3.6	LTC exposure measure and change in counterfactual real wages in Brazilian microregions. 1% shock in each sector.	109
3.7	Counterfactual wage changes in Brazilian microregions resulting from trade liberalization.	110
3.8	Exposure measures in Brazilian microregions	111
3.9	Observed nominal wage changes and market access changes in Brazilian microregions 1991-2000	112
3.10	Predicted wage changes and market access changes in Brazilian microregions 1991-2000	113
4.1	Share of workers by educational attainment (%)	135
4.2	Skill premium in Brazilian microregions: 1991 and 2000	136
4.3	Unskilled wages in Brazilian microregions: 1991 and 2000	136
4.4	Skilled wages in Brazilian microregions: 1991 and 2000	137
4.5	Skill premium changes in Brazilian microregions: 1991-2000	137
4.6	Skilled and unskilled wage changes in Brazilian microregions: 1991-2000	138
4.7	Skill intensity by sector in 1991	139

List of Abbreviations

<i>BRA</i>	Brazil
<i>DMA</i>	Domestic market access
<i>EMC</i>	Employment-weighted import change
<i>ELTC</i>	Employment and linkage-weighted trade cost change
<i>ETC</i>	Employment-weighted trade cost change
<i>LTC</i>	Linkage-weighted trade cost change
<i>MA</i>	Market access
<i>ROW</i>	Rest of the World
CNAE	Brazilian National Classification System of Economic Activities
MERCOSUR	Common Market of the South
NAFTA	North American Free Trade Agreement
SCN	Brazilian National Accounts System
SECEX	Brazilian Secretary of Foreign Trade

Chapter 1: Introduction

In the past few years, there has been a concern among economists and policy makers that increased openness to international trade affects some regions in a country more than others. Recent research, which started with the work by Topalova (2007) on India and was fueled by the influential work by Autor et al. (2013) on the U.S., has found that local labor markets more exposed to import competition through their initial employment composition experience worse outcomes in several dimensions such as employment, wages, and poverty.¹

However, there is evidence that regions within a country exhibit variation not only in their employment composition but also in the intensity with which they trade with each other and with other countries.² In this context, a nationwide shock that increases import competition leads to a direct decrease in domestic demand for a region's products that is larger the more intensely the region exports to domestic

¹President Obama addressed globalization-induced regional disparities in his 2012 State of the Union Address: "What's happening in Detroit can happen in other industries. It can happen in Cleveland and Pittsburgh and Raleigh." The study by Autor et al. (2013) was cited in 2012 by Alan Krueger as a Chairman of the Council of Economic Advisers and in the 2015 Economic Report of the President.

²The literature on the *home bias* documents that trade within a country is larger than international trade (Anderson and Van Wincoop (2003)). For instance, in Brazil during the 1990s trade between states was about five times larger than international trade. Several empirical studies find that intra-country trade flows follow a gravity structure which implies large variation in trade intensity across regions (see Fally et al. (2010) for Brazil, Bartelme (2014) and Monte et al. (2015) for the U.S., and Anderson and Van Wincoop (2003) and for the U.S. and Canada).

regions for which the rest of world is an important supplier. This decline in demand, triggers a series of general equilibrium effects, whose magnitude also depends on how intensely regions trade with each other. Nevertheless, trade linkages have been ignored in empirical analyses of the regional effects of trade, which focus on differences in employment composition.

In this dissertation, I investigate how local labor markets' trade linkages shape the response of wages to international trade shocks. My analysis shows that within a standard model of trade, the exposure of a region to a trade shock – understood as a direct effect, holding other variables constant – is a function of the initial configuration of its trade linkages rather than a function of its initial employment composition as previous literature assumes. Based on the standard model, I propose an estimation strategy that uses exposure as an instrument for changes in market access, an endogenous variable that summarizes the real demand faced by a region and that drives wage changes in the model. Applying this approach to the Brazilian trade liberalization episode, I find that previous estimation strategies overstate the negative impact of tariff reductions on regional wages.

In the second chapter, I lay out a standard multi-sector general equilibrium model of trade, where domestic regions trade with each other and with the rest of the world. Using this benchmark, I decompose a region's wage change resulting from a national import cost shock into a direct effect on prices and a series of general equilibrium effects. I argue that the direct effect provides a natural measure of exposure within the model since it summarizes the effect of the shock on a region's wage as a function of its initial conditions. These initial conditions are a weighted

average of the rest of the world's share in the imports of each domestic region, where the weights are given by the share of each domestic region in the exports of the region of interest. Therefore, a region is more exposed if the domestic markets that represent an important share of its exports, tend to have a high import share from the rest of the world. I call my proposed measure linkage exposure while I refer to the measures used in previous studies as employment exposure.

My theoretical analysis also assesses the assumptions used in previous studies. Importantly, I show that the assumptions previous studies make on trade linkages are not consistent with the standard trade model. Finally, I simulate the model to illustrate my main arguments in a simple example with two domestic regions.

In the third chapter, I take the model to the data, I analyze counterfactual scenarios to increases in import competition, and I propose an empirical strategy to estimate the impact of import shocks on wages. By using several sources of publicly available data, I calibrate the model to the Brazilian economy in 1991, at the beginning of a period of trade liberalization. I perform two counterfactual exercises, in which the unit of analysis is a Brazilian microregion, a collection of municipalities that can be considered a local labor market.

In the first exercise, I perform a series of experiments to quantify how much of the cross-regional variation in wage changes generated using the model is explained by employment exposure, the most commonly used measure in the empirical literature, which is given by a region's initial employment share in the affected sector. In each experiment, I reduce the Brazilian import cost by 1 percent in a single sector, while holding constant all other parameters, and I obtain counterfactual wage

changes in all the microregions. I find that regions with higher employment exposure do not exhibit larger wage losses in all the experiments, which goes against the intuition behind previous approaches. In addition, I find initial employment shares are weakly related to counterfactual wage changes. Over this set of experiments, employment shares explain 2 percent of the variation in counterfactual wage changes for the median sector.

Given the weak relationship between employment exposure and counterfactual wage changes, I use as a measure of exposure the direct effect I obtained in the decomposition in chapter 2. This measure incorporates information about the trade linkages between each region and all other regions, including the rest of the world. I find that with this definition of exposure, regions more exposed to the shock exhibit larger wage losses in most of the experiments. Furthermore, over this set of experiments, the linkage-based measure explains 44 percent of the variation in counterfactual wage changes for the median sector. As a final exercise, I let employment and linkage exposure compete in a regression in which a microregion's counterfactual wage change is the dependent variable. I find that linkage exposure explains 75 percent of the variation explained by both measures.

In the real world, trade shocks tend to occur in several sectors simultaneously and with varying intensity. To analyze this case, I perform a trade liberalization experiment and I evaluate exposure measures in that context. More precisely, I reduce trade costs by the amount implied by the tariff changes that took place in Brazil during the early 1990s. I find employment exposure explains between 3 and

7 percent of the variation in counterfactual wage changes whereas linkage exposure explains between 11 and 26 percent of this variation.

The calibration of the model allows me to perform a welfare analysis of the effect of shocks. I obtain that counterfactual real wages, the measure of welfare in the model, increase by 0.11 log points on average as a consequence of trade liberalization. However, real wages do not increase in all regions, which implies that some regions lose from tariff reductions. I also find that linkage exposure can explain more of the cross-regional variation in real wages than employment exposure.

Having evaluated the predictive capacity of exposure measures using model-generated wage changes, I propose an estimation strategy that incorporates trade linkages in the analysis of the effects of trade on observed wages. In the model, changes in wages are completely determined by changes in market access. Market access is an endogenous variable that summarizes the real demand faced by a region and it is equal to a trade cost weighted average of real expenditures in all regions. Since changes in market access are not exogenous, we expect the estimates of the elasticity of wages with respect to market access to be biased. I argue that a measure of linkage exposure is a valid instrument for changes in market access within Brazil. By using this instrument and observed wage changes between 1991-2000, I estimate a wage elasticity with respect to domestic market access of 0.12. I perform several robustness checks that leave this main result essentially unaltered.

Finally, I compare the results of the market access approach I propose with the results of the reduced-form approach, which utilizes employment exposure to predict wage changes. My estimates imply that a region at the 25th percentile of

the change in domestic market access induced by trade liberalization, experiences a 0.6 log points larger wage decline (or smaller wage increase) than a region at the 75th percentile. The estimates from a regression of wages changes on employment exposure imply that a region at the 25th percentile of exposure experiences a 3 log points larger wage decline (or smaller wage increase) than a region at the 75th percentile.³ I conclude that reduced-form estimates based on employment exposure overstate the negative impact of trade liberalization on wages in Brazil.

In the fourth chapter, I extend the standard model to allow for two types of workers according to their education levels: skilled and unskilled. I show that there is substantial variation across Brazilian regions in the skill premium, defined as the log difference of wages between skilled and unskilled workers. I use the exogenous variation provided by tariff changes to estimate the impact of market access on the skill premium. I find that decreased domestic market access resulting from trade liberalization resulted in a higher skill premium and that this is explained by a higher elasticity of unskilled wages to market access compared to skilled wages. I propose a mechanism to explain this result: that the manufacturing sector is relatively more intensive in unskilled labor and I show empirical evidence that supports this hypothesis.

My dissertation contributes to several strands of the literature. In the first place, it relates to the aforementioned empirical studies on the effects of trade shocks on local labor markets' outcomes. In her work, Topalova (2007) and Topalova

³Since tariff changes are negative, a lower (more negative) exposure leads to a higher wage decline.

(2010) find that rural districts in India that were more exposed to trade liberalization through their employment composition experienced a lower decline in poverty than less exposed regions. McLaren and Hakobyan (2010) find that U.S. locations more vulnerable to NAFTA experienced lower wage growth than locations that had no protection against Mexico. Kovak (2013) finds that Brazilian regions that experienced a larger price decline due to a change in tariffs experienced a larger wage decline. Autor et al. (2013) estimate the impact of the increase in Chinese import penetration in the U.S. and find that rising imports from China cause higher unemployment, lower labor force participation, and lower wages in commuting zones that are relatively more specialized in import-competing manufacturing industries. Chiquiar et al. (2015) apply this empirical framework to Mexico and find heterogeneity between border and non-border states, with effects of exposure on the labor market being higher in states that share a border with the US.

Among these studies, only Kovak (2013) and Autor et al. (2013) derive their econometric specifications from a trade model. Kovak (2013) develops a specific-factors model that delivers that the wage change in a region is a weighted average of goods prices changes with weights that depend on the fraction of regional labor allocated to each sector. In this model, trade linkages do not play a role in determining price changes since goods are homogeneous and there are no transport costs. Therefore, trade flows do not follow a gravity structure. Autor et al. (2013) use a conventional trade model that yields a gravity equation. However, they make a set of simplifying assumptions to arrive to an estimating equation and leave no role for trade linkages between local labor markets. I analyze these assumptions in section

2.4. I contribute to this literature in several ways: first, by deriving an exposure measure that takes trade linkages into account, second, by analyzing the predictive power of employment-based exposure measures in the context of a standard model, and third, by reinterpreting exposure measures as instruments for domestic market access.

In the second place, my dissertation relates to the international trade literature that quantifies the effects of globalization. Several authors use the workhorse model of Eaton and Kortum (2002) to calculate gains from trade and to measure the impact of trade shocks on labor markets. For example, Caliendo and Parro (2014) quantify the welfare and real wage effects of the NAFTA agreement and Caliendo et al. (2014) study the transmission of productivity shocks through trade linkages across U.S. states. Some studies allow for migration between labor markets, like Fan (2015), Redding (2012), Monte (2014) and Monte et al. (2015), where the last two incorporate commuting. In a recent paper, Caliendo et al. (2015) quantify the impacts of the China shock studied by Autor et al. (2013) using a dynamic trade model and find that this shock can explain 50 percent of the unexplained change in manufacturing employment.

The studies in this literature that are closely related to my dissertation are Monte (2014) and Monte et al. (2015). Both of them evaluate the performance of reduced-form approaches compared to the predictions of a model with linkages. Monte (2014) does so in a model with commuting but where trade costs are uniform across all regions in the country, so there is no role for gravity forces. In addition, he only shocks one industry while I conduct an analysis industry by industry. This

allows me to find that in some sectors the correlation between employment exposure and changes in wages goes in the wrong direction. In Monte et al. (2015) trade follows a gravity structure but the focus of the study is on the employment elasticity while the focus of my analysis is on the wage elasticity. In addition, their model is single-sector, so industry composition plays no role. Finally, both papers evaluate reduced form methods by using counterfactual wage changes, while I go one step further and propose an estimation strategy that uses observed wage changes.

In the third place, my dissertation relates to the literature that studies the influence of geography on factor prices. Harris (1954) first introduced the concept of market potential defined *ad hoc* as the summation of markets accessible to a region divided by their distances from that region.⁴ Fifty years later, Fujita et al. (1999) derived from first principles the *wage equation*: a relationship between factor prices and market potential. Based on their model, Redding and Venables (2004) propose a two-step procedure to estimate the wage equation using cross-country variation in income. Hanson (2005) also uses Fujita et al. (1999) framework to study the spatial correlation between wages and consumer purchasing power across U.S. counties for the period 1970-1990. In a related study, Kumar (2007) estimates the impact of domestic and foreign market access on regional wages in India during a period of trade liberalization. More recently, Bartelme (2014) estimates the elasticity of wages and employment with respect to market access, an analogous notion to market potential, using U.S. MSA level data for the periods 1990-2000 and 2000-2007.

⁴Although the notion of market potential is usually attributed to Harris, in his 1954 paper he attributes the concept to Colin Clark.

I contribute to this literature by proposing an estimation strategy that uses changes in tariffs to identify the elasticity of wages to market access. The works by Kumar (2007) and Bartelme (2014) are the most closely related to my dissertation. My strategy differs from Kumar (2007) in two main aspects. First, he uses an *ad hoc* measure of market access, in the spirit of Harris (1954), whereas I construct market access using the structure of the model. Second, he measures the effect of trade liberalization by interacting initial market access with a liberalization dummy whereas I use variation in tariff changes to construct an instrument for market access. My strategy differs from Bartelme (2014) in that he uses an *ad hoc* Bartik instrument for market access while I derive my instrument from a trade model.

The influence of geography on relative factor prices or on the skill premium is a much less studied phenomenon. Venables and Limão (2002) develop a spatial Heckscher-Ohlin model in which the relationship between relative factor returns and distance is non-monotonic across different zones. Thus, depending on the pattern of specialization, relative factor returns can increase or decrease with remoteness. Chiquiar (2008) develops a simplified version of Venables and Limão (2002) to analyze if there were Stolper-Samuelson effects in Mexican regions after trade liberalization and finds that U.S. bordering and North states had an increase in unskilled wages and a drop in the skill premium. Redding and Schott (2003) extend Fujita et al. (1999) to allow for two types of workers and they find that educational attainment has a positive correlation with market access across countries. However, they do not estimate the impact of market access on the skill premium. My dissertation

contributes to this literature by providing an estimate of the elasticity of the skill premium to market access.

Finally, my dissertation relates to the studies on the impact of trade liberalization on wage inequality in developing countries. Goldberg and Pavcnik (2007) document an increase in inequality in several developing countries during the 1980s and 1990s. Galiani and Sanguinetti (2003), Attanasio et al. (2004), and Gonzaga et al. (2006) use variation in the industry skill premium to estimate the impact of increased trade openness on wage inequality. Galiani and Sanguinetti (2003) use Argentine data and find that sectors in which import penetration increased, also experienced an increase in wage inequality but that this channel explains a small share of the observed rise in wage inequality. Attanasio et al. (2004) find that industry affiliation is an important determinant of workers' earnings in Brazil but the structure of industry wage premia is relatively stable over time. Gonzaga et al. (2006) document that skilled labor earnings differentials decreased during trade liberalization and they find that employment shifted from skilled to unskilled intensive manufacturing sectors, that each sector increased its relative share of skilled labor, and that relative prices fell in skill-intensive sectors. Following a different strategy, Dix-Carneiro and Kovak (2014) use regional variation in employment composition to analyze the effect of the Brazilian trade liberalization on regional wage inequality and find small effects of trade liberalization on the skill premium between 1991 and 2000. My dissertation contributes to this literature by providing a rationale to the tendency towards an increase in the skill premium in developing countries:

unilateral trade liberalizations reduces domestic market access and if the non-traded sector is relatively skill intensive, then the skill premium increases.

My dissertation bridges the gap between parallel literatures on the impacts of trade on labor markets' outcomes by reinterpreting the reduced form approach based on employment composition within a standard trade framework. I directly contribute to the reduced form literature by deriving a measure of exposure from a trade model and by showing this measure explains more of the variation in model-generated wage changes than measures based on employment composition. Furthermore, I propose an estimation strategy that uses my measure of exposure as an instrument for changes in market access, which are the driver of wage changes in the model. My estimates suggest that the employment-based measures used in the reduced form literature overstate the negative impact of trade liberalization on wages in Brazil.

Chapter 2: Theory

2.1 Introduction

Trade linkages between regions in a country generate complex interactions between endogenous variables after a nationwide shock. To understand these mechanisms, in section 2.2 I lay out a standard multi-region and multi-sector general equilibrium trade model and in section 2.3 I decompose regional wage changes after a sectoral trade cost shock. In order to derive an employment exposure measure from this decomposition, it is necessary to make a set of simplifying assumptions that I state explicitly in section 2.4. Finally, to illustrate some of my main arguments I simulate the model in a 2-region and 2-sector case in section 2.5.

2.2 Environment, behavior, and equilibrium

Environment The world economy consists of two main countries: the Home country, denoted by H , and the Rest of the World (ROW), denoted by R . The Home country, is a collection of regions $i = 1 \dots N$ that trade with each other as well as with ROW. There are multiple sectors in each region: one non-traded sector and several traded sectors $k = 1, \dots, K$ that use labor as the only a factor of production.

Labor is supplied inelastically in each region and is perfectly mobile between sectors but immobile between regions. The non-traded good is produced using a constant returns to scale technology. The traded goods are produced using an increasing returns to scale technology and are subject to *iceberg* transport costs and, if they cross international borders, to tariffs.¹

Consumer The representative consumer from region j (from now on j denotes the importing region and i the exporting region) has preferences given by a Cobb-Douglas upper-tier utility function

$$U_j = \prod_{k=0}^K (C_j^k)^{\mu_k} \quad (2.1)$$

with $\sum_{k=0}^K \mu_k = 1$, where C_j^0 denotes consumption of the non-traded good (indexed with 0) and C_j^k ($k = 1 \dots K$) corresponds to a CES consumption index of traded goods of differentiated varieties.²

Expenditure minimization by consumers in each region j yields an ideal price index for each sector $k = 1, \dots, N$ given by

$$(P_j^k)^{1-\sigma_k} = \sum_i n_i^k (p_{ij}^k)^{1-\sigma_k} \quad (2.2)$$

¹This setup is based on Autor et al. (2013). The main difference lies in the assumptions on trade deficits and on the non-traded sector technology. See footnote 5 for more detail. In addition, I analyze the effect of tariff shocks while their focus is on foreign productivity shocks.

²In equilibrium, all the varieties of an industry from region i are equally priced and therefore, they are consumed in the same amount by consumers from region j , which yields: $(C_j^k)^{\frac{\sigma_k-1}{\sigma_k}} = \sum_i n_i^k (c_{ij}^k)^{\frac{\sigma_k-1}{\sigma_k}}$, $\sigma_k > 1$, $k = 1, \dots, N$. Where σ_k is the (constant) elasticity of substitution between varieties in sector k and n_i^k is the number of varieties of good k .

where $\sigma_k > 1$ is the (constant) elasticity of substitution between varieties in sector k , n_i^k is the number of varieties of good k , and p_{ij}^k is the price of any variety of good k produced in region i and consumed in region j .³

By Sheppard's Lemma, region j 's demand for any variety of good k produced in region i is

$$c_{ij}^k = (p_{ij}^k)^{-\sigma_k} \mu_k E_j (P_j^k)^{\sigma_k - 1} \quad (2.3)$$

where E_j is region's j total expenditure, which includes expenditure on both manufactures and non-traded goods.⁴ Demand for any good produced by i is: the decreasing in its c.i.f. price (p_{ij}^k), increasing in expenditure in the destination market (E_j), and decreasing in competition in the destination market (i.e. increasing in P_j^k).

Non-traded sector The non-traded good is produced with a CRS technology that employs labor as the only factor of production

$$q_i^0 = \eta (L_i^0) \quad (2.4)$$

where L_i^0 is the amount of labor used in production of the non-traded good and η is a productivity parameter that does not vary across regions.⁵

³The equality in equation 2.2 relies on symmetry of prices in equilibrium.

⁴ $E_j = p_j^0 C_k^0 + \sum_{k=1}^N \sum_i n_i^k p_{ij}^k c_{ij}^k$, where p_j^0 is the price of the non-traded good.

⁵Autor et al. (2013) assume decreasing returns to scale in the production of the non-traded good so that any shock that expands employment in this sector exerts a downward pressure on wages in the region. However, the assumption of a Cobb-Douglas upper tier utility function implies employment in this sector is constant. To guarantee that non-traded sector employment changes after a shock, the authors allow for exogenous trade deficits. Since the focus of my analysis is on wages and not on non-traded employment, I assume balanced trade. I leave the analysis of trade imbalances and changes in non-traded employment for future research.

Producers in this sector take their good price and the wage as given so profit maximization implies

$$w_i = \eta p_i^0 \quad (2.5)$$

where w_i is the wage in region i and p_i^0 is the price of the non-traded good in region i .

Traded sectors The traded goods production side of the model follows Krugman (1980). These type of goods are produced with an increasing returns to scale technology and are traded at a cost $\kappa_{ij}^k \equiv d_{ij}^k b_{ij}^k \tau_{ij}^k$. The first component, $d_{ij}^k > 1$ is the *iceberg* transport cost of shipping good k from i to j . This cost is incurred when shipping goods between regions in *Home* and when shipping goods between *Home* and *ROW*, and viceversa. It is also incurred when any region trades with itself which means, $d_{ii} > 1$.⁶ The second component, $b_{ij}^k > 1$ is equal to 1 plus the tariff equivalent of non-tariff border-related costs. The third component, $\tau_{ij}^k \equiv 1 + t_{ij}^k$ is the tariff factor, with t_{ij}^k the ad-valorem tariff region j applies to imports from region i and sector k . Note that $b_{ij}^k = \tau_{ij}^k = 0$ when both i and j belong to *Home* and when *ROW* trades with itself.

The market structure is characterized by monopolistic competition: each firm in sector k faces a downward sloping demand curve with perceived price elasticity equal to σ_k , the elasticity of substitution in the consumption of varieties of good k .

⁶ d_{ij}^k units of the good must be shipped at the origin for one unit to arrive at the destination, which implies that a fraction $(d_{ij}^k - 1)/d_{ij}^k$ of the good is lost in transit.

The amount of labor required in each traded sector to produce q_i^k units of the good is given by

$$l_i^k = F^k + q_i^k / \beta_i^k \quad (2.6)$$

where F^k is a fixed cost that does not vary across regions and β_i^k is the productivity specific to sector k and region i . Variation in β_i^k reflects regional variation in natural comparative advantage.⁷

Under Mill pricing, the f.o.b. price $p_i^k = p_{ij}^k / d_{ij}^k$ does not depend on the destination market. Therefore, producers in each traded sector solve for one price: p_i^k . Due to CES preferences and Cobb-Douglas upper-tier utility, the optimal pricing rule is a constant mark-up over marginal costs

$$p_i^k = \left(\frac{\sigma_k}{\sigma_k - 1} \right) w_i / \beta_i^k \quad (2.7)$$

To close the production side of the model, I assume free entry into the traded sector. This drives firms' profits to zero and determines the number of firms in each traded sector. Since each firm's perceived demand elasticity is constant and equal to σ_k , the optimal scale of the firm depends only on parameters that are not affected by shocks to prices: $q_i^k = (\sigma_k - 1) F^k \beta_i^k$. Therefore, all of the adjustment to a sectoral

⁷The assumption that fixed costs do not vary across regions is for ease of interpretation and without loss of generality. Allowing for F^k to vary across regions would only change the interpretation of the calibrated parameters, β_i^k , since β_i^k and F^k enter multiplicatively in the expression for expenditure shares (equation 2.11) but since F^k is assumed not to vary across regions it is simplified and does not appear in the expression.

trade shock occurs at the extensive margin through changes in the number of firms (varieties) and not through changes in the size of firms.

Finally, free entry implies the following linear relationship between the number of firms and sectoral employment

$$n_i^k = \frac{L_i^k}{\sigma_k F_k} \quad (2.8)$$

Since the number of firms is generally not observed but sectoral employment is, this expression proves useful to bring the model to the data.

Gravity In this model trade flows are governed by a gravity equation

$$X_{ij}^k = n_i^k \mu_k E_j \left(\frac{P_j^k}{p_i^k \kappa_{ij}^k} \right)^{\sigma_k - 1} \quad (2.9)$$

where X_{ij}^k is region j 's expenditure on sector k goods produced in region i . This formula can be derived by multiplying the number of varieties of good k produced in region i , n_i^k , by prices paid in j for those varieties, p_{ij}^k , and by the demand for those varieties in region j , c_{ij}^k .

X_{ij}^k is increasing in the number of varieties produced in i in sector k , n_i^k , increasing in total expenditure in region j , E_j , decreasing in competition in market j , P_j^k , and decreasing in the prices at which i sells its varieties to j , p_{ij}^k .

The gravity equation is one of the most successful empirical applications of trade theory.⁸ In the quantitative section of this paper I use this equation to estimate sectoral trade costs.

Equilibrium Under balanced trade, total expenditure in region j is given by $E_j = w_j L_j$.⁹ The non-traded good market clearing condition is given by $p_j^0 X_j^0 = \mu_0 w_j L_j$, which together with equations 2.4 and 2.5 implies labor in the non-traded sector is a fixed share of total labor $L_j^0 = \mu_0 L_j$.

Labor market clearing in each region j requires that the sum of labor demand for each type of labor across sectors is equal to labor supply, which is fixed since I assume workers are immobile across regions

$$\sum_{k=1}^K L_j^k = (1 - \mu_0) L_j \quad (2.10)$$

where L_j is region j 's fixed endowment of labor and I used that labor in the non-traded sector is constant and equal to $\mu_0 L_j$.

To express the traded goods market clearing conditions in a simple way it is useful to define the expenditure share of region j on region i varieties of good k

$$\pi_{ij}^k = \frac{X_{ij}^k}{E_j^k} = \frac{L_i^k \left(w_i \kappa_{ij}^k / \beta_i^k \right)^{1-\sigma_k}}{\sum_h L_h^k \left(w_h \kappa_{hj}^k / \beta_h^k \right)^{1-\sigma_k}} \quad (2.11)$$

⁸See Anderson and Yotov (2012).

⁹This formula only holds under zero tariff revenue. I assume that the central government runs a balanced budget so that the decrease in tariff revenue after a reduction in tariffs is compensated by an equal increase in government transfers. I leave the analysis of the effect of changes in tariff revenue for future research.

where X_{ij}^k is region j 's expenditure on region i in sector k , and E_j^k is total expenditure in region j in sector k . The second equality in equation 2.11 uses equation 2.9.

The traded goods market equilibrium conditions are given by

$$w_i L_i^k = \sum_j \pi_{ij}^k \mu_k w_j L_j \quad k = 1, \dots, K \quad (2.12)$$

where the left-hand side is the wage bill of sector k , which is equal to the sector's revenue due to zero profits in equilibrium. The right hand side is equal to the sum of expenditures of all destination markets j on sector k goods proceeding from region i , where $\pi_{ij}^k \mu_k w_j L_j = X_{ij}^k$.¹⁰

Equations 2.10 and 2.12 form a system of $N + N \times K$ equations (N labor market clearing conditions and $N \times K$ goods market equilibrium conditions) in $N + N \times K$ variables (N regional wages and $N \times K$ sectoral labor allocations). The equilibrium of this economy is defined next.

Definition 1. An equilibrium is a set of wages $\{w_i, \dots, w_N\}$ and sectoral labor allocations $\{L_i^k, \dots, L_N^k\}_{k=1}^K$ given parameters $\{L_i, \kappa_{ij}^k, \beta_i^k\}_{i=1\dots N; k=1\dots K}$ such that:

1. Labor markets clear (equation 2.10 holds)
2. Goods market clear (equation 2.12 holds)

By Walras' Law, one equilibrium condition in a region is redundant. Therefore, the model can be solved up to a normalization.

¹⁰It can easily be shown that labor market clearing (equation 2.10) together with manufacturing goods market equilibrium (equation 2.12) imply balanced trade.

2.3 Decomposing regional wage changes

A region's wage elasticity to a trade shock is a parameter of interest to evaluate the impact of trade on local economies. For instance, Autor et al. (2013) find that U.S. Commuting Zones at the 75th percentile of exposure to Chinese import growth experience a 0.8 % larger fall in log earnings in the period 2000-2007 compared to Commuting Zones at the 25th percentile. Kovak (2013) finds that a 10 percentage-points larger price decline in a region—induced by tariff reduction—implies a 4 percentage-points larger wage decline. As I explained in the Introduction, these estimations do not take linkages between regions into account.¹¹ In this section I use the model in section 2.2 to decompose the wage change to a sector-specific tariff shock. This decomposition incorporates regional trade linkages and it takes into account the changes in endogenous variables in every region.

Consider a change in τ^s , the ad-valorem tariff factor that *Home* applies to *ROW* in sector s . Totally differentiating the equilibrium conditions in equation 2.12 yields an implicit formula for wage changes. To simplify notation, $\hat{x} = d \ln x$ denotes the log change of x . I decompose the wage change of any region that produces good s , into three effects

$$\hat{w}_i = DE_i + PE_i + SE_i \tag{2.13}$$

¹¹Using the treatment-effects lexicon, this type of analysis relies on the stable unit treatment value assumption (or SUTVA). Therefore, it assumes away any linkages or spillovers between treated units (regions in this case).

The first one, which I call *direct effect*, is the effect of the tariff change that operates directly on prices, holding constant wages and sectoral employment in every region, including region i . I denote the second component, *partial effect*, which is the effect of the tariff change that operates through changes in wages and in sectoral employment in region i , holding constant wages and sectoral employment in other regions $\neq i$. The third one, which I call *spillover effect*, comprises the effect on the wages of region i of changes in wages and sectoral employment in every region except for region i itself.¹² Each of these effects is given by

$$DE_i = (\sigma_s - 1) \left(\sum_{j \neq R} \xi_{ij}^s \pi_{Rj}^s \right) \hat{\tau}_s \quad (2.14)$$

$$PE_i = -(\sigma_s - 1) \left[\sum_j \xi_{ij}^s (1 - \pi_{ij}^s) \right] \hat{w}_i + \xi_{ii}^k \hat{E}_i - \left(\sum_j \xi_{ij}^s \pi_{ij}^s \right) \hat{L}_i^s \quad (2.15)$$

$$SE_i = (\sigma_s - 1) \sum_{h \neq i} \left(\sum_j \xi_{ij}^s \pi_{hj}^s \right) \hat{w}_h + \sum_{h \neq i} \xi_{ih}^s \hat{E}_h - \sum_{h \neq i} \left(\sum_j \xi_{ij}^s \pi_{hj}^s \right) \hat{L}_h^s \quad (2.16)$$

where $\xi_{ij}^s = \frac{\pi_{ij}^s \mu_k w_j L_j}{w_i L_i^s}$ is the share of sales to location j in location i 's total

revenue in sector s .¹³ The *direct effect* is a price effect that is not mediated by any

¹²Equation 2.14 also applies to isomorphic models like Armington or Eaton and Kortum (2002) in which expenditures shares do not depend on sectoral labor L_i^k .

¹³Note that the decomposition is valid as long as region i produces good s , the good affected by the shock. If region i does not produce the good, then $\xi_{ij}^k = 0$ for all j and all terms are equal to zero. However, the wage elasticity is not zero in this case since the changes in endogenous variables have an effect on region i wages through the goods market clearing conditions in the rest of the

change in wages or employment. Suppose the tariff that *Home* applies to products from ROW in sector s falls, this implies a fall in the prices that Home consumers pay for imported products in sector s , generating a substitution from domestic to imported goods. For a producer in region i , this generates a loss in her market share in every region at *Home*, including region i itself. This direct marginal loss in revenue is equal to the sum of the expenditure shares that each domestic market j allocates to sector s goods imported from the ROW, π_{Rj}^s , weighted by the share of sales to each domestic market j in location i 's total revenue in sector s , ξ_{ij}^s , with $\sum_j \xi_{ij}^s = 1$. This effect is higher when domestic locations that have high import penetration from the rest of the World in sector s (this is, high π_{Rj}^s) also tend to have a high weight in region i 's revenue in sector s (this is, high ξ_{ij}^k).¹⁴ To satisfy the zero profit condition, and given that all other endogenous variables are held constant, wages in region i must fall. The *partial effect* is the sum of three effects. The first term is negative because when region i wages decrease, region i goods become cheaper, therefore increasing demand from all markets to which it sells its products. This effect is greater the smaller is the share of region i in region j 's expenditure (the smaller is π_{ij}^s) weighted by ξ_{ij}^s .¹⁵ The second term is an expenditure effect. Since expenditures are monotonically related to wages, a decrease in wages in region i implies a decrease in expenditure. This affects region i 's demand for sector s goods more negatively if region i producers sell a high proportion of goods

sectors in which region i is a producer. See Appendix A.1 for the total differentiation of the system of equations that define an equilibrium.

¹⁴In addition, this effect increases with the elasticity of substitution σ_k since a higher σ_k implies a higher price elasticity of demand.

¹⁵The size of the effect is decreasing in π_{ij}^s because π_{ij}^s is concave in w_i .

to their local market (high ξ_{ii}^k). Finally, the last term is the effect of changes in hours worked in sector s , which are monotonically related to the number of varieties (firms) produced in the sector. If the number of varieties in sector s in region i decrease, competition for producers decreases dampening the initial effect of tariffs on wages

The intuition behind the spillover effect is analogous to the intuition behind the partial effect, with the exception that the sign of the wage change term is negative. When wages in a region different from i decrease, this generates more competition for region i producers and, other things equal, wages in region i have to decrease to restore zero profits.

I further decompose the direct and partial effects into own-region effects and the effect of linkages to other regions:

$$DE_i = \underbrace{(\sigma_s - 1) \xi_{ii}^s \pi_{Ri}^s}_{own} \hat{\tau}_s + \underbrace{(\sigma_s - 1) \sum_{j \neq i, R} \xi_{ij}^s \pi_{Rj}^s}_{linkages} \hat{\tau}_s \quad (2.17)$$

$$PE_i = \underbrace{-(\sigma_s - 1) [\xi_{ii}^s (1 - \pi_{ii}^s)] \hat{w}_i + \xi_{ii}^k \hat{E}_i - (\xi_{ii}^s \pi_{ii}^s) \hat{L}_i^s}_{own} \quad (2.18)$$

$$\underbrace{-(\sigma_s - 1) \left[\sum_{j \neq i} \xi_{ij}^s (1 - \pi_{ij}^s) \right] \hat{w}_i - \left(\sum_{j \neq i} \xi_{ij}^s \pi_{ij}^s \right) \hat{L}_i^s}_{linkages}.$$

In the case of the direct effect, the term labeled “own” reflects that a tariff reduction makes producers in region i face more competition in their own region.

This effect is more negative the higher is the import share from the ROW in region i , π_{Ri} , and the higher is the share of exports of region i to itself, ξ_{ii} . The term labeled “linkages” captures that when tariffs decrease, producers from region i face more competition in other domestic regions.¹⁶ In the extreme case that region i does not export to any other domestic region, ξ_{ii} would be equal to 1 and the “linkages” term would be zero.

The same type of reasoning can be applied to the partial effect in equation 2.18. A part of the effect can be attributed to changes in region i that affect region i through sales to itself. For example, the first term in the term labeled “own” captures that a decrease in wages, makes region i ’s goods cheaper for consumers in region i , which increases sales of region i to itself and counteracts the initial negative effect on wages of the tariff shock. The first term of the “linkages” term captures that a decrease in wages makes region i ’s goods cheaper for consumers in other domestic regions. The own effect has an extra term that the linkages effect does not have, which is an income effect. When wages decrease due to the tariff change, incomes in region i decrease and this depresses sales of producers from i to their own region even further.

2.4 The wage elasticity and employment composition

In this section I derive a reduced form employment composition measure from the decomposition formula for regional wage changes in equations 2.14-2.16. I define a

¹⁶Mathematically, this effect is derived from the differentiation of π_{ij}^s ($j \neq R, i$) with respect to τ^s . Therefore, the linkage effect takes place even if the trade cost between regions i and j does not change.

parameter called the partial-equilibrium wage elasticity and I state the necessary assumptions to go from that parameter to an employment weighted average of the shock. I show these assumptions are incompatible with assumptions on trade costs.

2.4.1 The partial-equilibrium wage elasticity

I define a parameter that I call *partial-equilibrium elasticity*.¹⁷ This is the elasticity of wages in region i setting the change in all the endogenous variables in other regions to zero: $dw_h = dL_h^k = 0, \forall h \neq i$. I denote $\tilde{x} = \frac{dx}{x} \frac{\tau_s}{d\tau_s}$ the elasticity of a variable, x , with respect to a tariff change in sector s . The partial equilibrium wage elasticity is given by

$$\tilde{w}_i^{PE} = \frac{(\sigma_s - 1) \sum_{j \neq R} \xi_{ij}^s \pi_{Rj}^s - \left(\sum_j \xi_{ij}^s \pi_{ij}^s \right) \tilde{L}_i^s}{1 + (\sigma_s - 1) \sum_j \xi_{ij}^s (1 - \pi_{ij}^s) - \xi_{ii}^s} \quad (2.19)$$

where $\tilde{w}_i^{PE} = \left. \frac{dw_i}{d\tau^s} \frac{\tau^s}{w_i} \right|_{dw_h = dL_h^k = 0 \forall h \neq i}$. The different terms in equation 2.19 are related to the direct and partial effects in equations 2.14 and 2.15. The first term in the numerator, analogously to the direct effect in equation 2.14, captures that the partial-equilibrium wage elasticity is higher if markets that are an important source of revenue for i tend to have a high expenditure shares on ROW goods. The second term in the numerator captures a competition effect.¹⁸ If tariffs decrease, the number of varieties (firms) in sector s in region i decrease, which decreases competition for producers in that sector and that region and dampens the negative effect of a reduction of tariffs on wages. The term given by $(\sigma_s - 1) \sum_j \xi_{ij}^s (1 - \pi_{ij}^s)$

¹⁷I borrow this terminology from Monte et al. (2015).

¹⁸Sectoral employment and varieties are linearly related by equation 2.8.

in the denominator is what multiplies region i 's wage changes in equation 2.15 and captures that lower wages in region i —as a consequence of a decrease in tariffs—make region i 's producers more competitive, dampening the initial negative effect of the tariff change on wages. Finally, ξ_{ii} multiplies region i 's expenditure changes in equation 2.15 and captures that lower lower wages in region i as a consequence of a decrease in tariffs, decrease expenditure of region i 's consumers, amplifying the initial negative effect of the tariff change on wages.

Equation 2.19 is not in closed form since it depends on the sectoral elasticities of hours, \tilde{L}_i^s . With a high number of sectors, a closed form expression becomes large. Therefore, in the next paragraph I study a two sectors case.

Two Sectors Case To obtain a closed form solution for the partial elasticity, I analyze the case where there are two traded sectors, $k = \{1, 2\}$, and a tariff shock at *Home* that only affects sector 1.¹⁹ The system of equations that determines the solution to the partial wage elasticity is given by

$$\begin{cases} \tilde{w}_i^{PE} = \frac{\left[(\sigma_s - 1) \sum_{j \neq R} \xi_{ij}^s \pi_{Rj}^s \right] \mathbf{1}[k=s] - \left(\sum_j \xi_{ij}^k \pi_{ij}^k \right) \tilde{L}_i^s}{1 + (\sigma_s - 1) \sum_j \xi_{ij}^s (1 - \pi_{ij}^s) - \xi_{ii}^s} & k = 1, 2 \\ \sum_{k=1,2} L_i^k \tilde{L}_i^k = 0 \end{cases} \quad (2.20)$$

where $\mathbf{1}[k = s]$ is the indicator function. Therefore, the first term is only active for sectors that experience a direct tariff shock. The equation in the second row is the total derivative of the labor market clearing condition equation 2.10. Solving

¹⁹Autor et al. (2013) model has two manufacturing sectors as well.

the 3×3 system in equation 2.20, the **partial-equilibrium wage elasticity** in region i equals:

$$\tilde{w}_i^{PE} = g_i^1 \frac{L_i^1}{L_i} (\sigma_1 - 1) \left(\sum_{j \neq R} \xi_{ij}^1 \pi_{Rj}^1 \right) \quad (2.21)$$

where $g_i^1 = \frac{a_i^2}{a_i^2 b_i^1 (L_i^1/L_i) + a_i^1 b_i^2 (L_i^2/L_i)}$, $a_i^k = \sum_j \xi_{ij}^k \pi_{ij}^k$, and $b_i^k = 1 + (\sigma_k - 1) \sum_j \xi_{ij}^k (1 - \pi_{ij}^k) - \xi_{ii}^k$.

The partial equilibrium wage elasticity is increasing in $\sum_{j \neq R} \xi_{ij}^1 \pi_{Rj}^1$, which captures the direct effect of the shock. In addition, the elasticity is increasing in L_i^1/L_i , the share in total employment of the sector affected by the shock. The higher is this share, more workers reallocate from sector 1 to sector 2 after the shock. Finally, the elasticity is decreasing in $\sum_j \xi_{ij}^1 \pi_{ij}^1$ and increasing in $\sum_j \xi_{ij}^2 \pi_{ij}^2$. These terms multiply the elasticity of sectoral employment in equation 2.19 and, since employment and varieties are linearly related, they determine the magnitude of the impact due to the change in the number of varieties in each sector. When there is a decrease in tariffs in sector 1, the number of varieties in the sector decreases, which decreases competition in the sector, dampening the initial impact of tariffs. To satisfy labor market clearing, the number of varieties in sector 2 has to increase and this is the reason why the partial wage elasticity is increasing in $\sum_j \xi_{ij}^2 \pi_{ij}^2$.

With data on sectoral elasticities of substitution, expenditure shares and revenue shares, the partial elasticity of regional wages with respect to tariffs can be computed for every region without the need of solving the full model. With data on tariff changes, wage changes can be predicted. However, this is not the approach

taken in the empirical literature which resorts to a regression of wage changes on an employment-weighted average of shocks.

2.4.2 From partial-equilibrium elasticities to employment-weighted averages of shocks

As I reviewed in Chapter 1, several authors use a regression approach based on variation in regional *trade exposure* to assess the impact of trade shocks on regional wages. I formulate a set of assumptions on expenditure and revenue shares that, starting from the expression of a partial equilibrium elasticity, allow Autor et al. (2013) to arrive to an estimating equation.²⁰ These assumptions pertain to endogenous objects of the model, like expenditure and revenue shares. I show that assumptions on fundamentals like trade costs can yield the same outcomes for expenditure shares but not for revenue shares.

From equation 2.21 I define the *partial-equilibrium* wage change due to a tariff shock in sector k as

$$\hat{w}_i^{PE} = g_i^k \frac{L_i^k}{L_i} (\sigma_1 - 1) \left(\sum_j \pi_{Rj}^k \xi_{ij}^k \right) \hat{\tau}^k \quad (2.22)$$

where $\hat{x} = dx/x$ denotes the proportional change in variable x . The partial-equilibrium wage change, \hat{w}_i^{PE} , is equal to the partial wage elasticity multiplied by the tariff change.

Assumption 1. $g_i^k = \alpha$, $i \in Home$

²⁰Some of the assumptions are not made explicit in their paper.

g_i^k is referred by Autor et al. (2013) as a “general equilibrium adjustment factor”. Assumption 1 states it does not vary across regions in the Home country. When tariff shocks occur in multiple sectors, it is also required that it does not vary across sectors either.

Assumption 2. $\pi_{Rj}^k = \pi_{RH}^k, \quad j \in Home$

Assumption 2 states expenditures shares π_{Rj}^k are constant across domestic regions ($j \in H$). This implies that every domestic region (j) allocates the same share of expenditure in sector k to goods proceeding from the ROW.

Assumption 3. $\xi_{ij}^k = \xi_{iH}^k, \quad i, j \in Home$

Assumption 3 states revenue shares ξ_{ij}^k are constant across domestic destinations ($j \in H$). This implies that the share of revenue received by producers in region i in sector k that corresponds to goods sold to region j is the same across all domestic destinations j .

Assumptions 1 to 3 yield:

$$\begin{aligned} \hat{w}_i &= \alpha \frac{L_i^k}{L_i} \pi_{RH}^k \xi_{iH}^k (\sigma_1 - 1) \hat{\tau}^k \\ &= \alpha \frac{L_i^k}{L_i} \frac{X_{RH}^k}{X_H^k} \frac{R_{iH}^k}{R_i^k} (\sigma_1 - 1) \hat{\tau}^k \end{aligned} \tag{2.23}$$

where R_{iH}^k is the revenue received by region i in sector k that corresponds to goods sold to any region $j \in Home$ and R_i^k is total revenue received by region i in sector k .

Assumption 4. $R_{iH}^1/X_H^1 \approx L_i^1/L_H^1$, $i \in Home$

The share of region i in total Home purchases in industry k can be approximated by the share of region i in Home employment in industry k

Two properties of the model are: $\frac{R_{iH}^k}{R_i^k} = \frac{q_{iH}^k}{q_i^k}$ (the revenue shares are equal to the shares of shipments) and L_i^k/q_i^k equals a constant. These properties in addition to the previous assumptions 1 to 4 yield:

$$\hat{w}_i = -\tilde{\alpha} \frac{L_i^k}{L_H^k} \frac{X_{RH}^k \hat{A}_R^k}{L_i^k} \quad (2.24)$$

where, using Autor et al. (2013) notation, $\hat{A}_R^k = -(\sigma_1 - 1) \hat{\tau}^k$ is the change in the export supply capability of the ROW due to a change in the tariff Home applies to ROW goods in sector k .²¹

Finally, assume:

Assumption 5. $\Delta X_{RH}^k = X_{RH}^k \hat{A}_R^k$

The change in Home imports in sector k equals initial imports times the change in the export supply capability of the ROW. This proportionality between changes in tariffs and changes in expenditures does not hold in the model and has to be assumed. Using the gravity equation $X_{ij}^k = n_i^k \left(\frac{p_i^k d_{ij}^k \tau_{ij}^k}{P_j^k} \right)^{1-\sigma_k} \mu_k E_j$ we can see that keeping prices, varieties and expenditure constant it is the case that $\hat{X}_{ij}^k = -(\sigma_k - 1) \hat{\tau}_{ij}^k$, but these variables vary endogenously after the change in tariffs. Assumption 5 is needed even if assumptions 1 to 4 hold. To see why, consider

²¹The model can accommodate different sources of shocks like: changes in tariffs, sectoral employment, tariffs or productivity of the ROW so the general formula used in ADH for this variable is: $\hat{A}_R^k = \hat{L}_R^k - (\sigma_k - 1) (\hat{w}_R + \hat{\beta}_R^k + \hat{\tau}^k)$

the case where the home country consists of one region, so there is no within country heterogeneity in revenue and expenditure shares. Even in this extreme case, the gravity equation implies that changes in imports are not linearly related to changes in tariffs.²²

Assumptions 1 to 5 yield the following expression:

$$\hat{w}_i = -\tilde{\alpha} \left(\frac{L_i^k}{L_H^k} \frac{\Delta X_{RH}^k}{L_i} \right) \quad (2.25)$$

To arrive to an estimating equation it is necessary to assume that there are other factors unobservable to the econometrician that affect wages at the same time as the shock occurs, which are captured by a stochastic error term (ε_i):

$$\hat{w}_i = \alpha_0^{emc} + \alpha_1^{emc} EMC_i^k + \varepsilon_i \quad (2.26)$$

where $EMC_i^k = \frac{L_i^k}{L_H^k} \frac{\Delta X_{RH}^k}{L_i}$ is an employment-weighted change in imports, as defined in Autor et al. (2013) .

Equation 2.26 can be generalized to tariff changes in several sectors by replacing EMC_i^k with $EMC_i = \sum_k \frac{L_i^k}{L_H^k} \frac{\Delta X_{RH}^k}{L_i^k}$, which is a weighted sum of sectoral trade shocks. The weights are employment shares $\frac{L_i^k}{L_H^k}$.²³

²²Autor et al. (2013) are not explicit about Assumption 5 . However, it is needed to go from the formula for \hat{A}_R^k to an estimating equation based on changes in import penetration as they estimate.

²³The system in equation 2.20 can be generalized to accommodate shocks in multiple sectors. As I explain when I state assumption 1, when the shocks occur in several sectors, the general equilibrium adjustment factor, g_i^k , has to be equal not only across regions but also across sectors. Under this modified assumption and assumptions 2 to 5, it can be shown that equation 2.26 holds. See Appendix A.2 for a proof.

Other authors, like Kovak (2013) and Topalova (2010) have used trade liberalization episodes, so the shock is directly observable. In Autor et al. (2013) this is not the case since the source of the shock is an unobservable change in foreign productivity. When the shock is observable, this version of equation 2.26 can be estimated:

$$\hat{w}_i = \alpha_0^{etc} + \alpha_1^{etc} ETC_i^k + \varepsilon_i \quad (2.27)$$

where $ETC_i = \sum_k \frac{L_i^k}{L_i^k} \hat{\tau}^k$, is an employment-weighted tariff change in region i .

The assumptions I stated involve endogenous variables like expenditure and revenue shares. An important question is if it is possible to arrive to the same outcomes making assumptions on fundamentals of the model instead, such as trade costs. The following proposition states that assuming that trade costs do not vary across regions at Home does yield constant trade shares.

Definition 2. Trade costs in sector k are **homogeneous** if: the cost of shipping good k from region i to region j is the same for all Home regions i, j (this is, $\kappa_{ij}^k = \kappa_H^k$ for all i, j) and the cost of shipping good k from the Rest of the World to a Home region i is the same for all Home regions (this is, $\kappa_{Ri}^k = \kappa_R^k$ for all i).

Proposition 3. *Suppose trade costs in sector k are homogeneous, then Assumption 2 holds. This is, $\pi_{Ri}^k = \pi_{RH}^k$. In addition, $\pi_{ij}^k = \pi_{ih}^k$, $i, j, h \in Home$.*

Proof. Under homogeneity of trade costs, the expenditure share of region i on ROW goods in sector k is given by:

$$\pi_{Ri}^k = \frac{L_R^k \left(w_R \kappa_R^k / \beta_R^k \right)^{1-\sigma_k}}{\sum_{j \in Home} L_j^k \left(w_j \kappa_H^k / \beta_j^k \right)^{1-\sigma_k} + L_R^k \left(w_R \kappa_R^k / \beta_R^k \right)^{1-\sigma_k}} \quad (2.28)$$

which is the same for any region $i \in Home$. So $\pi_{Rj}^k = \pi_{RH}^k$.

In addition, homogeneity implies the expenditure share of region j on region i 's goods in sector k is given by:

$$\pi_{ij}^k = \frac{L_i^k \left(w_i \kappa_H^k / \beta_{Ri}^k \right)^{1-\sigma_k}}{\sum_{h \in Home} L_h^k \left(w_h \kappa_H^k / \beta_h^k \right)^{1-\sigma_k} + L_R^k \left(w_R \kappa_R^k / \beta_R^k \right)^{1-\sigma_k}} \quad (2.29)$$

which is the same for any region $j \in Home$. So $\pi_{ij}^k = \pi_{ih}^k$. \square

Proposition 1 states that assuming homogeneous trade costs across domestic regions, implies that expenditure shares are the same for all domestic regions. Thus, if region i imports 25% of sector k goods from region j , then region h also imports 25% of sector k goods from region j . However, the following proposition shows homogeneity of trade costs is not enough to ensure the same result holds for revenue shares.

Proposition 4. *Homogeneity of trade costs in sector k is not sufficient for Assumption 3 to hold.*

Proof. The ratio of revenue shares in sector k in region i from sales to regions j and h is given by:

$$\frac{\xi_{ij}^k}{\xi_{ih}^k} = \frac{w_j L_j}{w_h L_h} \quad (2.30)$$

where I used that, by proposition 1: $\pi_{ij}^k = \pi_{ih}^k$. Assumptions on trade costs are not sufficient to ensure that $\xi_{ij}^k = \xi_{ih}^k$ since this would require that $w_j L_j = w_h L_h$. \square

As the proof of Proposition 2 shows, equality of revenue shares ($\xi_{ij}^k = \xi_{ih}^k$) requires equality of nominal incomes ($w_j L_j = w_h L_h$). Aside from knife-edge combinations of productivity and size that would yield this result, assuming equal size of regions is not enough since wages would vary across regions due to differences in productivity. Assuming productivities do not vary across regions, would yield that labor shares do not vary across regions, which is precisely the type of variation used in the exposure approach.

2.5 An illustration using simulations

In this section I explore how accurately partial elasticities measure the impact of trade shock on local wages. In other words, I analyze how costly is it to ignore spillovers between regions. The difference between partial and general equilibrium elasticities depends on the specific parameter configuration of the Home economy and the Rest of the World. Therefore, I compare partial equilibrium and general equilibrium elasticities under different initial geography configurations.

For simplicity, I analyze the case of two regions in the Home country and a region called Rest of the World, denoted $i = \{1, 2, R\}$, and two sectors, denoted by $k = \{1, 2\}$. My procedure is the following. Given initial values of parameters $\{L_i, \sigma_k, \beta_i^k, \kappa_{ij}^k\}$, I solve the model for endogenous wages and sectoral employment: $\{w_i, L_i^k\}$. With this vector I calculate expenditure and revenue shares that, together

with initial values of parameters, are sufficient to obtain the partial-equilibrium elasticities $\{\tilde{w}_i^{PE,k}\}$ using equation 2.21. Then, I shock the Home economy with an aggregate tariff reduction. I solve the model again and I obtain the new values: $\{(w_i)', (L_i^k)'\}$. I use the new wages the model predicts to calculate the general-equilibrium elasticity, which I compare to the partial equilibrium elasticity.

A comparison between predicted wages using partial elasticities and predicted wages using the exposure approach is not possible in this simple example since the exposure approach requires running a regression, and that is not possible with only two regions in the Home country.²⁴ However, we can compare the differences in employment shares with the differences in partial elasticities. I show that partial equilibrium elasticities can be different in the two regions at home while labor shares are the same. This suggests there are cases in which labor shares cannot capture the variation in the actual partial effect of the shock.²⁵

The reduced form approach focuses on regional variation in the sectoral composition of employment as the only factor leading to differences in the sensitivity of shocks. This approach neglects other factors, like distance to other markets as possible explanations. If trade costs vary with distance, then distance affects wages and expenditure shares, and therefore, general equilibrium and partial equilibrium elasticities. To get a sense of how distance affects these elasticities, I start with symmetric regions and progressively increase the distance of region 2 to both region 1

²⁴The purpose of a regression approach is to evaluate how much of the variation in wage changes can be explained by the variation in initial conditions as captured, for example, by employment composition. With only two regions, there is not enough variation in initial conditions. I perform this type of exercise in sections 3.3.2 and 3.3.3 when I take the model to the Brazilian data.

²⁵In Appendix A.4 I provide a formal proof that a regression of wage changes on employment exposure gives a biased estimator of the partial-equilibrium elasticity.

and to the ROW. This makes trade between region 2 and the other regions costlier. For each level of distance, I solve for wages and sectoral employment and then I shock trade costs to obtain the general equilibrium elasticity and compare it to the partial-equilibrium elasticity. I do this both for aggregate shocks (that affect both regions at Home) and for region-specific shocks that affect one region at a time.

I report the parameter values I use in the simulations in Appendix A.3. I start with symmetric regions that are at a distance of 1 km. from one another and then progressively increase distance to a maximum of 3 km. The cost of shipping goods within region 2 is kept constant. In addition, both sectors are symmetric, the elasticities of substitution, Cobb-Douglas shares, parametrization of trade costs, and productivities are the same for all region-sectors. This implies that as the distance of region 2 to the other regions increases, equilibrium wages change but not equilibrium sectoral employment.

Figure 2.1 shows the response to a national-level shock in sector 1. The shock consists of a tariff reduction of 5% in both regions 1 and 2. When regions are symmetric (i.e. when distance on the x-axis equals 1 km.), regions 1 and 2 have the same values of the wage elasticity. As region 2 becomes more remote, it has a higher actual elasticity. This result goes against the intuition in some reduced form analyses that incorporate distance as a factor that dampens the impact of trade shocks (see for example, Chiquiar et al. (2015)). In addition, the figure shows that partial elasticities do not provide a correct ranking of regions in terms of sensitivity to the shock when distance is greater than 1.5. In this interval, the true elasticity is higher for region 2 while the partial elasticity is higher for region 1. Why does the more

remote region has a higher sensitivity to tariff changes? The intuition behind this result can be understood with Figure 2.2: the difference between regional elasticities is smaller in the left panel figure. This means that region 2 is more affected by wage changes in region 1 than region 1 is by wage changes in region 2. Also, as region 2 becomes more remote the elasticities to a shock in region 1 increase while the elasticities to a shock in region 2 decrease. The reason is that as region 2 becomes more remote the equilibrium relative wage of region 1 versus region 2 increases, so region 1 becomes a bigger market.

In the previous exercise, the employment composition does not vary as region 2 becomes more remote: the share of sector 1 in total employment is $1/2$ in both regions. Therefore, if we rank the regions' sensitivity to shocks based on employment composition we obtain that both regions are equally sensitive while the simulation of the model shows that region 2 is, in fact, more sensitive.

The previous example has parameter values such that regions are diversified (i.e. they produce both goods 1 and 2). It is interesting to analyze cases where the regions are completely specialized in one sector, because this shows there is still a spillover to the region that does not produce the good affected by the shock. Figure 2.3 shows a completely specialized home economy. Region 1 is 10 times more productive in good 1 than in good 2, and region 2 is 10 times more productive in good 2.²⁶ An approach based on employment composition, would conclude that region 2 is not exposed at all to the shock. The figure shows, as expected, that

²⁶I choose these values to guarantee that for any value of distance in the x-axis, region 1 is completely specialized in good 1 while region 2 is completely specialized in good 2.

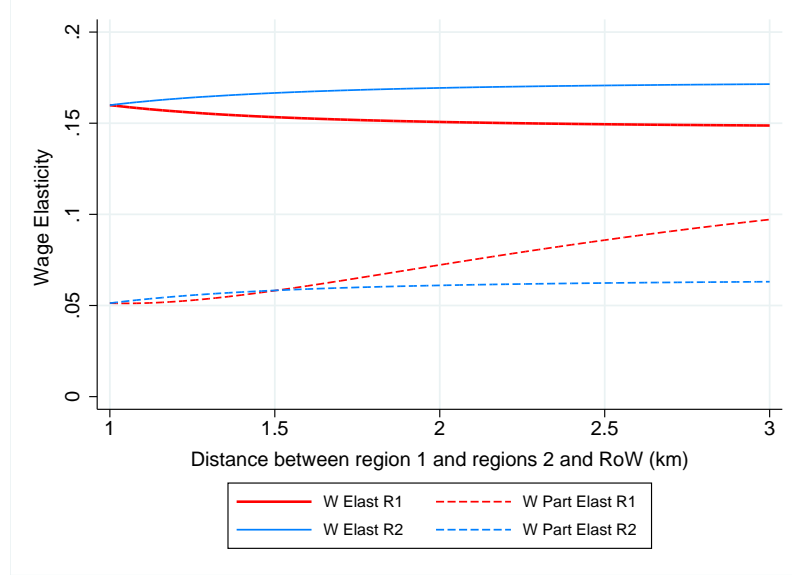
region 2 is less affected than region 1. However, its wage elasticity is far from zero, and the difference between both elasticities decreases with the remoteness of region 2 because, as I explained before, region 1 becomes a more important market.

2.6 Conclusion

In this section I laid out a standard trade model to understand the mechanics of wage changes when regions in a Home country trade with each other and with the ROW. I decomposed wage changes after a sector-specific trade cost shock in several channels and I stated the assumptions that are necessary to go from this decomposition to an employment-composition based measure. Importantly, I showed these assumptions are incompatible with assumptions on trade costs. Finally, I used simulations to show that reduced form measures can yield, under certain parameter configurations, an incorrect ordering of which region is more affected.

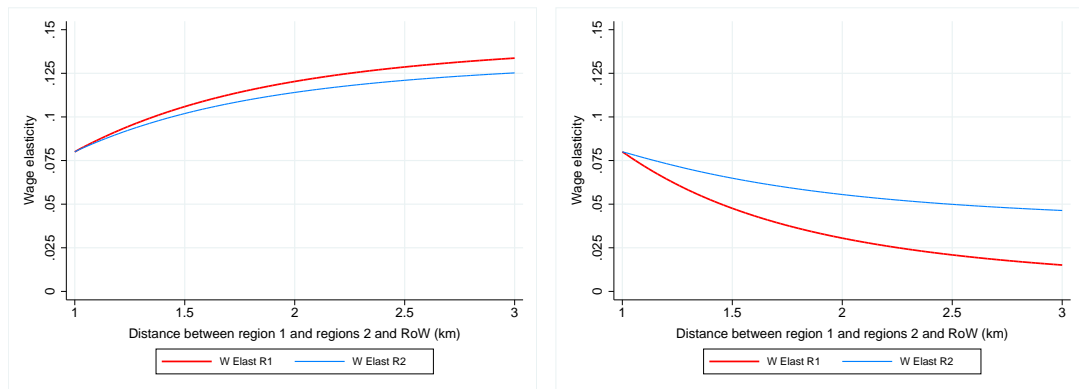
2.7 Figures

Fig. 2.1: Simulated wage response to a national shock to sector 1



Notes: Wage elasticity as a result of a 5% decrease in Home tariffs applied to ROW in sector 1 as a function of the distance between region 2 and regions 1 and ROW (in km). All regions are completely diversified. The ratio of employment in sector 1 to employment in sector 2 is equal to 1/2 in both Home regions.

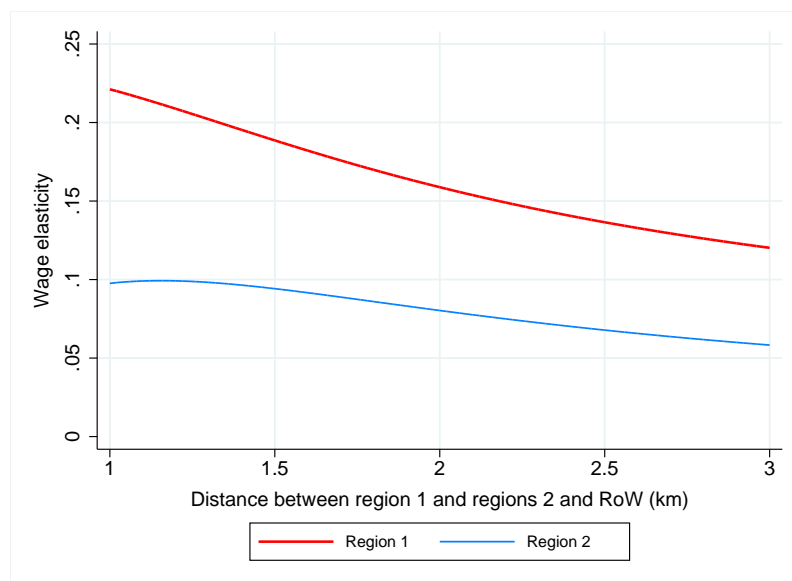
Fig. 2.2: Simulated wage response to region-specific shocks to sector 1



Notes: Wage elasticity to a 5% decrease in region 1 tariffs applied to ROW in sector 1 as a function of the distance between region 2 and regions 1 and ROW (in km).

Notes: Wage elasticity to a 5% decrease in region 2 tariffs applied to ROW in sector 1 as a function of the distance between region 2 and regions 1 and ROW (in km).

Fig. 2.3: Simulated wage response to a shock to sector 1 under complete specialization



Notes: Wage elasticity a result of 5% decrease in Home tariffs applied to ROW in sector 1 as a function of the distance between region 2 and regions 1 and ROW (in km). Region 1 is completely specialized in sector 1 and region 2 is completely specialized in sector 2.

Chapter 3: Trade shocks, linkages, and wages in Brazil

3.1 Introduction

The simulations in Chapter 2 suggest that employment composition can be a poor predictor of which regions in a country are more affected by a nationwide trade shock. However, we may wonder how problematic this is for an actual country. To answer this, I calibrate the model to the Brazilian economy in 1991, at the beginning of a massive trade liberalization process. In section 3.3 I perform a series of counterfactual exercises using the structure of the model and I show that, in fact, employment composition explains a relatively small share of the variation in counterfactual wage changes. Therefore, I propose a linkage exposure measure derived from the model and I show it explains a significantly higher share of the variation in counterfactual wage changes. In section 3.4 I propose an estimation strategy closely related to the model and I show trade exposure is a valid instrument for changes in domestic market access. Next, I estimate the elasticity of wages to domestic market access and I compare the predictions of different reduced form measures with the predictions using the estimated elasticity and counterfactual market access changes.

3.2 Data sources and measurement

This section describes the data sources I use for the quantitative exercises and the econometric estimations in this chapter and in Chapter 4. The reader can find more details about each variable in Appendix B. The main unit of analysis is a Brazilian microregion, which is a subdivision of Brazilian states defined by the Brazilian Institute of Geography and Statistics (IBGE). A microregion groups together several municipalities in proximity that share common social and economic characteristics. There are a total of 494 microregions in Brazil that can be defined in a geographically consistent way across the 1991 and 2000 Demographic Censuses using the mapping by Kovak (2013). Throughout the analysis, I also refer to microregions simply as “regions”. Table 3.1 shows there is a large degree of heterogeneity across microregions in area, distance to sea ports, employment, and wages. Figure 3.1 shows there is also high variation in the shares of manufacturing employment across microregions.

Wages and employment data are from the Brazilian Demographic Censuses of 1991 and 2000, conducted by the Brazilian Institute of Geography and Statistics (IBGE).^{1,2} Individuals are asked to report, among other variables, the monthly

¹These datasets are also employed by Kovak (2013), Dix-Carneiro and Kovak (2014), and Adao (2015), among others.

²Another source for data on earnings is the household annual survey (Pesquisa Nacional de Amostras por Domicilio, PNAD), which is also conducted by the IBGE. I do not use this dataset since its sample size is not large enough to conduct an analysis at the local labor market level. A solution to the small sample size would be to use several cross sections of states. I do not follow this approach since it would imply using short-run variation in tariffs. Although in section 3.4.1 I argue that tariff changes between the beginning and the end of trade liberalization were independent of political economy considerations, this argument may not apply to short run fluctuations in tariffs. However, using long-run changes has the downside that other aggregate shocks can be confounded with tariffs. I address these risks in section 3.4.1.

earnings from their main job and weekly hours worked, their industry of employment, and the municipality where they live.³ With this data, I assign individuals to regions and I obtain L_i^k as the sum of yearly hours worked across all workers in region i and sector k and w_i as the average hourly wage across all workers in region i .⁴

The industry classifications are not uniform across the different data sources, so I use several mappings to arrive to my own classification which consists of one non-traded sector and 12 traded sectors. Both the crosswalks and my classification are described in detail in Appendix B.

Average *ad-valorem* tariffs applied by Brazil at the sectoral level for the period 1990-1998 are from Kume et al. (2000).⁵ Tariffs were reduced substantially between 1987 and 1998, as Figure 3.2 shows. The high levels before 1990 should be interpreted with caution since tariff redundancy was very high due to special customs regimes that waived or reduced import duties for particular goods, as Kume et al. (2000) explains.⁶ Trade liberalization reduced both tariff levels and their dispersion, as Table 3.2 shows.⁷

³Since the industry classification varies across censuses, I use the crosswalk in Kovak (2013) as described in Appendix B.

⁴My measure of regional wages differs from Kovak (2013), since he uses region fixed effects from a Mincer regression. In section 3.4.4 I report results using his measure as a robustness check.

⁵This dataset is also used by Gonzaga et al. (2006) and Kovak (2013).

⁶The authors also report that during 1990 the list of forbidden imports was abolished along with the majority of the special customs regimes. Simultaneously, tariff levels were adjusted to provide equivalent levels of protection as those existing before the reform. By the middle of the year, tariff redundancy was eliminated so tariffs became the main instrument to protect the local industry.

⁷At the tariff line level, the simple average tariff was reduced from 50.6% (1988) to 14.2% (1993), and the standard deviation was reduced from 26.2 to 9.5 Kume et al. (2000)

I obtain the elasticities of substitution between varieties in each sector, σ_k , from the data in Broda and Weinstein (2006).⁸ As Table B.1 shows, they range from 2.53 in Paper to 3.65 in Chemicals without considering Footwear and Leather, which has a value of 14.64.

Trade flows between Brazilian states at the sector level are from Vasconcelos and Oliveira (2006) and trade flows between Brazil and the ROW at the sector level are obtained from the AliceWeb System (supported by the Secretary of International Trade, SECEX). The interstate trade data corresponds to the year 1999. Unfortunately, there is no data on interstate trade at the sector level for previous years.

I calculate geographic distances between regions in Brazil and between regions and their nearest sea port using the Haversine distance formula. The formula for the distance of a region to itself, d_{ii} , is described in Appendix B.1.5.

I obtain ROW wages and sectoral hours worked in 1991 and 2000 using value added at the sector level for 50 countries from UNIDO Industrial Statistics, and total hours worked and GDP from Penn World Tables version 8.1.

To calculate Cobb Douglas shares I first obtain the share of non-traded sector in total value added, μ_0 , from the Brazilian National Accounts (IBGE), which is equal to 0.7 in 1991 and 0.76 in 2000.⁹ I calculate the Cobb Douglas shares for traded sectors ($k = 1...K$) as the domestic absorption in a sector divided by total value

⁸Appendix B.1.3 and B.1.7 contain details about the industry crosswalks between the data in Broda and Weinstein (2006) and the Brazilian Census.

⁹Since in the model the share of labor in the non-traded sector is constant across regions and equal to μ_0 (see Section 2.2, Equilibrium), I adjust manufacturing hours in all Brazilian regions and in the ROW so that $L_i^0 = \mu_0 L_i$.

added, using the formula: $\mu_{k,BRA} = (1 - \mu_0) \frac{VA_{BRA}^k + M_{BRA}^k - X_{BRA}^k}{\sum_{k=1}^{12} VA_{BRA}^k}$, where $VA_{BRA}^k = \sum_{i \in BRA} w_i L_i^k$ is the value added by labor in sector k in Brazil and is defined as in the model and calculated using the Census data, and M_{BRA}^k and X_{BRA}^k are Brazilian imports and exports in sector k , obtained from data source mentioned above and adjusted by the share of the value of production accruing to wages in each sector.¹⁰ I obtain the ROW Cobb-Douglas shares in traded sectors in a consistent way so that markets clear, using the formula: $\mu_{k,ROW} = (1 - \mu_0) \frac{VA_{ROW}^k + M_{ROW}^k - X_{ROW}^k}{\sum_{k=1}^{12} VA_{ROW}^k}$, where $M_{ROW}^k = X_{BRA}^k$ and $X_{ROW}^k = M_{BRA}^k$. The Cobb-Douglas shares for Brazil and for ROW are displayed in Table B.1 and the sectoral share of imports and exports of value added are shown in Table B.2.

3.3 Counterfactuals

I perform several counterfactual exercises for Brazil. These experiments allow me first, to quantify the importance of the different channels through which a reduction in tariffs can affect regional wages by using the decomposition derived in section

¹⁰Since in the model the only factor of production is labor and there are no intermediates, I measure the value of production using the data on labor income from the Census. Given that my calibration targets the observed ratio of imports to value added (which equals the value of production in the model), I calculate the Cobb-Douglas shares so that they are consistent with total absorption in the data. Otherwise, markets would not clear. Ideally, I should use only imports and exports of final goods to obtain the Cobb Douglas shares. However, there is no data on Brazilian exports and imports discriminated by final and intermediate goods at the level of sectoral aggregation I use. Therefore, I adjust imports and exports by multiplying them by the share of the total value of production accruing to wages in each sector (this share is obtained from the input-output tables by IBGE). I also adjust sector level exports to eliminate the aggregate trade deficit, since in the model trade is balanced. Other papers in the literature either use data on intermediate exports and imports, such as the work by Costinot and Rodríguez-Clare (2014) who use OECD data, or they model intermediates and use input-output data, such as Caliendo and Parro (2014) and Caliendo et al. (2015).

2.4, and second, to evaluate the capacity of different exposure measures to predict wage changes.

3.3.1 Approach

To generate the counterfactuals I calibrate the parameters of the model to match the initial equilibrium of the Brazilian economy in 1991. Next, I apply a trade cost shock and I solve for regional nominal wages and sectoral employment in the new equilibrium. I also solve for the change in the price index, which allows me to obtain real wage changes. In several applications, the calibration of the model is not a necessary stage for obtaining counterfactual wages. For example, the method of solving the model in changes, developed by Dekle et al. (2008), would only require data on initial wages, sectoral employment, and expenditure shares (π_{ij}^k) to solve for regional wage changes. However, there are many interesting applications in which the full matrix of expenditure shares is not available. This typically occurs when bilateral trade is not observed at the level of the unit of analysis. For instance, in the case of Brazil bilateral trade is available at the level of states but not at the level of smaller units like microregions.¹¹ When this is the case, we can use the structure of the model and a parametrization of trade costs to calibrate unobserved parameters and predict trade between units. A non-exhaustive list of authors employing procedures similar to the one I use is: Fan (2015), Bartelme (2014), and Monte et al. (2015).

¹¹For the level of disaggregation in the empirical application of this paper (12 industries), there is no data on production at the state level. Therefore, the full matrix of expenditure shares is not observed at the state level.

I use 1991 as the initial year due to data availability. As I explained in section B.1, the process of trade liberalization began in 1988 but it was not until 1990 that tariff changes started to reflect actual changes in protection. Ideally, 1990 should be the base year for the calibration. If observed wages in 1991 already reflect some anticipation of future tariff reduction then, the counterfactual exercise in which I reduce tariffs by the magnitude observed in 1991-1995, would overstate the change in wages. Therefore, the absolute magnitude of wage changes should be interpreted with caution. However, the main purpose of this section is to analyze how much of the regional variation in wage changes across regions is explained by exposure measures. Therefore, as long as anticipatory effects affect all regions in the same way, the main results in this section would be unaffected.

Equations 2.12 and 2.10, together with equation 2.11 completely characterize the equilibrium of the model. Given the values for the trade costs (κ_{ij}^k), region-sector specific productivities (β_i^k), and Cobb Douglas shares (μ_k), wages and sectoral employment in every region can be solved for. The solution algorithm is described in Appendix A.5. In the next paragraphs I describe the procedure I use to calibrate productivities and to parametrize trade costs.

I use the market clearing conditions in equation 2.12 together with 2.11 to solve for $N \times S$ productivities (β_i^k) in 1991. For each sector there is a set of N equations. Given regional wages (w_i) in 1991, hours worked in each region and sector (L_i^k) in 1991, elasticities of substitution (σ_k), Cobb-Douglas shares (μ_k), and a

parametrization of trade costs (κ_{ij}^k) as a function of distance, each set of N equations can be solved for a unique vector $\{\beta^k\}_{i=1}^N$ up to a normalization.¹²

I obtain regional wages, region-sector employment, elasticities of substitution, and Cobb-Douglas shares from the data as described in section 3.2. Regarding trade costs, I follow the standard gravity literature by parametrizing trade costs between regions as a function of distance. For any two regions i, j in Brazil, the domestic trade cost has the following functional form $D_{ij}^k = (dist_{ij})^{\delta_k}$ where $dist_{ij}$ is the bilateral distance between regions i and j .¹³ I estimate the elasticity of domestic trade costs with respect to distance, δ , with a gravity equation for each sector using data on trade flows between Brazilian states. I do not use the gravity equation to estimate an international border effect but rather, I calibrate the border effect together with productivities as I describe below. Taking logs of equation 2.9 in stochastic form yields for each sector k

$$\ln X_{ij}^k = exp_i^k + imp_j^k + \psi_k \ln dist_{ij} + \varepsilon_{ij}^k \quad i, j \in BRA, k = 1 \dots K \quad (3.1)$$

where $exp_i^k = \ln \left[n_i^k (p_i^k)^{1-\sigma_k} \right]$ is an exporter fixed effect, $imp_j^k = \ln \left[E_j^k (P_j^k)^{\sigma_k-1} \right]$

is an importer fixed effect, $\psi_k = (1 - \sigma_k) \delta_k$, and ε_{ij}^k is a stochastic error term. With

¹²I refer the reader to the proof of Proposition 3 in Monte et al. (2015).

¹³This parametrization assumes no border effect between Brazilian states. Since I use this parametrization to predict trade between microregions, including a border effect in the trade cost would imply assuming that the state level border effect is equal to the microregion level border effect. Although there is evidence that there are border effects at the state and province level for the U.S. and Canada (Anderson and Van Wincoop (2003), Anderson and Yotov (2010)), there is not equivalent evidence for trade at a disaggregated level, such as a microregion. Therefore, I exclude within country border effects from my parametrization.

data on σ_k , and estimates of ψ_k , the distance elasticity $\delta_k = \psi_k / (1 - \sigma_k)$ can be recovered. The origin fixed effect, exp_i^k , controls for region i 's number of varieties and Mill prices in sector k . The destination fixed effect, imp_j^k , controls for region j 's aggregate expenditure in sector k , E_j^k , and its price index in sector k , P_j^k . I estimate ψ_k for each sector k using the Poisson Maximum Likelihood estimator developed by Silva and Tenreyro (2006) and I cluster standard errors by pairs of states. The dependent variables are interstate trade flows in 1999. Since I use this parametrization of trade costs to predict trade shares in 1991, the maintained assumption is that between 1991 and 1999 the cost of shipping goods within Brazil did not change.¹⁴ The first column of Table 3.3 shows the coefficient of the log of distance, ψ_k , which ranges from -0.35 in Footwear and Leather Apparel to -2.2 in Mining. The second column reports the implied value of δ_k , the distance elasticity in the trade cost function, given the sectoral elasticities of substitution displayed in Table B.1.¹⁵

With the estimated domestic distance elasticity we can predict trade costs between any pair of Brazilian regions. To predict the trade cost between a Brazilian region and ROW, I assume international trade costs for a Brazilian region i have the following form: $I_{i,ROW}^k = (distport_i)^{\delta_k} \chi^k$, where $distport_i$ is the distance between region i and the nearest sea port in Brazil and χ^k is equal to one plus the tariff equivalent of all border related costs, including tariffs. Note that both distance and the border effect are symmetric, so $I_{i,ROW}^k = I_{ROW,i}^k$. Since goods shipped to the port

¹⁴This assumption is supported by empirical studies that find that distance elasticities are stable over time. See, for example, Anderson and Yotov (2010) on Canadian province trade.

¹⁵The number of observations differs across industries because not all states export all goods. Since in the data there are 27 importing states but only 21 exporting states, the maximum number of observation is $546 = 21 \times (27 - 1)$. See Appendix B for more details on the interstate trade data.

are transported by land, I assume the distance elasticity is δ_k , the value I obtained by estimating the gravity equation.

I calibrate a border effect for each sector χ^k to exactly match the Brazilian observed ratio of imports to value added in each sector in 1991.¹⁶ Note that the border effect could be estimated with a gravity equation by including trade flows between states and ROW and a border dummy (Anderson and Van Wincoop (2003) and Anderson and Yotov (2010)). However, the predicted import penetration at the country-sector level using an estimated border effect would not necessarily match the one observed in the data. Since the effect of trade on wages is driven by expenditure shares, it is important that expenditure shares match the data as closely as possible. The calibrated border effect factors, χ_k , are shown in Table B.3. These are higher than other estimates from the literature, such as those in Anderson and Yotov (2010). The reason is that I measure distance to the ROW as the distance to the nearest sea port, so I do not take into account the distance from the sea port to destination countries. If I only used the distance coefficient, this would predict a much higher level of imports (and exports) than is observed in the data. Therefore, to match the observed ratio of imports to value added, the border effect has to be higher.

I study asymmetric trade shocks that decrease the cost of shipping goods from ROW to Brazil. In the first set of experiments, I decrease trade costs $\kappa_{ROW,BRA}^k$ in a one sector at a time by 1%. In the second experiment, I decrease trade costs in

¹⁶By doing this, I also match the ratio of exports to value added since Cobb Douglas shares are defined by $\mu_k = \frac{VA^k + M^k - X^k}{\sum_k VA^k}$.

all sectors at the same time by the amount implied by the tariff reduction that took place in Brazil between 1991 and 1995, the period of trade liberalization. For this purpose I use the formula

$$\kappa_{ROW,BRA,95}^k = \kappa_{ROW,BRA,91}^k \frac{(1 + t_{95}^k)}{(1 + t_{91}^k)} \quad (3.2)$$

where $\kappa_{ROW,BRA,91}^k$ are initial trade costs in 1991, calculated as described previous paragraphs using the formula $\kappa_{ROW,BRA,91}^k = (distport_i)^{\delta_k} \chi_{91}^k$, and t_{91}^k and t_{95}^k are the ad-valorem Brazilian tariffs in sector k in 1991 and 1995 respectively. Equation 3.2 implies that a 1% reduction in the Brazilian tariff factor in sector k , yields a reduction of 1% in the trade cost, $\kappa_{ROW,BRA}^k$. The corresponding reduction in trade flows is, by equation 2.9, of $(1 - \sigma_k)\%$. A way to rationalize equation 3.2 is to assume trade costs in 1995 have the form $\kappa_{ROW,BRA,95}^k = (distport_{i,i})^{\delta_k} (1 + b_{ROW,BRA,95}^k) (1 + t_{ROW,BRA,95}^k)$, where $b_{ROW,BRA,95}^k$ is the tariff equivalent of all border related costs other than tariffs and is not symmetric.¹⁷

Table 3.4 shows the level of tariffs in both 1991 and 1995 and the ratio of tariff factors, $\frac{(1+t_{95}^k)}{(1+t_{91}^k)}$. Note that all sectors experienced tariff reductions between 1991 and 1995, except for Agriculture. The highest tariff reductions took place in Apparel (21 log percentage points) and Textiles (13 log percentage points). Between 1995-1998

¹⁷One advantage of assuming an initially symmetric border effect is that I do not need data on applied tariffs by the ROW to perform the calibration and I only need applied tariff changes by Brazil to calculate the trade shock. As a robustness check, I assumed that the initial total border effect is asymmetric: $\chi_{BRA,ROW}^k = (1 + b^k) (1 + t_{BRA,ROW}^k)$ and $\chi_{ROW,BRA}^k = (1 + b^k) (1 + t_{ROW,BRA}^k)$, so that the non-tariff border effect, b^k , is symmetric but the overall initial border effect, χ^k , is not. I calibrated a symmetric non-tariff border effect b^k using data on average applied ad-valorem tariffs by the ROW to Brazil from TRAINS. This dataset has a high proportion of missing values in the early 1990s, so I did not use it for the benchmark calibration. Qualitative results do not change using this alternative methodology. Results are available upon request.

there was a small reversal of trade reform to meet demands for more protection and to keep imports at levels consistent with balanced trade. The government raised the import tariffs for cars, motorcycles, bicycles, tractors, consumer electronics, fabrics, blankets, and shoes. However, the Common External tariff of the MERCOSUR prevented further changes in the tariff structure. Therefore, this period was the one in which tariffs were more stable since the start of the trade reform process. Throughout the paper, I consider tariff changes between 1991-1995, which is the period with the most substantial change in tariffs.^{18,19}

With the data on regional wages (w_i) in 1991, hours worked in each region and sector (L_i^k) in 1991, elasticities of substitution (σ_k), Cobb-Douglas shares (μ_k), the calibrated productivities (β_i^k) and the updated trade costs after a shock ($\kappa_{ROW,BRA,95}^k$), I use the equilibrium conditions of the model to solve for the counterfactual regional wages after the shock normalizing the ROW wage to one both in the initial and the final equilibrium. Therefore, counterfactual nominal wage changes should be interpreted as normalized by ROW wage changes.²⁰

¹⁸Kovak (2013) considers tariff changes between 1990 and 1995. Since the data for wages is from 1991, I use 1991 as the start year for tariff changes. None of the qualitative results of the paper change when considering tariff changes between 1991 and 1998. For the counterfactual section, results are available upon request. For the estimation, I report results in Section 3.4.4.

¹⁹During this period Brazil signed the MERCOSUR agreement with Argentina, Paraguay, and Uruguay. This agreement implied not only a reduction in applied tariffs by Brazil but also a decrease in Brazilian export costs to member countries. In this dissertation, I focus on the increase in import competition caused by Brazilian tariff reductions during the 1990s and I leave the analysis of the consequences of increased export penetration to foreign markets for future research.

²⁰The solution algorithm is described in Appendix A.5.

3.3.2 A 1% import cost reduction in each sector

In this section I report the results for the first set of counterfactual experiments that consist in reducing the trade costs for Brazil of importing goods from the ROW by 1%, one sector at a time, so $\hat{\kappa}_{ROW,BRA}^k = -0.01$.²¹ For each industry, I compute employment and linkage exposure. The former is standard in the literature and I derive the latter by decomposing wage changes.

3.3.2.1 Nominal wage decomposition

Table 3.5 shows in its first column nominal wage changes due to a 1% trade cost reduction with the ROW in each sector.²² This shock causes nominal wages to decline, since Brazilian regions face more competition from abroad. However, wage responses vary depending on the sector that is affected by the shock. For example, a 1% trade cost shock in Footwear and Leather causes a wage decline of 0.02 log points while a 1% trade cost shock in Mineral Manufacturing or in Apparel causes a wage decline two orders of magnitude smaller.

The next three columns report the decomposition of wage changes into its components using equations 2.14, 2.15, and 2.16. The direct effect is in every case negative, as expected, but the partial and spillover effects are positive. This means that the adjustment in endogenous variables, namely wages and sectoral employ-

²¹The reason for doing a 1% reduction is that the decomposition in equations 2.13, 2.17, and 2.18 is only valid in a neighborhood of the initial equilibrium.

²²The procedure I used is the following. I performed 12 different experiments in which I shocked each of the sectors at a time. I solved for regional wage changes in every Brazilian microregion and then I averaged across all regions weighting by the initial share of total hours worked in each region. Therefore, results shown in Table 3.5 are weighted averages.

ment, overturns the direct effect of the trade cost change on prices. The 5th column shows that the share of the direct effect overturned by the sum of partial and spillover effects ranges from 92.3% in Food to 99.7% in Mineral Manufacturing. Note that not all the terms in equations 2.18 and 2.16 have the same sign. The ones with a positive sign operate in the same direction as the direct effect—they tend to reduce wages—while the ones with a negative sign work in the opposite direction. The fact that in Table 3.5, PE and SE are positive means that the effects with a negative sign dominate. These are the wage and sectoral employment terms in 2.18 and the sectoral employment term in 2.16.²³

To account for the importance of linkages to other regions, the last column of Table 3.5 shows the share of the own region effect in the sum of the partial and the direct effects, in equations 2.17 and 2.18. This share is never higher than 12% in any sector. For example, this share is 11.3% in Paper. This means that 88.7% of the direct and partial effects comes from linkages to other regions rather than the own region. In Agriculture, Footwear and Leather, and Food this share is negative because the own-region effect is positive, as shown in the column to the left.²⁴ The reason is that the terms with a negative sign in the own-region term in equation 2.18, overturn the effect of the term with a positive sign.

To my knowledge, the only other study to perform a decomposition of wage changes in a general equilibrium trade model is Monte (2014). However, his model

²³Sectoral employment changes, capture a variety effect (equation 2.8). This effect is specific to the Krugman (1980) model and does not occur in other isomorphic models like Eaton and Kortum (2002), unless we assume there are externalities. I leave a comparison of this decomposition using an isomorphic model for future research.

²⁴Note from columns 2 and 3 in Table 3.5 that the sum of DE and PE is negative in every sector.

has free trade within the country and a uniform trade cost with the ROW, as well as commuting links, which my model does not have. Therefore, the formula for the wage decomposition is not directly comparable. It would be interesting in future research to decompose wage changes in a context with both costly trade and commuting or migration linkages.

3.3.2.2 Wages, exposure, and linkages

This section evaluates the contribution of two different measures of exposure to the variation in counterfactual wage changes. The first measure, denoted by ETC , is an employment-weighted trade cost change, the type of measure used in Topalova (2010) and Kovak (2013).²⁵ The second measure is a linkage-weighted trade cost change, LTC . This measure is equal to the direct effect on wages of a trade cost change with the rest of the World, as defined in equation 2.14. The formulas for each Brazilian region i and sector k are given by

$$ETC_i^k = \frac{L_i^k}{L_i} \hat{\kappa}_{ROW,BRA}^k, \quad (3.3)$$

$$LTC_i^k = (\sigma_k - 1) \left(\sum_{j \neq R} \xi_{ij}^k \pi_{Rj}^k \right) \hat{\kappa}_{ROW,BRA}^k. \quad (3.4)$$

²⁵Autor et al. (2013) measure exposure as an employment-weighted average of changes in import penetration. I introduce a measure analogous to their measure in section 3.3.3 when I shock several industries at the same time. In the context of a shock to only one industry, this measure and the weighted trade cost change are collinear.

ETC weights trade cost changes by the share of total traded employment in region i that corresponds to employment in sector k .²⁶ Even though *ETC* can be derived from theory, it is not derived from the class of models that yield a gravity equation for trade flows. The specific-factors model in Kovak (2013) assumes perfect competition and no transport costs. Therefore, changes in regional prices are equal to changes in tariffs. This is not the case in the model with a gravity structure for trade flows.

The theory-consistent way of distributing the trade cost shock across regions is using linkages as in *LTC*. This measure takes into account the structure of linkages that the model predicts has a first order effect on local wages. Regardless, it is a reduced-form measure since it does not take into account the feedback from changes in endogenous variables: w_i and L_i^k in all regions.

To my knowledge, there are two other studies that use linkage-based measures of exposure. One is Stumpner (2015) who defines the trade demand shock faced by a US region as a weighted sum of household leverage in all other regions, where the weights are given by the share of exports to that region. My measure is different from Stumpner (2015) since he considers the demand shock as exogenous, and therefore, he only weights by export shares whereas I also weight by import shares. The other is Monte et al. (2015) who use the partial elasticity of employment to productivity shocks to predict changes in regional employment. My measure is different from theirs in three aspects: first, I consider only the direct effect while they incorporate

²⁶I use as a denominator total traded employment in the region and not total employment to be consistent with the assumption I made for calibrating μ_0 : that the share of non-traded employment is the same for all Brazilian regions. See section 3.3.1 and footnote 9 for details.

other partial-equilibrium effects, second, since their model has commuting and migration, their measure incorporates other channels that are absent in my analysis, and third, the type of shock they study is a region-specific productivity shock, so the import weights are different. I also contribute to the analyses in those studies by providing an interpretation of exposure measures as instruments for domestic market access in Section 3.4.

Figures 3.3, and 3.4 show scatter plots of changes in nominal wages in the vertical axis and the exposure measures in the horizontal axis. We expect the correlation to be positive between *ETC* and *LTC* and wages since a decrease in tariffs, increases competition and therefore, decreases wages. The solid lines are the fit from a linear regression of nominal wage changes on exposure for each experiment. The first thing to note from these figures is that the correlation between exposure and wage changes does not always have the expected sign. Second, the employment-based measure captures a small share of the variation in wage changes, as measured by the R-squared of the regressions. *ETC* explains a median of 2% of the variation in counterfactual wage changes and a maximum of 17% of this variation, in the Agriculture sector, although the sign is negative. In some sectors, like Mining, Mineral Manufacturing, Textile, and Food the explanatory power of this variable is almost zero. The linkage-based measure, *LTC*, performs better in predicting wage changes. This measure is, as expected, positively and strongly related to wage changes except in two sectors: Mining and Plastic and Rubber. In these sectors, general equilibrium effects, given by the terms *PE* and *SE* in equation 2.13, overturn the sign of the correlation between *LTC* and changes in wages. But in the rest of the

10 sectors it holds that regions with more negative exposure as measured by *LTC*, have higher declines in wages. *LTC* explains a median of 44% of the variation in counterfactual wage changes and a maximum of 86% of this variation, in the Footwear and Leather sector.

To evaluate how much each measure contributes to the overall variation in counterfactual wage changes, I estimate for each sector a linear model on a 3rd-degree polynomial of all measures and decompose the R-squared coefficient. This synthetic regression is given by

$$\hat{w}_{i,k} = a_1 + \mathbf{a}_2 \text{Pol}(ETC_i^k) + \mathbf{a}_3 \text{Pol}(LTC_i^k) + e_i \quad k = 1, \dots, 12 \quad (3.5)$$

where $\hat{w}_{i,k}$ is the counterfactual nominal wage change in region i due to a 1% trade cost reduction in sector k and $\text{Pol}(x_i^k) = x_i^k + (x_i^k)^2 + (x_i^k)^3$, and $\mathbf{a}_2, \mathbf{a}_3$ are (3×1) vectors of coefficients. Table 3.6 shows the results of the Shorrocks-Shapley R-squared decomposition (see Shorrocks (1982) and Shorrocks (2013)).²⁷ Model 1 is simply a linear model where $\text{Pol}(x_i^k) = x_i^k$. For ten of the twelve sectors *LTC* explains at least 75% of the variation explained by the model. *ETC* explains most of the variation in Agriculture and Chemicals. Model 2 is a 3rd-degree polynomial in each variable. Each column groups the contribution of the polynomial of each

²⁷In cooperative game theory, the Shapley value provides a rule to distribute the total gains of a game by calculating the marginal impact of each of the factors as they are eliminated successively, and then averaging them over all the possible elimination sequences. In the case of the R-squared, the Shapley value is the marginal contribution to the R-squared averaging over all possible sub-models. In a model with two regressors, x_1 and x_2 , one way to compute the marginal contribution of x_1 is to subtract from the overall R-squared the R-squared obtained when x_1 is omitted from the regression. The other way is to estimate the R-squared of a regression including x_1 and not x_2 , and subtract the R-squared of a regression from which both variables were omitted, which is, zero. Averaging these two possibilities gives the Shapley contribution of x_1 .

variable. In this case, for 11 sectors *LTC* explains at least 70% of the variation explained by the model.

The exercises presented in this section suggest a link-based measure of exposure is better at explaining the variation of counterfactual wage changes generated by the model. The advantage of a measure like this is that it with trade data at the regional level, it can be calculated without the need to solve the model and it can be used to predict which regions are more affected by a trade cost shock.

To my knowledge, the only existing study that compares wage changes predicted using a model with linkages with those using employment exposure measures is Monte (2014). However, his analysis is different because he focuses on a shock to either all sectors, removing all trade barriers with the ROW or on a single NAICS 3-digit sector, "Computer and Electronics Product Manufacturing". Consistent with my findings he finds a weak explanatory power of exposure. However, he finds the sign is the expected one. By shocking each sector at a time, I find that in some sectors the correlation can be negative. In addition, I propose a linkage-based measure, *LTC*.

3.3.3 Trade liberalization

The previous experiments consisted in reducing trade costs one sector at time. However, in reality, these shocks frequently occur in several sectors simultaneously and the magnitudes of the shocks vary across sectors. In this section I perform a trade liberalization experiment and evaluate different exposure measures in that context.

I reduce the trade cost of shipping goods from the ROW to Brazil by the actual change in tariffs observed between 1991 and 1995 in Brazil (see Table 3.4).

The first row of Table 3.7 shows the distribution of counterfactual nominal wage changes after the trade liberalization experiment. Nominal wages decrease in every region, with a median reduction of -0.26 log points. The hours-weighted mean across all regions is also equal to -0.26 log points.

Panel A of Figure 3.7 shows the geography of the counterfactual wage changes. The most affected regions are the ones in the South and South East of the country. These regions are located near Sao Paulo, Rio de Janeiro, and Minas Gerais, the biggest centers of economic activity in the country. Regions in the North West of the country are more isolated from the shock, since they have fewer linkages to other regions, and therefore, suffer smaller wage changes in the counterfactual exercise.

To evaluate the reduced form measures of exposure in this context, we need to aggregate the measures from the sector and region level to the region level. The model yields that the direct effect of each sectoral shock on wages is given by linkage exposure, LTC_i^k . I propose an *ad hoc* aggregation scheme for linkage exposure, which is given by an unweighted sum, LTC_i

$$LTC_i = \sum_{k=1}^K (\sigma_k - 1) \left(\sum_{j \neq R} \xi_{ij}^k \pi_{Rj}^k \right) \hat{\kappa}_{ROW, BRA}^k \quad (3.6)$$

A theory-based way of aggregating the sectoral shocks to the region level is by taking into account that in the data wages vary across sectors. Therefore, we can start from an expression where wages vary at the region-sector level

$$\hat{w}_i^k \approx \sum_{k=1}^K (\sigma_k - 1) \left(\sum_{j \neq R} \xi_{ij}^k \pi_{Rj}^k \right) \hat{\kappa}_{ROW, BRA}^k \quad k = 1 \dots K, \quad (3.7)$$

and aggregate it to the region level using employment weights:

$$\hat{w}_i \approx \sum_{k=1}^K \frac{L_i^k}{L_i} (\sigma_k - 1) \left(\sum_{j \neq R} \xi_{ij}^k \pi_{Rj}^k \right) \hat{\kappa}_{ROW, BRA}^k \quad (3.8)$$

where I used that $\hat{w}_i = \sum_k \frac{L_i^k}{L_i} \hat{w}_i^k$ and the approximation comes from the fact that all the changes in endogenous variables are set to be equal to zero.

I denote by $ELTC_i$ this employment and linkage based measure of exposure, which is given by

$$ELTC_i = \sum_{k=1}^K \frac{L_i^k}{L_i} (\sigma_k - 1) \left(\sum_{j \neq R} \xi_{ij}^k \pi_{Rj}^k \right) \hat{\kappa}_{ROW, BRA}^k \quad (3.9)$$

I compare LTC_i and $ELTC_i$ with the standard employment exposure, which is an employment-weighted average of tariff shocks (Topalova (2007, 2010), Kovak (2013))

$$ETC_i = \sum_{k=1}^K \frac{L_i^k}{L_i} \hat{\kappa}_{ROW, BRA}^k \quad (3.10)$$

I also consider a third exposure measure, EMC , which is an employment-weighted import change

$$EMC_i = \sum_{k=1}^K \frac{L_i^k}{L_i} \left(\frac{\widehat{M}}{VA} \right)_{BRA}^k \quad (3.11)$$

where $\left(\widehat{\frac{M}{VA}}\right)_{BRA}^k$ is the counterfactual log change in the imports to value added ratio in sector k in Brazil.

EMC is the measure most closely related to the definition of exposure in Autor et al. (2013). In EMC the source of the shock is the change in nationwide imports while in ETC and LTC it is the change in trade costs as a consequence of the change in tariffs. Autor et al. (2013) use changes in imports because the shock they analyze, the increase in Chinese productivity, is not directly observable. Since import penetration changes are not observable at the commuting zone level, the authors use the following measure: $\sum_{k=1}^K \frac{L_i^k}{L_{Home}^k} \frac{\Delta M_{Home}^k}{L_i}$ to allocate the change in imports per worker in a region $\frac{\Delta M_{Home}^k}{L_i}$, proportionally to the share of that region in national employment in that sector $\frac{L_i^k}{L_{Home}^k}$. To calculate this measure, I use the counterfactual increase in the ratio of Brazilian imports to value added instead of using total initial employment as they do. Therefore, I replace $\frac{\Delta M_{Home}^k}{L_{Home}^k}$ with $d \ln \frac{M_{Home}^k}{VA_{Home}^k}$. The reason is that the counterfactual value of imports does not have a meaningful interpretation in the model since what drives changes in wages are changes in import penetration.

Table 3.8 shows the correlation coefficient of the four exposure measures. EMC is negatively correlated with the rest since decreases in trade costs drive increases in import penetration. $ELTC$ is more correlated with ETC than with LTC .

I analyze the contribution of each measure to the variation in wage changes using the following synthetic regression

$$\hat{w}_i = b_0 + \mathbf{b}_1 \text{Pol}(Exposure_i) + e_i \quad (3.12)$$

with $Exposure_i = \{ETC_i, EMC_i, LTC_i, ELTC_i\}$ and Pol is a polynomial that is either 1-degree (linear) or 3rd- degree. Table 3.9 displays the regression output for the standard exposure measures ETC and EMC and their polynomials. In the linear specification in columns 1 and 5, ETC has a positive sign and EMC a negative one, as expected. When included on their own, ETC and EMC explain 3.1% and 6.7% of the variation in wage changes respectively, which is small taking into account that this regression uses counterfactual wage changes after a change in tariffs keeping constant everything else. The R-squared of the regressions increases more than an order of magnitude when including state fixed effects. This reflects that in the model a great part of the variation in regional wages comes from geographic factors. However, both measures lose explanatory power and significance. Columns 3, 4, 7, and 8 show results for 3rd-degree polynomial in exposure. EMC remains significant but ETC does not.

The results for the linkage-based measures of exposure are displayed in Table 3.10. In the linear specification in columns 1 and 5, both LTC and $ELTC$ have a positive and significant coefficient. When included on their own, they explain 25.6% and 11.4% of the variation in wages respectively, more than the what the standard measures explain. As in the previous Table, controlling for state fixed effects increases the R-squared of the regressions but, unlike in the previous Table, the coefficients remain significant although with smaller coefficients (columns 2 and

6). Using a polynomial does not improve the fit of the regression substantially and even changes the sign of the coefficient for *LTC* .

Finally, Table 3.11 includes all the measures at the same time as regressors. *LTC* and *ELTC* remain positive and significant whereas *ETC* becomes negative and *EMC* loses significance. Table 3.12 reports the Shapley decomposition of the R-squared of a regression that includes all the measures of exposure in linear form, as in Column 1 of Table 3.11, and a third-degree polynomial of the measures. In the linear specification in Panel A, linkage-based measures, *LTC* and *ELTC* combined explain 30% of the variation in wage changes (83% of an R-squared equal to 0.36) and in the polynomial specification they explain 37% (75% of an R-squared of 0.503).

In sum, the results in this section suggest that weighted averages of link-based exposure measures are better suited for explaining cross-regional variation in wage changes than the standard measures. Among the standard measures, proxies for regional changes in import penetration (*EMC*) explain slightly more of the variation in regional wage changes than employment weighted tariff changes (*ETC*).

3.3.4 Real wages

The empirical literature on the impact of trade on local labor markets focuses on nominal wages instead of cost-of-living adjusted wages since price indices are usually not observed at a disaggregated level, such as commuting zones or microregions. The advantage of performing counterfactual exercises is that the change in microregion price indices can be solved for and this allows to calculate real wage changes. In the

model, a unilateral trade liberalization triggers both a reduction in nominal wages and in price indices. Depending on which of these two variables decreases more, real wages can increase or decrease.²⁸

The formula for the real wage change in region i is given by: $\hat{w}_i^R = \hat{w}_i - \hat{P}_i$, where \hat{w}_i is the change in the nominal wage and \hat{P}_i is the change in the aggregate price index in region i .²⁹

Table 3.7 shows in the second row that price indices decrease in more than 90% of the regions as a consequence of more competition from trade liberalization. Real wage changes range from a -0.26 log points decrease in the most affected region to a 1 log point increase in the most benefited region. The median change is equal to -0.034 log points. However, when weighting by the number of hours in each region, the average real wage gain in Brazil after trade liberalization is 0.11 log points. This fact indicates that bigger regions were on average more benefited by trade liberalization. In multisector Krugman-like models it is possible for some regions to lose from trade. This can be the case for regions with comparative disadvantage in sectors with strong scale effects, as explained in Costinot and Rodriguez-Clare (2014).

The geographic distribution of real wage changes differs from that of nominal wage changes. Figure 3.7 shows a map of each variable. A clear pattern emerges in Panel B of regions located in the coast being more benefited by trade liberalization and welfare gains becoming progressively negative as we move towards central re-

²⁸Note that tariffs in agriculture increased (see Table 3.4), so it can be possible for wages and/or price indices to actually increase in some regions.

²⁹Details of the calculation of the change in price indices are in Appendix A.6.

gions. By comparing Panels A and B, we can see that in some regions changes in price indices overturned changes in nominal wages. This is the case for example of some coastal regions in the South East that have nominal wage losses that are in the top quintile of the country but experience real wage gains.

Figures 3.5 and 3.6 are scatter plots by sector of the relationship between real wage change measures. Although, none of the measures are meant to explain real wages, it is interesting to analyze their correlation. *ETC* is negatively correlated with real wage changes, for almost every sector. *LTC* is positively correlated with real wage changes in some sectors and negatively correlated with them in others. The fit of the regression is particularly good in Metals, Chemicals, Textiles, and Food. Results for employment exposure measures are in line with findings in Monte (2014) who finds a weak and negative correlation between an exposure measure like *ETC* and real wage changes.

3.4 Estimating the wage elasticity with respect to market access

The analysis of the counterfactual experiments in the previous section showed that standard exposure measures are not as good explaining the variation of wage changes as linkages-based measures of exposure, even in the absence of other shocks. One caveat is that the previous section used model-generated counterfactual wages and in the model, it is precisely the linkages measures that constitute the direct effect of trade shocks on wages. In this section, I propose an estimation strategy that

incorporates trade linkages in the analysis of the effects of trade on observed wages. In addition, I show that exposure measures can be interpreted as an instrument for the endogenous change in market access, which in the model drives nominal wage changes. Finally, I compare the results of my approach with the results of an exposure approach.

3.4.1 Exposure as an instrument for changes in market access

I start by deriving the structural equation that relates wage changes to changes in market access, a trade cost-weighted average of the surrounding regions' real demand. To derive this equation, sum demand (equation 2.3) across all destination markets and use the pricing equation (2.7) and the optimal scale of the firm.³⁰ This yields what Fujita et al. (1999) first called a *wage equation* that holds for each sector. These are N zero profit conditions that determine the maximum wages region i can afford to pay given their market access in each industry, MA_i^k

$$w_i = \zeta_i^k \left(MA_i^k \right)^{1/\sigma_k} \quad k = 1 \dots K \quad (3.13)$$

where $\zeta_j^k \equiv \left[(\sigma_k - 1) \beta_j^k \right]^{\frac{\sigma_k - 1}{\sigma_k}} \left[F^k (\sigma_k)^{\sigma_k} \right]^{-1} > 0$ is a combination of parameters in the model and $MA_i^k \equiv \sum_j \left(\kappa_{ij}^k \right)^{1 - \sigma_k} \mu_k w_j L_j \left(P_j^k \right)^{\sigma_k - 1}$ is region i 's market access in sector k , a measure of real demand. Region i 's overall market access in industry k

³⁰Set $c_{ji}^k = x_{ji}^k / d_{ji}^k$ in the demand for manufactures, equation (2.3), to capture that due to iceberg trade costs, when region j ships x_{ji}^k units of the manufactured good k , region i consumes x_{ji}^k / d_{ji}^k .

is a sum of all of its destinations' capacities to buy its goods appropriately weighted by trade costs.³¹ Regions that are near high-income (high $w_j L_j$) or low-competition (high P_j^k) markets have higher market access and are able to pay higher wages, other things equal.

If we allow for productivities to change over time, total differentiation of equation 3.13 yields

$$\hat{w}_i = \frac{1}{\sigma_k} \widehat{MA}_i^k + \left(\frac{\sigma_k - 1}{\sigma_k} \right) \hat{\beta}_i^k \quad k = 1 \dots K \quad (3.14)$$

In the data, however, both market and wages vary across sectors. To be consistent with the model in Section 4.2 and with the literature that estimates the impact of trade on local labor markets, the left hand side cannot vary across sectors. Therefore, we can start from an expression where the left-hand side varies at the region-sector level

$$\hat{w}_i^k = \frac{1}{\sigma_k} \widehat{MA}_i^k + \left(\frac{\sigma_k - 1}{\sigma_k} \right) \hat{\beta}_i^k \quad k = 1 \dots K, \quad (3.15)$$

and aggregate it to the region level using employment weights:

$$\hat{w}_i = \sum_k \frac{L_i^k}{L_i} \frac{1}{\sigma_k} \widehat{MA}_i^k + \sum_k \frac{L_i^k}{L_i} \left(\frac{\sigma_k - 1}{\sigma_k} \right) \hat{\beta}_i^k \quad (3.16)$$

where I used that $\hat{w}_i = \sum_k \frac{L_i^k}{L_i} \hat{w}_i^k$.

From equation 3.16 we can derive the following estimating equation

³¹Note that region i itself is also included in the category “region i 's destination markets”.

$$\widehat{w}_i = \alpha_0 + \alpha_1 \widehat{MA}_i + \nu_i, \quad (3.17)$$

where $\widehat{MA}_i = \sum_k \frac{L_i^k}{L_i} \widehat{MA}_i^k$ is an aggregate of changes in sectoral market access, α_1 is the elasticity of regional wages with respect to aggregate market access and ν_i is an unobservable error term, that contains random factors that affect regional wages and also includes unobserved changes in productivity, as equation 3.16 shows. This equation constrains the elasticity of substitution σ to be equal across sectors, so that $\alpha_1 = 1/\sigma$. Otherwise, each sectoral σ_k would be included in the formula for \widehat{MA}_i and α_1 would be constrained to be equal to 1.

There are two challenges to estimate α_1 . First, \widehat{MA}_i is not directly observed, and second, the estimate of α_1 is biased due to structural correlation between unobserved changes in region-sector productivities, β_i^k , and MA_i^k . I describe how I deal with both issues in the following paragraphs.

I obtain MA_i^k for every region and sector from the following set of systems of equations that are derived from the model

$$\begin{cases} MA_i^k &= \sum_j \mu_k w_j L_j (P_j^k)^{\sigma_k-1} (\kappa_{ij}^k)^{1-\sigma_k} \\ (P_j^k)^{1-\sigma_k} &= \sum_h w_h L_h^k (MA_h^k)^{-1} (\kappa_{hj}^k)^{1-\sigma_k} \end{cases} \quad k = 1 \dots K \quad (3.18)$$

These systems are the same ones used by Anderson and Yotov (2010) to estimate buyer's and seller's incidence using Canadian and U.S. data. With data on wages, sectoral employment, elasticities of substitution, Cobb-Douglas shares and a

parametrization of trade costs, each of the systems that can be solved independently, up to a normalization, for a unique solution for MA_i^k and P_i^k .

Totally differentiating the system in equation 3.19, it is possible to decompose the change in a Brazilian region i 's market access in the following way

$$\begin{aligned} \widehat{MA}_i^k = & \underbrace{\sum_{j \in BRA} \xi_{ij}^k \widehat{w}_i - \sum_{j \in BRA} \sum_{h \in BRA} \xi_{ij}^k \pi_{hj}^k \left(\widehat{w}_h + \widehat{L}_h^k - \widehat{MA}_h^k \right)}_{\text{endogenous}} + \\ & \underbrace{(\sigma_k - 1) \sum_{j \in BRA} \xi_{ij}^k \pi_{Rj}^k \widehat{\kappa}_{ROW, BRA}^k + \xi_{iROW}^k \left[(\sigma_k - 1) \left(\widehat{P}_{ROW}^k - \widehat{\kappa}_{BRA, ROW}^k \right) \right]}_{\text{exogenous}} \end{aligned} \quad (3.19)$$

The endogenous component of changes in region i 's market access include the changes in market access, wages, and sectoral employment in all the Brazilian microregions, including region i . The exogenous component includes changes in trade costs from the ROW to Brazilian regions, $\widehat{\kappa}_{ROW, BRA}^k$, and from Brazilian region i to the ROW, $\widehat{\kappa}_{BRA, ROW}^k$, and changes in the price index of the ROW, \widehat{P}_{ROW}^k . Consider a unilateral tariff reduction by Brazil. In this case, $\widehat{\kappa}_{BRA, ROW}^k = 0$ but $\widehat{\kappa}_{ROW, BRA}^k \neq 0$. Since tariffs are set at the country level, I denote them without microregion subscripts: τ_{BRA}^k . Therefore, we can express the exogenous component of MA changes due to tariff changes as $(\sigma_k - 1) \sum_{j \in BRA} \xi_{ij}^k \pi_{Rj}^k \widehat{\tau}_{BRA}^k$.³² This is exactly the formula for the direct effect in equation 2.14, which I denoted by LTC in equation 3.4. This direct effect, only affects MA_i^k through changes in the price index, without the mediation of changes in other endogenous vari-

³²I used that $\kappa_{ij}^k \equiv d_{ij}^k b_{ij}^k \tau_{ij}^k$, and therefore $\widehat{\kappa}_{ij}^k \equiv \widehat{d}_{ij}^k + \widehat{b}_{ij}^k + \widehat{\tau}_{ij}^k$

ables. Therefore, $LTC_i^k = (\sigma_k - 1) \left(\sum_{j \neq R} \xi_{ij}^k \pi_{Rj}^k \right) \hat{\tau}_{BRA}^k$ is an instrument for \widehat{MA}_i^k . Using the same formula as above to aggregate it to the region level, we have: $ELTC_i^k = \sum_k \frac{L_i^k}{L_i} (\sigma_k - 1) \left(\sum_{j \neq R} \xi_{ij}^k \pi_{Rj}^k \right) \hat{\tau}_{BRA}^k$, which can be used as an instrument for \widehat{MA}_i .

Among the exogenous components in equation 3.19, some are observable-like tariff changes—and some are not-like changes in border barriers other than tariffs. Any measurement error in the changes of border effects, can contaminate the measure of market access. Therefore, for the estimation below, I focus on the elasticity of wages with respect to market access within Brazil, which at the sector level is given by

$$DMA_i^k = \sum_{j \in BRA} \mu_k w_j L_j \left(P_j^k \right)^{-1} \left(\kappa_{ij}^k \right)^{1-\sigma_k} \quad (3.20)$$

where DMA_i^k stands for domestic market access. The change in this variable can be aggregated to the region level using the formula: $\widehat{DMA}_i = \sum_k \frac{L_i^k}{L_i} \widehat{DMA}_i^k$. $ELTC_i$ is a valid instrument for changes in domestic market access \widehat{DMA}_i , by the arguments presented above. In fact, Brazilian tariff changes can only affect regional wages through changes in Brazilian (i.e. domestic) market access, assuming Brazil is small enough for changes in its wages to not affect World prices.³³ To control for changes in foreign market access, I include in my preferred specifications the distance to the nearest sea port and the distance to Sao Paulo, which is the largest economic center in the country.

³³The two other papers that, to my knowledge, estimate a wage equation in changes using regional data, also focus on domestic market access (see Hanson (2005) and Bartelme (2014)).

The identifying assumption is that $Cov(LTC_i^k, \hat{\beta}_i^k) = 0$ and is not testable. This assumption is violated if tariff changes are correlated with sector-region productivity changes, $\hat{\beta}_i^k$. This could occur if regions that expect to have a negative productivity shock in a sector, lobby for smaller tariff reductions in that sector to mitigate the negative effect shock. On the one hand, tariff policy is set at the country level so it is unlikely that a single microregion can have such lobby power. On the other hand, since sectors tend to be clustered in space and productivity shocks can be spatially correlated, it is possible that a group of microregions colludes to obtain smaller tariff reductions. However, the tariff reduction process that took place in Brazil in the early 1990s was arguably not driven by lobbies. It has been argued by a number of authors that tariff cuts in Brazil were driven more by the government as a way to reduce distortions and to deal with macroeconomic problems rather than by interest groups (Abreu (2004), Gonzaga et al. (2006), Kovak (2013), and Kume et al. (2000)). In addition, the MERCOSUR common external tariff provided another exogenous force for tariff cuts. Furthermore, as Kovak (2013) shows, the correlation between the pre-liberalization tariff levels and changes in tariff is -0.90, which gives support to the hypothesis that one of the main goals of the tariff cuts was a reduction in cross-sectoral variation. However, there is still potential endogeneity if pre-liberalization tariff levels were set according to productivity considerations. Gonzaga et al. (2006) argue that after 1974 tariff and non-tariff barriers were increased to cope with the macroeconomic instability caused by the oil shocks in the late 1970s and the debt crisis in the 1980s and not to protect sectors in which Brazil had comparative disadvantage. The authors show

evidence that tariff levels in 1988 had a weakly positive correlation with sectoral skill intensity, but with considerable variance.³⁴

Another potential threat to identification are aggregate shocks that occurred during trade liberalization. During the 1990s Brazil suffered from large devaluations and high inflation. A devaluation in the model is equivalent to a uniform tariff increase across sectors. Since in the model a uniform tariff change only affects wages indirectly through changes in market access, not considering exchange rate shocks does not invalidate my instrument, but it can make it weaker. However, inflation can affect wages directly through bargaining. The maintained identifying assumption is that sectoral variation in tariffs is uncorrelated with sectoral variation in inflation rates. This seems reasonable given that one of the main goals of the tariff cuts was a reduction in cross-sectoral variation. In the case of other aggregate shocks that affect regional wages uniformly, their effect is captured by the intercept of the wage equation.

A final concern is that wages had already adjusted—at least partially—before 1991 due in part to timing, since tariff changes started before 1991, and in part to anticipatory behavior by economic agents. If this was the case, the fall in market access between 1991 and 2000 would be less correlated with tariff changes than in a

³⁴The evidence for Brazil contrasts with that of Colombia. Karacaovali (2011) finds that in Colombia more productive sectors were more protected and that the sectors with higher productivity gains were liberalized less. However, the author acknowledges that Colombia may be a special case among developing countries that unilaterally liberalized trade: “Although some authors [...] acknowledge the potential for endogeneity, they argue that it may not be such an issue in their studies given that the tariffs were reduced uniformly or proportionally across sectors. This is not true for Colombia; the liberalization was not uniform.”

scenario with no anticipation. This would make my instrument weak. However, as I show in the next section, the F-statistic of the first stage rejects the null hypothesis.

3.4.2 Estimation results

I estimate the following equation

$$d \ln w_{i,91-00} = \alpha_0 + \alpha_1 d \ln DMA_{i,91-00} + \nu_i^k \quad (3.21)$$

where $d \ln w_{i,91-00}$ are observed log wage changes in region i between 1991 and 2000 and the right hand side variable is given by $d \ln DMA_{i,91-00} = \sum_{k=1}^K \frac{L_{i,t}^k}{L_{i,t}} d \ln (DMA_{i,91-00}^k)$, an aggregate of sector-region log changes in domestic market access.

Panel A of Table 3.13 reports the results of the estimation of equation 3.21 with OLS. My preferred specification is the one in column (4) which includes state fixed effects and controls for the regional share of employment in manufacturing as well as for the log of the distance to Sao Paulo and to the nearest sea port. The elasticity of wages with respect to market access is around 0.05 and is significant at the 1%. Since this equation is estimated in first differences, it removes any locational advantages or amenities that are constant across time. However, there remains a potential source of endogeneity in the estimated OLS coefficient: changes in productivity. As I explained in the previous section, there is a structural correlation between productivity and market access that would induce a positive covariance between changes in market access and omitted changes in productivity, creating a positive bias in the estimated OLS coefficient. However, if productivity changes are strongly

negatively spatially correlated, the covariance between domestic market access and productivity can be negative.³⁵

To deal with endogeneity, I use $ELTC_i$ as an instrument for $d \ln DMA_{i,91-00}$

$$ELTC_i = \sum_k \frac{L_{i,91}^k}{L_{i,91}} (\sigma_k - 1) \left(\sum_{j \neq R} \xi_{ij,91}^k \pi_{Rj,91}^k \right) \hat{\tau}_{91,95}^k, \quad (3.22)$$

where $\hat{\tau}_{91,95}^k$ is the log change in the Brazilian tariff factor between 1991 and 1995 and is given by $\hat{\tau}_{91,95}^k = \ln \left(\frac{1+t_{95}^k}{1+t_{91}^k} \right)$, where t_y^k are ad-valorem tariffs in sector k in year $y = 1991, 1995$.

Panel B in Table 3.13, reports the results of estimating equation 3.21 with 2SLS using $ELTC$ as an instrument for changes in domestic market access.³⁶ The elasticity of wages with respect to domestic market access in my preferred specification is around 0.12 and significant at the 1%. This value is greater than the OLS estimate and suggests a negative covariance between unobserved changes in productivity and changes in market access. The Hausman test of exogeneity, rejects the null hypothesis at the 1% of significance. Compared to other estimates from the literature, these values are similar to values obtained for Brazil and smaller than values for the U.S. Fally et al. (2010) estimate a wage elasticity with respect to domestic market access of 0.16 using cross-sectional data at the state-sector level for Brazil in 1999. For the U.S., Hanson (2005) estimates a value of 0.5 using county-

³⁵Bartelme (2014) provides a similar argument.

³⁶Table 3.21 shows summary stats for regressions.

level data for the period 1970-1990 and Bartelme (2014) estimates a value of 0.6 using data at the MSA level for the periods 1990-2000 and 2000-2007.³⁷

Although *ELTC* is my preferred instrument since it is derived from the model, I explore results using other exposure measures as instruments. This is useful for two reasons. First, there might be cases where linkages between regions are not observed and it is of interest to ask if an employment-weighted average of shocks, like *ETC*, can be used as an instrument. In addition, there can be cases where the shock itself is not observed and we may wonder if a measure using changes in imports, like *EMC*, can be used as an instrument. Second, by including other instruments, I can perform tests of overidentifying restrictions.

In Table 3.14 I use the other exposure measures as instruments. *ETC* and *LTC* are constrained versions of the more general and theory consistent, *ELTC*. In particular, *ETC* assumes no variation in linkages and elasticities of substitution, and *LTC* assumes no variation in employment shares. *EMC* does not use information in tariff changes and uses instead variation in sectoral import penetration between 1991 and 2000. This measure would be more appropriate in cases where the shock is not observable. All measures yield positive and significant estimates that are greater than those estimated with OLS, which gives support to the hypothesis that the OLS coefficient is biased downwards. *EMC* yields a negative and non-significant coefficient. Comparing these estimates with the ones obtained using the model-based measure *ELTC*, *ETC* gives similar values, of around 0.09, but

³⁷Kumar (2007) estimates an elasticity of regional wages with respect to domestic market access of around 0.4 for India.

LTC gives much estimates for the elasticity, of around 0.3 and the precision is very low when including controls. This means that discarding the variation in linkages does not alter the estimate of the elasticity as much as discarding the information in employment shares. The Sargan test of overidentifying restrictions of a specification including *ELTC* and *ETC* as instruments cannot reject the null hypothesis, and the same occurs with a specification including *ELTC* and *EMC*. However, this is not the case for *LTC*, the Sargan test of a specification including *ELTC* and *LTC* as instruments rejects the null hypothesis, suggesting *LTC* is not a valid instrument.

Table 3.15 shows the results from the first stage regression for all the instrumental variables. The instruments are highly correlated with market access and the F-statistics are well above the rule of thumb value of 10. *LTC* is the only instrument that is negatively related to changes in domestic market access. However, in the second stage it provides the correct sign for the coefficients on market access. Intuitively, we would expect the instruments to be positively associated with changes in market access. However, it is important to keep in mind that these variables give a measure of a direct effect of changes in tariffs on changes in market access, and it is possible that other effects overturn the sign of this relationship. In fact, I show this is a possible scenario in the sector-by-sector analysis in Section 3.3.2.

Finally, Table 3.16 reports the results from the reduced form estimations. These correspond to the type of specification in the exposure approach, where changes in wages are a function of exposure. Both *ELTC* and *ETC* yield positive and significant coefficients. *LTC* has a negative coefficient, which can seem counter-intuitive but, as I mentioned in the previous paragraph, measures of direct

effects can be overturned by other effects. The coefficients for *EMC*, the measure most similar to the one by Autor et al. (2013), are always negative, as expected, and significant. The coefficients in Panel B, columns 1 and 2, can be compared to results in Kovak (2013). He obtains a value of 0.4 when controlling for state fixed effects while I obtain a much larger value of 1.2. The reason is that both my measure of wages and my measure of exposure are different from his. As I mentioned in Section 3.2, I use average wages and he uses region fixed effects from a Mincer regression. Also, my measure of exposure is different since I work at a more aggregated level: I have 12 sectors whereas he has 19.³⁸ In section 3.4.4 I compare my results with his.

3.4.3 Economic significance across different approaches

I showed how exposure can serve as instrument for changes in market access and I estimated an elasticity of wages to domestic market access of 0.12. But, how does this estimate compares with the reduced form estimates in Table 3.16?

There are two issues that complicate a straightforward comparison. First, reduced form estimates are not designed to recover the elasticity of wages to market access, so the values of the estimates cannot be directly compared across approaches. In the empirical literature the reduced form estimates have been interpreted in a difference-in-difference way as measuring how a relative increase in a region's exposure with respect to another region would change its wages. Second, there is not an explicit function relating changes in tariffs to changes in market access.

³⁸The interstate trade data I use to calibrate the model and predict trade shares forces me to work at a higher level of aggregation. See Appendix B for details.

Market access changes in a single region are a function of changes in endogenous variables in all Brazilian regions. I tackle these issues in the following way. First, I simulate the change in domestic market access using the system of equations in equation 3.18 that would take place under the observed change in tariffs, holding everything else constant. I call this variable, counterfactual domestic market access, DMA^C . To compare across approaches, I calculate the interquartile range change in either exposure or market access and I obtain the estimated wage change by multiplying it by the corresponding estimated coefficients I obtained in the previous sections.

The results from this exercise are in Table 3.20. Panel A contains the results from increasing market access and exposure from the 25th to 75th percentile and Panel B and C decompose this change into the change from the 25th to 50th percentile and the change from the 50th to the 75th. Columns 1 and 2 show the difference in the effect when I use observed changes in market access with counterfactual ones. Since domestic market access could change due to a variety of factors, the measure DMA^C is a way of using the structure of the model to obtain the change in market access that would result from a change of tariffs while keeping everything else constant. Moving from the 25th to the 75th percentile in this variable implies a wage change of 0.62%. Note that trade liberalization decreased market access, so this number should be interpreted as: an interquartile range decrease in domestic market access, implies a decrease in wages of 0.62%. Columns 3 and 4 report that an interquartile wage increase in exposure as measured by $ELTC$ and ETC leads to a change in wages of around 3%. Since exposure is a linear function of tariffs and tar-

iffs decreased during the period, this result says that an interquartile range decrease in weighted-tariff changes, leads to a 3% decrease in wages. The implied value for EMC is similar except that the sign is the opposite since trade liberalization lead to an increase in imports and EMC is a weighted average of imports.

Figure 3.10 shows wage changes predicted with a market access approach—by using each regions' DMA^C and the estimated elasticity of 0.12—with those predicted with employment exposure—by using each regions' ETC and the estimated coefficient of 0.71. Although some regions change quintiles depending on the approach, the most affected regions are in the South East, surrounding the cities of Belo Horizonte, Curitiba, Rio de Janeiro, and Sao Paulo. These regions have a decrease in predicted wages of between -0.7 and -3.3 log points using the market access approach (Panel A) and a strictly more negative decrease of between -3.4 and -8.4 using the employment exposure approach.

In conclusion, reduced form measures of exposure overstate the impact of tariff changes on wages compared to an approach based on market access.

3.4.4 Robustness

In this section I deal with potential concerns about the previous estimates. First, I re-estimate the wage equation using tariff changes between 1991 and 1998 instead of using those between 1991 and 1995. As I explained in Section 3.2, the bulk of tariff reductions occurred at the beginning of the decade but there was a small reversal of trade reform between 1995 and 1998. Second, I re-calculate link-based

instruments leaving out the elasticities of substitution. The main reason to do this is consistency. Estimating a unique wage elasticity to market access requires to constrain the elasticity of substitution to be equal across sectors. Third, I re-run the reduced form regressions using Kovak (2013) measure of wages. This is to check how much the estimated reduced form coefficient changes due to a higher level of sectoral aggregation. Fourth, I address concerns about migration flows. If workers arbitrage by moving to microregions experiencing higher market access, my estimates of the wage elasticity should be interpreted as a lower bound. Finally, I use a measure of domestic market access that does not consider own region wages, to avoid any mechanical correlation between changes in domestic market access and wage changes.

Panel A of Table 3.17 shows results are robust to considering tariff changes between 1991 and 1998, the estimates of the wage elasticities using *ELTC*, *ETC*, and *LTC* are similar to those obtained using changes between 1991 and 1995.

Panel B of Table 3.17 reports the wage elasticities using as instruments link-based exposure measures that discard variation in the elasticities of substitution. The estimates using *ELTC*, my preferred instrument, are still significant but lower than in the baseline specification. The wage elasticity is 0.07 when including a full set of controls.

Table 3.18 shows results using Kovak’s measure of regional wages. This measure corresponds to the change between 1991 and 2000 in the region fixed effect of

a Mincer regression that controls for individual characteristics.³⁹ Columns 1 and 2 in Panel B are directly comparable to his estimates. I estimate a coefficient of 0.58, which is higher than the estimate of 0.4 that he obtains. This can be attributed to both my higher level of sector aggregation and the fact that I use tariff changes between 1991 and 1995 and he uses tariff changes between 1990 and 1995. However, when including the full set of controls the coefficient is no longer significant.

I estimate the effect of market access on the in-migration rate of a microregion. For this purpose I use a question in the Census questionnaire that asks workers the municipality they lived in 5 years ago. The dependent variable is $\frac{Migrants_{ji,95-00}}{Natives_{i,95}}$ where the numerator are the hours worked by workers living in microregion i in 2000 that report having live in microregion $j \neq i$ in 1995 and the denominator are the hours worked by workers that lived in microregion i both in 2000 and in 1995. Table 3.19 shows that in most of the cases the estimates are not significant. Column 2 in Panel A shows that when using *ELTC*, my preferred instrument for market access, the coefficient for the change in market access is negative and significant at the 10%, but the coefficient is no longer significant when including the full set of controls. These results suggest migration flows do not respond to changes in domestic market access.

Finally, my results are unaltered when I exclude own-region wages from the computation of domestic market access.⁴⁰

³⁹I weight the regression using the standard error of the fixed effects estimates as in Kovak (2013).

⁴⁰I do not show the results here for the sake of brevity.

3.5 Conclusion

In this Chapter I took the theoretical model in Chapter 3 to the Brazilian data in two ways. First, I calibrated the model in the initial year, 1991 and I performed a set of counterfactual exercises. These exercises show employment-based exposure measures are poor predictors of wage changes generated by the model while linkage-based exposure measures explain more of the variation in counterfactual wages. Second, I proposed an estimation strategy and an instrument based on the model. I showed a linkage-based exposure measure can be interpreted as an instrument for changes in domestic market access and I estimated an elasticity of wages with respect to market access of 0.12. Using this elasticity I showed that reduced form approaches overstate the impact of trade openness on wages.

3.6 Tables

Tab. 3.1: Descriptive statistics of microregion level variables.

Variable	mean	s.d.	min	max
Area (km2)	17270.3	37377.7	18.4	423951.7
Distance to port (km)	318.3	305.5	5.9	1537.9
Distance to Sao Paulo (km)	1259.3	840.7	18.1	3348.5
Hourly wage (usd)	0.90	0.37	0.31	2.58
Hours worked (mill.)	179.6	526.7	1.5	8457.4
Full-time equivalent workers (thou.)	86.4	253.2	0.7	4066.0

Notes: Values corresponding to the 494 Brazilian microregions in 1991. Distance to port is the minimum distance of a microregion's centroid to the nearest sea port. Distance to Sao Paulo is the distance of a microregion's centroid to the centroid of the microregion containing the city of Sao Paulo. Microregions' wages in current US dollars and hours worked are calculated as described in Section 3.2. Full time equivalent workers are equal to total hours divided by the number of hours worked in a year, assuming full-time workers works 40 hours per week.

Sources: Areas: IBGE. Distances: own calculation using publicly available shape files from IBGE. Wages and hours: Demographic Census, IBGE. See Appendix B for details.

Tab. 3.2: Descriptive statistics of ad-valorem tariffs.

year	mean	median	s.d.	min	max	N
1988	38.61	39.25	17.06	5.63	80.93	53
1989	31.54	28.89	17.73	1.88	77.33	53
1990	28.59	28.35	15.49	3.33	78.68	53
1991	22.77	20.75	13.46	1.67	63.65	53
1992	15.07	14.20	8.38	0.56	38.97	53
1993	13.01	12.50	6.26	2.14	33.97	52
1994	10.68	9.50	5.94	2.21	31.19	52
1995	11.90	10.74	6.64	1.48	41.00	53

Notes: Ad-valorem tariffs applied by Brazil at the 4-digit National Accounts classification ("Nivel 80"). N corresponds to the number of 4-digit products over which statistics were calculated in each year.

Source: Kume et al. (2003). See Appendix B for details.

Tab. 3.3: Gravity equation: 1999. PPML estimates.

Dependent variable: log bilateral exports

	log dist	s.e.	obs.	dist elasticity
<i>Sector</i>	(1)	(2)	(3)	(4)
Agric.	-1.217***	0.120	546	0.529
Mining	-2.196***	0.260	520	0.845
Mineral Manuf.	-1.387***	0.104	546	0.821
Metal, Mach., Elect. & Autom	-0.410***	0.078	546	0.193
Paper	-1.051***	0.127	546	0.688
Rubber & Plastic	-0.857***	0.117	546	0.371
Chemic., Pharm. & Petrol.	-0.894***	0.132	546	0.541
Textiles	-0.498***	0.101	520	0.234
Apparel	-0.735***	0.060	546	0.328
Footwear & Leather	-0.345***	0.089	494	0.025
Food	-1.214***	0.065	546	0.473
Wood & Others	-0.621***	0.085	546	0.308

Notes: Column 1 reports PPML estimates of the log distance coefficient in sector-level gravity equations in 1999. Robust standard errors are reported in column 2. Standard errors are clustered by pair of states. Column 4 reports δ_k , the implied elasticity of trade costs with respect to distance. N=494. All specifications include an intercept. Significant at *10%, **5%, ***1%.

Tab. 3.4: Tariff levels and tariff changes by sector: 1991-1995 and 1991-1998

<i>Sector</i>	1991	1995	1998	$\frac{(1+t_{95})}{(1+t_{91})}$	$\frac{(1+t_{98})}{(1+t_{91})}$
Agric.	5.100	7.400	9.900	1.022	1.046
Mining	3.202	1.237	2.828	0.981	0.996
Mineral Manuf.	19.600	10.200	13.600	0.921	0.950
Metal, Mach., Elect. & Autom	29.035	17.592	18.422	0.911	0.918
Paper	13.400	9.800	14.200	0.968	1.007
Rubber & Plastic	32.532	14.301	16.942	0.862	0.882
Chemic., Pharm. & Petrol.	16.270	5.936	10.316	0.911	0.949
Textiles	30.600	14.900	19.400	0.880	0.914
Apparel	48.300	19.800	22.800	0.808	0.828
Footwear & Leather	24.800	17.900	17.200	0.945	0.939
Food	27.217	12.595	15.983	0.885	0.912
Wood & Others	23.514	11.886	15.016	0.906	0.931

Notes: Ad-valorem tariffs applied by Brazil aggregated to 12 traded sectors as described in Appendix B. The last two columns report the ratio of the tariff factors in 1995 and 1998 with respect to the level in 1991.

Source: Kume et al. (2003)

Tab. 3.5: Counterfactual wage change decomposition.

<i>Sector</i>	wage change	wage change decomposition				<i>DE</i> + <i>PE</i> decomposition	
		<i>DE</i>	<i>PE</i>	<i>SE</i>	$\frac{PE+SE}{ DE }$	<i>own</i>	$\frac{own}{DE+PE}$
Agric.	-0.0015	-0.0424	0.0047	0.0361	0.964	0.0020	-0.0533
Mining	-0.0153	-0.4203	0.1128	0.2921	0.964	-0.0264	0.0859
Mineral Manuf.	-0.0002	-0.0608	0.0135	0.0472	0.997	-0.0056	0.1181
Metal, Mach., Elect. & Autom	-0.0127	-0.4898	0.0600	0.4171	0.974	-0.0121	0.0282
Paper	-0.0012	-0.1175	0.0240	0.0923	0.990	-0.0105	0.1125
Rubber & Plastic	-0.0012	-0.2839	0.0458	0.2369	0.996	-0.0270	0.1132
Chemic., Pharm. & Petrol.	-0.0057	-0.4535	0.0958	0.3520	0.987	-0.0329	0.0919
Textiles	-0.0018	-0.1995	0.0150	0.1827	0.991	-0.0085	0.0463
Apparel	-0.0002	-0.0301	0.0027	0.0272	0.995	-0.0028	0.1028
Footwear & Leather	-0.0203	-1.3741	0.3600	0.9937	0.985	0.2328	-0.2296
Food	-0.0169	-0.2196	0.0602	0.1425	0.923	0.0166	-0.1044
Wood & Others	-0.0035	-0.2465	0.0209	0.2221	0.986	-0.0094	0.0416

Notes: decomposition of the counterfactual wage change due to 1% shock in the trade cost of importing goods from ROW in each sector. Wage changes in log points. *DE*, *PE*, *SE*, and *own* correpond to the direct, partial, spillover, and own-region effects defined in Section 2.3.

Tab. 3.6: R-squared decomposition from synthetic OLS regressions. 1% shock in each sector.

Dependent variable: log counterfactual wage change						
<i>1% shock to sector:</i>	Model 1. Linear			Model 2. Polynomial		
	R^2	ETC	LTC	R^2	ETC	LTC
Agric.	0.200	0.553	0.447	0.235	0.515	0.485
Mining	0.002	0.052	0.948	0.208	0.092	0.908
Mineral Manuf.	0.179	0.004	0.996	0.280	0.147	0.853
Metal, Mach., Elect. & Autom	0.366	0.162	0.838	0.625	0.095	0.905
Paper	0.403	0.196	0.804	0.442	0.251	0.749
Rubber & Plastic	0.518	0.026	0.974	0.536	0.046	0.954
Chemic., Pharm. & Petrol.	0.056	0.863	0.137	0.290	0.191	0.809
Textiles	0.480	0.255	0.745	0.488	0.301	0.699
Apparel	0.645	0.077	0.923	0.731	0.201	0.799
Footwear & Leather	0.855	0.001	0.999	0.918	0.007	0.993
Food	0.542	0.017	0.983	0.689	0.011	0.989
Wood & Others	0.444	0.030	0.970	0.469	0.049	0.951

Notes: Results from a 1% decrease in the trade cost of importing from the ROW in each sector. Shapley decomposition of the R-squared of a synthetic regression of counterfactual log wage changes on: exposure measures ETC and LTC (Panel I) and 3rd-degree polynomial of exposure measures (Panel II). Column 1 in each panel reports the R-squared and the three following columns report the share of the R-squared explained by each variable (Panel I) or group of variables (Panel II).

Tab. 3.7: Distribution of counterfactual wage changes and price index changes. Trade liberalization.

Variable	min	p5	p25	p50	p75	p90	max	mean	wgt. mean
\hat{w}	-0.386	-0.313	-0.272	-0.255	-0.228	-0.119	-0.002	-0.242	-0.256
\hat{P}	-0.997	-0.646	-0.435	-0.245	-0.153	-0.009	0.034	-0.294	-0.362
$\hat{w} - \hat{P}$	-0.262	-0.214	-0.108	-0.034	0.187	0.477	0.995	0.052	0.106

Notes: values correspond to the trade liberalization counterfactual. Wage changes, \hat{w} , price index changes, \hat{P} , and real wage changes $\hat{w} - \hat{P}$ in log points. Weighted mean value corresponds to the average of each variable weighted by the total number of hours worked in the microregion.

Tab. 3.8: Correlation coefficient between exposure measures

	<i>ETC</i>	<i>EMC</i>	<i>LTC</i>	<i>ELTC</i>
<i>ETC</i>	1			
<i>EMC</i>	-0.506	1		
<i>LTC</i>	0.224	-0.252	1	
<i>ELTC</i>	0.862	-0.521	0.253	1

Tab. 3.9: Standard exposure measures and change in counterfactual wages in Brazilian microregions. Trade liberalization. OLS estimates.

Dependent variable: log counterfactual wage change								
Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
<i>ETC</i>	0.354*** (0.105)	-0.0195 (0.102)	0.375 (0.261)	-0.266 (0.282)				
<i>ETC</i> ²			3.857 (9.501)	2.999 (8.696)				
<i>ETC</i> ³			41.97 (76.59)	70.48 (66.15)				
<i>EMC</i>					-0.00117*** (0.000207)	-1.44e-05 (0.000192)	-0.00124*** (0.000365)	5.78e-05 (0.000320)
<i>EMC</i> ²							-1.82e-05*** (6.82e-06)	-8.31e-06 (5.56e-06)
<i>EMC</i> ³							-1.37e-08 (1.91e-07)	-7.42e-08 (1.51e-07)
R-squared	0.031	0.396	0.032	0.406	0.067	0.396	0.085	0.398
State FE		X		X		X		X

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. The dependent variable is the counterfactual wage change resulting from trade liberalization. N=494. All specifications include an intercept. Significant at *10%, **5%, ***1%.

Tab. 3.10: Linkages-based exposure measures and change in counterfactual wages in Brazilian microregions. Trade liberalization. OLS estimates.

Dependent variable: log counterfactual wage change								
Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
<i>LTC</i>	0.457*** (0.0498)	0.310*** (0.0567)	-2.807*** (0.792)	-1.709*** (0.533)				
<i>LTC</i> ²			-15.01*** (3.196)	-8.694*** (2.435)				
<i>LTC</i> ³			-20.98*** (4.113)	-11.49*** (3.445)				
<i>ELTC</i>					2.322*** (0.331)	0.909*** (0.304)	2.656** (1.208)	0.729 (1.108)
<i>ELTC</i> ²							-4.996 (53.74)	-20.58 (46.07)
<i>ELTC</i> ³							-322.6 (518.0)	-342.0 (437.9)
R-squared	0.256	0.479	0.315	0.496	0.114	0.407	0.120	0.409
State FE		X		X		X		X

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. The dependent variable is the counterfactual wage change resulting from trade liberalization. N=494. All specifications include an intercept. Significant at *10%, **5%, ***1%.

Tab. 3.11: Exposure measures and change in counterfactual wages in Brazilian microregions. Trade liberalization. OLS estimates.

Dependent variable: log counterfactual wage change		
Variable	(1)	(2)
<i>ETC</i>	-0.939** (0.377)	-0.728** (0.328)
<i>EMC</i>	-0.000350* (0.000186)	0.000256 (0.000212)
<i>LTC</i>	0.399*** (0.0489)	0.308*** (0.0531)
<i>ELTC</i>	4.060*** (1.490)	2.780** (1.265)
R-squared	0.359	0.515
State FE		X

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. The dependent variable is the counterfactual wage change resulting from trade liberalization. N=494. All specifications include an intercept. Significant at *10%, **5%, ***1%.

Tab. 3.12: R-squared decomposition from synthetic OLS regressions. Trade liberalization.

Dependent variable: log counterfactual wage change				
R^2	share of R^2 explained by:			
	ETC	EMC	LTC	$ELTC$
Panel A. Linear				
0.359	0.099	0.072	0.581	0.248
Panel B. Polynomial				
0.498	0.182	0.071	0.484	0.262

Notes: Shapley decomposition of the R-squared of a synthetic regression of counterfactual log wage changes on: exposure measures ETC , EMC , and LTC (Panel A) and 3rd-degree polynomial of exposure measures (Panel B). Column 1 in each panel reports the R-squared and the three following columns report the share of the R-squared explained by each variable (Panel A) or group of variables (Panel B). Regressions do not include state fixed effects.

Tab. 3.13: Changes in market access and changes in wages in Brazilian microregions. OLS and 2SLS estimates. 1991-2000

Dependent variable: $d \ln wage_{91-00}$				
Variable	(1)	(2)	(3)	(4)
Panel A. OLS				
$d \ln DMA_{91-00}$	0.0886*** (0.0230)	0.0779*** (0.0142)	0.0625*** (0.0203)	0.0458*** (0.0153)
R-squared	0.055	0.329	0.079	0.365
Panel B. 2SLS (IV: $ELTC$)				
$d \ln DMA_{91-00}$	0.173*** (0.0620)	0.166*** (0.0363)	0.138** (0.0572)	0.118*** (0.0447)
R-squared	0.006	0.288	0.051	0.343
Controls			X	X
State FE		X		X

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. N=494. All specifications include an intercept. Controls: share of hours worked in manufacturing in 1991, log distance to Sao Paulo, and log distance to the nearest sea port. Significant at *10%, **5%, ***1%.

Tab. 3.14: Changes in market access and changes in wages in Brazilian microregions.
Other IVs. 2SLS estimates. 1991-2000

Dependent variable: $d \ln wage_{91-00}$				
Variable	(1)	(2)	(3)	(4)
Panel A. 2SLS (IV: <i>ETC</i>)				
$d \ln DMA_{91-00}$	0.154*** (0.0437)	0.168*** (0.0322)	0.102** (0.0467)	0.0868** (0.0370)
R-squared	0.025	0.287	0.071	0.358
Panel B. 2SLS (IV: <i>LTC</i>)				
$d \ln DMA_{91-00}$	0.287*** (0.0750)	0.335*** (0.0888)	12.12 (65.55)	-2.107 (21.80)
R-squared				
Panel C. 2SLS (IV: <i>EMC</i>)				
$d \ln DMA_{91-00}$	0.142*** (0.0337)	0.117*** (0.0156)	0.115*** (0.0317)	0.0642*** (0.0202)
R-squared	0.035	0.320	0.065	0.364
Controls			X	X
State FE		X		X

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. N=494. All specifications include an intercept. Controls: share of hours worked in manufacturing in 1991, log distance to Sao Paulo, and log distance to the nearest sea port. Significant at *10%, **5%, ***1%.

Tab. 3.15: Changes in market access and changes in wages in Brazilian microregions.
First stage estimates. 1991-2000

Dependent variable: $d \ln DMA_{91-00}$				
Variable	(1)	(2)	(3)	(4)
Panel A. <i>ELTC</i>				
<i>ELTC</i>	26.28*** (4.211)	23.24*** (5.755)	28.24** (10.53)	26.58** (12.05)
R-squared	0.375	0.456	0.470	0.517
F-stat	38.96	16.30	71.59	155.4
Panel B. <i>ETC</i>				
<i>ETC</i>	7.842*** (0.784)	7.047*** (0.904)	8.532*** (0.917)	8.230*** (1.086)
R-squared	0.396	0.481	0.465	0.518
F-stat	99.98	60.75	120.9	63.09
Panel C. <i>LTC</i>				
<i>LTC</i>	-3.239*** (0.839)	-2.496*** (0.797)	-0.125 (0.706)	0.136 (1.309)
R-squared	0.105	0.313	0.322	0.408
F-stat	14.89	9.800	23.72	30.91
Panel D. <i>EMC</i>				
<i>EMC</i>	-0.792*** (0.0481)	-0.769*** (0.0474)	-0.832*** (0.0462)	-0.851*** (0.0548)
R-squared	0.594	0.646	0.622	0.664
F-stat	270.7	263.2	104.8	129.3
Controls			X	X
State FE		X		X

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. N=494. All specifications include an intercept. Significant at *10%, **5%, ***1%.

Tab. 3.16: Exposure measures and changes in wages in Brazilian microregions. OLS (reduced form) estimates. 1991-2000

Dependent variable: $d \ln wage_{91-00}$				
Variable	(1)	(2)	(3)	(4)
Panel A. <i>ELTC</i>				
<i>ELTC</i>	4.536*** (1.125)	3.869*** (0.675)	3.908*** (1.001)	3.135*** (0.637)
R-squared	0.079	0.334	0.080	0.367
Panel B. <i>ETC</i>				
<i>ETC</i>	1.205*** (0.388)	1.181*** (0.225)	0.870** (0.392)	0.715** (0.291)
R-squared	0.066	0.340	0.071	0.362
Panel C. <i>LTC</i>				
<i>LTC</i>	-0.928*** (0.181)	-0.836*** (0.181)	-1.512*** (0.384)	-0.286 (0.504)
R-squared	0.061	0.335	0.127	0.357
Panel D. <i>EMC</i>				
<i>EMC</i>	-0.112*** (0.0242)	-0.0903*** (0.0139)	-0.0958*** (0.0279)	-0.0547*** (0.0171)
R-squared	0.084	0.334	0.088	0.364
Controls			X	X
State FE		X		X

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. N=494. All specifications include an intercept. Significant at *10%, **5%, ***1%.

Tab. 3.17: Changes in market access and changes in wages in Brazilian microregions.
Robustness checks. 2SLS estimates. 1991-2000

Dependent variable: log wage change 1991-2000									
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Panel A. Tariff changes 1991-1998									
Variable	I. IV: <i>ELTC</i>			II. IV: <i>ETC</i>			III. IV: <i>LTC</i>		
$d \ln DMA_{91-00}$	0.182** (0.0707)	0.181*** (0.0510)	0.143* (0.0795)	0.155*** (0.0440)	0.164*** (0.0304)	0.0846** (0.0337)	0.231*** (0.0760)	0.307*** (0.0845)	-0.380 (1.506)
R-squared	0.273			0.326			0.290		
Controls	X			X			X		
State FE	X			X			X		
Panel B. No σ_k in exposure measures									
	I. IV: <i>ELTC</i>			II. IV: <i>LTC</i>					
$d \ln DMA_{91-00}$	0.145*** (0.0403)	0.127*** (0.0203)	0.0676*** (0.0205)	0.331*** (0.0710)	0.360*** (0.106)	-3.203 (46.88)			
R-squared	0.033			0.363					
Controls	X			X			X		
State FE	X			X			X		
<i>Notes:</i> Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. N=494. All specifications include an intercept. Significant at *10%. **5%. ***1%.									

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. N=494. All specifications include an intercept. Significant at *10%, **5%, ***1%.

Tab. 3.18: Changes in market access and changes in wages in Brazilian microregions. Robustness checks. OLS estimates. 1991-2000

Dependent variable: Kovak's log wage change 1991-2000				
Variable	(1)	(2)	(3)	(4)
Panel A. <i>ELTC</i>				
<i>ELTC</i>	3.321*	2.403***	-0.104	0.856
	(1.775)	(0.529)	(1.213)	(0.504)
R-squared	0.106	0.733	0.185	0.745
Panel B. <i>ETC</i>				
<i>ETC</i>	0.589	0.583***	-0.658	-0.0621
	(0.646)	(0.200)	(0.507)	(0.224)
R-squared	0.037	0.725	0.200	0.744
Panel C. <i>LTC</i>				
<i>LTC</i>	0.0620	-0.316***	0.197	-0.191
	(0.440)	(0.107)	(0.392)	(0.406)
R-squared	0.000	0.709	0.187	0.744
Panel D. <i>EMC</i>				
<i>EMC</i>	-0.0689*	-0.0505***	-0.0399	-0.0256
	(0.0336)	(0.00595)	(0.0322)	(0.0207)
R-squared	0.122	0.738	0.197	0.747
Controls			X	X
State FE		X		X

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. N=494. All specifications include an intercept. Significant at *10%, **5%, ***1%.

Tab. 3.19: Changes in market access and in-migration rates in Brazilian microregions. Robustness checks. 2SLS estimates.

Dependent variable: in-migration rate in 2000				
Variable	(1)	(2)	(3)	(4)
Panel A. IV: <i>ELTC</i>				
$d \ln DMA_{91-00}$	-0.00214 (0.0128)	-0.0139* (0.00770)	0.00101 (0.00862)	0.00841 (0.00583)
R-squared		0.299	0.668	0.726
Panel B. IV: <i>ETC</i>				
$d \ln DMA_{91-00}$	-0.00556 (0.0157)	-0.0212 (0.0132)	-0.00566 (0.00836)	0.00606 (0.00926)
R-squared		0.296	0.669	0.727
Panel C. IV: <i>LTC</i>				
$d \ln DMA_{91-00}$	0.0515 (0.0363)	0.0176 (0.0315)	-0.0295 (0.337)	-0.593 (4.832)
R-squared		0.286	0.656	
Panel D. IV: <i>EMC</i>				
$d \ln DMA_{91-00}$	-0.00186 (0.0141)	-0.0121 (0.00808)	-0.00590 (0.00636)	0.00643* (0.00371)
R-squared		0.300	0.669	0.727
Controls			X	X
State FE		X		X

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. N=494. Controls: in-migration rate in 1991 and the same set of controls as in previous tables. Significant at *10%, **5%, ***1%.

Tab. 3.20: Estimated wage change implied by an interquartile range change in market access and exposure

	(1)	(2)	(3)	(4)	(5)
	<i>DMA</i>	<i>DMA^C</i>	<i>ELTC</i>	<i>ETC</i>	<i>EMC</i>
Panel A. 25th to 75th percentile					
Δvar	0.39	0.052	0.0094	0.042	0.29
coeff	0.12	0.12	3.2	0.71	-0.085
wage change (%)	4.71	0.62	3.01	2.97	-2.45
Panel B. 25th to 50th percentile					
Δvar	0.087	0.034	0.0061	0.028	0.25
coeff	0.12	0.12	3.2	0.71	-0.085
wage change (%)	1.04	0.41	1.95	1.98	-2.13
Panel C. 50th to 75th percentile					
Δvar	0.20	0.018	0.0033	0.014	0.14
coeff	0.12	0.12	3.2	0.71	-0.085
wage change (%)	2.42	0.21	1.06	0.99	-1.20

Notes: DMA^C is the increase in domestic counterfactual market access as described in the text. Δvar is the change in the variable in either $DMA, DMA^C, ELTC, ETC,$ or EMC in the specified percentile range. Coeff is the estimated coefficient for that variable in the preferred specification. The wage change is equal to the change in the variable times the estimated coefficient.

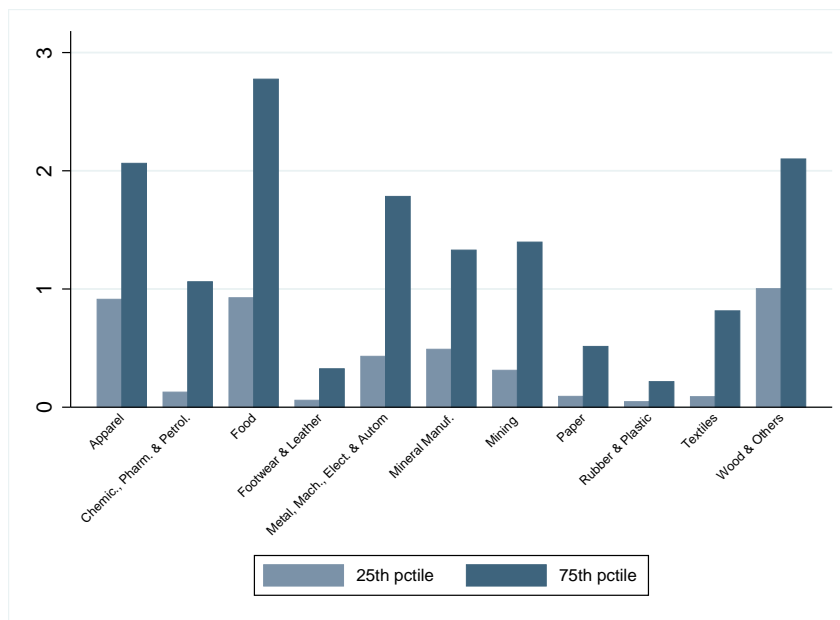
Tab. 3.21: Descriptive statistics of variables in wage regressions

	mean	s.d.	min	max
<u>1991</u>				
$\ln wage$	-0.1857	0.4147	-1.1838	0.9476
DMA	-11.3903	1.6442	-20.5610	-6.9890
In-migration rate	0.1135	0.1106	0.0108	1.5371
<u>2000</u>				
$\ln wage$	0.2689	0.3604	-0.7402	1.2940
DMA	-12.6065	1.2998	-19.2831	-7.9309
In-migration rate	0.0968	0.0677	0.0196	0.7342
<u>Changes 1991-2000</u>				
$d \ln wage_{91-00}$	0.4546	0.1406	-0.0223	0.9794
$d \ln DMA_{91-00}$	-1.4205	0.3735	-2.9435	-0.4070
<u>Exposure in 1991</u>				
$ELTC$	-0.0092	0.0087	-0.0773	0.0010
ETC	-0.0233	0.0300	-0.1172	0.0216
LTC	-0.3422	0.0374	-0.4026	-0.1593
EMC	0.338	0.363	0.045	2.090

Notes: N=494. Wages are in current US dollars. The i-migration rate corresponds to the ratio of hours worked by workers living in microregion i in t that report having live in microregion in $t - 5$ and hours worked by workers that lived in microregion i both in t and in $t - 5$

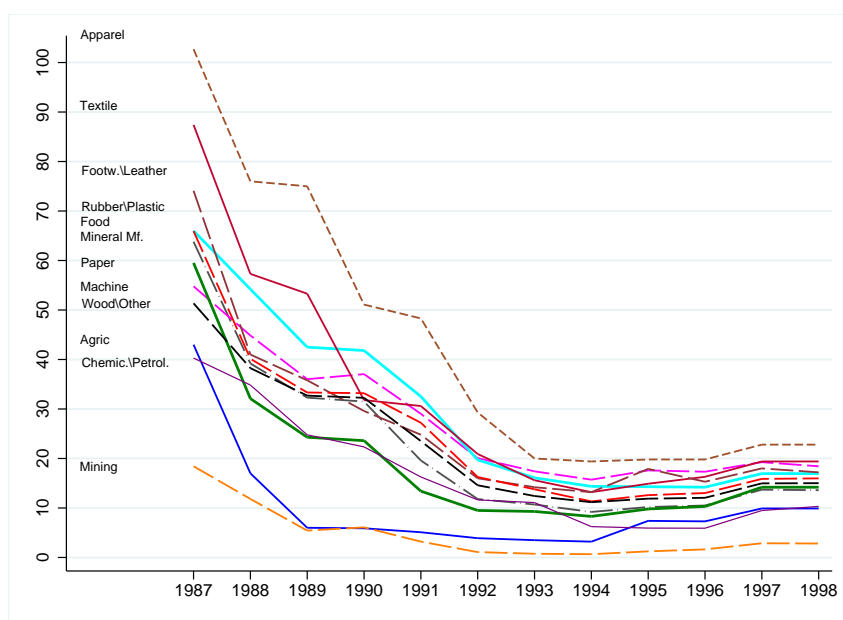
3.7 Figures

Fig. 3.1: Regional share of hours in manufacturing sectors (in %): 25th and 75th pctile



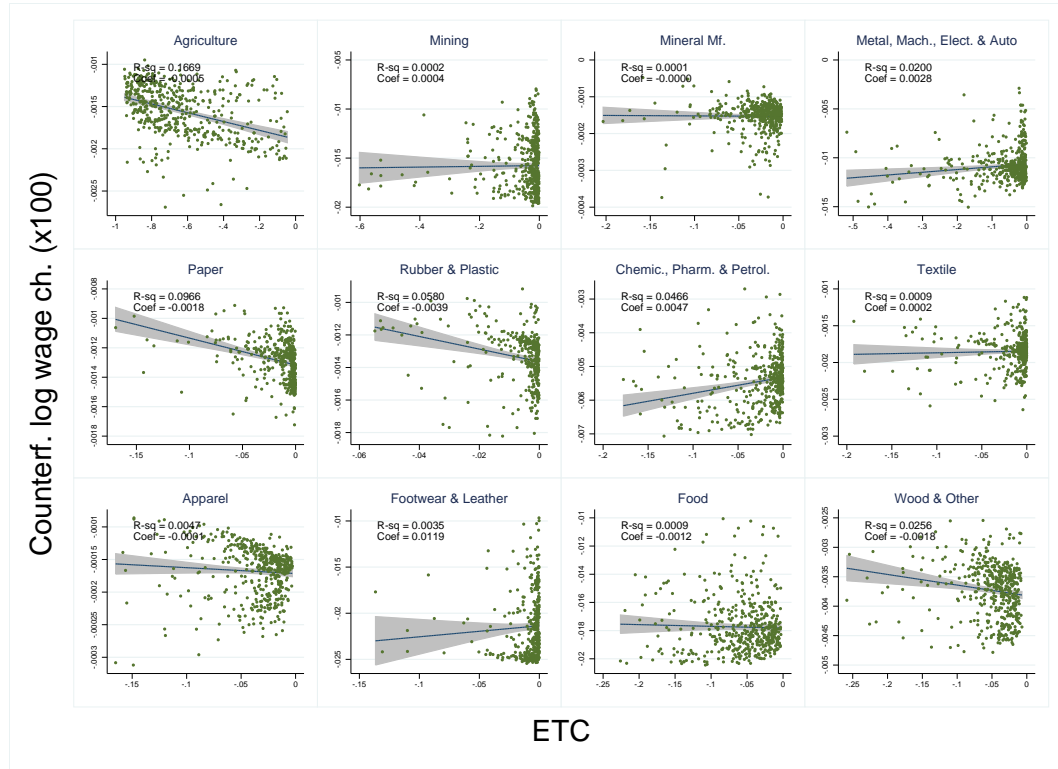
Notes: Share of hours worked in the 10 manufacturing sectors at the 25th percentile and at the 75th percentile of each sector's distribution across Brazilian microregions.

Fig. 3.2: Evolution of applied Brazilian tariffs: 1987-1998



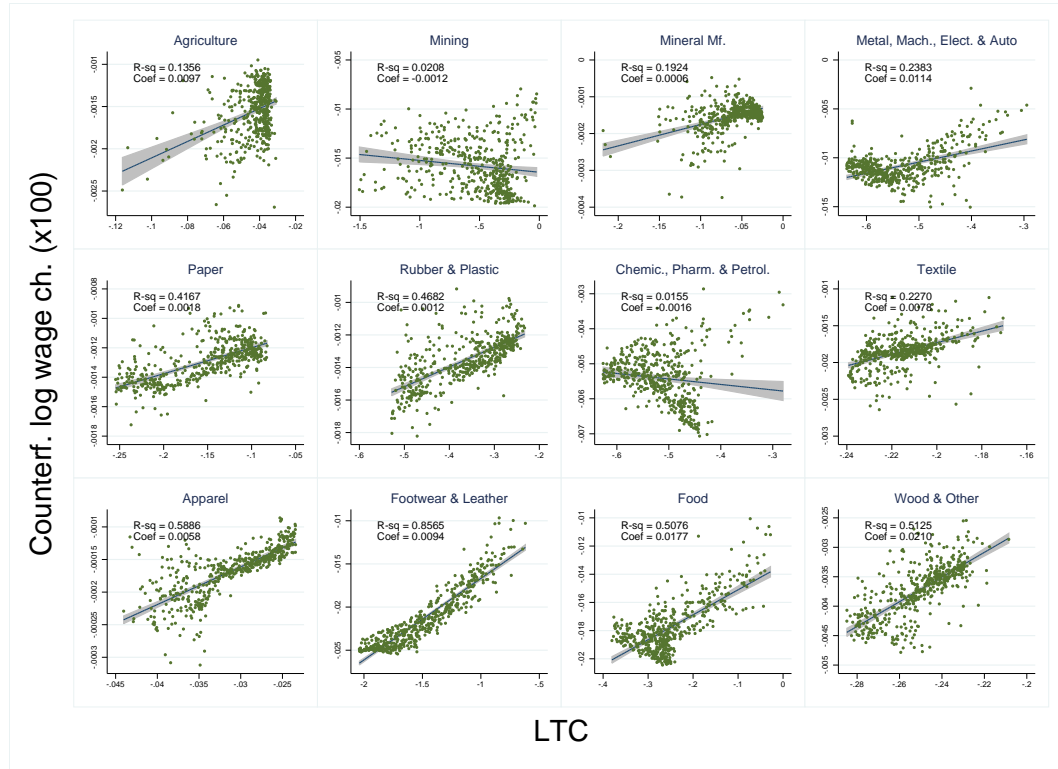
Notes: Ad-valorem tariffs applied by Brazil aggregated to 12 traded sectors as described in Appendix B. Source: Kume et al. (2003)

Fig. 3.3: ETC exposure measure and change in counterfactual wages in Brazilian microregions. 1% shock in each sector.



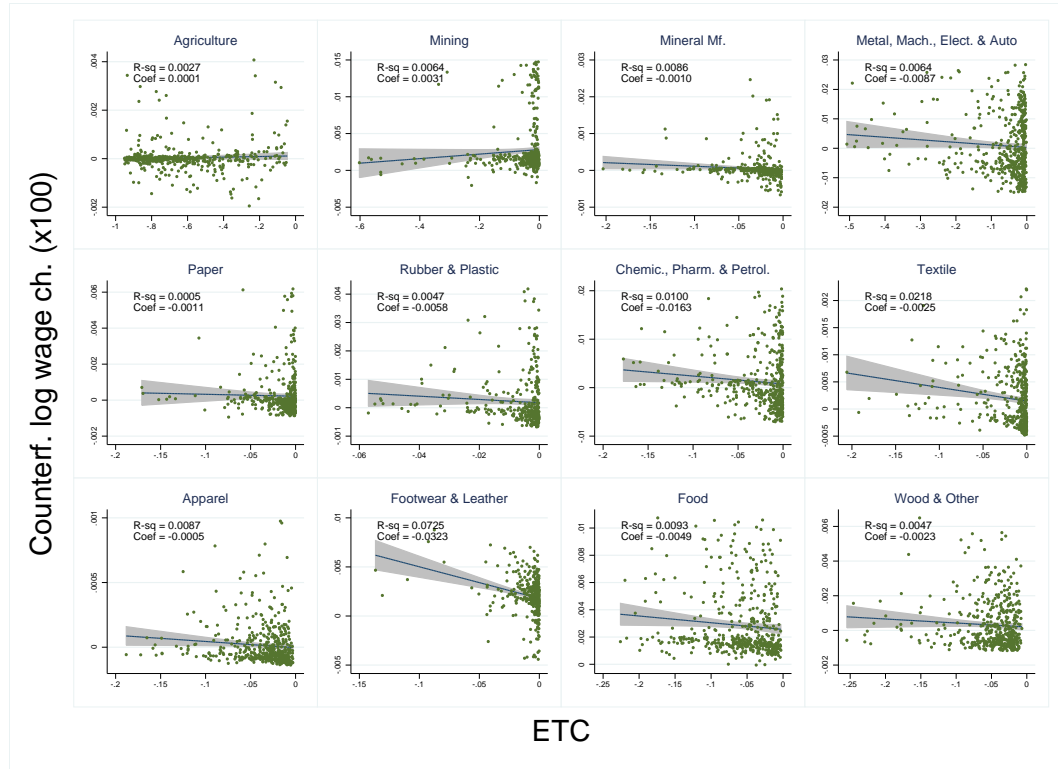
Notes: counterfactual wage changes in log points due to a 1% reduction in the import cost from ROW in each sector. Top and bottom percentile of each variable are trimmed. Fitted OLS regression line in red.

Fig. 3.4: LTC exposure measure and change in counterfactual wages in Brazilian microregions. 1% shock in each sector.



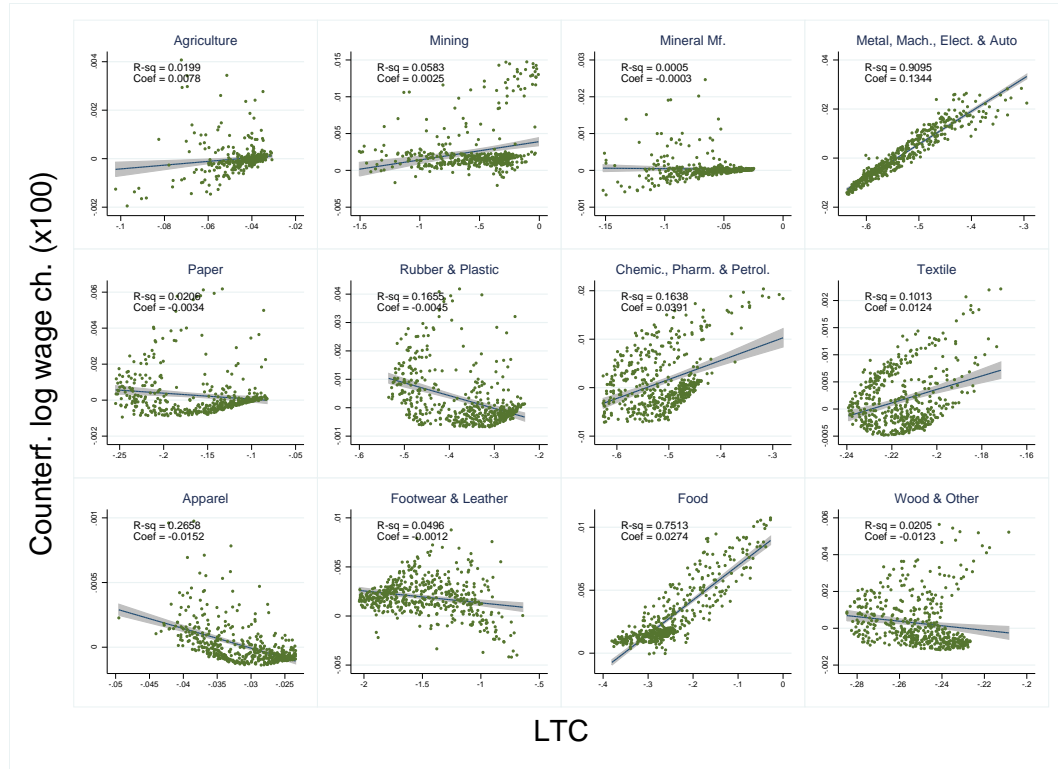
Notes: counterfactual wage changes in log points due to a 1% reduction in the import cost from ROW in each sector. Top and bottom percentile of each variable are trimmed. Fitted OLS regression line in red.

Fig. 3.5: ETC exposure measure and change in counterfactual real wages in Brazilian microregions. 1% shock in each sector.



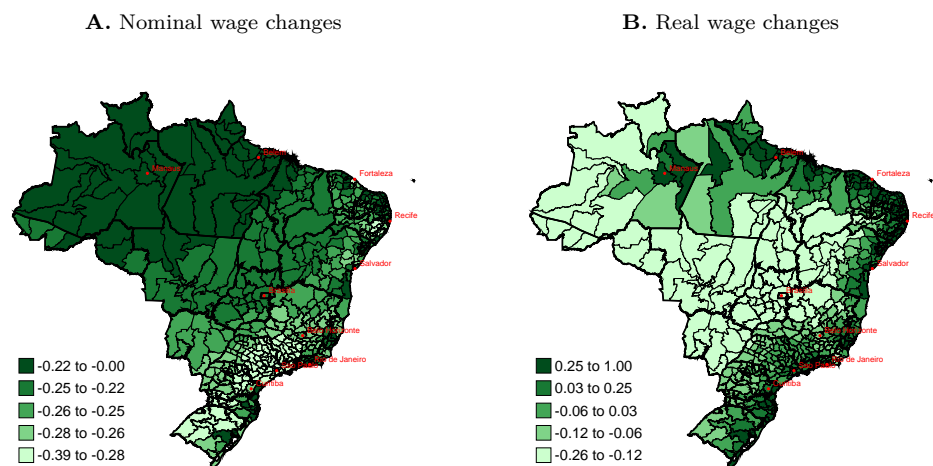
Notes: counterfactual wage changes in log points due to a 1% reduction in the import cost from ROW in each sector. Top and bottom percentile of each variable are trimmed. Fitted OLS regression line in red.

Fig. 3.6: LTC exposure measure and change in counterfactual real wages in Brazilian microregions. 1% shock in each sector.



Notes: counterfactual wage changes in log points due to a 1% reduction in the import cost from ROW in each sector. Top and bottom percentile of each variable are trimmed. Fitted OLS regression line in red.

Fig. 3.7: Counterfactual wage changes in Brazilian microregions resulting from trade liberalization.



Notes: quintiles of counterfactual nominal (Panel A) and real wage changes (Panel B). Wage changes are expressed in log points. 494 microregions. Bold lines delimit the 26 states and the federal district.

Fig. 3.8: Exposure measures in Brazilian microregions

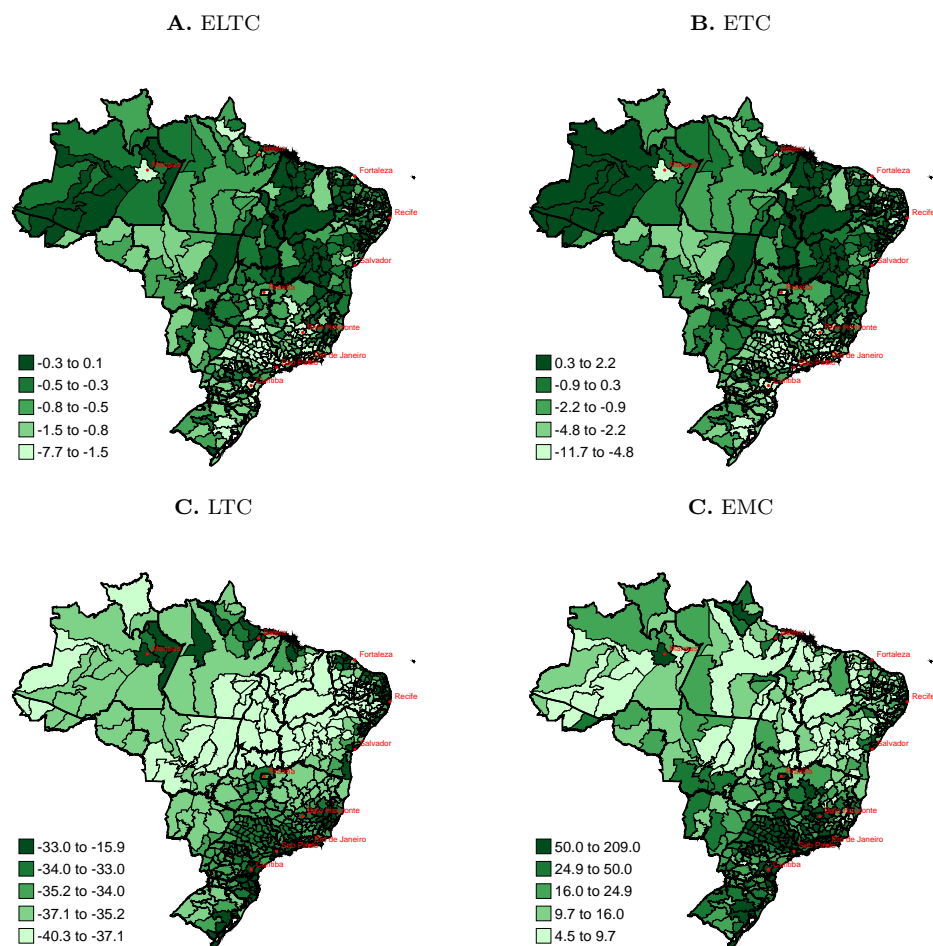
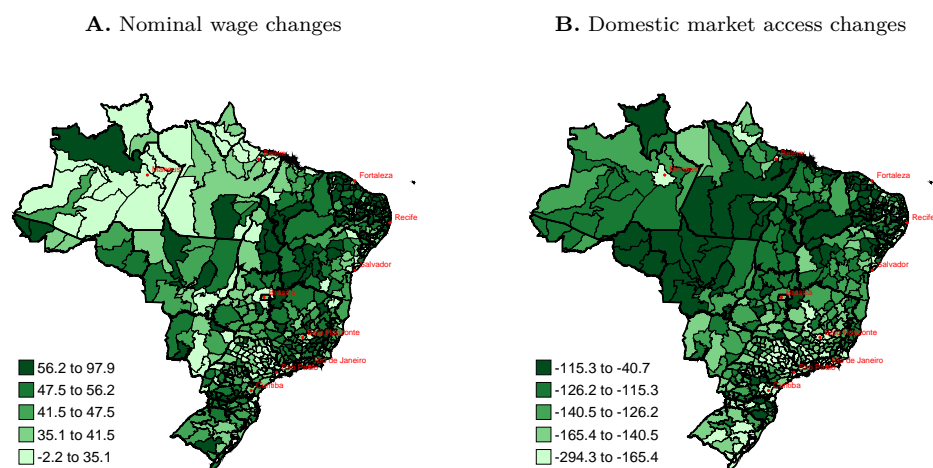
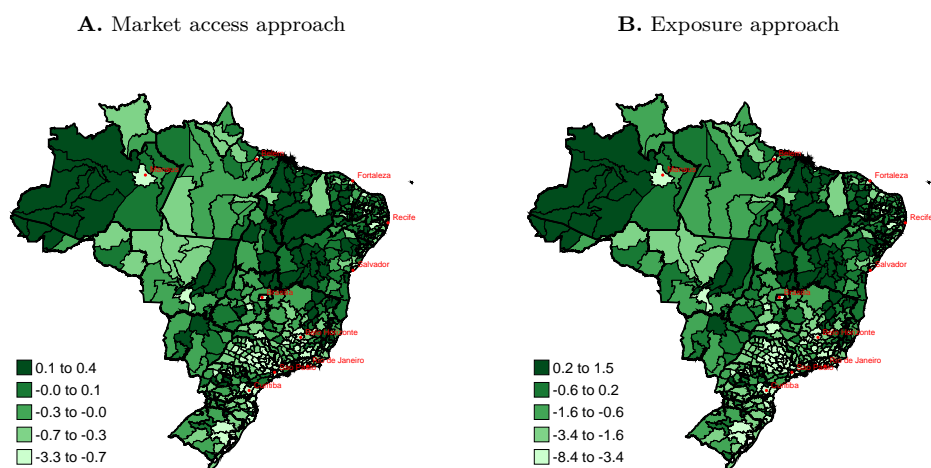


Fig. 3.9: Observed nominal wage changes and market access changes in Brazilian microregions 1991-2000



Notes: quintiles of observed nominal wage changes (in current US dollars). Wage changes are expressed in log points. 494 microregions. Bold lines delimit the 26 states and the federal district.

Fig. 3.10: Predicted wage changes and market access changes in Brazilian microregions 1991-2000



Notes: quintiles of predicted log wage changes. Panel A uses counterfactual market access and the elasticity of wages with respect to market access that is estimated using the IV approach. Panel B uses a regression of wage changes on employment exposure. Wage changes are expressed in log points. 494 microregions. Bold lines delimit the 26 states and the federal district.

Chapter 4: Market access and the skill premium in Brazil

4.1 Introduction

The influence of geographic location on wages—or more generally, on incomes—is a feature shared by several trade models. Regions with better access to markets, save more on transport costs and are able to pay higher wages. However, the influence of location on **relative** wages is much less studied. In section 4.2 I extend the model in Chapter 2 to incorporate two types of workers: skilled and unskilled. In section 4.3 I show descriptive evidence on the skill premium in Brazil and in section 4.4 I propose an estimation strategy. I find regions where domestic market access fell more, experienced a lower decrease in the skill premium. I propose an explanation to this finding based on the relatively higher skill intensity of the non-traded sector in Brazil.

4.2 Model

I extend the model in Chapter 2, which considered homogeneous workers, to incorporate two types of workers: skilled and unskilled. When taking the model to the data I consider high school graduates as skilled and high school dropouts as unskilled.¹

The environment is the same as in Chapter 2, unless otherwise noted. Each region i has a fixed endowment of skilled and unskilled workers given, respectively, by S_i and U_i . For simplicity, I assume that even though there are two types of workers, there is a representative consumer that has the same preferences as in Chapter 2. This allows me to keep track of only price index for each region throughout the analysis. The main difference in this new setup is the production function in traded sectors, which in this version employs two factors. This alters the pricing equation and consequently, the *wage equation*.

Traded sectors As in Chapter 2, traded goods are produced using an increasing returns to scale technology and are subject to classic iceberg trade costs that are composed of shipping costs, Brazilian tariffs, and other trade barriers. The market structure is characterized by monopolistic competition: each firm in each sector k faces a downward sloping demand curve with perceived price elasticity equal to σ_k , the elasticity of substitution in the consumption of varieties of the good.

¹See Section 4.3 for details.

With two types of workers, the cost function of an individual traded good producer in region i and in sector k is given by

$$C(w_i^S, w_i^U, q_i^k) = (w_i^S)^\alpha (w_i^U)^{1-\alpha} (F + q_i^k / \beta_i^k) \quad (4.1)$$

The profit maximization problem faced by the producer of any variety of the traded good k in region i is given by:

$$\begin{aligned} \max_{p_i^k} \quad & \sum_j \frac{p_{ij}^k q_{ij}}{\kappa_{ij}^k} - (w_i^S)^\alpha (w_i^U)^{1-\alpha} (F + q_i / \beta_i^k) \\ \text{s.t.} \quad & q_{ij} = (p_{ij}^k)^{-\sigma} \mu_k E_i P_i^{\sigma-1} \end{aligned} \quad (4.2)$$

where q_{ij}^k is the volume of sales to region j .

As in the one-factor version of the model, the optimal pricing rule is a constant mark-up over marginal costs

$$p_i^k = \left(\frac{\sigma_k}{\sigma_k - 1} \right) (w_i^S)^\alpha (w_i^U)^{1-\alpha} / \beta_i^k, \quad (4.3)$$

the difference with the one factor version is that with two factors prices depend on both the skilled and the unskilled wage rate through the price of the composite input, $(w_i^S)^\alpha (w_i^U)^{1-\alpha}$.

Free entry pins down the scale of each firm which is, as in the one-factor version, equal to: $q_i^k = (\sigma_k - 1) F^k \beta_i^k$.

Finally, aggregate demand for each type of worker in each traded sector is equal to the number of firms in each sector, n_i^k times the labor demand of an individual firm, s_i^k and u_i^k ,

$$S_i^k = n_i^k s_i^k \quad (4.4)$$

$$U_i^k = n_i^k u_i^k. \quad (4.5)$$

Traded sectors equilibrium The traded goods market equilibrium conditions in region i are given by:

$$w_i^S S_i^k + w_i^U U_i^k = \sum_j \pi_{ij}^k \mu_k (w_j^S S_j + w_j^U U_j) \quad (4.6)$$

where the left hand side is the total wage bill paid by sector k to both skilled and unskilled workers which, by free entry, should equal the right hand side given by total revenues summed across all markets j .

Market access, wages, and the skill premium A great part of the empirical literature is concerned with the role of spatial frictions in determining regional wages or incomes. Intuitively, a region that is more isolated spends a higher proportion of its income on transport costs and this imposes a penalty on wages. In a world with more than one factor of production although remoteness imposes a penalty on total income, its effect on relative factor rewards depends on the relative skill intensities across different sectors.

As in the one-factor case, to derive a relationship between wages and market access, sum demand across all destination markets and use the pricing equation (4.3) and the optimal scale of the firm to obtain

$$(w_i^S)^{\alpha_k} (w_i^U)^{1-\alpha_k} = \zeta_i^k (MA_i^k)^{\frac{1}{\sigma_k}} \quad \text{all } k \quad (4.7)$$

where $\zeta_j^k \equiv [(\sigma_k - 1) \beta_j^k]^{\frac{\sigma_k - 1}{\sigma_k}} [F^k (\sigma_k)^{\sigma_k}]^{-1} > 0$ is a combination of parameters and $MA_i^k \equiv \sum_j (\kappa_{ij}^k)^{1-\sigma_k} E_j^k (P_j^k)^{\sigma_k - 1}$ is market access. In relative changes, this equation is given by

$$\alpha_k \hat{w}_i^S + (1 - \alpha_k) \hat{w}_i^U = \frac{1}{\sigma_k} \widehat{MA_i^k} \quad \text{all } k \quad (4.8)$$

Comparing equation 4.7 to the *wage equation* derived in Chapter 2 (equation 3.14), we see it no longer determines the regional average wage but it determines the price of the composite labor input. Equation 4.7 implies that regions with higher market access are able to afford paying higher wages to both skill and unskilled workers. However, this equation is silent about relative factor rewards, w_i^S/w_i^U .

Suppose there is a constant returns to scale non-traded sector, as in Chapter 2. I add two assumptions to those in Chapter 2. First, I assume the non-traded sector is relatively skilled-intensive compared to traded sectors and second, I allow for exogenous trade deficits. If market access decreases, wages of skilled and unskilled workers decrease proportionally but keeping relative wages constant. If we allow for trade deficits, aggregate expenditures do not decrease proportionally to the decrease in wages, triggering an increase in the output of the non-traded sector.² If the non-traded sector is skilled-intensive relative to traded sectors, it absorbs relatively more skilled workers than the traded sector releases, and therefore, for labor markets to

²See Appendix A.7 for a proof.

clear, the relative wage of the skilled has to increase.^{3,4} This resembles a Stolper-Samuelson effect: an increase in market access -and a consequent increase in the relative price of the traded good- benefits the factor used relatively more intensively in the traded sector.⁵

In the following analysis I use the trade liberalization shock in Brazil as an exogenous source of variation in market access to estimate how relative wages respond to changes in market access and I check if the assumption that the non-traded sector is skilled-intensive holds in the data.

4.3 Descriptive evidence

The data sources and variable definitions used in this Chapter are the same as in Chapter 2, except otherwise noted.⁶

I define a worker as skilled if she is a High School Graduate, which in Brazil is equivalent to 11 years of education and as unskilled if she is a High School Dropout. Figure 4.1 shows the share of workers by educational attainment in both census years. High School Dropouts represent 76% of the population in 1991 and 67% in

³If there are no trade deficits, the output of the non-traded sector is pinned down by the magnitude of the relative wage between skilled and unskilled, as I show in equation A.12 in Appendix A.7. Therefore, in equilibrium, an increase in market access increases skilled and unskilled wages in the same proportion, $\hat{w}_i^S = \hat{w}_i^U = \widehat{MA}_i^k$, leaving the skill premium unchanged.

⁴Redding and Schott (2003) assume there is a residual freely-traded sector that is unskilled-intensive in production. This implies skilled and unskilled wages have to change in opposite directions. As I show in Appendix A.7, increases in market access would lead to an increase in the wage of the skilled and a decrease in the wage of the unskilled.

⁵However, there is no “magnification effect”. This is, skilled wages do not decline, they just grow less than unskilled wages. Francois and Nelson (1998) show in a 2x2 Heckscher-Ohlin model that introducing non-homogeneous goods breaks down the classical “magnification effect”.

⁶A complete description of sources and definitions can be found in Section 3.2 and in Appendix B.

2000 and High School Graduates represent 24% of the population in 1991 and 33% in 2000. The proportion of workers with at least some college education is low: 9% in 1991 and 11% in 2000. To minimize the problem of having microregions with a small number skilled workers, my definition of a skilled worker is a High School Graduate instead of a worker with Some College.

The new variables I include in this section for each Brazilian microregion i are: skilled and unskilled wages, w_i^S and w_i^U , total hours worked by the skilled and the unskilled, S_i and L_i , and hours worked by sector, S_i^k and L_i^k . These variables come from the Population Censuses of 1991 and 2000 and I calculate them as follows. First, I defined an individual worker as skilled if she has completed at least 11 years of education, and as unskilled otherwise. Second, I assign individuals to microregions and I obtain S_i^k and L_i^k as the sum of yearly hours worked across skilled and unskilled workers in region i and sector k and w_i^S and w_i^U , as the average hourly wage across each type of worker in region i . Finally, for each region I obtain the skill premium, this is the log of the relative wage of the skilled and the unskilled.

Table 4.6 shows the wages of the skilled and the unskilled have substantial dispersion across regions both in 1991 and in 2000. Panel A shows unskilled workers have in 1991 wages that range from 0.3 dollars per hour (usd/h) to 1.6 usd/h in 1991 with an average of 0.7 usd/h. Skilled workers have substantially higher wages ranging from 0.5 usd/h to 4.6 usd/h with an average of 1.9 usd/h. Variation in skilled and unskilled wages across regions is not surprising given the results in Chapter 2: market access varies substantially across regions and is an important determinant of wages. What is surprising is that **relative** wages vary considerably

across regions. The first row of the table shows the skill premium in 1991 had an average value of 102 log points. This is, skilled workers have wages in the average microregion more than twice as high as unskilled workers. However, the values range from 33 log points in the microregion with the lowest skill premium to 214 log points in the region with the highest skill premium. The maps in Figure 4.2 shows that the skill premium tends to be higher in the South East and North East regions and lower in the more remote West and Central regions. Considering unskilled and skilled wages separately, figures 4.3 and 4.4 show that both skilled and unskilled wages tend to be lower in the North East region.

In terms of changes, Panel C of Table 4.6 shows that both skilled and unskilled wages increased on average between 1991 and 2000. However, the average skill premium decreased by 8.5 log points. This evidence is consistent with Gonzaga et al. (2006), who find a decrease in the skill premium during this decade using household survey data. Figure 4.5 shows the skill premium decreased in most regions, especially in the North West and South East regions. The regions where the skill premium increased are mostly concentrated in the Central and North East regions. Figure 4.6 shows that skilled and unskilled wages did not change in the same proportion across regions.

To shed light into the determinants of changes in the regional skill premium, in the next section I estimate the impact of changes in market access on the skill premium and on skilled and unskilled wages using the trade liberalization episode as a source of exogenous variation in market access.

4.4 Estimation

4.4.1 Empirical approach

Based on equation 4.8 I estimate the following

$$d \ln Y_{i,91-00} = \beta_0 + \beta_1 d \ln DMA_{i,91-00} + \nu_i^k \quad (4.9)$$

where $Y_{i,91-00}$ is the change in the outcome of interest in region i between 1991 and 2000 and $d \ln DMA_{i,91-00}$ is the change in domestic market access as defined in Chapter 2 equation 3.20. The outcome of interest can be the change in the skill premium, $\widehat{SP}_i \equiv d \ln(w_i^S/w_i^U)$, the change in skilled wages, \widehat{w}_i^S , or the change in unskilled wages, \widehat{w}_i^U .

As I explained in Chapter 2 Section 3.4, changes in domestic market access are endogenous and need to be instrumented for. The instrument I propose is

$$ELTC_i = \sum_k \frac{L_{i,91}^k}{L_{i,91}} (\sigma_k - 1) \left(\sum_{j \neq R} \xi_{ij,91}^k \pi_{Rj,91}^k \right) \widehat{\tau}_{91,95}^k, \quad (4.10)$$

where L_i is equal to total hours worked in a region, including both skilled and unskilled hours.

4.4.2 Results

Table 4.1 shows the estimation of equation 4.9. The OLS estimates in Panel I are insignificant. However, they become negative and significant when I use $ELTC$ as

an instrument and I control for state fixed effects. This is my preferred specification for three reasons. First, by being in first differences it eliminates any locational advantage that is constant in time. Second, by using *ELTC* as an instrument it eliminates any bias coming from unobserved changes in regional productivity. Third, by controlling for state fixed effects it controls for any variation in policies that vary across states and that affect wages. This specification suggests a strong and negative effect of increases in domestic market access in regional wage inequality. A doubling of domestic market access reduces the skill premium by 13.4 log points. For the average region this implies a reduction from 102 log points to 88.6 log points. One caveat when interpreting this results is that changes in market access are not exogenous. A given shock to trade costs increases market access differently in each region.

To understand the source of the negative elasticity of the skill premium to market access, I estimate the impact of changes in market access on skilled and unskilled wages separately. Table 4.2 shows the results. My preferred specification, in column (4), shows that the unskilled are more affected by changes in market access than the skilled.⁷ So, a doubling of domestic market access increases the wages of the unskilled by 21 log points (Panel B) while increasing the wages of the skilled by 7.5 log points, which results in a decrease of the skill premium. Columns 1 to 3 show similar elasticities for skilled and unskilled workers and this is why in Table 4.1, the effect of changes of market access on the skill premium is insignificant. The difference

⁷The null hypothesis that the elasticities are the same for both specification is rejected at a 1% level of confidence.

between the OLS estimate the 2SLS estimate is much larger for the unskilled. This means that unobserved factors that affected the wages of the unskilled between 1991 and 2000 were strongly negatively correlated with changes in domestic market access during the same period.

As I explained in section 3.4.3, market access does not change in proportion to tariffs. Therefore, to evaluate the economic significance of the estimates, I simulate the change in domestic access that results from a change in tariffs as the one observed between 1991 and 1999. A change from the 75th to the 25th percentile in the distribution of the change in *DMA* implied by the tariff shock, yields an increase in the skill premium of 0.7 log points. Since the skill premium decreased on average in Brazil during the 1990s, this result means that regions where domestic market access fell more as a result of trade liberalization, experienced a smaller decline in the skill premium.

Tables 4.3 and 4.4 show estimation results for the skilled premium and skilled and unskilled wages using the other three instrumental variables proposed in Chapter 2. Panel I of Table 4.3 reports that using *ETC*, the standard employment-weighted tariff change, yields a negative skill premium elasticity with respect to market access changes that is of a similar magnitude to the one I obtained using my preferred instrument, *ELTC*. The other two instruments: *EMC* and *LTC* do not yield significant estimates for this elasticity. Panel A of Table 4.4 shows that none of the instruments yields a significant elasticity estimate for the skilled while Panel B shows that the elasticity of unskilled wages with respect to domestic market access is positive and significant with values that range from 0.2 to 0.36. Therefore,

considering other instruments does not alter the baseline result that an increase in market access reduces the skill premium in Brazil.⁸

Finally, Table 4.5 reports the results from the reduced form regressions. These regressions resemble the ones estimated in the exposure approach. The dependent variable is the outcome of interest and it is regressed on the different measures of exposure. Note that more exposure to a decrease in tariffs as measured by *ETC* and *ELTC* yields an increase in the skill premium and a decrease in unskilled wages more pronounced than the decrease in skilled wages.⁹ This is consistent with the market access results in the previous paragraphs: a decrease in tariffs decreases market access leading to an increase in the skill premium through a decrease in the wages of the unskilled higher than the decrease in the wages of the skilled.

Comparing these results with the literature, Dix-Carneiro and Kovak (2014) find a small but negative effect of trade liberalization on the skill premium, which is opposite to my findings. There are two main differences between their empirical approach and mine. First, they measure the regional skill premium by estimating returns to education that vary across regions, and predicting wage gaps using national-level returns to other worker characteristics. With this methodology they find that the aggregate skill premium increased in Brazil, which is opposite to the findings in Gonzaga et al. (2006) who find a decrease in the national skill premium during the same period using Household Survey data. Second, their measure of exposure weights by the difference in the share of skilled and unskilled workers em-

⁸The first stages estimates are the same as the ones reported in Table 3.15.

⁹Since tariff changes were negative, a decrease in weighted tariffs implies an increase in the skill premium and a decrease in skilled wages.

ployed in each sector while my instrument does not. The reason is that their measure is derived from a specific factors model with no trade costs and homogeneous goods. I employ a market access approach which differs fundamentally to their approach by taking into account trade linkages between regions.

4.4.3 Mechanism

The results in the previous section are consistent with the findings in Chapter 2: a decrease in market access leads to a decrease in the wages of both skilled and unskilled workers. However, the wages of the unskilled decrease more. Why is this the case?

One possible explanation for the positive elasticity of the skill premium with respect to market access is that the non-traded sector is skilled-intensive. Trade liberalization in Brazil generated a decrease in domestic market access and a shrinkage of the manufacturing sector relative to the non-traded sector. If the non-traded sector is skilled-intensive relative to manufacturing, the relative wage of the skilled must increase. For this to be a possible mechanism, the evidence must show two things: that it is indeed the case that the non-traded sector is skilled-intensive relative to manufacturing and that the non-traded sector increased in size after trade liberalization.

I measure the skill intensity of a sector k in two alternative ways. The first measure is an hours-based measure, SI_k^H , given by

$$SI_k^H = \frac{S_k}{S_k + U_k}. \quad (4.11)$$

This measure implies that sector k is relatively more skilled-intensive than sector s if the share of total hours worked by skilled workers is higher than in sector s .

The second measure is a wage bill-based measure, SI_k^W , given by

$$SI_k^W = \frac{w_k^S S_k}{w_k^S S_k + w_k^U U_k}. \quad (4.12)$$

This measure implies that sector k is relatively more skilled-intensive than sector s if the share of the total sector's wage bill paid to skilled workers is higher than in sector s .

Figure 4.7 shows the skill intensity by sector in 1991 measured in these two ways. The non-traded sector has a skill intensity of 0.28 as measured in hours and of 0.55 as measured in terms of the wage bill. Using the hours-based measure, there are only two individual sectors that are more skilled-intensive than the non-traded sector: Chemicals, with a skill intensity of 0.29 and Paper, with a skill intensity of 0.31. Using the wage bill measure, only Chemicals is more skilled-intensive than non-traded, with an intensity of 0.63. Since the trade liberalization shock affected the whole manufacturing sector at the same time, the relevant comparison is with the skill intensity of the whole manufacturing sector.¹⁰ The last two columns show

¹⁰I leave Agriculture and Mining out of the comparison for two reasons: first they are relatively skill un-intensive, so in any case they would lower the traded sector average, and second tariffs in agriculture increased and tariffs in Mining decreased only mildly. So the bulk of the reduction in domestic market access is due to changes in manufacturing tariffs (see Table 3.4).

that using both measures, the non-traded sector is more skilled-intensive than the manufacturing sector.

Table 4.7 shows the change in each sector's share of total hours and of total wage bill in the economy between 1991 and 2000. The non-traded sector increased its relative size both measured as its share of hours and its share of the wage bill at the expense of the rest of sectors in the economy, Agriculture, Mining, and Manufacturing. Although this increased importance of the non-traded sector could be due to other factors, like changes in preferences or technology, it is consistent with the mechanism I proposed: increased competition as a consequence of trade liberalization led to an increase in the size of the non-traded sector, a reduction in the size of manufacturing, and therefore, an increase in the skill premium.

Finally, one aspect of the empirical strategy to take into account is that it is not meant to explain the country wide evolution of the skill premium but the relative evolution across regions. The country-wide skill premium decreased in Brazil, which could be due to other factors. The model I estimate simply says that besides the overall decreasing trend in the skill premium, regions that experienced a higher decrease in domestic market access as a result of increased competition from the rest of the World, experienced a smaller decrease in the skill premium than they would have otherwise experienced. So that increased openness to international trade contributed to an increase in within-region inequality as measured by the skill premium.

4.5 Conclusions

In this Chapter, I extended the model in Chapter 2 to allow for different levels of education across workers. I proposed an estimation strategy to explore the impact of changes in market access on changes in the skill premium. I found that decreased domestic market resulting from trade liberalization resulted in a higher skill premium in Brazilian microregions and that this is explained by a higher elasticity of unskilled wages to market access compared to skilled wages. I show empirical evidence that supports the non-traded sector is relatively skilled-intensive compared to the rest of the sectors in Brazil, which can explain my findings.

4.6 Tables

Tab. 4.1: Changes in domestic market access and changes in the skill premium in Brazilian microregions. OLS and 2SLS estimates. 1991-2000

Dependent variable: log skill premium change 1991-2000				
Variable	I. OLS		II. 2SLS (IV: <i>ELTC</i>)	
	(1)	(2)	(3)	(4)
$d \ln DMA_{91-00}$	0.0113 (0.0204)	-0.0198 (0.0187)	-0.0272 (0.0291)	-0.134*** (0.0418)
R-squared	0.000	0.136		0.103
State FE		X		X

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. N=494. All specifications include an intercept. Significant at *10%, **5%, ***1%.

Tab. 4.2: Changes in domestic market access and changes in wages in Brazilian microregions. OLS and 2SLS estimates. 1991-2000

	I. OLS		II. 2SLS (IV: <i>ELTC</i>)	
	(1)	(2)	(3)	(4)
Panel A. Dependent variable: log skilled wages change 1991-2000				
$d \ln DMA_{91-00}$	0.0936*** (0.0327)	0.0691*** (0.0175)	0.144** (0.0720)	0.0749** (0.0306)
R-squared	0.030	0.254	0.021	0.254
Panel B. Dependent variable: log unskilled wages change 1991-2000				
$d \ln DMA_{91-00}$	0.0822*** (0.0276)	0.0889*** (0.0120)	0.171** (0.0789)	0.208*** (0.0549)
R-squared	0.044	0.342		0.274
State FE		X		X

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. N=494. All specifications include an intercept. Significant at *10%, **5%, ***1%.

Tab. 4.3: Changes in domestic market access and changes in the skill premium in Brazilian microregions. 2SLS estimates. 1991-2000

Dependent variable: log skill premium change 1991-2000						
Variable	I. IV: <i>ETC</i>		II. IV: <i>EMC</i>		III. IV: <i>LTC</i>	
	(1)	(2)	(3)	(4)	(5)	(6)
$d \ln DMA_{91-00}$	-0.0453 (0.0342)	-0.143*** (0.0322)	-0.130 (0.0962)	-0.215 (0.142)	0.141 (0.174)	0.393 (0.573)
R-squared		0.097		0.038		
State FE		X		X		X

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. N=494. All specifications include an intercept. Significant at *10%, **5%, ***1%.

Tab. 4.4: Changes in domestic market access and changes in wages in Brazilian microregions. 2SLS estimates. 1991-2000

Variable	I. IV: <i>ETC</i>		II. IV: <i>EMC</i>		III. IV: <i>LTC</i>	
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. Dependent variable: log skilled wage change 1991-2000						
$d \ln DMA_{91-00}$	0.0966 (0.0677)	0.0570 (0.0428)	0.123 (0.0807)	0.142 (0.121)	-0.0554 (0.310)	0.256 (0.279)
R-squared	0.030	0.254	0.027	0.241		0.168
Panel B. Dependent variable: log unskilled wage change 1991-2000						
$d \ln DMA_{91-00}$	0.142** (0.0552)	0.200*** (0.0282)	0.253*** (0.0833)	0.358*** (0.103)	-0.196 (0.382)	-0.137 (0.418)
R-squared	0.021	0.283				0.100
State FE		X		X		X

Notes: Robust standard errors reported in parentheses below coefficient estimates. Standard errors clustered by state. N=494. All specifications include an intercept. Significant at *10%, **5%, ***1%.

Tab. 4.5: Exposure and changes in the skill premium and wages in Brazilian microregions. Reduced form estimates. 1991-2000

	(1)	(2)	(3)	(4)
Panel A. Dependent variable: log change in skill premium 1991-2000				
<i>ELTC</i>	-3.104*** (0.486)			
<i>ETC</i>		-1.009*** (0.263)		
<i>LTC</i>			0.537 (0.358)	
<i>EMC</i>				-0.0801 (0.0518)
R-squared	0.456	0.481	0.313	0.270
Panel B. Dependent variable: log skilled wage change 1991-2000				
<i>ELTC</i>	1.740** (0.708)			
<i>ETC</i>		0.401 (0.311)		
<i>LTC</i>			-0.356 (0.312)	
<i>EMC</i>				-0.0522 (0.0615)
R-squared	0.246	0.245	0.246	0.243
Panel C. Dependent variable: log unskilled wage change 1991-2000				
<i>ELTC</i>	4.844*** (0.624)			
<i>ETC</i>		1.410*** (0.221)		
<i>LTC</i>			-0.893*** (0.213)	
<i>EMC</i>				0.0280 (0.0598)
R-squared	0.358	0.360	0.344	0.305

Notes: Robust standard errors reported in parentheses below coefficient estimates. N=494. All specifications include an intercept and state fixed effects. Standard errors clustered by state. Significant at *10%, **5%, ***1%.

Tab. 4.6: Descriptive statistics of variables in regressions

	mean	s.d.	min	max
<u>Panel A: 1991</u>				
$\ln(w^S/w^U)$	1.0187	0.2037	0.3342	2.1346
w^S in usd	1.9316	0.6893	0.5112	4.5144
w^U in usd	0.6919	0.2447	0.2951	1.6226
DMA	-11.3903	1.6442	-20.5610	-6.9890
<u>Panel B: 2000</u>				
$\ln(w^S/w^U)$	0.9339	0.1857	0.4603	1.5723
w^S in usd	2.5454	0.8180	1.0578	6.8336
w^U in usd	0.9954	0.2995	0.3842	1.9687
DMA	-12.6065	1.2998	-19.2831	-7.9309
<u>Panel C: Changes 1991-2000</u>				
$\ln(w^S/w^U)$	-0.0848	0.1998	-0.9416	0.9064
w^S in usd	0.2941	0.2028	-0.5435	1.2125
w^U in usd	0.3789	0.1471	-0.1599	0.9287
$d \ln DFMA_{91-00}$	-1.4205	0.3735	-2.9435	-0.4070
<u>Panel D: Exposure in 1991</u>				
ETC	-0.0233	0.0300	-0.1172	0.0216
EMC	0.9667	0.1455	0.0353	1.3625
LTC	-0.3422	0.0374	-0.4026	-0.1593
$ELTC$	-0.0092	0.0087	-0.0773	0.0010

Notes: The number of observations is 494 for every variable. Wages are in current US dollars.

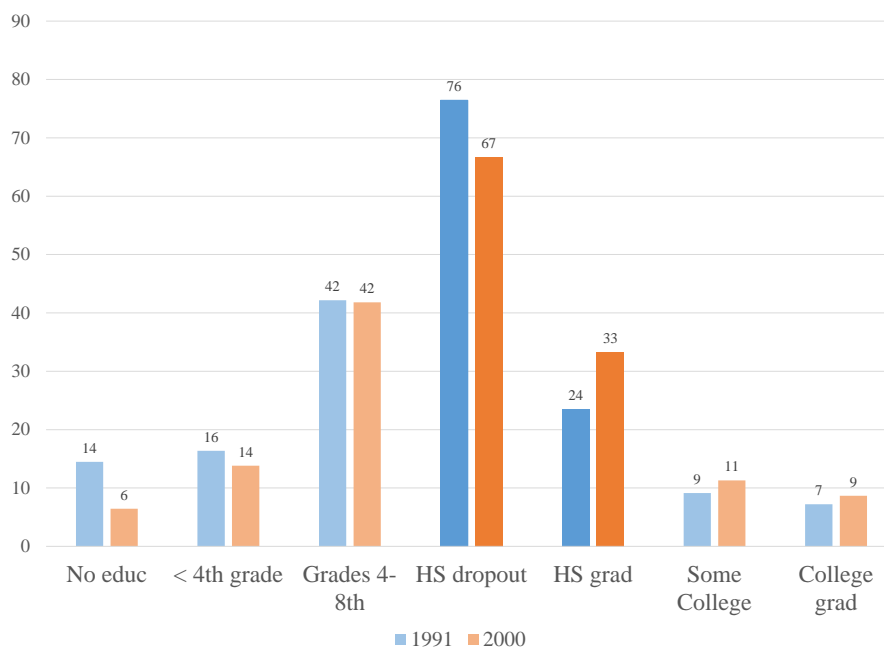
Tab. 4.7: Share of hours and share of wage bill by sector, 1991 and 2000

	share of hours (%)		share of wage bill (%)	
	1991	2000	1991	2000
Agriculture	23.84	18.27	8.90	6.99
Mining	1.07	0.50	1.12	0.49
Manufacturing	16.47	14.77	18.99	15.31
Non-Traded	58.61	66.46	70.99	77.21

Notes:

4.7 Figures

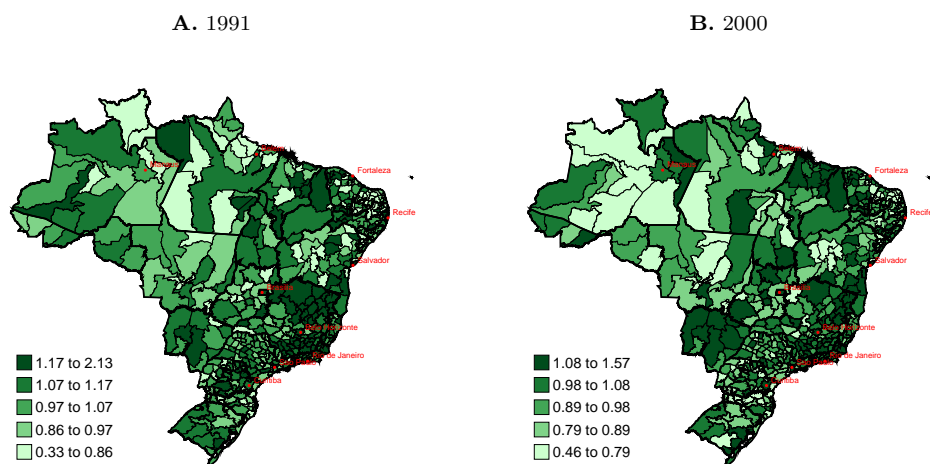
Fig. 4.1: Share of workers by educational attainment (%)



Notes: Workers aged 18-55 not enrolled in school. Years of education for each category: No education: 0; Less than 4th grade: 1-3; Grades 4-8th: 4-8; High School dropout: 0-10; High School grad: 11+; Some College: 12+; College graduate: 15+. The share of individuals who started High School but did not finish it is 3% in 1991 and 5% in 2000.

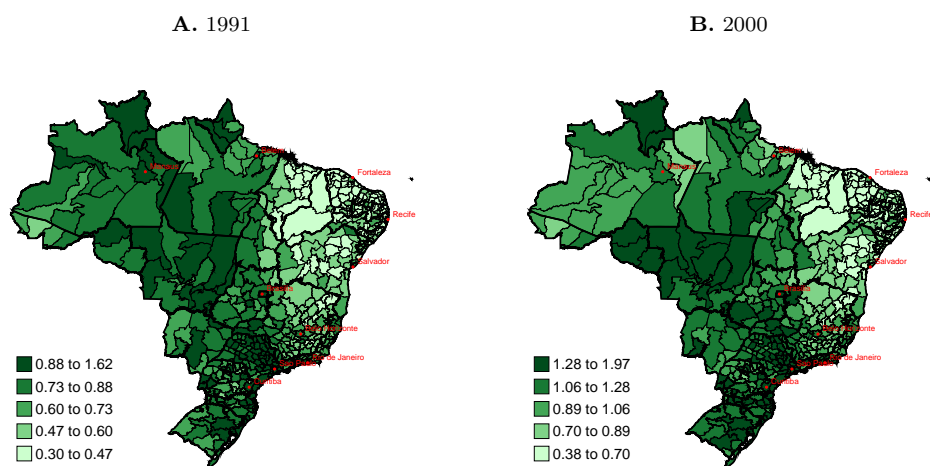
Source: Brazilian Population Census 1991 and 2000 (IBGE).

Fig. 4.2: Skill premium in Brazilian microregions: 1991 and 2000



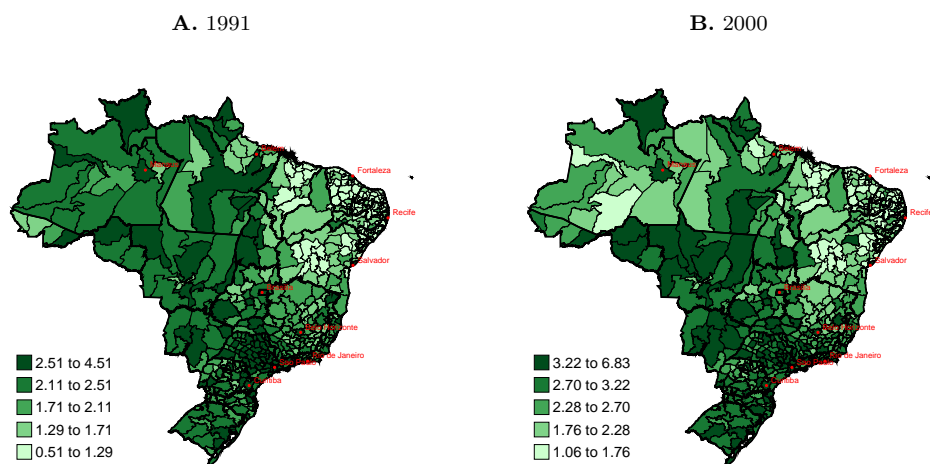
Notes: quintiles of the skill premium in 1991 (Panel A) and 2000 (Panel B). 494 microregions. Bold lines delimit the 26 states and the federal district.

Fig. 4.3: Unskilled wages in Brazilian microregions: 1991 and 2000



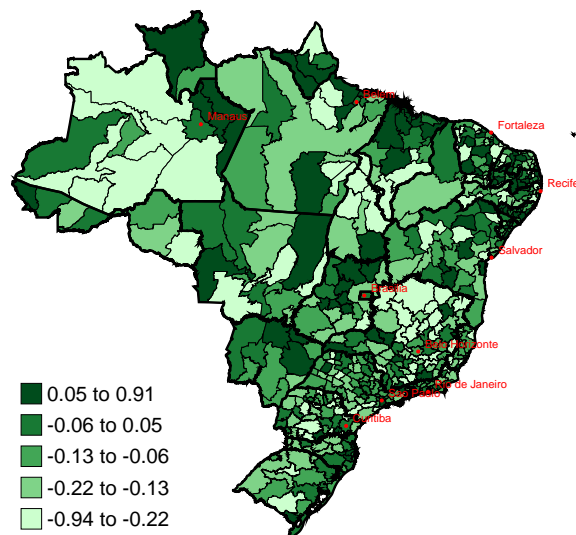
Notes: quintiles of the unskilled wages in 1991 (Panel A) and 2000 (Panel B). 494 microregions. Bold lines delimit the 26 states and the federal district.

Fig. 4.4: Skilled wages in Brazilian microregions: 1991 and 2000



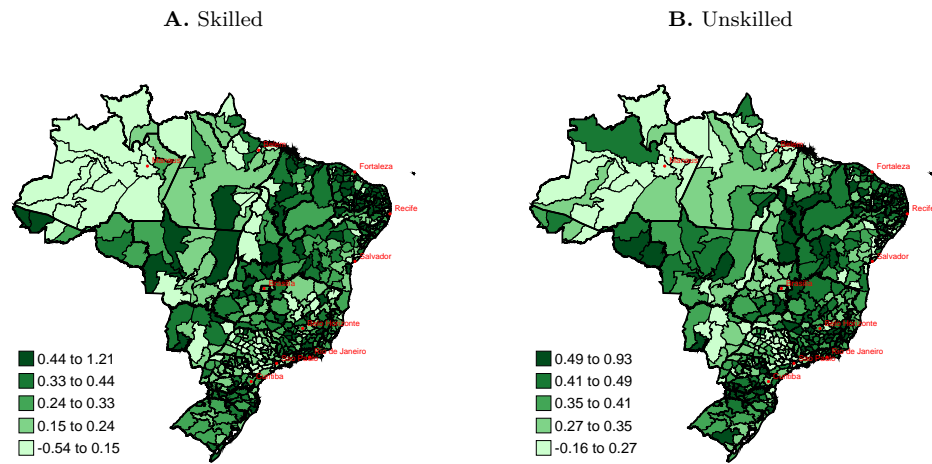
Notes: quintiles of the skilled wages in 1991 (Panel A) and 2000 (Panel B). 494 microregions. Bold lines delimit the 26 states and the federal district.

Fig. 4.5: Skill premium changes in Brazilian microregions: 1991-2000



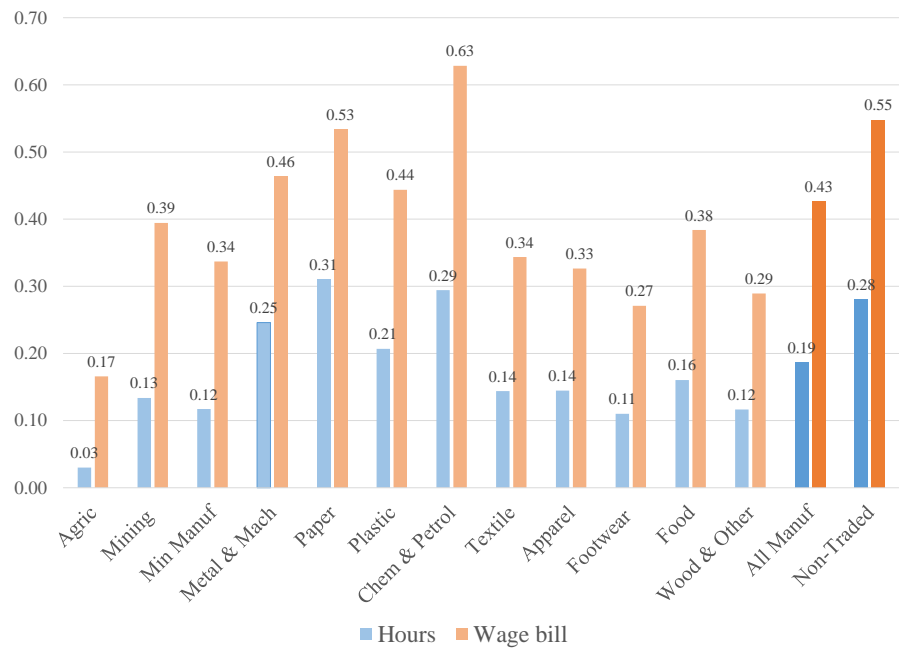
Notes: quintiles of skill premium changes expressed in log points. 494 microregions. Bold lines delimit the 26 states and the federal district.

Fig. 4.6: Skilled and unskilled wage changes in Brazilian microregions: 1991-2000



Notes: quintiles of skilled (Panel A) and unskilled (Panel B) wage changes between 1991 and 2000. 494 microregions. Bold lines delimit the 26 states and the federal district.

Fig. 4.7: Skill intensity by sector in 1991



Notes: Source: Brazilian Population Census 1991 (IBGE).

Chapter 5: Conclusion

In this dissertation I analyze how trade linkages across local labor market shape the response of local wages to nationwide trade shocks. In doing so, I reinterpret the reduced form approach within a standard model of trade. I derive a measure of exposure from the model and I show this measure explains more of the variation in counterfactual wage changes than measures based on employment composition. I also propose an estimation approach with which I show that employment-based measures of exposure used in the reduced form literature overstate the negative impact of trade liberalization in Brazil.

However, my analysis has several limitations. First, I assumed away any migration between regions. Even though I showed changes in market access do not have a significant effect on the flow of migrants to a region, it would be useful to incorporate migration in the model and analyze how it can potentially alter the results. This can be even more important in the fourth chapter where I consider workers of different skills since it is widely known that mobility varies across skill groups. Second, I assumed there are no input-output linkages across sectors. These linkages have the potential to generate spillovers across regions through changes in input prices. In addition, it is likely that a great part of trade within the country is

trade in intermediates. Although data limitations prevent me from distinguishing between intermediate and final goods trade, this is still an important issue that should be addressed in future research. Third, a limitation of the standard trade model I use, is that it assumes full employment. One of the biggest burdens for workers after trade reform is unemployment. It would be interesting to incorporate this friction in the model and use variation in regional unemployment rates to estimate the impact of tariff changes in unemployment. Finally, a limitation of my empirical strategy is that I only consider changes in import competition without analyzing the effect that the reduction of Brazil's trading partners' tariffs had on wages. I leave the analysis of changes in foreign market access for future research.

In a world that is becoming increasingly interconnected, understanding which regions are most sensitive to negative trade shocks is crucial to designing the right policies to mitigate the negative consequences of trade.

APPENDIX

Appendix A: Theory

A.1 Derivation of the wage elasticity

Consider a region $i \in Home$, total differentiation of the equilibrium system (equations 2.12 and 2.10) yields the following system of differential equations

$$\left\{ \begin{array}{l} w_i L_i^s \left(\frac{dw_i}{w_i} + \frac{dL_i^s}{L_i^s} \right) = \sum_{j \neq R} (\sigma_s - 1) \pi_{ij}^s \pi_{Rj}^s w_j L_j \frac{d\tau^s}{\tau^s} - (\sigma_s - 1) \sum_j (1 - \pi_{ij}^s) \pi_{ij}^s E_j^s \frac{dw_i}{w_i} + \pi_{ii}^s \mu_s E_i \frac{dE_i}{E_i} + \\ \quad (\sigma_s - 1) \sum_h \sum_j \pi_{ij}^s \pi_{hj}^s \mu_s w_j L_j \frac{dw_h}{w_h} + \sum_h \pi_{ih}^s \mu_s E_h \frac{dE_h}{E_h} + \\ \quad \sum_j (1 - \pi_{ij}^s) \pi_{ij}^s \mu_s w_j L_j \frac{dL_i^s}{L_i^s} - \sum_h \sum_j \pi_{ij}^s \pi_{hj}^s \mu_s w_j L_j \frac{dL_h^s}{L_h^s} \\ w_i L_i^k \left(\frac{dw_i}{w_i} + \frac{dL_i^k}{L_i^k} \right) = -(\sigma_k - 1) \sum_j (1 - \pi_{ij}^k) \pi_{ij}^k E_j^k \frac{dw_i}{w_i} + \pi_{ii}^k \mu_k E_i \frac{dE_i}{E_i} + (\sigma_k - 1) \sum_h \sum_j \pi_{ij}^k \pi_{hj}^k \mu_k w_j L_j \frac{dw_h}{w_h} \\ \quad \sum_h \pi_{ih}^k \mu_k E_h \frac{dE_h}{E_h} + \sum_j (1 - \pi_{ij}^k) \pi_{ij}^k \mu_k w_j L_j \frac{dL_i^k}{L_i^k} - \sum_h \sum_j \pi_{ij}^k \pi_{hj}^k \mu_k w_j L_j \frac{dL_h^k}{L_h^k} \quad k \neq s \\ \sum_{k=1,2} L_i^k \tilde{L}_i^k = 0 \end{array} \right.$$

where s is the sector in which tariffs change. Note that if region i does not produce good s , then the first equation terms are equal to 0 and the wage effect comes entirely from the set of equations for sectors $k \neq s$.

Multiplying by $\frac{\tau^s}{d\tau^s}$, using that $\frac{dE_i}{d\tau^s} \frac{\tau^s}{E_i} = \frac{dw_i}{d\tau^s} \frac{\tau^s}{w_i}$, defining $\xi_{ij}^s = \frac{\pi_{ij}^s \mu_k w_j L_j}{w_i L_i^s}$, and changing notation to $\tilde{x} = \frac{dx}{d\tau^s} \frac{\tau^s}{x}$ yields

$$\begin{aligned}\tilde{w}_i &= (\sigma_s - 1) \sum_{j \neq R} \pi_{Rj}^s \xi_{ij}^s - (\sigma_s - 1) \left[\sum_j \xi_{ij}^s (1 - \pi_{ij}^s) \right] \tilde{w}_i + \xi_{ii}^k \tilde{E}_i - \left(\sum_j \pi_{ij}^s \xi_{ij}^s \right) \tilde{L}_i^s \\ &\quad + (\sigma_k - 1) \sum_h \left(\sum_{j \neq i} \pi_{hj}^k \xi_{ij}^k \right) \tilde{w}_h + \sum_{h \neq i} \xi_{ih}^k \tilde{E}_h - \sum_h \left(\sum_{j \neq i} \pi_{hj}^k \xi_{ij}^s \right) \tilde{L}_h^s\end{aligned}$$

where for the terms containing $\frac{dL_i^k}{L_i}$ I used that $\sum_j \xi_{ij} = 1$.

Rearranging terms containing \tilde{w}_i yields

$$\begin{aligned}\tilde{w}_i &= \left[1 + (\sigma_s - 1) \sum_j (1 - \pi_{ij}^s) \xi_{ij}^s - \xi_{ii}^k \right]^{-1} \left\{ (\sigma_s - 1) \sum_{j \neq R} \pi_{Rj}^s \xi_{ij}^s - \left(\sum_j \pi_{ij}^s \xi_{ij}^s \right) \tilde{L}_i^s \right. \\ &\quad \left. + (\sigma_k - 1) \sum_h \left(\sum_{j \neq i} \pi_{hj}^k \xi_{ij}^k \right) \tilde{w}_h + \sum_{h \neq i} \xi_{ih}^k \tilde{E}_h - \sum_h \left(\sum_{j \neq i} \pi_{hj}^k \xi_{ij}^s \right) \tilde{L}_h^s \right\}\end{aligned}$$

A.2 Shock affecting the two sectors

When there are tariff shocks that affect the sectors 1 and 2 simultaneously, the system in equation 2.20 becomes

$$\begin{cases} \hat{w}_i^{PE} &= \frac{[(\sigma_1 - 1) \sum_{j \neq R} \xi_{ij}^1 \pi_{Rj}^1] - (\sum_j \xi_{ij}^1 \pi_{ij}^1) \hat{L}_i^1}{1 + (\sigma_1 - 1) \sum_j \xi_{ij}^1 (1 - \pi_{ij}^1) - \xi_{ii}^1} \hat{\tau}_1 \\ \hat{w}_i^{PE} &= \frac{[(\sigma_2 - 1) \sum_{j \neq R} \xi_{ij}^2 \pi_{Rj}^2] - (\sum_j \xi_{ij}^2 \pi_{ij}^2) \hat{L}_i^2}{1 + (\sigma_2 - 1) \sum_j \xi_{ij}^2 (1 - \pi_{ij}^2) - \xi_{ii}^2} \hat{\tau}_2 \\ \sum_{k=1,2} L_i^k \hat{L}_i^k &= 0 \end{cases} \quad (\text{A.1})$$

Solving the 3×3 system in equation A.1, the **partial-equilibrium wage change** in region i equals:

$$\hat{w}_i^{PE} = g_i^1 \frac{L_i^1}{L_i} (\sigma_1 - 1) \left(\sum_{j \neq R} \xi_{ij}^1 \pi_{Rj}^1 \right) \hat{\tau}_1 + g_i^2 \frac{L_i^2}{L_i} (\sigma_2 - 1) \left(\sum_{j \neq R} \xi_{ij}^2 \pi_{Rj}^2 \right) \hat{\tau}_2 \quad (\text{A.2})$$

where $g_i^1 = \frac{a_i^2}{a_i^2 b_i^1 (L_i^1/L_i) + a_i^1 b_i^2 (L_i^2/L_i)}$, $g_i^2 = \frac{a_i^1}{a_i^2 b_i^1 (L_i^1/L_i) + a_i^1 b_i^2 (L_i^2/L_i)}$, $a_i^k = \sum_j \xi_{ij}^k \pi_{ij}^k$, and $b_i^k = 1 + (\sigma_k - 1) \sum_j \xi_{ij}^k (1 - \pi_{ij}^k) - \xi_{ii}^k$.

Under assumption 1, $g_i^1 = g_i^2 = \alpha$. Under assumptions 2 to 5, we can express the wage change as:

$$\hat{w}_i = -\tilde{\alpha} \sum_k \left(\frac{L_i^k}{L_H^k} \frac{\Delta X_{RH}^k}{L_i} \right) \quad (\text{A.3})$$

which corresponds to the estimating equation in Autor et al. (2013).

A.3 Parameter values for simulations

- Parameter values.

$$- L(\text{reg1}, \text{reg2}, \text{RoW}) = [50, 50, 100]$$

$$- \beta^1 = \beta^2 = [1, 1, 1]$$

$$- \sigma_1 = \sigma_2 = 2; \mu_1 = \mu_2 = 0.5$$

$$- \kappa_{ij}^1 = \begin{array}{c|ccc} & 1 & 2 & R \\ \hline 1 & 1 & d_{12}^{-2} & 2d_{1R}^{-2} \\ 2 & d_{12}^{-2} & 1 & 2d_{2R}^{-2} \\ R & 2d_{1R}^{-2} & 2d_{2R}^{-2} & 1 \end{array}; \kappa_{ij}^2 = 1.1 \kappa_{ij}^1$$

- The parametrization of trade costs assumes a gravity coefficient of -2 and an ad-valorem equivalent of the border effect between the Home country and the Rest of the World of 2.
- Trade shock: 5% fall in trade costs in sector 1 from RoW to regions 1 and 2 (alternatively and both at the same time).
- Experiment: regions are ex ante symmetric, and region 2 becomes more distant with respect to both region 1 and the RoW (d_{12} and d_{1R} increase sequentially).
- Due to symmetry in size and no comparative advantage, $Exposure = 0.5$ for all values of trade costs

A.4 Omitted variable bias

This section explores the econometric consequences of the exposure approach and treats it as an omitted variable problem.

The partial-equilibrium wage change derived from the model is given by:

$$\hat{w}_i^{PE} = \alpha_0 + \tilde{w}_i^{PE} \hat{\tau}^k + \varepsilon_i \quad (\text{A.4})$$

where $\hat{\tau}^k$ is the tariff change of good k and \tilde{w}_i^{PE} is the partial-equilibrium elasticity as defined in equation 2.21.

We can formalize the exposure approach as parametrizing this elasticity in the following way: $\tilde{w}_i^{PE} = \alpha_1 \frac{L_i^k}{L_i}$ where $\beta\alpha_1$ is unobserved by the econometrician. This allows to rewrite equation A.4 as:

$$\hat{w}_i = \beta_0 + \alpha_1 \frac{L_i^k}{L_i} \hat{\tau}^k + \zeta_i \quad (\text{A.5})$$

with the new error term $\zeta_i = \varepsilon_i + \left(\tilde{w}_i^{PE} - \beta_1 \frac{L_i^k}{L_i} \right) \hat{\tau}^k$. Even under no correlation between the structural error term and employment shares (this is, $Cov\left(\frac{L_i^k}{L_i}, \varepsilon_i\right) = 0$), employment shares will be correlated with ζ_i unless the partial elasticity is a linear function of employment shares $\rho_i^k = \alpha_1 \frac{L_i^k}{L_i}$. The formulation of the partial elasticity in equation , shows this is not the case.

A.5 Solution algorithm

The algorithm that solves for counterfactual wages and sectoral employment is based on Alvarez and Lucas (2007) but extended to solve for employment shares. Given the values for L_i , β_i^k , κ_{ij}^k , μ_k , and σ_k , this algorithm solves for N wages and $N \times K$ sectoral employment values using the following steps

1. Propose an initial guess for wages, $\{w_{i,0}\}$, and sectoral employment shares, $\{\lambda_{i,0}^k\}$, with $\lambda_i^k = \frac{L_i^k}{L_i}$ and $\sum_k \lambda_i^k = 1$.
2. Update employment shares using goods markets clearing in each sector, equation 2.12

$$\lambda_{i,1}^k = \frac{\sum_j \pi_{ij,0}^k \mu_k w_{j,0} L_j}{w_{i,0} L_i} \quad k = 1, \dots, K$$

$$\text{with } \pi_{ij,0}^k = \frac{\lambda_{i,0}^k L_i (w_{i,0} \kappa_{ij}^k / \beta_i^k)^{1-\sigma_k}}{\sum_h \lambda_{h,0}^k L_h (w_{h,0} \kappa_{hj}^k / \beta_h^k)^{1-\sigma_k}}.$$

3. Compare $\lambda_{i,1}^k$ and $\lambda_{i,0}^k$, if the difference is greater than an arbitrary $v_L \sim 0$, update the guess using the formula

$$\lambda_{i,2}^k = s\lambda_{i,0}^k + (1-s)\lambda_{i,1}^k$$

where $s \in (0,1)$ is a smoothing parameter. Repeat steps 2 and 3 until

$$|\lambda_{i,t}^k - \lambda_{i,t-1}^k| < v_L.$$

4. Use updated employment shares, $\lambda_{i,t}^k$, and the initial guess for wages, $w_{i,0}$, to check if aggregate trade is balanced in each region

$$TB_{i,0} = \sum_k \sum_j \pi_{ij,t}^k \mu_k w_{j,0} L_j - \sum_k \sum_j \pi_{ji,t}^k \mu_k w_{i,0} L_i$$

5. If $|TB_{i,0}| < v_w$ with $v_w \sim 0$, then $\{w_{i,0}\}, \{\lambda_{i,t}^k\}$ is a solution.
6. If $|TB_{i,0}| > v_w$, update wages using

$$w_{i,1} = w_{i,0} (1 + \phi TB_{i,0})$$

where $\phi > 0$ is an adjustment factor. As in Alvarez and Lucas (2007), I exploit that the equation for the trade balance is a contraction in wages. If $TB_i > 0$ ($TB_i < 0$), then there is an excess demand (supply) for region i 's products and therefore, region i 's wages must increase (decrease) to re-establish trade balance. Repeat steps 2-4 using $w_{i,1}$ and $\lambda_{i,t}^k$ as an initial guess.

A.6 Price index changes

The aggregate price index formula, in levels, is given by

$$P_i = \left(\frac{p_i^0}{\mu_0} \right)^{\mu_0} \prod_{k=1}^N \left(\frac{P_i^k}{\mu_k} \right)^{\mu_k} \quad (\text{A.6})$$

where p_i^0 is the price of the non traded good and P_i^k is the ideal price index of each traded sector given by equation 2.2. Equation A.6 follows from the fact that the upper tier utility function has a Cobb-Douglas functional form.

The level of the sectoral price indices, P_i^k , cannot be pinned down since its calculation requires the knowledge of the number of varieties in each sector n_i^k and therefore, of the fixed cost, F^k , in equation 2.6. To avoid the calibration of these parameters and since the level of the price index is irrelevant—only its change matters for calculating the real wage change—, I obtain the price index change by using the formula

$$\frac{P_{i,t+1}}{P_{i,t}} = \left(\frac{w_{i,t+1}}{w_{i,t}} \right)^{\mu_0} \prod_{k=1}^N \left(\frac{\sum_j L_{j,t+1}^k \left(w_{j,t+1} \kappa_{ji}^k / \beta_j^k \right)^{1-\sigma_k}}{\sum_j L_{j,t}^k \left(w_{j,t} \kappa_{ji}^k / \beta_j^k \right)^{1-\sigma_k}} \right)^{\mu_k} \quad (\text{A.7})$$

where I use that the change in the non-traded good price is equal, by equation 2.5, to the change in wages: $\frac{p_{i,t+1}^0}{p_{i,t}^0} = \frac{w_{i,t+1}}{w_{i,t}}$; and that $\left(P_{i,t}^k \right)^{1-\sigma_k} = v_k \sum_j L_{j,t}^k \left(w_{j,t} \kappa_{ij}^k / \beta_j^k \right)^{1-\sigma_k}$ with $v_k = (F_k \sigma_k)^{-1} \left(\frac{\sigma_k}{\sigma_k - 1} \right)^{1-\sigma_k}$, which follows from equations 2.7 and 2.8. Note that wages and sectoral employment in equation A.7 are directly obtained from the data in t . Their values in $t + 1$ correspond to counterfactuals obtained solving the model

for the new tariffs. The parameters, κ_{ji}^k , β_j^k , σ_k , μ_k are calibrated in t (1991) and are assumed to remain constant after the trade shock.

A.7 Assumptions on residual sectors and implications for relative wages

A.7.1 Constant returns to scale non-traded sector. Balanced trade.

Suppose the non-traded good is produced in region i with a constant returns to scale Cobb-Douglas technology that uses skilled and unskilled labor

$$X_i^N = \left(S_i^N\right)^\eta \left(U_i^N\right)^{(1-\eta)} \quad (\text{A.8})$$

with $\eta < 1$ and where S_i^N and U_i^N are, respectively, the amounts of skilled and unskilled labor used in production, measured in hours.

Perfect competition in this sector implies producers take the price of their good and wages as given. Therefore, first order conditions are given by

$$w_i^S = p_i^N \eta \left(U_i^N / S_i^N\right)^{1-\eta} \quad (\text{A.9})$$

$$w_i^U = p_i^N (1 - \eta) \left(S_i^N / U_i^N\right)^\eta \quad (\text{A.10})$$

where w_i^S and w_i^U are, respectively, the hourly wage paid to skilled and unskilled labor.

The non-traded good market clearing condition in region i is given by

$$X_i^N = \frac{\mu_0 [w_i^S S_i + w_i^U U_i]}{\eta w_i^S + (1 - \eta) w_i^U} \quad (\text{A.11})$$

where $\mu_0 [w_i^S S_i + w_i^U U_i]$ is the share of the non-traded good in total regional expenditure. I used that by perfect competition in the non-traded good sector, price equals unit costs and therefore: $p_i^N = \eta w_i^S + (1 - \eta) w_i^U$.

Dividing numerator and denominator by unskilled wages yields

$$X_i^N = \frac{\mu_0 [(w_i^S/w_i^U) S_i + U_i]}{\eta (w_i^S/w_i^U) + (1 - \eta)} \quad (\text{A.12})$$

This equation shows that the given the values of parameters, μ_0 and η , endowments, $\{S_i, U_i\}$, and relative wages, w_i^S/w_i^U , the output of the non traded sector is fixed. Therefore, an increase in the level of wages that keeps relative wages constant, does not affect output in the non-traded sector.

A.7.2 Constant returns to scale non-traded sector. Unbalanced trade.

Suppose that each region has an exogenous trade deficit given by D_i . Regional aggregate expenditure is now given by $w_i^S S_i + w_i^U U_i + D_i$ so the non-traded sector equilibrium condition is given by

$$X_i^N = \frac{\mu_0 [w_i^S S_i + w_i^U U_i + D_i]}{\eta w_i^S + (1 - \eta) w_i^U}. \quad (\text{A.13})$$

Dividing by unskilled wages yields

$$X_i^N = \frac{\mu_0 \left[(w_i^S/w_i^U) S_i + U_i + D_i/w_i^U \right]}{\eta (w_i^S/w_i^U) + (1 - \eta)} \quad (\text{A.14})$$

Therefore, a constant relative wage no longer implies a constant size of the non-traded sector. Suppose both skilled and unskilled wages increase in the same proportion, the size of the non-traded sector decreases since income does not increase in the same proportion as costs.

A.7.3 Freely traded residual sector

Consider a freely traded good that is produced in every region i with a constant returns to scale technology which uses both skilled and unskilled labor. Perfect competition in this sector yields a zero profit condition that states that the price of this good equals its unit cost of production:

$$1 = (w_i^S)^\eta (w_i^U)^{1-\eta} \quad (\text{A.15})$$

where the price of the freely traded good is normalized to one, without loss of generality.

Total differentiation of equation (A.15) yields

$$0 = \eta \widehat{w}_i^S + (1 - \eta) \widehat{w}_i^U \quad (\text{A.16})$$

Solving for changes in unskilled wages, this equation implies skilled and unskilled wages have to change in opposite directions: $\widehat{w}_i^U = -\frac{\eta}{1-\eta} \widehat{w}_i^S$. Substituting this expression in the wage equation in changes equation (4.8) yields:

$$\sigma_k \left[\frac{\alpha_k - \eta}{1 - \eta} \right] \widehat{w}_i^S = \widehat{MA}_i^k \quad (\text{A.17})$$

If the traded sectors are skilled intensive relative to the freely traded sector, then $\frac{\alpha_j}{1-\alpha_j} > \frac{\eta}{1-\eta}$ and the term in square brackets is positive. Therefore, an increase in equilibrium market access leads to an increase in skilled wages and a decrease in unskilled wages.

Note that this is a partial equilibrium result since by definition, equilibrium market access of region i depends on every other region's wages though equilibrium expenditures and price indices which are themselves functions of wages in every location. Any exogenous shock to region's i market access has a direct impact on region's i wages through equation (A.17). In turn, the change in region i 's wages has a feedback effect on region's i market access and on all of its trading partner's market access.

Appendix B: Data

B.1 Data sources and definitions

B.1.1 Wages and employment

Wages and employment data are from the Brazilian Demographic Censuses of 1991 and 2000 conducted by the Brazilian Institute of Geography and Statistics (IBGE). The censuses are carried out in July of the corresponding year. As in Kovak (2013), I restrict the sample to workers aged 18–55 who were not enrolled in school at the time they were surveyed.

Based on the reported municipality, I assign workers to microregions. Microregions are defined in a geographically consistent way across the 1991 and 2000 Demographic Censuses using a mapping provided in Kovak (2013).

I obtain the worker’s hourly wage by dividing the monthly income in their main job by the weekly hours worked in their main job times 4.33, which corresponds to the average number of weeks in a month. The yearly income of a worker is simply the reported monthly income in the main job times 12. I weight both the hourly wage and the income by the individual’s sampling weight, which is provided in the census data. I then obtain individual annual hours worked by dividing the annual

income by the hourly wage. I convert the hourly wage to current US dollars using exchange rates from the World Development Indicators dataset.

Finally, I aggregate the data at the region-sector level. I obtain hours worked in a region-sector (L_i^k) as the sum of yearly hours of workers in that region-sector and the regional wage (w_i) as the simple average of hourly wages across all workers in a region.

B.1.2 Tariffs

Brazilian tariff data are from Kume et al. (2000). The authors start from data at the tariff-line level for the period 1988-1998 and aggregate it using simple averages up to SCN (Sistema de Contas Nacionais), the Brazilian National Accounts sector classification, at the 4-digit level (also called “Nível 80” or Level 80). Then, they aggregate the data to the 2-digit level (also called SCN 43) using industry value-added weights. I aggregate their data from SCN 43 to my own 12-sector industry classification using value added weights from the Input Output matrix in the year 1991, obtained from IBGE.¹

B.1.3 Elasticities of substitution

The elasticities of substitution between varieties in each sector are from Broda and Weinstein (2006).² I use import demand elasticities for Brazil, at the HS 3-digit level. I use a converter from the Harmonized System to the 4-digit SCN from

¹Available at http://downloads.ibge.gov.br/downloads_estatisticas.htm.

²Available at <http://www.columbia.edu/~dew35/TradeElasticities/TradeElasticities.html>.

IBGE.³ At the HS 3-digit level, there are many-to-many mappings between both classifications, so I keep the match with the highest imports. To aggregate from SCN to my classification of 12 sectors, I use the median elasticity.

B.1.4 Interstate and international trade

Trade flows between Brazilian states at the sector level are from Vasconcelos and Oliveira (2006). This data corresponds to the year 1999 and is reported using the 2-digit CNAE classification (59 codes).

The original data corresponds to exports by Brazilian states to other states. Of a total of 26 states plus the Federal District, only 22 had data that, according to the authors, was of enough quality to be used. In addition, one state did not report its exports by destination state.⁴ This leaves me with interstate export data for 21 states. Since these states report exports to the 6 missing states, interstate import flows can be recovered for all 27 states. Therefore, the number of observations for each industry is $546 = 21 \times (27 - 1)$. The 6 missing states have an estimated importer fixed effect in the gravity estimation but not an estimated exporter fixed effect.

Trade flows between Brazil and the ROW are obtained from the AliceWeb System (supported by the Secretary of International Trade, SECEX). This dataset reports the f.o.b. value of exports and imports in current US dollars at the 10-digit Brazilian Nomenclature (NBM, Nomenclatura Brasileira de Mercadorias) sector classification in 1991 and at the 8-digit Mercosur's Common Nomenclature in

³available at <http://concla.ibge.gov.br/classificacoes/correspondencias/atividades-economicas>.

⁴The states of Acre, Amapá, Maranhão, Rio Grande do Norte, and Roraima, did not have data of enough quality and Ceará only reported aggregate exports.

2000 (NCM96, Nomenclatura Comum do Mercosul). I use converters from both classifications to the CNAE classification from IBGE.

B.1.5 Distances

I calculate distances between the microregions' centroids using the Haversine formula. For the gravity estimation, which is done at the state level, I calculate Haversine distances between the states' centroids.⁵

I calculate distances between the microregions and the ROW using the Haversine formula between a microregion's centroid and its nearest sea port. For the gravity estimation, I calculate distances between the states and the ROW as the simple average distance of the microregions in that state to the nearest sea port.

I define the distance of a microregion or of a state to itself as the average distance in a circular region of the same area, using the formula $dist_{ii} = 0.66(area/\pi)^{1/2}$. The data on the areas of the microregions and the states is from the Brazilian Institute for Applied Economic Research (IPEA).⁶

B.1.6 Rest of the world wages and employment

I obtain the ROW hourly wages using data on hours worked and GDP from Penn World Tables version 8.1 with the following procedure. I calculate the ratio of hours worked in Brazil to hours worked in the ROW, which is .115 in 1991 (and 0.117 in 2000). I calculate total hours worked in the ROW as: $hours_{ROW} =$

⁵Shapefiles are available at http://downloads.ibge.gov.br/downloads_geociencias.htm.

⁶Available at <http://www.ipeadata.gov.br/>

$\frac{hours_{BRA}}{.115} - hours_{BRA}$, where $hours_{BRA}$ are total hours worked in Brazil according to the population Census data and $hours_{ROW}$ are total hours worked in the Rest of the World excluding Brazil . Then, I obtain the ratio of the GDP of Brazil to the ROW from PWT, which is 0.021 in 1991 (and 0.027 in 2000), and I calculate the wagebill of the ROW as $wagebill_{ROW} = \frac{wagebill_{BRA}}{.021} - wagebill_{BRA}$ where $wagebill_{BRA}$ is the Brazilian wagebill obtained from the population Census. Finally, I obtain the hourly wage of the ROW as: $wage_{ROW} = \frac{wagebill_{ROW}}{hours_{ROW}}$.

I obtain the ROW sectoral employment using data on sector's value added from UNIDO Industrial Statistics and UN Statistics Division National Accounts Main Aggregates Database. First, I calculate ROW total value added in current US dollars for 1991 and 2000 from UN National Accounts data for 213 countries, excluding Brazil. I obtain Agricultural, Manufacturing, and Mining shares in VA from UN National Accounts. The data is in national currencies, so I obtain the ROW aggregate shares weighting countries shares by their share in World VA in US dollars. Agricultural and Mining shares correspond directly to sectors 1 and 2 of my classification, Manufacturing corresponds to the rest of the sectors. To obtain the share of the rest of the manufacturing sectors, I use data from UNIDO. To minimize missing data I use the version INDSTAT3 for 1991 and the version INDSTAT4 for 2000. I restrict the sample to the 50 countries that appear in both versions of the data, excluding Brazil. INDSTAT3 uses the ISIC Rev2 classification and INDSTAT4 uses the ISIC Rev3 classification. I convert both industry classifications to my own classification using a converter from ISIC Rev3 to CNAE from IBGE.⁷

⁷For INDSTAT3 I first convert to ISIC Rev 3, and then to CNAE.

Then, I calculate the share of each manufacturing sector in manufacturing value added. Finally, I obtain VA in US dollars for each sector by multiplying these shares by the share of manufacturing in total VA and by the ROW total VA in US dollars.

B.1.7 Industry crosswalks

The data at the sector level used in this paper is originally reported using different industry classifications. The 1991 Census report individuals' industry of employment using the *atividade* classification, which is specific to Brazil; the 2000 Census uses the CNAE *domiciliar*, a specific branch of Brasil's National Classification of Economic Activities (CNAE, Classificação Nacional das Atividades); Brazilian imports and exports at the product level are reported using Mercosur's Common Nomenclature 1996 (NCM96, Nomenclatura Comum do Mercosul); interstate trade is reported using the 2-digit National Classification of Economic Activities prevailing in Brazil (CNAE); elasticities of substitution are reported using the Harmonized System at the 3-digit level; tariffs are reported using the National Accounts System at the two digit level (SCN43, Sistema de Contas Nacionais), and the value added of the Rest of the World is reported using ISICRev2 at the 3-digit level for 1991 and ISICRev3 at the 4-digit level for 2000.

The variables that restrict the possible mappings the most are tariffs (2-digit SCN43) and interstate trade (2 digit-CNAE). The SCN classification is convertible to the CNAE classification. The mapping is one to many going from 2-digit SCN

to 4-digit CNAE. However, going to 2-digit CNAE the mapping becomes many-to-many. To deal with this issue, I create a coarser industry classification described in table B.4. This is the classification that implied the minimum number of arbitrary mappings.⁸

In order to make compatible the rest of the variables, I assign them to either the 2-digit CNAE or to the 2-digit SCN. I convert hours and wages, and international trade to the 2-digit CNAE classification while I convert elasticities of substitution and the Rest of the World's value added to the 2-digit SCN.

⁸In fact, the only arbitrary mapping I made was to assign sector 14 in CNAE (Non-metallic minerals extraction) to sector 2 in the SCN classification (Mineral Extraction, except fuel) when it was also assigned to sector 31 in SCN (Other food and beverage industries).

B.2 Supplementary tables

Tab. B.1: Elasticities of substitution and Cobb-Douglas shares

Sector	σ_k	$\mu_{k,BRA}$	$\mu_{k,ROW}$
Agric.	3.30	0.310	0.137
Mining	3.60	0.054	0.069
Mineral Manuf.	2.69	0.030	0.030
Metal, Mach., Elect. & Autom	3.12	0.241	0.368
Paper	2.53	0.039	0.081
Rubber & Plastic	3.31	0.024	0.033
Chemic., Pharm. & Petrol.	2.65	0.086	0.100
Textiles	3.13	0.032	0.026
Apparel	3.24	0.060	0.016
Footwear & Leather	14.64	0.010	0.005
Food	3.56	0.048	0.101
Wood & Others	3.01	0.065	0.035

Notes: Parameters and targets by sector used in the calibration in 1991. σ_k is the elasticity of substitution across varieties, $\mu_{k,BRA}$ and $\mu_{k,ROW}$ are the calibrated Cobb-Douglas shares for Brazil and the ROW respectively.

Sources: σ_k : Broda et al. (2006), $\mu_{k,BRA}$ and $\mu_{k,ROW}$: own calculation based on IBGE, SECEX, PWT, and UNIDO. See Data Appendix for details.

Tab. B.2: Imports and exports to value added ratio, 1991 and 2000.

<i>Sector</i>	1991		2000	
	<i>M/VA</i>	<i>X/VA</i>	<i>M/VA</i>	<i>X/VA</i>
Agric.	0.018	0.008	0.055	0.154
Mining	0.705	0.294	1.289	0.966
Mineral Manuf.	0.034	0.041	0.161	0.340
Metal, Mach., Elect. & Autom	0.361	0.353	1.773	1.472
Paper	0.101	0.190	0.196	0.415
Rubber & Plastic	0.163	0.114	0.907	0.678
Chemic., Pharm. & Petrol.	0.510	0.160	1.800	0.641
Textiles	0.127	0.186	0.280	0.250
Apparel	0.011	0.015	0.022	0.041
Footwear & Leather	0.171	0.722	0.196	0.765
Food	0.164	0.545	0.269	1.028
Wood & Others	0.151	0.098	0.050	0.201

Notes: M/VA (X/VA) are Brazilian sectoral imports (exports) divided by sectoral value added in the respective year.

Source: Trade flows: SECEX. Value added: Population census, IBGE. See Section 3.2 and Data Appendix for details.

Tab. B.3: Calibrated border effects: 1991 and 2000.

<i>Sector</i>	1991	2000
Agric.	101.0	36.0
Mining	43.0	9.0
Mineral Manuf.	451.0	111.0
Metal, Mach., Elect. & Autom	9.0	1.0
Paper	231.0	91.0
Rubber & Plastic	26.0	8.0
Chemic., Pharm. & Petrol.	56.0	16.0
Textiles	16.0	10.0
Apparel	46.0	29.0
Footwear & Leather	1.3	1.2
Food	16.0	1.0
Wood & Others	23.0	22.0

Notes: Calibrated border effect factors, χ_k , at the microregion level in 1991 and 2000. See section 3.3 for details.

Tab. B.4: Correspondence between industrial classifications.

Sector	Description	SCN	CNAE
1	Agriculture	1	1, 2, 5
2	Mineral Mining	2, 3	10, 11, 13, 14
4	Non-Metallic Mineral Manufacturing	4	26
5	Metal, Machinery, Electrical Equipment, and Automobile Manufacturing	5-13	27-35
15	Paper Manufacturing, Publishing, and Printing	15	21,22
16	Rubber and Plastic Products Manufacturing	16, 21	25
17	Chemical and Pharmaceutical Manufacturing and Petroleum Refining	17-20	23, 24
22	Textiles Manufacturing	22	17
23	Apparel Manufacturing	23	18
24	Footwear and Leather Products Manufacturing	24	19
25	Food Processing	25-31	15, 16
32	Wood Products, Furniture, and Other Products Manufacturing	14, 32	20, 36, 37
99	Non-Traded Goods and Services	33, 43	40-73

Notes: Mapping between 2-digit national accounts (SCN) classification and 2-digit National Classification of Economic Activities (CNAE) and my own classification (first and second columns).

Tab. B.5: Data Sources and Classifications.

Variable	Data Source	Classification
Elasticities of substitution	Broda et al. (2006)	HS-3 digit
Hours and wages	Demographic Census 1991 and 2000	Atividade and CNAE Dom
International trade	Aliceweb (SECEX)	NCM96
Interstate trade	Vasconcelos and Oliveira (2006)	2-digit CNAE
Tariffs	Kume et al. (2000)	2-digit SCN43
Value added (ROW)	UNIDO INDSTAT3 and INDSTAT4	ISIC Rev2 and ISIC Rev3

Bibliography

- Abreu, M. d. P. (2004). *Trade liberalization and the political economy of protection in Brazil since 1987*. Buenos Aires: IDB-INTAL.
- Adao, R. (2015). Worker heterogeneity, wage inequality, and international trade: Theory and evidence from Brazil. *Unpublished*.
- Alvarez, F. and R. E. Lucas (2007). General equilibrium analysis of the Eaton–Kortum model of international trade. *Journal of Monetary Economics* 54(6), 1726–1768.
- Anderson, J. E. and E. Van Wincoop (2003). Gravity with gravitas: A solution to the border puzzle. *The American Economic Review* 93(1), 170–192.
- Anderson, J. E. and Y. V. Yotov (2010). The changing incidence of geography. *American Economic Review* 100(5), 2157–2186.
- Anderson, J. E. and Y. V. Yotov (2012). Gold standard gravity. *National Bureau of Economic Research Working Paper* 17835.
- Attanasio, O., P. K. Goldberg, and N. Pavcnik (2004). Trade reforms and wage inequality in Colombia. *Journal of development Economics* 74(2), 331–366.
- Autor, D. H., D. Dorn, and G. H. Hanson (2013). The China syndrome: Local labor market effects of import competition in the United States. *American Economic Review* 103(6), 2121–68.
- Bartelme, D. (2014). Trade costs and economic geography: Evidence from the US.
- Broda, C., J. G. and D. Weinstein (2006). From groundnuts to globalization: A structural estimate of trade and growth. *National Bureau of Economic Research Working Paper* 12512.
- Caliendo, L., M. Dvorkin, and F. Parro (2015). The impact of trade on labor market dynamics. *National Bureau of Economic Research Working Paper* 21149.
- Caliendo, L. and F. Parro (2014). Estimates of the trade and welfare effects of NAFTA. *The Review of Economic Studies*.

- Caliendo, L., F. Parro, E. Rossi-Hansberg, and P.-D. Sarte (2014). The impact of regional and sectoral productivity changes on the US economy. *National Bureau of Economic Research Working Paper 20168*.
- Chiquiar, D. (2008). Globalization, regional wage differentials and the Stolper–Samuelson theorem: evidence from Mexico. *Journal of International Economics* 74(1), 70–93.
- Chiquiar, D., E. Covarrubias, and A. Salcedo (2015). Labor market consequences of trade openness and competition in foreign markets. *Unpublished*.
- Costinot, A. and A. Rodriguez-Clare (2014). Trade theory with numbers: Quantifying the consequences of globalization. *Handbook of International Economics* 4, 197.
- Dekle, R., J. Eaton, and S. Kortum (2008). Global rebalancing with gravity: measuring the burden of adjustment. *National Bureau of Economic Research Working Paper 13846*.
- Dix-Carneiro, R. and B. K. Kovak (2014). Trade liberalization and the skill premium: A local labor markets approach. *National Bureau of Economic Research Working Paper 20912*.
- Eaton, J. and S. Kortum (2002). Technology, geography, and trade. *Econometrica* 7(5), 1741–1779.
- Fally, T., R. Paillacar, and C. Terra (2010). Economic geography and wages in Brazil: Evidence from micro-data. *Journal of Development Economics* 91(1), 155–168.
- Fan, J. (2015). Internal geography, labor mobility, and the distributional impacts of trade. *Unpublished*.
- Francois, J. and D. Nelson (1998). Trade, technology, and wages: General equilibrium mechanics. *The Economic Journal* 108(450), 1483–1499.
- Fujita, M., P. Krugman, and A. J. Venables (1999). *The spatial economy: cities, regions and international trade*. MIT Press.
- Galiani, S. and P. Sanguinetti (2003). The impact of trade liberalization on wage inequality: Evidence from Argentina. *Journal of Development Economics* 72(2), 497–513.
- Goldberg, P. K. and N. Pavcnik (2007). Distributional effects of globalization in developing countries. *Journal of Economic Literature* 45(1), 39–82.
- Gonzaga, G., N. Menezes Filho, and C. Terra (2006). Trade liberalization and the evolution of skill earnings differentials in Brazil. *Journal of International Economics* 68(2), 345–367.

- Hanson, G. (2005). Market potential, increasing returns and geographic concentration. *Journal of International Economics* 67(1), 1–24.
- Harris, C. D. (1954). The market as a factor in the localization of industry in the united states. *Annals of the association of American geographers* 44(4), 315–348.
- Karacaovali, B. (2011). Productivity matters for trade policy: Theory and evidence. *International Economic Review* 52(1), 33–62.
- Kovak, B. K. (2013). Regional effects of trade reform: What is the correct measure of liberalization? *The American Economic Review* 103(5), 1960–1976.
- Krugman, P. (1980). Scale economies, product differentiation, and the pattern of trade. *The American Economic Review* 70(5), 950–959.
- Kumar, U. (2007). *Essays on Inequality and Reforms: Evidence from India*. Ph. D. thesis, University of Maryland.
- Kume, H., G. Piani, and C. F. Souza (2000). A política brasileira de importação no período 1987-98: descrição e avaliação. *Rio de Janeiro: IPEA*.
- McLaren, J. and S. Hakobyan (2010). Looking for local labor market effects of NAFTA. *National Bureau of Economic Research Working Paper* 16535.
- Monte, F. (2014). The local incidence of trade shocks. *Available at SSRN* 2337270.
- Monte, F., S. J. Redding, and R.-H. Esteban (2015). Commuting, migration and local employment elasticities. *Unpublished*.
- Redding, S. and P. K. Schott (2003). Distance, skill deepening and development: will peripheral countries ever get rich? *Journal of Development Economics* 72(2), 515–541.
- Redding, S. and A. J. Venables (2004). Economic geography and international inequality. *Journal of International Economics* 62(1), 53–82.
- Redding, S. J. (2012). Goods trade, factor mobility and welfare. *National Bureau of Economic Research Working Paper* 18008.
- Shorrocks, A. F. (1982). Inequality decomposition by factor components. *Econometrica* 50(1), 193–211.
- Shorrocks, A. F. (2013). Decomposition procedures for distributional analysis: a unified framework based on the Shapley value. *Journal of Economic Inequality* 11(1), 99–126.
- Silva, J. S. and S. Tenreyro (2006). The log of gravity. *The Review of Economics and Statistics* 88(4), 641–658.

- Stumpner, S. (2015). Trade and the geographic spread of the great recession. *Unpublished*.
- Topalova, P. (2007). Trade liberalization, poverty, and inequality. *in: Globalization and Poverty*, 291–336.
- Topalova, P. (2010). Factor immobility and regional impacts of trade liberalization: Evidence on poverty from India. *American Economic Journal: Applied Economics* 2(4), 1–41.
- Vasconcelos, J. R. d. and M. A. d. Oliveira (2006). Análise da matriz por atividade econômica do comércio interestadual no Brasil-1999. *IPEA, Texto para Discussão 1159*.
- Venables, A. J. and N. Limão (2002). Geographical disadvantage: a Heckscher–Ohlin–von Thünen model of international specialisation. *Journal of International Economics* 58(2), 239–263.