Local Labor Market Effects of Trade Policy: Evidence from Brazilian Liberalization*

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Abstract

This paper measures the effects of Brazil's 1987-1995 trade liberalization on local labor market wages and internal migration patterns. I develop a specific-factors model of regional economies to examine the impact of national price changes on local labor markets. In the model, a region's industry mix determines the local impact of liberalization, with larger wage declines in regions where workers are concentrated in industries facing the largest tariff cuts.

I find that regions whose output faced a 10% larger liberalization-induced price decline experienced a 9.4% larger wage decline. In addition, liberalization resulted in a shift in migration patterns. The most affected Brazilian states gained or lost approximately 0.5% of their populations as a result of liberalization-induced shifts in migration patterns. These results demonstrate the importance of considering the local effects of national trade liberalization and represent the first systematic evaluation of the effects of liberalization on internal migration.

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1 Introduction

Over the last forty years, trade barriers around the world have fallen to historically low levels. As part of this process, many developing countries abandoned import substituting industrialization policies by sharply lowering trade barriers, motivating a large literature examining the effects of trade liberalization on various national labor market outcomes such as poverty and inequality.[1] The focus on national outcomes follows the predictions of classical trade theory, which takes the country as the geographic unit of analysis. In this paper, I develop a specific-factors model of regional economies to examine the relationships between trade liberalization and local labor market outcomes at the sub-national level. I use the model's predictions to measure the effect of Brazil's trade liberalization on regional wages, finding substantial heterogeneity across different locations. I also show that workers responded to the geographically distinct impacts of liberalization on wages by migrating toward more positively impacted labor markets. Together, these results imply that although workers migrated in response to changing incentives across locations, the migration flows were not sufficient to equalize the local impacts of liberalization.

Brazil presents an excellent context in which to study the local effects of trade liberalization. Brazilian liberalization involved drastic reductions in overall trade restrictions and a decrease in the variation of trade restrictions across industries. Average tariffs fell from 54.9% in 1987 to 10.8% in 1995, and the standard deviation of tariffs across industries fell from 21.3 to 7.4, implying substantial cross-industry variation in tariff cuts. Additionally, the industrial composition of the labor force varies substantially across Brazilian regions. These two sources of variation combine to identify the effect of liberalization on local wages. The model implies that a region's wage change is determined by the weighted average of liberalization-induced price changes across industries, where the weights depend on the size of each industry in the region. Intuitively, liberalization's effect on a given region's wages depends primarily on tariff cuts in the region's most important industries.

The empirical results confirm the model's prediction. I find that local labor markets whose workers were concentrated in industries facing the largest tariff cuts were negatively impacted by liberalization relative to markets facing smaller cuts. Regions whose output faced a 10% larger liberalization-induced price decline experienced a 9.4% larger wage decline, relative to other regions. Moreover, I find that migration flows shifted away from

[1] See Winters, McCulloch and McKay (2004) and Goldberg and Pavcnik (2007) for summaries of the literature.

regions whose labor force faced the largest tariff cuts and toward regions facing smaller cuts. The most affected Brazilian states gained or lost approximately 0.5% of their populations as a result of liberalization-induced shifts in migration patterns. Both of these findings confirm the importance of considering sub-national effects of liberalization and support the theoretical predictions of the specific-factors model.

This paper makes two main contributions. First, it presents a model in which national trade policies have disparate effects across different regions of a country. By considering many regions and many industries, including a nontraded sector, the model's predictions are directly estimable in the data. The model's weighted average prediction for liberalization's effect on regional wages closely resembles the estimating equations used in recent empirical studies of the local effects of liberalization (Topalova 2007, Edmonds, Pavcnik and Topalova 2010, Hasan, Mitra and Ural 2007, Hasan, Mitra, Ranjan and Ahsan forthcoming, McCaig 2009, Topalova 2010, McLaren and Hakobyan 2010). In fact, under particular technological and labor market restrictions, the model yields an estimating equation that differs from the prior literature only by a positive scale factor, which leaves sign tests of the effects across regions unaffected.[2] The restrictions imposed by the model provide a number of additional practical benefits beyond motivating the weighted-average approach. The model suggests that liberalization affects labor markets by changing prices faced by producers, which can be examined empirically. This clarifies the channel through which liberalization affects wages and gives the results an intuitive scale interpretation: the percent change in regional wage for a percent change in the price of regional output. The model yields predictions for both the sign and magnitude of liberalization's effect across regions, both of which are borne out in the empirical analysis. Finally, the model clarifies the treatment of the nontraded sector, about which various prior analyses disagree.

Second, this is to my knowledge the first study to systematically evaluate the effects of national trade policy on internal migration. Recent papers studying the dynamic adjustments to trade liberalization focus on interindustry adjustment rather than geographic labor market adjustment through migration, and the large literature examining interregional migration has not considered the impact of trade policy.[3] The most closely related paper in this regard is Aguayo-Tellez, Muendler and Poole (2009) which shows that in the post-liberalization

[2] The restrictions are i) identical Cobb-Douglas production functions across industries and ii) all regions must employ an identical fraction of the labor force in the nontraded sector.

[3] Dix-Carneiro (2010) and Cosar (2010) study the interindustry adjustment process in the context of Brazilian liberalization, while Artuc, Chaudhuri and McLaren (2010) examine interindustry adjustment of laborers in the U.S. context.

period of 1996-2001 Brazilian workers at exporting firms were less likely to migrate, and that migrants tended to choose destinations with a high concentration of foreign-owned firms.

Since the specific-factors model of regional economies is driven by price changes across industries, it is not limited to examining liberalization. It can be applied to any situation in which national price changes drive changes in local labor demand. As an example, consider the U.S. local labor markets literature, in which which researchers use local industry mix to measure the effects of changes in national industry employment on local labor markets (Bartik 1991, Blanchard and Katz 1992, Bound and Holzer 2000). In the Brazilian context, changes in national industry employment are driven by plausibly exogenous trade policy variation.[4] If price changes across industries similarly drove the changes in national industry employment in the U.S., the specific factors model would provide a theoretical foundation for using local industry mix in that context as well.

The remainder of the paper is organized as follows. Section 2 develops a specific-factors model of regional economies in which industry price changes at the national level have disparate effects on wages in the country's different regional labor markets. Section 3 describes the data sets used, and Section 4 describes the specific trade policy changes implemented in Brazil's liberalization along with evidence supporting the exogeneity of the tariff changes to industry performance. Section 5 presents an empirical analysis of the effects of trade liberalization on wages across local labor markets, and Section 6 demonstrates liberalization's impact on changes in interstate migration patterns in Brazil, both supporting the predictions of the model and finding economically significant effects of liberalization across regions. Section 7 concludes.

2 Specific-Factors Model of Regional Economies

2.1 Price Changes' Effects on Regional Wages

Each region within a country is modeled as a Jones (1975) specific-factors economy.[5] Consider a country with many regions, indexed by r. The economy consists of many industries, indexed by i. Production uses two inputs. Labor, L, is assumed to be mobile between in-

[4] See Figure 2 below.

[5] The specific-factors model is generally used to model a country rather than a region. The current model could be applied to a customs union in which all member countries impose identical trade barriers and face identical prices.

dustries, is supplied inelastically, and is fully employed. Labor is immobile between regions in the short run, but may migrate between regions in the long run, as considered below. The second input, T, is not mobile between industries or regions. This input represents fixed characteristics of a region that increase the productivity of labor in the relevant industry. Examples include natural resource inputs such as mineral deposits, fertile land for agriculture, regional industry agglomerations that increase productivity (Rodriguez-Clare 2005), or fixed industry-specific capital.[6] All regions have access to the same technology, so production functions may differ across industries, but not across regions within each industry. Further, assume that production exhibits constant returns to scale. Goods and factor markets are perfectly competitive. All regions face the same goods prices, P_i, which are taken as given (endogenous nontradables prices are considered below).

When labor is immobile across regions, this setup yields the following relationship between regional wages and goods prices. All theoretical results are derived in Appendix A (the following expression is (A13) with labor held constant).

$$\hat{w}_r = \sum_i \beta_{ri} \hat{P}_i \qquad \forall r, \tag{1}$$

$$\text{where} \qquad \beta_{ri} = \frac{\lambda_{ri} \frac{\sigma_{ri}}{\theta_{ri}}}{\sum_{i'} \lambda_{ri'} \frac{\sigma_{ri'}}{\theta_{ri'}}}. \tag{2}$$

Hats represent proportional changes, $\lambda_{ri} = \frac{L_{ri}}{L_r}$ is the fraction of regional labor allocated to industry i, σ_{ri} is the elasticity of substitution between T and L, and θ_{ri} is the cost share of the industry-specific factor T in the production of good i in region r. Note that each $\beta_{ri} > 0$ and that $\sum_i \beta_{ri} = 1 \ \forall r$, so the proportional change in the wage is a weighted average of the proportional price changes.

Equation (1) describes how a particular region's wage will be impacted by changes in goods prices. If a particular price P_i increases, the marginal product of labor will increase in industry i, thus attracting labor from other industries until the marginal product of labor in other industries equals that of industry i. This will cause an increase in the marginal product of labor throughout the region and will raise the wage. In order to understand what drives the magnitude of the wage change, note that for a constant returns production function, the

[6] An alternative interpretation of T is as a multiplicative productivity term on a concave production function taking L as an input. If production is assumed to be Cobb-Douglas, i.e. $Y = AT^{\alpha} L^{1-\alpha}$, one can see that variation in T^{α} is isomorphic to variation in the productivity term A.

labor demand elasticity equals $\frac{\sigma}{\theta}$.[7] The magnitude of the wage increase resulting from an increase in P_i will be greater if industry i is larger or if its labor demand is more elastic. Large industries and those with very elastic labor demand will need to absorb large amounts of labor from other industries in order to decrease the marginal product of labor sufficiently to restore equilibrium. Thus, price changes in these industries have more weight in determining equilibrium wage changes. For further intuition, see the graphical treatment in Appendix A.2.

The relationship described in (1) captures the essential intuition behind this paper's analysis. Although all regions face the same set of price changes across industries, the effect of those price changes on a particular region's labor market outcomes will vary based on each industry's regional importance. If a region's workers are relatively highly concentrated in a given industry, then the region's wages will be heavily influenced by price changes in that regionally important industry.

2.2 Nontraded Sector

This subsection introduces a nontraded sector in each region, demonstrating that nontraded prices move with traded prices. This finding guides the empirical treatment of nontradables, which generally represent a large fraction of the economy under study. As above, industries are indexed by $i = 1...N$. The final industry, indexed N, is nontraded, while other industries ($i \neq N$) are traded. The addition of the nontraded industry does not alter the prior results, but makes it necessary to describe regional consumers' preferences to determine the nontraded good's equilibrium price in each region. I assume that all individuals have identical Cobb-Douglas preferences, permitting the use of a representative regional consumer who receives as income all wages and specific factor payments earned in the region.[8]

When labor is immobile across regions, this setup yields the following relationship between the regional price of nontradables and tradable goods prices (the following expression is (A22) with labor held constant).

$$\hat{P}_{rN} = \sum_{i \neq N} \xi_{ri} \hat{P}_i, \tag{3}$$

[7] Denoting the production function $F(T, L)$, and noting that T is fixed by definition, the labor demand elasticity is $\frac{-F_L}{F_{LL}L}$. Constant returns and Euler's theorem imply that $-F_{LL}L = F_{LT}T$. The elasticity of substitution for a constant returns production function can be expressed as $\sigma = \frac{F_T F_L}{F_{LT} F}$. Substituting the last two expressions into the first yields the desired result.

[8] CES consumer preferences yield very similar results, available upon request.

$$\text{where} \qquad \xi_{ri} = \frac{\frac{\sigma_{rN}}{\theta_{rN}}(1-\theta_{rN})\beta_{ri} + \varphi_{ri}}{\sum_{i'\neq N}\left[\frac{\sigma_{rN}}{\theta_{rN}}(1-\theta_{rN})\beta_{ri'} + \varphi_{ri'}\right]}, \tag{4}$$

where φ_{ri} is the share of regional production value accounted for by industry i. Note that each $\xi_{ri} > 0$ and that $\sum_{i\neq N}\xi_{ri} = 1\ \forall r$, so the proportional change in the nontraded price is a weighted average of the proportional price changes for traded goods.

To gain some intuition for this result, consider a simplified model with one traded good and one nontraded good. Assume the traded good's price rises by 10%, and the nontraded good's price stays fixed. The wage in the traded industry will rise, drawing in laborers, increasing traded output and decreasing nontraded output. In contrast, consumers shift away from traded goods and toward nontraded goods. This cannot be an equilibrium, since production shifts away from the nontraded good and consumption shifts toward it. The only way to avoid this disequilibrium is for the nontraded price to grow by the same proportion as the traded price. Appendix A.3 extends this intuition to the case with many traded goods, yielding (3) and (4).

This finding is important in guiding the empirical treatment of the nontraded sector. Previous empirical studies of trade liberalizations' effects on regional labor markets pursue two different approaches. The first approach sets the nontraded term in (1) to zero, since trade liberalization has no direct impact on the nontraded sector.[9] In the context of the present model, this is equivalent to assuming no price change for nontraded goods. This approach is not supported by the model presented here, which predicts that nontraded prices move with traded prices. Setting the price change to zero in the large nontraded sector would greatly understate the scale of liberalization's impact on regional wages. However, this difference does not necessarily invalidate the previous literature's conclusions, even if the present model is correct. Under additional technological and labor market restrictions, setting the nontraded price change to zero is equivalent to multiplying the full weighted average by a positive scalar.[10] This difference will have no effect on the sign tests implemented in the previous literature, but will only affect the size of the estimates. If the additional restrictions hold, conclusions regarding the effects on liberalization across regions remain

[9] This approach is used in Edmonds et al. (2010), McCaig (2009), McLaren and Hakobyan (2010), Topalova (2007), and Topalova (2010).

[10] If all industries use identical Cobb-Douglas technology ($\theta_i = \theta\ \forall i$), and all regions allocate an identical fraction of their workforce to the nontraded sector ($\lambda_{rN} = \lambda_N\ \forall r$), then setting the nontraded price change to zero is equivalent to multiplying the full weighted average by $(1-\lambda_N)$.

largely unaffected.

The second approach removes the nontraded sector from the weighted average in (1) and rescales the weights for the traded industries in (2) such that they sum to one.[11] This approach more closely conforms to the model just described. If the nontraded price changes by approximately the same amount as the average traded price, as described in (3), then dropping the nontraded price from (1) will have very little effect upon the overall average.[12] Ideally, one would simply calculate the terms in (4) using detailed data on production values across industries at the regional level and substitute the result into (1). However, when data on regional output by industry are unavailable, as is the case in the empirical analysis below, the model implies that dropping the nontraded sector is likely to provide a very close approximation to the ideal calculation.

2.3 Interregional Migration

Following a change in goods prices, the disparate wage effects across regions will change workers' incentives to locate in different regions. Workers can benefit by moving from regions whose wages were relatively negatively impacted and toward regions that were relatively positively impacted. These interregional migrants act as arbitrageurs, tending to equalize the impact of the price change across regions. This equalizing effect of migration can be seen by examining the effect of an increase in labor on a region's wage while holding traded goods' prices constant (the following is (A13) with $\hat{P}_i = 0 \ \forall i \neq N$).

$$\hat{w}_r = \frac{-\hat{L}_r}{\sum_{i'} \lambda_{ri'} \frac{\sigma_{ri'}}{\theta_{ri'}}} + \beta_{rN} \hat{P}_{rN} \tag{5}$$

There are two channels through which an increase in regional labor can affect wages. The first channel directly lowers wages through a decrease in the marginal product of labor, holding nontraded prices fixed. (5) shows that the size of this effect depends on the overall regional labor demand elasticity, which is a weighted average of each industry's labor demand elasticity. The second effect operates through labor's effect on nontraded goods prices, which may be positive or negative. Although a potential increase in nontraded prices may act to

[11] This approach is used in Hasan et al. (forthcoming) and Hasan et al. (2007), presented as a robustness check in McCaig (2009), and used as an instrumental variable in Edmonds et al. (2010), Topalova (2007), and Topalova (2010).

[12] Appendix A.4 describes the conditions under which the nontraded sector will have exactly no affect on the overall average and can be omitted. In particular, identical Cobb-Douglas technology ($\theta_i = \theta \ \forall i$) is a sufficient condition.

increase wages, Appendix A.5 shows that the direct effect always dominates and that an increase in regional labor will decrease the regional wage.

Therefore migration away from relatively negatively impacted regions and toward relatively positively affected regions will decrease the wage gaps between locations that would have been observed in the absence of equalizing migration. In practice, migration costs and other frictions make it unlikely that the cross-region wage variation generated by price changes will be entirely equalized. This expectation is supported by the analyses presented in Sections 5 and 6, which find evidence of equalizing migration but not enough to completely equalize cross-region wage impacts of liberalization.

Migration in the presence of nontraded goods poses two additional potential complications. First, when nontraded goods are present, each region's consumers face a unique price level, and workers' migration decisions depend on the real wage change in a given location rather than the nominal change. Under the restrictions necessary to drop the nontraded sector from the weighted average in (1) described in Appendix A.4, when a given region experiences a nominal wage decline relative to another region, it will also experience a real wage decline relative to the comparison region.[13] In this situation nominal wage comparisons are sufficient to reveal real wage differences across regions, and the migration analysis can proceed using expressions for nominal wage changes as in (1). Second, the change in total income to residents of a given location determines the price change for regional nontradables. If specific factor owners migrate, it becomes very difficult to keep track of specific factor income transfers across regions. For simplicity, the analysis presented here assumes that migrants do not own specific factors, earning only wage income.

3 Data

The preceding section described a specific-factors model of regional economies, which yields predictions for the effects of changes in tradable goods' prices on regional wages, the prices of nontraded goods, and the incentives to migrate between regions. This framework can be

[13]In particular, the proportional change in a region's real wage, ω_r, can be expressed as follows:

$$\hat{\omega}_r = (1 - \mu_N)\hat{w}_r - \sum_{1 \neq N} \mu_i \hat{P}_i$$

where μ_i is industry i's share of consumption. The second term on the right hand side does not vary across regions and is irrelevant to interregional comparisons, while the first term is the nominal wage change scaled by the traded goods' share of consumption.

used to measure the local impacts of any event in which a country faces price changes that vary exogenously across industries. In the remainder of the paper, I apply the model to the analysis of the regional impacts of trade liberalization in Brazil, requiring the combination of various industry-level and individual-level data sources.

The model is driven by exogenous changes in prices across tradable industries. In order to apply the model in the context of trade liberalization, I estimate the impact of trade policy changes on industry prices, yielding a measure of liberalization-induced price changes. Trade policy data at the Nível 50 industrial classification level (similar to 2-digit SIC) come from researchers at the Brazilian Applied Economics Research Institute (IPEA) (Kume, Piani and de Souza 2003). Kume et al. (2003) also calculated effective rates of protection (ERP) from nominal tariffs and the Brazilian input-output tables, accounting for the effect of tariffs on final goods as well as tariffs on imported intermediate inputs. Given that ERP's account for intermediate inputs, the results use the ERP as the preferred measure of protection. All results were also generated using nominal tariffs without any substantive differences from those presented here. Since Brazil does not calculate a producer price index (Muendler 2003b), I use the wholesale price index, IPA-OG maintained by Fundação Getulio Vargas and distributed by IPEA. As a proxy for world prices, U.S. prices for manufactures come from the BLS Producer Price Index and agriculture prices from the USDA-NASS All Farm Index. As demonstrated below and in earlier work on Brazilian liberalization, the effect of a tariff change on the relevant price depends on industry import penetration.[14] Industry import penetration was calculated from Brazilian National Accounts data available from the Brazilian Census Bureau (Instituto Brasileiro de Geografia e Estatistica - IBGE). Following Gonzaga, Filho and Terra (2006), I measure import penetration as imports divided by the sum of imports and domestic production.

Wage, employment, and migration data come primarily from the long form Brazilian Demographic Censuses (*Censo Demográfico*) for 1991 and 2000 from IBGE. Throughout the analysis, local labor markets are defined as microregions. Each microregion is a grouping of economically integrated contiguous municipalities with similar geographic and productive characteristics (IBGE 2002).[15] Wages are calculated as earnings divided by hours. The

[14] See section 5.2 for a detailed discussion.

[15] To account for changing administrative boundaries between 1991 and 2000, I use information on municipality border changes described by Reis, Pimentel and Alvarenga (2007) to generate consistent areas over time by aggregating microregions when necessary. The original 558 microregions were aggregated to yield 494 consistent microregions. Details of the aggregation, including descriptive maps and GIS files are available upon request.

Census also reports employment status and industry of employment, which permits the calculation of the industrial distribution of labor in each microregion. Migration information in each Census is based on individuals' current municipality and their municipality of residence five years earlier. In the wage and migration analyses, I restrict the sample to individuals aged 18-55 who are not currently enrolled in school in order to focus on people who are most likely to be tied to the labor force. The wage analysis in Section 5 further restricts the sample to those receiving nonzero wage income. While it would be ideal to have wage and employment information in 1987, just prior to liberalization, I use the 1991 Census as the baseline period under the assumption that wages and employment shares adjusted slowly to the trade liberalization. An alternative annual household survey, the *Pesquisa Nacional por Amostra de Domicílios* (PNAD), is available in 1987, but only reports state-level geographic information, making it impossible to identify local markets. I therefore use the Census when analyzing the effects of liberalization on local wages and migration, and use the PNAD for a few descriptive figures in which geographic detail is unimportant.[16]

In order to utilize these various data sets in the analysis, it was necessary to construct a common industry classification that is consistent across data sources. The final industry classification consists of 21 industries, including agricultural and nontraded goods. A crosswalk between the various industry classifications is presented in Appendix B, along with more detail on the data sources, variable construction, and auxiliary results.

4 Trade Liberalization in Brazil

4.1 Context and Details of Brazil's Trade Liberalization

From the 1890's to the mid 1980's Brazil pursued a strategy of import substituting industrialization (ISI). Brazilian firms were protected from foreign competition by a wide variety of trade impediments including very high tariffs, quotas, and other non-tariff barriers (Abreu 2004a, Kume et al. 2003). Although systematic data on non-tariff barriers are not available, tariffs alone provide a clear picture of the high level of protection in 1987, just before liberalization. The average tariff level in 1987 was 54.9%, with values ranging from 15.6% on oil, natural gas, and coal to 102.7% on apparel. This tariff structure, characterized by high average tariffs and large cross-industry variation in protection, reflected a tariff

[16] Earlier versions of this paper used the PNAD to examine liberalization's effects on state wages and interstate migration. The results were qualitatively similar to those presented here, but much less precisely estimated, likely due to the noise introduced by aggregating across heterogeneous local labor markets.

system first implemented in 1957, with small modifications (Kume et al. 2003).

While Brazil's ISI policy had historically been coincident with long periods of strong economic growth, particularly between 1930 and 1970, it became clear by the early 1980's that the policy was no longer sustainable (Abreu 2004a). Large amounts of international borrowing in response to the oil shocks of the 1970's followed by slow economic growth in the early 1980's led to a balance of payments crisis and growing consensus in government that ISI was no longer a viable means of generating sufficient economic growth. Between 1986 and 1987, Brazil ended a posture of obstruction in trade negotiations and began to seek concessions from trading partners in return for reductions in its own trade barriers (Abreu 2004b). It appears that this shift in trade policy came from within government rather than from the private sector. There is no evidence of political support from consumers of imported goods or of resistance from producers of goods losing protection (Abreu 2004b).

Tariff reforms began in late 1987 with a governmental Customs Policy Commission (Comissão de Politica Aduaneira) proposal of a sharp tariff reduction and the removal of many non-tariff barriers.[17] In June of 1988 the government adopted a weaker reform that lowered tariffs and removed some non-tariff barriers. In March 1990 import bans were eliminated, and firm-level import restrictions were removed in July 1991, so that by the end of 1991 tariffs represented the primary means of import protection. Between 1991 and 1994, phased tariff reductions were implemented, with the goal of reducing average tariff levels and reducing the dispersion of tariffs across industries in hopes of reducing the gap between internal and external costs of production (Kume et al. 2003). Following 1994, there was a slight reversal of the previous tariff reductions, but tariffs remained essentially stable following this period.

4.2 Exogeneity of Tariff Changes to Industry Performance

The empirical analysis below utilizes variation in tariff changes across industries. In order to interpret the subsequent empirical results as reflecting the causal impact of trade liberalization, the tariff changes must have been uncorrelated with counterfactual industry performance. Such a correlation may arise if trade policy makers impose different tariff cuts on strong or weak industries or if stronger industries are able to lobby for smaller tariff cuts (Grossman and Helpman 1994).

There are a number of reasons to believe that these general concerns were not realized in

[17] See Kume et al. (2003) for a detailed account of Brazil's liberalization, from which this paragraph is drawn.

the specific case of Brazil's trade liberalization. Qualitative analysis of the political economy of liberalization in Brazil indicates that the driving force for liberalization came from government rather than from the private sector, and that private sector groups appear to have had little influence on the liberalization process (Abreu 2004a, Abreu 2004b). The 1994 tariff cuts were heavily influenced by the Mercosur common external tariff (Kume et al. 2003). Argentina had already liberalized at the beginning of the 1990's, and it successfully negotiated for tariff cuts on capital goods and high-tech products, undermining Brazil's desire to protect its domestic industries (Abreu 2004b). Thus, a lack of private sector interference and the importance of multilateral trade negotiations decrease the likelihood that the tariff cuts were managed to protect industries based on their strength or competitiveness.

More striking support for exogeneity comes from the nature of the tariff cuts during Brazil's liberalization. It was a stated goal of policy makers to reduce tariffs in general, and to reduce the cross-industry variation in tariffs to minimize distortions relative to external incentives (Kume et al. 2003). This equalizing of tariff levels implies that the tariff changes during liberalization were almost entirely determined by the pre-liberalization tariff levels, as shown in Figure 1. Industries with high effective rates of protection before liberalization experienced the greatest cuts, with the correlation between the pre-liberalization ERP level and change in ERP equaling -0.94. The pre-liberalization tariff regime was based upon a tariff schedule developed in 1957 (Kume et al. 2003). Since the liberalization policy imposed cuts based on the tariff level that was set decades earlier, it is very unlikely that the tariff cuts were manipulated to induce correlation with counterfactual industry performance or with industrial political influence.

Additional suggestive evidence supporting the exogeneity of tariff changes comes from their relationship with industry employment growth. This relationship is demonstrated in Figure 2. As expected, industries facing larger tariff cuts shrank in terms of the number of workers employed in the industry, while those facing smaller tariff cuts grew. It is possible that certain industries were simply declining over time while others were growing, and that trade policy makers' choices were influenced by this observation. However, this interpretation can be tested by observing the pattern of industrial reallocation during the time period immediately preceding liberalization. If trade policy choices were related to industrial performance, there would be a correlation between pre-liberalization industry employment growth and subsequent tariff changes. As shown in Figure 3, this is not the case. There is no relationship between the pre-liberalization employment growth and the subsequent tariff changes, supporting the argument that tariff changes were not related to industry performance and can be considered exogenous in the empirical analysis below.

5 The Effect of Liberalization on Regional Wages

Given the previous section's evidence supporting the exogeneity of tariff changes, I move to analyzing the effect of liberalization on regional wages as predicted by the model in (1). I first calculate the necessary terms and then test the model's prediction that regions facing larger tariff cuts experience larger wage declines relative to other regions.

5.1 Regional Wage Changes

The model described in Section 2 considers homogenous labor, in which all workers are equally productive and thus receive identical wages in a particular region. In reality, wages differ systematically across individuals, and the observed wage change in a given region could be due changes in individual characteristics or changing returns to those characteristics. In order to net out these effects, I calculate regional wage changes as follows. In 1991 and 2000 I separately estimate a standard wage equation, regressing the log of real wages on demographic and educational controls, industry fixed effects, and microregion fixed effects.[18] I then normalize the microregion fixed effects relative to the average log wage change and calculate the associated standard errors based on Haisken-DeNew and Schmidt (1997).

Figure 4 shows the resulting estimated regional wage changes in each microregion of Brazil. States are outlined in bold while each smaller area outlined in gray is a microregion. Microregions that are lighter experienced the largest wage declines during the 1991-2000 time period, while darker regions experienced the largest wage increases, relative to the national average. As the scale indicates, some observations are quite large in magnitude, though only 7 observations fall outside the $\pm 30\%$ range, and these are all in sparsely populated areas with imprecise estimates that receive little weight in subsequent analysis.[19]

[18] The results of these regressions are reported in Appendix B Table B2.

[19] The substantial wage variation across regions is not an artifact of the demographic adjustment procedure. As shown in Appendix B Figure B3, unconditional regional wage changes are very similar (0.93 correlation) and exhibit somewhat larger amounts of variability, with 17 observations outside the $\pm 30\%$ range, again in sparsely populated ares.

5.2 Industry Price Changes

The model described in Section 2 concerns the impact of industry price changes on regional wages. In order to apply the model to the study of liberalization, it is necessary to measure the effect of trade liberalization on prices faced by producers in Brazil. Denote the tariff rate in industry i as τ_i. It is common in cross-industry studies of liberalization to assume that $d\ln P_i = d\ln(1+\tau_i)$ and substitute out prices faced by producers. Gonzaga et al. (2006) have rejected this assumption in the Brazilian context, showing that in spite of the large differences in tariff changes across industries they are unrelated to price changes in the relevant industries. Table 1 reproduces this result by regressing the change in log price between 1987 and 1995 on the change in $\ln(1+\tau_i)$ over the same time period, measured using the effective rate of protection. There is no bivariate relationship between these two variables in column (1), and the lack of relationship continues in column (2), which controls for the change in U.S. price as a proxy for the change in world prices.

However, Gonzaga et al. (2006) have shown that price changes do relate to tariff changes adjusted for the degree of import penetration in the industry. This adjustment is based on the idea that tariff changes pass through into prices faced by domestic producers more strongly in import-intensive industries. Gonzaga et al. (2006) support this intuition using an aggregation model in which some goods in each industry face import competition, while others do not. In an industry with very few locally produced goods facing import competition, even a very large change in tariff will have a small effect on the price level in that industry. In Appendix C, I demonstrate that a similar aggregation result holds in the context of the multi-good specific factors model. Thus, following the empirical approaches in Gonzaga et al. (2006) and Ferreira, Leite and Wai-Poi (2007), columns (3) and (4) of Table 1 relate the change in log price to the import penetration adjusted change in $\ln(1+\tau_i)$, denoting import penetration as γ_i. In column (3) the positive and statistically significant relationship indicates that industries facing larger import penetration adjusted tariff cuts faced larger price declines. The magnitude of the relationship is quite large, resulting from the fact that import penetration is generally low, 5.3% on average. It therefore appears that import penetration does capture cross-industry differences in the pass-through from tariff changes to price changes, but understates the average amount of pass-through. Column (4) controls for the change in U.S. prices and reports a relationship that is nearly identical in magnitude, though less precisely estimated.

Given these estimates, I define the "liberalization-induced price change" as the predicted values from the regression in column (3) of Table 1 minus the average price change across industries. This measure is referred to below as $d\ln\hat{(P_i)}$, where the hat represents an estimate. Figure 5 shows the liberalization-induced price changes resulting from this calculation.[20] Since this measure is normalized relative to the overall change in price level, it may be positive or negative in individual industries even though all tariffs were cut. Given the cross-sectional nature of the empirical exercises, this normalization is only for convenience of interpretation and has no substantive impact on the results.

5.3 Region-Level Tariff Changes

Based on (1), trade liberalization's effect on a region's wages is determined by a weighted average of liberalization-induced price changes. In what follows, I call this weighted average the "region-level tariff change." Calculating the β_{ri} terms in (1) requires information for each region on the allocation of labor across industries and on labor demand elasticities in each industry. The industrial allocation of labor is calculated for each microregion from the 1991 Census. There exist no credible estimates of labor demand elasticities by Brazilian industry and region; in fact, I am unaware of any estimates of industry-specific labor demand elasticities for any country, even restricting the elasticities to be constant across regions. Given this limitation, for the empirical analysis I assume that production in all industries is Cobb-Douglas, and that the factor shares may vary across industries, implying that $\sigma_{ri} = 1$ and $\theta_{ri} = \theta_i$. I calculate θ_i, as one minus the wagebill share of industry value added using national accounts data from IBGE. Given these restrictions I calculate the region-level tariff change (RTC) for each microregion as follows.

$$RTC_r = \sum_{i \neq N} \beta_{ri} d\ln\hat{(P_i)} \tag{6}$$

$$\text{where} \qquad \beta_{ri} = \frac{\lambda_{ri}\frac{1}{\theta_i}}{\sum_{i' \neq N} \lambda_{ri'}\frac{1}{\theta_{i'}}}. \tag{7}$$

Recall from Section 2.2 that ideally one would directly measure the nontraded price in each region or model them using the traded goods prices as in (3). Given that neither nontraded prices nor output by industry are available by region in Brazil, these ideal approaches are not feasible in this case. Instead, I drop the nontraded sector from the weighted average in (6) based on the conclusion that nontraded prices move with traded prices, following the discussion in Section 2.2.

[20] Detail on the other components of the regressions in Table 1 are presented in Appendix B.

The results of this calculation appear in Figure 6. Lighter microregions faced the most negative region-level tariff changes, while darker microregions faced more positive price changes. Recall that the liberalization-induced price changes are calculated relative to the overall price level, so although all tariffs were cut, the region-level tariff changes may be positive or negative. Figure 7 demonstrates the underlying variation driving differences in the region-level tariff changes by comparing the weights, β_{ri}, for the microregion with the most negative region-level tariff change, São José dos Campos, to those in the microregion with the most positive region-level tariff change, Sinop. The industries on the x-axis are sorted from the most negative to most positive liberalization-induced price change. São José dos Campos has more weight in the left side of the diagram, particularly in the Auto, Transport, Vehicles industry, due to the presence of aircraft producer Embraer. Sinop produces agricultural goods and lumber almost exclusively, all of which faced quite positive liberalization-induced price changes. Thus, although all regions faced the same set of liberalization-induced price changes across industries, variation in the weight applied to those industries in each region generates the substantial variation seen in Figure 6.

5.4 Wage-Tariff Relationship

Given empirical estimates of the regional wage changes and region-level tariff changes, it is possible to examine the effect of tariff changes on regional wages predicted by the specific-factors model. I form an estimating equation from (1) as

$$d\ln(w_r) = \zeta_0 + \zeta_1 RTC_r + \epsilon_r, \tag{8}$$

where $d\ln(w_r)$ is the regional wage change described in Section 5.1. Since these wage changes are estimates, I weight the regression by the inverse of the standard error of the estimates based on Haisken-DeNew and Schmidt (1997). ζ_0 captures the regional effect of liberalization on real wages between 1991 and 2000. In the model without migration, theory predicts that $\zeta_1 = 1$. As discussed in Section 2.3 any interregional mobility in response to liberalization will smooth out the regional wage variation that would have been observed on impact. In the extreme case of costless, instant worker mobility, all liberalization-induced wage variation would be immediately arbitraged away by worker migration and there would be no relationship between region-level tariff changes and regional wage changes. Since Brazil's population is particularly mobile (inter-state migration rates are similar to those in the U.S.), I expect some equalizing migration over the 9 year period being observed and thus expect that $0 < \zeta_1 < 1$. Finally, the error term ϵ_r captures any unobserved drivers of wage change that are unrelated to liberalization.

Table 2 presents the results of regressing regional wage changes on region-level tariff changes under various alternate specifications. Each specification is reported with and without state fixed effects, and all standard errors are clustered at the state level, accounting for remaining covariance in the error terms across microregions in the same state.[21] All specifications omit the city of Manaus, which is a free trade area, unaffected by liberalization. Columns (1) and (2) present the main specification as described above. As expected, the relationship between wage changes and region-level tariff changes is positive. This implies that microregions facing the largest tariff declines experienced slower wage growth than regions facing smaller tariff cuts, as predicted by the model. The estimate in column (1) of 0.945 implies that a region facing a 10 percentage point larger liberalization-induced price decline experienced a 9.4 percentage point larger wage decline relative to other regions. The addition of state fixed effects in column (2) has almost no effect on the point estimate, but absorbs residual variance such that the estimate is now statistically significantly different from zero at the 1% level.

The remaining columns of Table 2 examine the effects of various deviations from the preferred specification in columns (1) and (2), reflecting empirical approaches implemented in the previous literature. Columns (3) and (4) omit the labor share adjustment, which in the context of the model is equivalent to assuming that the labor demand elasticities are identical across industries so that the weights in each region are determined only by the industrial distribution of workers. All of the papers in the previous literature follow this approach. In the Brazilian context, the omission of this adjustment has very little effect on the estimates, as they have little effect on the weights across industries. Taking a region x industry pair as an observation, the correlation between the weights with and without labor share adjustment is 0.996. Columns (5) and (6) include the nontraded sector in the regional tariff change calculations, setting the nontraded price change to zero. Footnote 9 lists papers using this approach.[22] This change also has little impact on the point estimates, but increases the standard errors. Columns (3) - (6) therefore suggest that these two differences between

[21] State-specific minimum wages were not introduced until 2002, and so do not affect the analysis.

[22] The previous literature does not explicitly make assumptions about the price of nontraded goods, but rather includes a zero term for the nontraded sector in the weighted averages used in their empirical analyses. In the context of the present model, that is equivalent to assuming zero price change for nontraded goods. However, if there exists a different unspecified model that justifies measuring the local effect of liberalization as a weighted average, the previous approach may reflect a different assumption regarding nontraded goods.

the empirical approach suggested by the model and the previous empirical literature do not substantially affect the results in the Brazilian context.

Columns (7) and (8) replace $d\ln\hat{(P_i)}$ in equation (6) with the change in protection level, $d\tau_i$. This approach is also used all of the previous literature. This change does sharply reduce the measured relationship between the region-level tariff change and the regional wage change. This is not surprising given that Section 5.2 has shown that there is little relationship between prices faced by producers and the unadjusted change in protection. Columns (9) and (10) combine the changes considered individually in columns (3) - (8) to approximate the approach implemented in much of the previous literature, with the region-level tariff change measured as

$$RTC_r^{\sim} = \sum_i \lambda_{ri}\ d\tau_i \qquad \text{where } d\tau_N = 0 \tag{9}$$

Similar to columns (7) and (8), the results in columns (9) and (10) are extremely weak, and the point estimate in column (10) with state fixed effects is even negative. Based on the findings in columns (3) - (8), the large difference between the main specification in columns (1) and (2) and the previous literature approach in columns (9) and (10) is likely due to using the change in protection level instead of the liberalization-induced price change. However, it is not clear that this difference would occur in different country contexts, particularly if disaggregate price and protection measures are available. Disaggregate data would probably exhibit a stronger link between unadjusted protection and prices than that observed in more aggregate Brazilian data. The lesson here is that when using the weighted average approach to measuring the local effects of liberalization, we need a measure of the change in protection that relates to price changes faced by producers.

One of the benefits of deriving the estimating equation (8) from the theoretical model in Section 2 is that the model predicts both the sign and magnitude of the coefficient ζ_1. As discussed in Section 2.3, the theory predicts a coefficient of 1 in the absence of equalizing interregional migration, a coefficient of 0 with costless and instant interregional migration, and a coefficient between 0 and 1 for the more realistic case of costly or slow equalizing migration. Consistent with these predictions, the coefficients in columns (1) and (2) of Table 2 are between 0 and 1, as are the coefficients for most of the alternative specifications. Table 2 also reports p-values testing the null hypothesis that the coefficient estimate is equal to 1. In the main specification, we fail to reject the null hypothesis of no equalizing migration, suggesting the presence of migration frictions across microregions. However, the fact that the point estimates generally are less than 1 weakly suggests the presence of some equalizing migration, which is analyzed directly below. The model's magnitude prediction can only be tested when the region-level tariff change is measured in terms of liberalization's proportional effect on prices. Note that in columns (7) - (10) when the dependent variable is measured in terms of the change in protection level, rather than its effect on prices, the coefficients are significantly different from 1.

Columns (11) and (12) report a robustness check in which the change in unadjusted regional average log wage is regressed on the weighted average of observed price changes. This exercise demonstrates that the model's prediction is still present when considering raw wage and price data, without regard to liberalization. The estimates are much closer to zero, though still significantly different from zero at conventional levels.

As a final check, I ran a permutation test in which the liberalization-induced price changes were randomly assigned to industries. Since the proposed mechanism linking liberalization to local wage changes hinges upon variation across industries, it should be very unlikely to observe the the actual estimate of ζ_1 if liberalization-induced price changes are assigned to incorrect industries. I implemented the test by generating 10,000 random assignments, recalculating the region-level tariff change and running the wage analysis to yield an estimate of ζ_1 corresponding to each assignment. I then compare the actual estimate of ζ_1, in which industries are properly assigned, to the distribution of estimates generated when assigning industries randomly. For specification (1) in Table 2 without fixed effects, only 5.33% of the random assignments yielded estimates greater than 0.945, and the mean of the randomly assigned coefficients was 0.044. For specification (2) in Table 2 with fixed effects, only 0.46% of the random assignments yielded estimates greater than 0.932, and the mean coefficient was 0.050. Thus, the cross-industry variation is essential to the results, which are not driven by some artefact of of the weight calculations across regions.

These results confirm the model's prediction, particularly in finding an estimate of the expected sign that is significantly different from zero, but just below one. This supports the assumption that cross-region differences in the effects of liberalization are correctly measured by the region-level tariff change and can be applied to other labor market outcomes of interest. The next section does this by examining the effects of liberalization on interregional migration.

6 The Effect of Liberalization on Interstate Migration

6.1 Location Choice Specification

This section derives a framework for estimating the effect of liberalization on individuals' location choices from a model of maximizing behavior. I use a static model of location choices in which individuals' current locations are taken as given and individuals make a single choice about whether and where to relocate.[23] I assume that individuals make location choices based on wages and other considerations such as local amenities, proximity to friends and relatives, and costs of moving to a particular location. These various aspects of location choice can be captured in the following additive random utility model.

$$U_{isdt} = V_{sdt} + \epsilon_{isdt} \tag{10}$$

$$V_{sdt} \equiv \alpha \ln w_{dt} + \mu_{sdt} + \eta_{sd} \tag{11}$$

U_{isdt} is the utility that individual i in source region s would experience from living in destination region d at time t. V_{sdt} represents the average utility across individuals, while ϵ_{isdt} represents individual idiosyncratic deviations from the average. The average utility in a given destination depends on wages, w_{dt}, and unobservable characteristics of the destination, some of which vary over time, μ_{sdt}, and some of which are fixed over time, η_{sd}. These unobservable terms capture location-specific amenities in potential destinations, and by allowing them to vary by source-destination pairs, they also capture moving costs associated with distance. Idiosyncratic variation in the utility of a particular location, due to the presence or absence of friends and relatives, desire for a change, or individual deviations from average preferences, is captured by the error term ϵ_{isdt}. The parameter of interest is α, the importance of wages in location decisions.

Individuals compare all regions and choose to live in the one that maximizes utility. Assuming that the ϵ_{isdt} are independently drawn from a Type I extreme value distribution, the probability π_{sdt} that an individual in source s chooses destination d at time t is

$$\pi_{sdt} = \frac{e^{V_{sdt}}}{D_{st}} \qquad \text{where} \quad D_{st} \equiv \sum_{d'} e^{V_{sd't}}. \tag{12}$$

[23] Estimating a dynamic location choice model based on the seminal work of Kennan and Walker (2011) is beyond the scope of the current paper and requires richer information on lifetime migration patterns of individuals. Since Brazil's trade liberalization represents a permanent shock, the static model should provide a reasonable approximation to a richer dynamic framework.

In the absence of the unobservable η_{sd} and μ_{sdt} terms in V_{sdt}, this expression would reduce to a standard conditional logit model. Given that these unobserved terms capture the effects of distance, amenities, and other important aspects of location choice, dropping them is an unattractive alternative. In particular, if wages are correlated with these unobserved terms, omitting them and estimating a standard conditional logit model would yield inconsistent estimates of α. Instead, I employ a strategy developed by Scanlon, Chernew, McLaughlin and Solon (2002) and adapted to the migration context by Cadena (2007) that differences out the time invariant unobserved characteristics by applying a first-order Taylor series approximation. This process, implemented in Appendix D, yields the following equation.

$$d\ln S_{sd} - d\ln S_{ss} \approx \alpha(d\ln w_d - d\ln w_s) + \left[(d\mu_{sd} - d\mu_{ss}) + d\left(\frac{\xi_{sd}}{\pi_{sd}}\right) - d\left(\frac{\xi_{ss}}{\pi_{ss}}\right)\right] \tag{13}$$

Observations are defined by source-destination pairs. S_{sd} is the observed share of individuals from region s choosing to locate in destination region d. S_{ss} is the share of people from region s choosing to stay there rather than relocate. Thus the left hand side of (13) is the change in the share of individuals from s who choose to locate in d, relative to the change in the share that choose to stay home. This difference-in-difference structure removes the time-invariant unobservables, η_{sd}. The independent variable of interest is the regional wage change in destination d, again relative to the same expression at home. Measuring wage changes directly would introduce simultaneity bias, since changing migration patterns affect wages, as discussed in Section 2.3. Instead I replace $d\ln w_d$ with the model's predicted effect of liberalization on wages without offsetting migration, RTC_d.[24]

$$d\ln S_{sd} - d\ln S_{ss} \approx \alpha(RTC_d - RTC_s) + \left[(d\mu_{sd} - d\mu_{ss}) + d\left(\frac{\xi_{sd}}{\pi_{sd}}\right) - d\left(\frac{\xi_{ss}}{\pi_{ss}}\right)\right] \tag{14}$$

The term in brackets represents the error term, consisting of two parts. The first is the change in time varying amenities. The presence of this expression in the error term makes clear the additional identification assumption necessary to estimate (14) in practice - changes in regional amenities must be uncorrelated with region-level tariff changes. This term also introduces a common error component across observations considering the same destination, so I calculate standard errors clustered by destination. ξ_{sd} represents random sampling error in measuring S_{sd}, generating heteroskedasticity. I therefore weight by the square-root of the number of observations used to calculate S_{sd} (Scanlon et al. 2002).

[24] This can be thought of as a reduced-form equation in an instrumental variables context.

I will also estimate versions of (14) that allow the importance of wages in location choice and the unobservable amenity terms to vary by demographic group. I do this for two reasons. First, previous studies of Brazilian internal migration have shown that different demographic groups regularly move to particular regions of Brazil. For example, migrants from the northeast region to the southeast region are younger on average than migrants moving in the opposite direction (Fiess and Verner 2003). These patterns likely reflect different preferences for regional amenities by age group. Second, different demographic groups generally respond differently to labor market conditions in their location choices. Beginning with Sjaastad (1962), studies of internal migration consistently show stronger responses of younger workers and those with lower migration costs. By estimating α_g separately by demographic group, I can check whether the expected groups responded more strongly to the liberalization-induced changes in regional wages by altering their migration patterns. The estimating equation when allowing these differences across demographic groups g is

$$d\ln S_{sgd} - d\ln S_{sgs} \approx \alpha(RTC_d - RTC_s) + \left[(d\mu_{sgd} - d\mu_{sgs}) + d\left(\frac{\xi_{sgd}}{\pi_{sgd}}\right) - d\left(\frac{\xi_{sgs}}{\pi_{sgs}}\right)\right] \quad (15)$$

Observations are now defined by source, demographic group, destination triples. S_{sgd} is the fraction of individuals living in source region s and in demographic group g choosing to move to destination d.

6.2 Location Choice Results

The migration pattern measures in (14) are calculated using migration data from the 1991 and 2000 Demographic Censuses. Table 3 presents summary statistics regarding interregional migration in Brazil among different demographic groups. The first and third columns present the fraction of the total population in each demographic group. The second and fourth columns describe the fraction of individuals in each demographic group who reported living in a different microregion 5 years earlier. Mobility in Brazil is very high. 29% of adults report having moved across states, which is nearly identical to the same figure in the U.S. (Dahl 2002). As a comparison to another large developing country, interregional migration in Brazil is much more common than in India. Topalova (2007) reports that only 3-4% of people migrated between Indian districts within a ten year time period, whereas 8% of Brazilians reported moving between microregions during a five year period. Districts in India are on average smaller than Brazilian microregions, so the difference in mobility is particularly striking.

Table 3 shows that, men are slightly more mobile than women, and that younger individuals are more likely to move. More educated individuals are more mobile, through mobility rates are constant through the middle of the education distribution. Whites and those of mixed heritage (reporting *Pardo*) are much more mobile than Blacks. Unmarried individuals and those with smaller families are more mobile, presumably due to lower migration costs. These observations provide insight into what portions of the population are likely to be most mobile and therefore most likely to respond to changing geographic incentives by moving to a new location. These expectations are largely borne out in the empirical results.

The results of estimating (14) and (15) are presented in Table 4. The first row estimates (14), without any distinctions across demographic groups. Thus, each observation represents a source-destination region pair.[25] The estimate of α in the first row of Table 4 is 1.308. This positive and statistically significant estimate shows that migration patterns shifted away from regions that were relatively negatively affected by liberalization and toward regions that were positively affected.

In order to assess the scale of this estimate, differentiate (12) with respect to $\ln w_{dt}$ for all d.

$$d\pi_{sd} = \alpha\pi_{sdt}\left((1-\pi_{sdt})d\ln w_d - \sum_{d'\neq d}\pi_{sd't}d\ln w_{d'}\right) \tag{16}$$

This expression describes how changes in wages across all regions affect the probability that an individual from state s will choose to locate in state d. Evaluating this expression at the estimate of α, the observed pre-liberalization migration fractions, and the liberalization-induced wage changes given by (6), it is possible to calculate $d\pi_{sd}$ for each source-destination state pair. Then, by multiplying each of these estimates of the change in migration fraction by the relevant source region population in 1991 and summing over all sources for a given destination, it is possible to calculate the number of people represented by liberalization-induced shifts in the interstate migration pattern over a five-year time horizon. The results of this exercise are shown in Table 5. The first column reports the number of people in each state accounted for by liberalization-induced shifts in migration patterns, the second column reports the 1991 state population, and the third column reports the liberalization-

[25]Since the share of individuals choosing to stay home has been differenced from each observation, and there are 493 microregions included in the analysis, the total number of potential observations is $493 * 492 = 242,556$. The analysis drops any region pairs in which the share term, S_{gd}, was estimated using less than five underlying observations, so the realized number of observations is 11,686. Similar results are obtained when dropping observations based on less than 10 underlying observations.

induced population change as a fraction of the baseline population. For the states whose microregions faced the largest and smallest tariff cuts, liberalization accounts for gains or losses up to 0.5% of the state's population. Although not so large as to be implausible, this represents an economically significant shift in the Brazilian population's geographic distribution. The final two columns of Table 5 report similar figures with respect to the 1991 inflow to each state.

The remaining rows in Table 4 estimate (15), allowing unobservable amenity valuations and the importance of wages in location choice to vary by demographic group. Observations now represent source-group-destination triples, which increases the number of observations relative to first specification. However, all standard errors are clustered by destination, so the number of clusters is not affected by the nominal increase in the number of observations. I split the sample into two groups for each demographic characteristic in order to avoid slicing the bins too finely and losing too many observations due to lack of data.

Observing the coefficient estimates in the second column, all are positive and all but one are significant at conventional levels. The third column reports the p-value testing the null hypothesis that the estimate of α_g is identical across demographic groups. When grouping by gender, males' migration choices respond much more strongly to liberalization-induced wage changes than females'. This is not surprising since men are slightly more mobile than females and are much more connected to the labor force.[26] Older workers responded more strongly than younger workers, which is surprising given their lower overall rates of mobility. Different education groups do not exhibit a statistically significant difference in their responses. Whites responded more strongly than non-whites, which may reflect their somewhat higher levels of baseline mobility. The most striking difference between demographic groups is seen between married and unmarried individuals, with a much stronger migration response among unmarried individuals, likely reflecting smaller migration costs for single individuals. Finally, those with smaller families respond more strongly than those with larger families, though the difference is not statistically significant. All of these estimates show that migration patterns shifted away from locations facing the largest liberalization-induced price declines in favor of places facing smaller cuts, and they generally confirm the intuition that demographic groups that are more mobile and more connected the labor market outcomes responded more strongly.

Section 5 argued that interregional migration would partly offset the wage variation across regions generated by liberalization, implying a relationship between the results of the wage

[26] In 1991 among non-students aged 18-55, 87% of men and 37% of women worked for pay.

and migration analyses presented here. The model predicts that in the absence of equalizing migration, the coefficient in the wage analysis estimate of (8) should not significantly differ from 1. I run the following consistency check based on this observation. I assume that the liberalization-induced population changes listed in Table 5 never occurred, and plug the implied population change into (5) to calculate counterfactual wage changes without the liberalization-induced population change. Note that this counterfactual wage change was calculated using only the information in the region-level tariff change and migration data. I then re-estimate (8) using these counterfactual wages that have been purged of the effects of equalizing migration. The results are shown in Table 6. In specifications with and without state fixed effects, the estimates are not significantly different from 1, as predicted by the model. Thus, the scale of the migration effects measured in this section are consistent with the separately estimated wage effects in the previous section.

7 Conclusion

This paper develops a specific-factors model of regional economies addressing the local labor market effects of national price changes, and applies the model's predictions in measuring the effects of Brazil's trade liberalization on regional wages and interregional migration. The model predicts that wages will fall in regions whose workers are concentrated in industries facing the largest tariff cuts, and workers will then migrate away from these regions in favor of areas facing smaller tariff cuts. These predictions are confirmed by the empirical analysis. Regions whose output faced a 10% larger liberalization-induced price decline experienced a 9.4% larger wage decline than other regions. Liberalization also caused a shift in migration patterns in which the most affected Brazilian states gained or lost approximately 0.5% of their populations as a result of liberalization.

These findings imply a link between national policy changes, such as liberalization, and local policy challenges involving migration, transportation, and housing, as individuals migrate to restore geographic equilibrium. National policy makers can use the specific-factors model to predict what areas are likely to experience wage increases and an influx of migrants hoping to gain employment in an area and can mobilize local services to respond during the transition. On a larger scale, the migration results demonstrate a channel through which a country may reap the production gains from trade liberalization. Production gains can only occur by reallocating factors, and in countries with geographically distinct industrial distributions, a large scale industrial reallocation of labor requires laborers to migrate from one part of the country to another. Thus, relocation, transportation, and retraining services likely play an important role when pursuing a change in national policy that requires substantial industrial and geographic reallocation.

Given these results, it seems likely that liberalization has different local effects on other outcomes that could be studied in future work. For example, the framework presented here assumes full employment, so that all adjustment occurs through wages. In order to study the impact of liberalization on employment, the opposite assumption could be incorporated by fixing wages in the short run and allowing employment to adjust. Alternatively, Hasan et al. (forthcoming) motivate their study of the effects of liberalization on local unemployment with a two-sector search model. An interesting avenue for future work would be to incorporate a search framework into a multi-industry model and directly derive an estimating equation relating changes in regional unemployment to tariff changes, paralleling the approach taken here. The model also suggests a novel channel through which liberalization could affect inequality. While the present analysis considered a homogenous labor force, future work could examine the impact of trade liberalization in a situation with laborers of different skill levels working in industries of varying factor intensities. More mobile groups of individuals will be able to smooth out regional wage variation by migrating while less mobile individuals will not, as discussed in Bound and Holzer (2000). If the two groups work in segmented labor markets, liberalization could greatly increase national wage dispersion for the immobile group while leaving the mobile group's wages relatively unchanged.

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Figure 1: Relationship Between Tariff Changes and Pre-Liberalization Tariff Levels

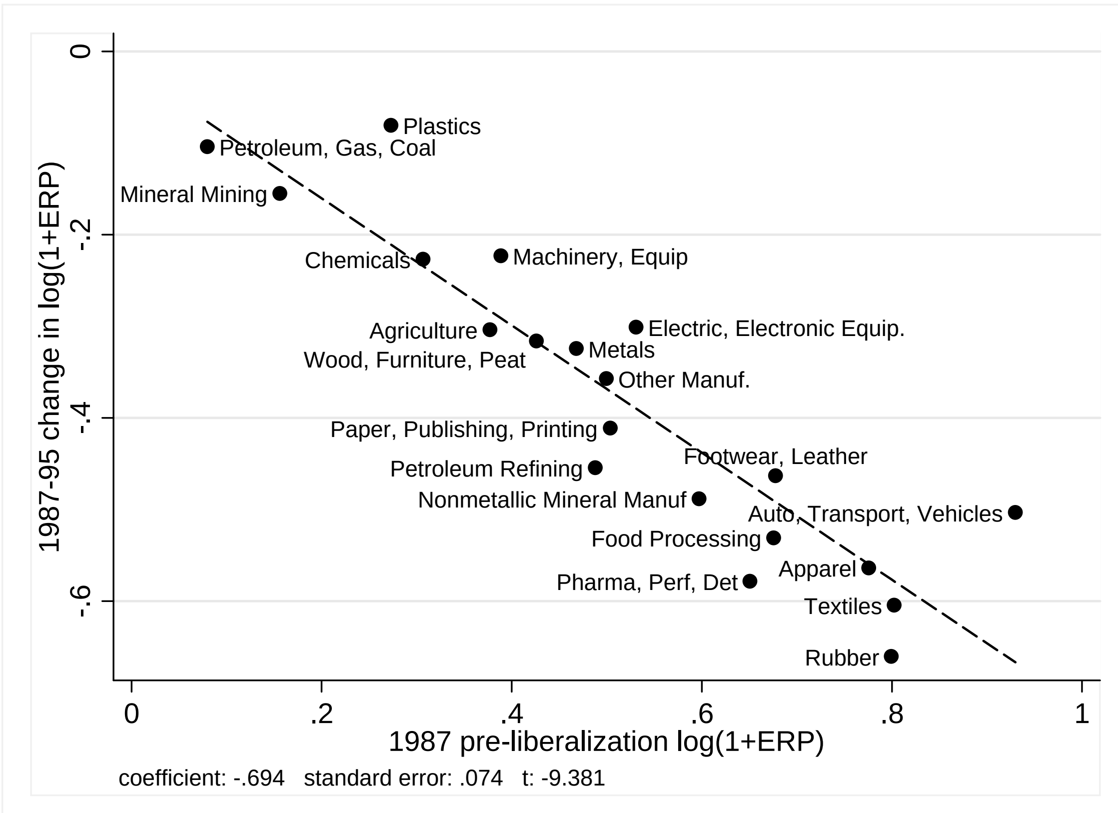


Figure 2: Industry Employment Growth and Tariff Changes

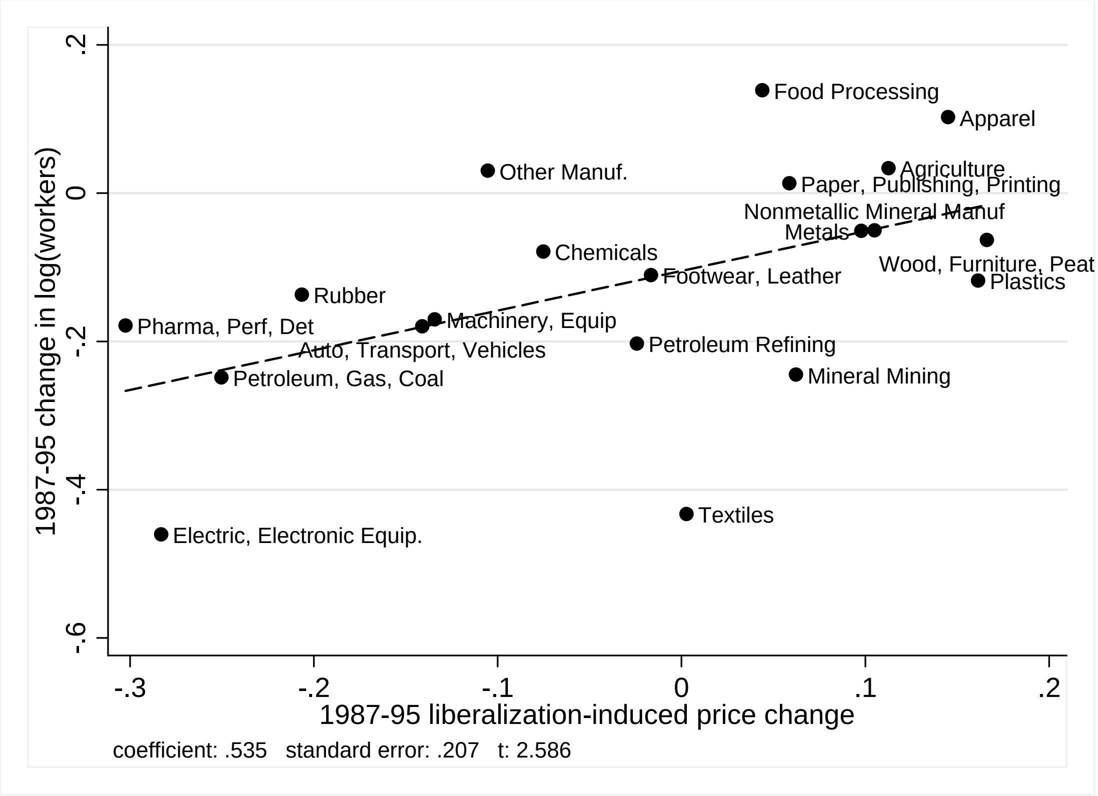


Figure 3: Pre-liberalization Industry Employment Growth and Tariff Changes

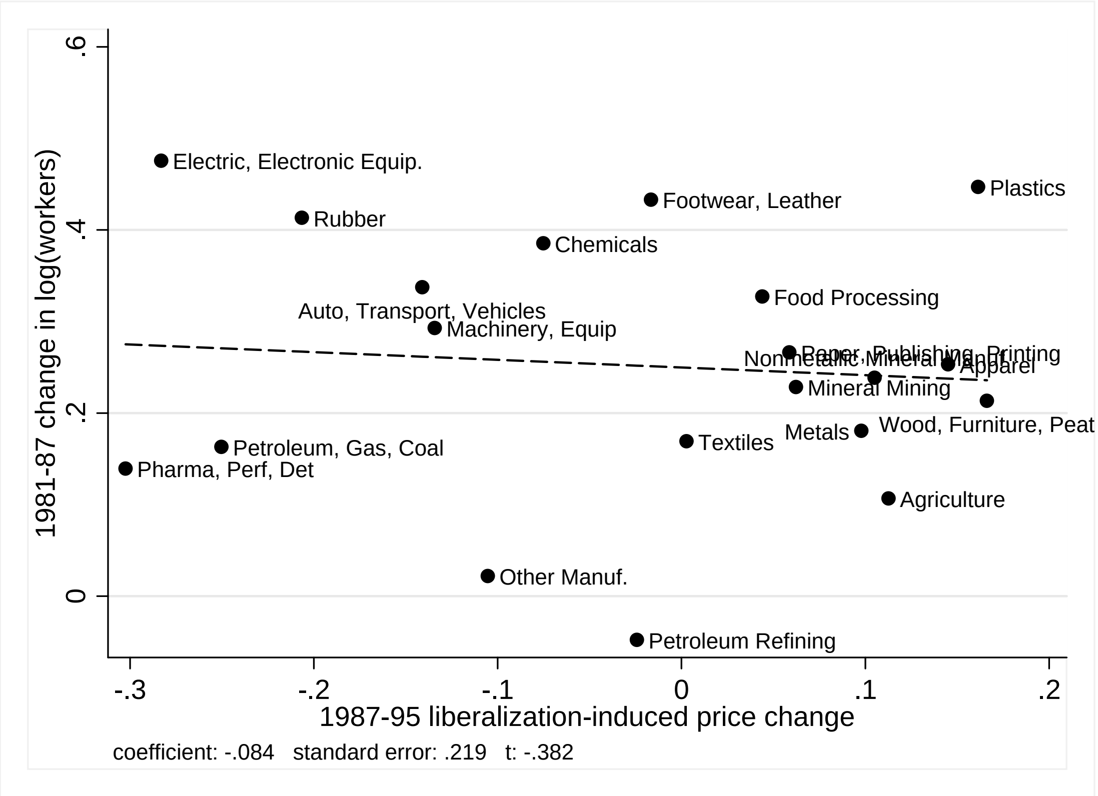
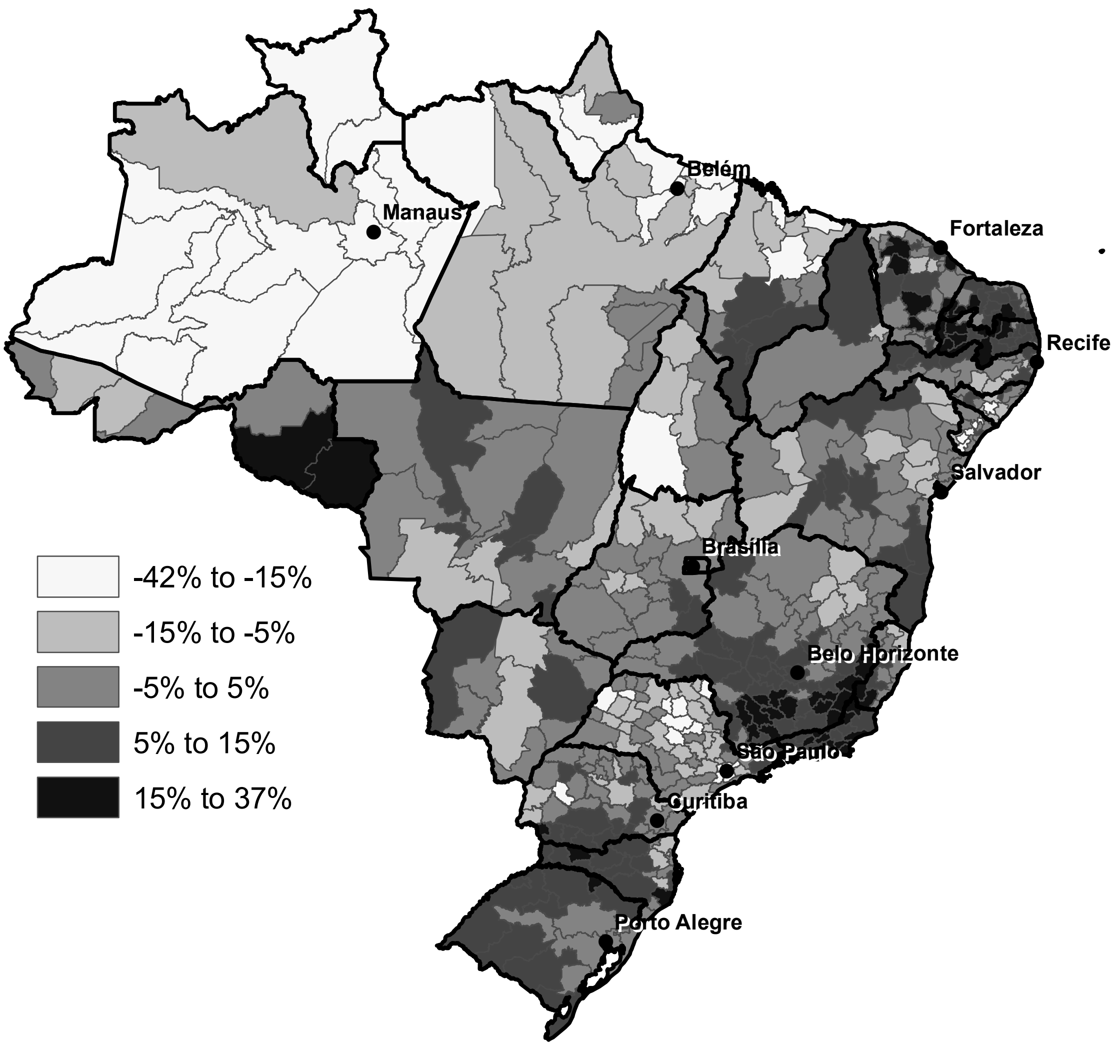
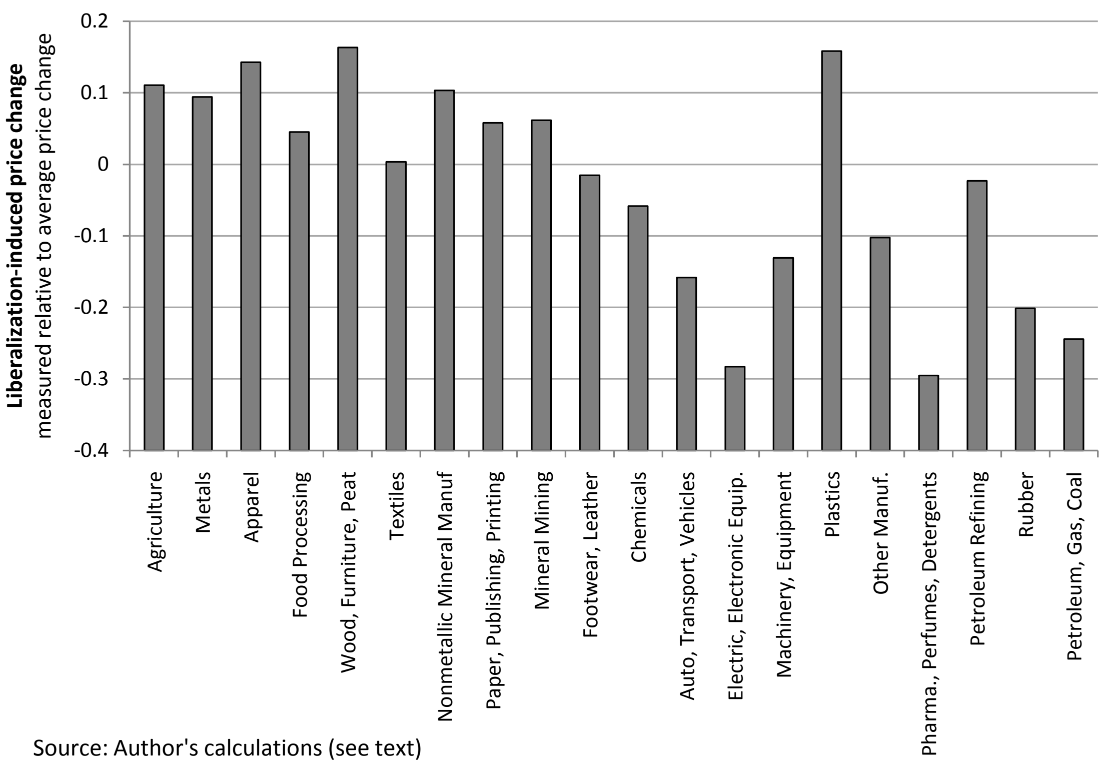


Figure 4: Regional Wage Changes



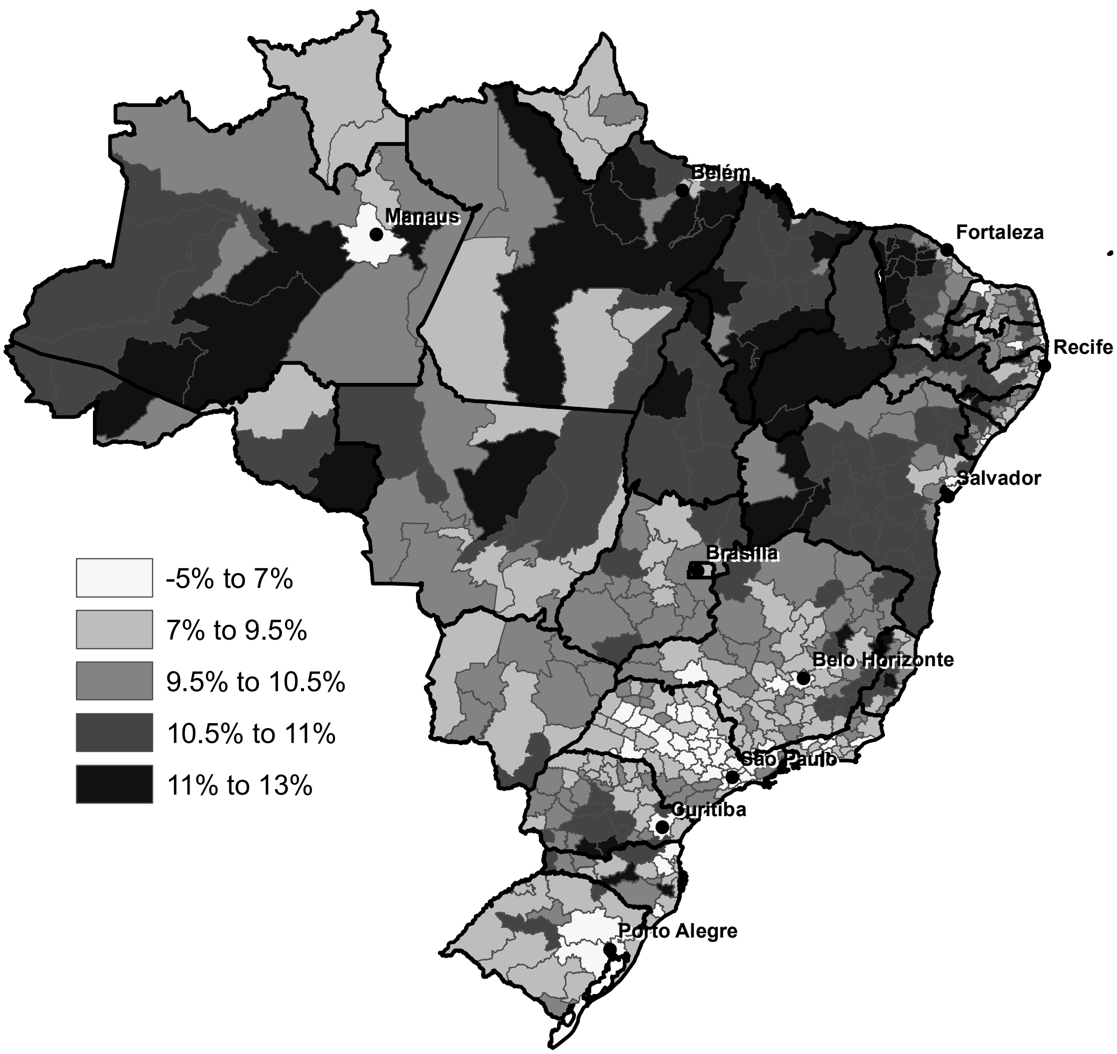
Proportional wage change by microregion - normalized change in microregion fixed effects from wage regression

Figure 5: Liberalization-Induced Price Changes



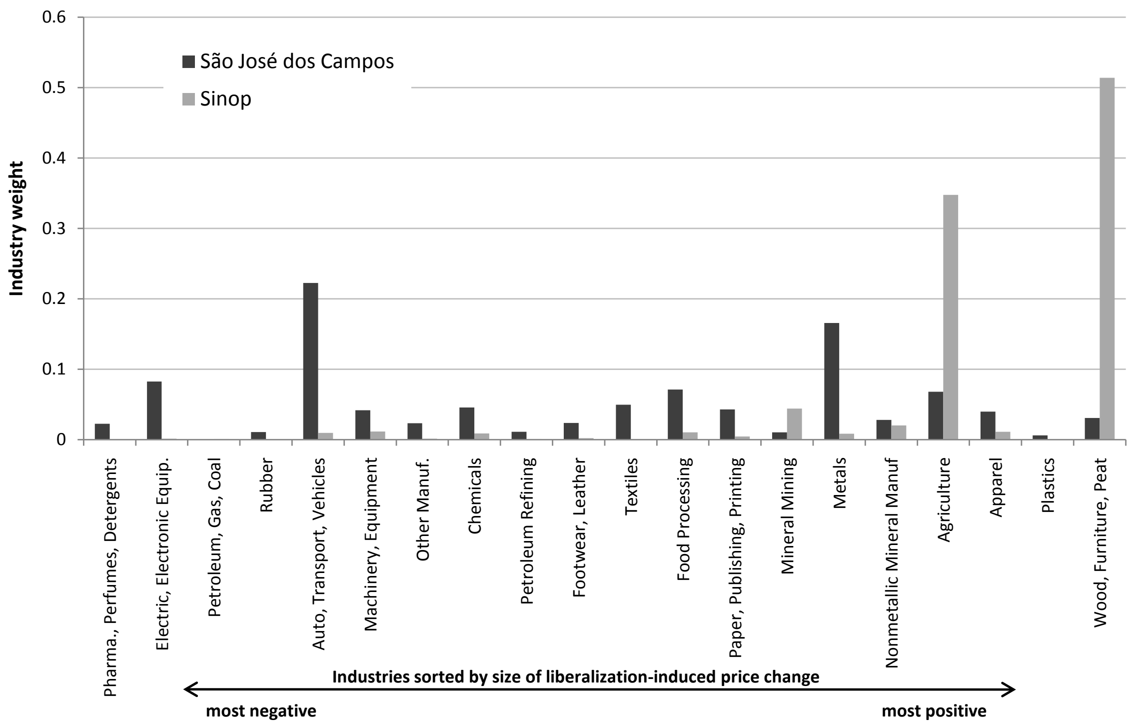
Source: Author's calculations (see text)
Sorted by national industry employment share in 1990 (largest to smallest)

Figure 6: Region-Level Tariff Changes



Weighted average of liberalization-induced price changes - see text for details.

Figure 7: Demonstration of Variation Behind Region-Level Tariff Changes



Source: Author's calculations (see text)

Table 1: The Effect of Tariff Changes on Price Changes

dependent variable: change in log wholesale price in Brazil

	(1)	(2)	(3)	(4)
$\Delta \ln(1 + \tau_i)$	0.069	0.147		
	(0.495)	(0.526)		
$\gamma i \, \Delta \ln(1 + \tau_i)$			12.409	12.258
			(5.614)*	(5.867)+
$\Delta \ln P_{US,i}$		0.452		0.108
		(0.854)		(0.745)
Constant	18.722	18.667	18.877	18.855
	(0.194)**	(0.224)**	(0.101)**	(0.185)**
R-squared	0.001	0.017	0.214	0.215

Standard errors in parentheses
+ significant at 10%; * significant at 5%; ** significant at 1%
20 industry observations
weighted by 1990 industry value added

Table 2: The Effect of Liberalization on Local Wages

	main		no labor share adjustment		nontraded price change set to zero		change in protection level		previous literature approach†		raw wage and price changes	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Regional tariff change	0.945	0.932	0.943	0.892	0.852	1.068	0.039	0.197	0.231	-0.033	0.292	0.314
standard error	(0.803)	(0.162)**	(0.770)	(0.150)**	(1.329)	(0.274)**	(0.226)	(0.077)*	(0.237)	(0.114)	(0.140)*	(0.048)**
p-value testing H_0: coefficient = 1	[0.946]	[0.678]	[0.942]	[0.479]	[0.912]	[0.807]	[0.000]	[0.000]	[0.003]	[0.000]	[0.000]	[0.000]
Dependent variable - regional wage change												
Regression-adjusted	X	X	X	X	X	X	X	X	X	X	.	.
Raw	.	.	.	.	.	.	.	.	.	.	X	X
Price change measure												
Liberalization-induced price change	X	X	X	X	X	X	.	.	.	.	.	.
Change in protection level	.	.	.	.	.	.	X	X	X	X	.	.
Raw price change	.	.	.	.	.	.	.	.	.	.	X	X
Nontraded sector												
Omitted	X	X	X	X	.	.	X	X	.	.	X	X
Price change set to zero	.	.	.	.	X	X	.	.	X	X	.	.
Labor share adjustment	X	X	.	.	X	X	X	X	.	.	.	.
State indicators (27)	.	X	.	X	.	X	.	X	.	X	.	X
R-squared	0.066	0.717	0.075	0.719	0.017	0.699	0.001	0.697	0.016	0.683	0.147	0.705

Standard errors adjusted for 27 state clusters (in parentheses)
p-value testing the null hypothesis that the regression coefficient on the regional tariff change equals 1 [in brackets]
+ significant at 10%, * significant at 5%, ** significant at 1%
493 micreregion observations (Manaus omitted)
Weighted by the inverse of the squared standard error of the estimated change in microregion wage, calculated using the procedure in Haisken-DeNew and Schmidt (1997)
† Implements the approach taken in much of the previous literature - see text for specific references and discussion

Table 3: Migration Summary Statistics

Demographic Group	Fraction of Population in Group 1991	Fraction of Group Migrated Between 1986-1991	Fraction of Population in Group 2000	Fraction of Group Migrated Between 1995-2000
All	100.00%	8.59%	100.00%	7.73%
Gender				
Female	50.73%	8.40%	50.69%	7.67%
Male	49.27%	8.79%	49.31%	7.80%
Age				
18-24	22.71%	11.66%	20.05%	10.36%
25-34	33.70%	9.72%	30.99%	8.97%
35-55	43.59%	6.12%	48.96%	5.87%
Education				
0	17.34%	7.39%	9.01%	6.27%
1-3	17.71%	8.79%	16.14%	7.25%
4-7	33.69%	8.78%	33.08%	7.91%
8-10	12.22%	8.91%	14.38%	8.30%
11-14	13.84%	8.76%	21.22%	7.75%
15+	5.19%	9.48%	6.17%	8.95%
Race				
White	53.32%	8.80%	54.81%	7.87%
Brown (Pardo)	40.66%	8.66%	37.53%	7.80%
Black	5.38%	6.55%	6.80%	6.33%
Asian	0.46%	6.91%	0.44%	6.89%
Indigenous	0.18%	4.52%	0.41%	7.31%
Marital Status				
Married	67.94%	8.49%	48.50%	7.13%
Unmarried	32.06%	8.80%	51.50%	8.30%
Family Size				
1-2	13.89%	12.01%	16.13%	10.13%
3-4	43.09%	8.66%	50.47%	7.57%
5-6	28.61%	7.81%	24.96%	6.82%
7+	14.41%	6.63%	8.44%	6.83%
Population	66,886,132	66,886,132	71,749,427	71,749,427
Observations	7,658,039	7,658,039	9,150,432	9,150,432

Source: Author's calculations based on 1991 and 2000 Brazilian Demographic Census
Sample: Age 18-55, not enrolled in school

Table 4: The Effect of Region-Level Tariff Changes on Location Choice

Demographic grouping	Coefficient estimate	H_0: Same Coef. Across Groups [p-value]	Observations
None	1.308		11,686
	(0.354)**		
Gender			
Female	0.871	[0.004]	15,247
	(0.315)**		
Male	1.851		
	(0.433)**		
Age			
Age 18-34	0.933	[0.032]	14,699
	(0.424)*		
Age 35-55	1.753		
	(0.326)**		
Education			
0-10 years	2.070	[0.108]	13,597
	(0.423)**		
11+ years	1.158		
	(0.501)*		
Race			
White	2.416	[0.000]	14,465
	(0.413)**		
Non-white	0.711		
	(0.417)+		
Marital Status			
Married	0.889	[0.000]	14,099
	(0.278)**		
Unmarried	3.937		
	(0.578)**		
Familiy Size			
4 or fewer	1.942	[0.247]	14,547
	(0.450)**		
5 or more	1.220		
	(0.448)**		

Sample: Age 18-55, not enrolled in school
Observations represent demographic group x source state x destination state triples
Standard errors clustered by 493 destination microregions
+ significant at 10%, * significant at 5%, ** significant at 1%
Dropped observations representing < 5 individuals in either period
Weighted by the square root of the number of observations in each cell in 2000

Table 5: Liberalization-induced Population Shifts

State	Liberalization-induced population change	Baseline population (1991)	Population change / Baseline population	Net migration inflow (1991)	Population change / Net inflow
Rondônia	2,957	496,683	0.60%	20,146	14.68%
Piauí	4,759	1,013,588	0.47%	-44,183	-10.77%
Tocantins	1,409	362,835	0.39%	8,882	15.86%
Mato Grosso do Sul	3,083	819,909	0.38%	13,958	22.09%
Paraíba	4,530	1,277,770	0.35%	-50,116	-9.04%
Ceará	8,995	2,647,901	0.34%	-80,459	-11.18%
Bahia	15,008	4,828,991	0.31%	-155,815	-9.63%
Mato Grosso	2,784	927,377	0.30%	71,075	3.92%
Alagoas	3,054	1,026,900	0.30%	-27,789	-10.99%
Acre	423	160,285	0.26%	-23	-1821.25%
Pernambuco	7,594	3,005,664	0.25%	-87,013	-8.73%
Pará	4,695	1,966,970	0.24%	19,516	24.06%
Maranhão	4,251	1,874,682	0.23%	-68,774	-6.18%
Paraná	8,845	4,051,315	0.22%	-118,002	-7.50%
Minas Gerais	13,265	7,361,171	0.18%	-70,165	-18.91%
Espírito Santo	1,991	1,210,929	0.16%	23,439	8.50%
Rio Grande do Norte	1,478	1,004,745	0.15%	-5,354	-27.60%
Goiás	2,184	1,891,203	0.12%	60,664	3.60%
Santa Catarina	871	2,245,776	0.04%	26,774	3.25%
Sergipe	123	611,871	0.02%	6,204	1.98%
Rio Grande do Sul	-4,033	4,539,675	-0.09%	-17,983	22.43%
Roraima	-125	96,031	-0.13%	16,536	-0.76%
Distrito Federal	-1,113	755,540	-0.15%	27,223	-4.09%
Rio de Janeiro	-11,127	6,360,260	-0.17%	-33,050	33.67%
Amapá	-362	110,499	-0.33%	9,423	-3.84%
São Paulo	-67,936	15,424,083	-0.44%	444,967	-15.27%

Source: Author's calculations based on 1991 and 2000 Brazilian Demographic Census

Sample: Age 18-55, not enrolled in school

Amazonas omitted due to presence of Manaus free-trade area

Table 6: Counterfactual Wage Analysis

	(1)	(2)
Regional tariff change	1.062	1.075
standard error	(0.809)	(0.165)**
p-value testing H_0: coefficient = 1	[0.940]	[0.651]
State indicators (27)	.	X
R-squared	0.081	0.722

A Specific Factors Model

A.1 Factor prices

This section closely follows Jones (1975), but deviates from that paper's result by allowing the amount of labor available to the regional economy to vary through migration. Consider a particular region, r, suppressing that subscript on all terms. Industries are indexed by $i = 1...N$. L is the total amount of labor and T_i is the amount of industry i-specific factor available in the region. a_{Li} and a_{Ti} are the respective quantities of labor and specific factor used in producing one unit of industry i output. Letting Y_i be the output in each industry, the factor market clearing conditions are

$$a_{Ti}Y_i = T_i \qquad \forall i, \tag{A1}$$

$$\sum_i a_{Li}Y_i = L. \tag{A2}$$

Under perfect competition, the output price equals the factor payments, where w is the wage and R_i is the specific factor price.

$$a_{Li}w + a_{Ti}R_i = P_i \qquad \forall i \tag{A3}$$

Let hats represent proportional changes, and consider the effect of price changes $\hat{P}_i$. θ_i is the cost share of the specific factor in industry i.

$$(1-\theta_i)\hat{w} + \theta_i\hat{R}_i = \hat{P}_i \qquad \forall i, \tag{A4}$$

which follows from the envelope theorem result that unit cost minimization implies

$$(1-\theta_i)\hat{a}_{Li} + \theta_i\hat{a}_{Ti} = 0 \qquad \forall i. \tag{A5}$$

Differentiate (A1), keeping in mind that T_i is fixed in all industries.

$$\hat{Y}_i = -\hat{a}_{Ti} \qquad \forall i \tag{A6}$$

Similarly, differentiate (A2), let $\lambda_i = \frac{L_i}{L}$ be the fraction of regional labor utilized in industry i, and substitute in (A6) to yield

$$\sum_i \lambda_i(\hat{a}_{Li} - \hat{a}_{Ti}) = \hat{L}. \tag{A7}$$

By the definition of the elasticity of substitution between T_i and L_i in production,

$$\hat{a}_{Ti} - \hat{a}_{Li} = \sigma_i(\hat{w} - \hat{R}_i) \qquad \forall i. \tag{A8}$$

Substituting this into (A7) yields

$$\sum_i \lambda_i\sigma_i(\hat{R}_i - \hat{w}) = \hat{L}. \tag{A9}$$

Equations (A4) and (A9) can be written in matrix form as follows.

$$\left[\begin{array}{cccc|c} \theta_1 & 0 & \dots & 0 & 1-\theta_1 \\ 0 & \theta_2 & \dots & 0 & 1-\theta_2 \\ \vdots & & \ddots & \vdots & \vdots \\ 0 & 0 & \dots & \theta_N & 1-\theta_N \\ \hline \lambda_1\sigma_1 & \lambda_2\sigma_2 & \dots & \lambda_N\sigma_N & -\sum_i \lambda_i\sigma_i \end{array}\right] \left[\begin{array}{c} \hat{R}_1 \\ \hat{R}_2 \\ \vdots \\ \hat{R}_N \\ \hline \hat{w} \end{array}\right] = \left[\begin{array}{c} \hat{P}_1 \\ \hat{P}_2 \\ \vdots \\ \hat{P}_N \\ \hline \hat{L} \end{array}\right] \tag{A10}$$

Rewrite this expression as follows for convenience of notation.

$$\left[\begin{array}{c|c} \Theta & \theta_L \\ \hline \lambda' & -\sum_i \lambda_i\sigma_i \end{array}\right] \left[\begin{array}{c} \hat{R} \\ \hline \hat{w} \end{array}\right] = \left[\begin{array}{c} \hat{P} \\ \hline \hat{L} \end{array}\right] \tag{A11}$$

Solve for $\hat{w}$ using Cramer's rule and the rule for the determinant of partitioned matrices.

$$\hat{w} = \frac{\hat{L} - \lambda'\Theta^{-1}\hat{P}}{-\sum_i \lambda_i\sigma_i - \lambda'\Theta^{-1}\theta_L} \tag{A12}$$

Note that the inverse of the diagonal matrix Θ is a diagonal matrix of $\frac{1}{\theta_i}$'s. This yields the effect of goods price changes and changes in regional labor on regional wages:

$$\hat{w} = \frac{-\hat{L}}{\sum_{i'} \lambda_{i'} \frac{\sigma_{i'}}{\theta_{i'}}} + \sum_i \beta_i \hat{P}_i \tag{A13}$$

$$\text{where} \qquad \beta_i = \frac{\lambda_i \frac{\sigma_i}{\theta_i}}{\sum_{i'} \lambda_{i'} \frac{\sigma_{i'}}{\theta_{i'}}} \tag{A14}$$

This expression with $\hat{L} = 0$ yields (1). Changes in specific factor prices can be calculated from wage changes by rearranging (A4).

$$\hat{R}_i = \frac{\hat{P}_i - (1-\theta_i)\hat{w}}{\theta_i} \tag{A15}$$

Plugging in (A13) and collecting terms yields the effect of goods price changes and changes in regional labor on specific factor price changes.

$$\hat{R}_i = \frac{(1-\theta_i)}{\theta_i} \frac{\hat{L}}{\sum_{i'} \lambda_{i'} \frac{\sigma_{i'}}{\theta_{i'}}} + \left(\beta_i + \frac{1}{\theta_i}(1-\beta_i)\right) \hat{P}_i - \frac{(1-\theta_i)}{\theta_i} \sum_{k \neq i} \beta_k \hat{P}_k \tag{A16}$$

Setting $\hat{L} = 0$ in (A13) and (A16) yields the equivalent expressions in Jones (1975).

A.2 Graphical representation of the model

The equilibrium adjustment mechanisms at work in the model are demonstrated graphically in Figure A1, which represents a two-region ($r = 1, 2$) and two-industry ($i = A, B$) version of the model.[27] Region 1 is relatively well endowed with the industry A specific factor. In each panel, the x-axis represents the total amount of labor in the country to be allocated across the two industries in the two regions, and the y-axis measures the wage in each region. Focusing on the left portion of panel (a), the curve labeled $P_A F_L^A$ is the marginal value product of labor in industry A, and the curve labeled $P_B F_L^B$ is the marginal value product of labor in industry B, measuring the amount of labor in industry B from right to left. Given labor mobility across sectors, the intersection of the two marginal value product curves determines the equilibrium wage, and the allocation of labor in region 1 between industries A and B, as indicated on the x-axis. The right portion of panel (a) is interpreted similarly for region 2. For visual clarity, the figures were generated assuming equal wages across regions before any price changes. This assumption is not necessary for any of the theoretical results presented in the paper.

Panel (a) of Figure A1 shows an equilibrium in which wages are equalized across regions. Since region 1 is relatively well endowed with industry A specific factor, it allocates a greater share of its labor to industry A. Panel (b) shows the effect of a 50% decrease in the price of good A, so the good A marginal value product curve in both regions moves down halfway toward the x-axis. Consistent with (1), the impact of this price decline is greater in region 1, which allocated a larger fraction of labor to industry A than did region 2. Thus, region 1's wage falls more than region 2's wage. Now workers in region 1 have an incentive to migrate to region 2. For each worker that migrates, the central vertical axis moves one unit to the left, indicating that there are fewer laborers in region 1 and more in region 2. As the central axis shifts left, so do the two marginal value product curves that are measured with respect to that axis. This shift raises the wage in region 1 and lowers the wage in region 2. Panel (c) shows the resulting equilibrium assuming costless migration and equalized wage across regions. Again, none of the theoretical results in the paper require this assumption.

A.3 Nontraded goods prices

As in Appendix A.1, consider a particular region, omitting the r subscript on all terms. Industries are indexed by $i = 1...N$. The final industry, indexed N, is nontraded, while other industries ($i \neq N$) are traded. The addition of a nontraded industry does not alter the results of the previous section, but makes it necessary to describe regional consumers' preferences to fix the nontraded good's equilibrium price.

Assume a representative consumer with Cobb-Douglas preferences over goods from each industry.[28] This implies the following relationship between quantity demanded by consumers, Y_i^c, total consumer income, m, and price, P_i.

$$\hat{Y}_i^c = \hat{m} - \hat{P}_i \tag{A17}$$

Consumers own all factors and receive all revenue generated in the economy; $m = \sum_i P_i Y_i^p$ where Y_i^p is the amount of good i produced in equilibrium. Applying the envelope theorem to the

[27]Figure A1 was generated under the following conditions. Production is Cobb-Douglas with specific-factor cost share equal to 0.5 in both industries. $\bar{L} = 10$, $T_{1A} = 1$, $T_{1B} = 0.4$, $T_{2A} = 0.4$, and $T_{2B} = 1$. Initially, $P_A = P_B = 1$, and after the price change, $P_A = 0.5$.

[28]A derivation with CES preferences, yielding very similar results, is available upon request.

revenue function for the regional economy, one can show[29]

$$\hat{m} = \eta_L \hat{L} + \sum_i \varphi_i \hat{P}_i. \tag{A18}$$

where η_L is the share of total factor payments accounted for by wages and φ_i is the share of regional production value accounted for by industry i. Plugging (A18) into (A17) for the nontraded industry and rearranging terms,

$$\hat{P}_N - \sum_{i \neq N} \frac{\varphi_i}{\sum_{i' \neq N} \varphi_{i'}} \hat{P}_i - \frac{\eta_L}{1-\varphi_N} \hat{L} = \frac{-1}{1-\varphi_N} \hat{Y}_N^c. \tag{A19}$$

Holding labor fixed, this expression shows that consumption shifts *away* from the nontraded good if the nontraded price increases relative to a weighted average of traded goods prices, with weights based on each good's share of traded sector output value.

A similar expression can be derived for the production side of the model. Wage equals the value of marginal product, so $\hat{w} = \hat{P}_N + \hat{F}_L^N$. The specific factor is fixed, so $\hat{Y}_N^p = (1-\theta_N)\hat{L}_N$. Combining these two observations with Euler's theorem and the definition of the elasticity of substitution as in footnote 7 yields

$$\hat{w} = \hat{P}_N - \frac{\theta_N}{\sigma_N(1-\theta_N)} \hat{Y}_N^p \tag{A20}$$

Substitute in the expression for $\hat{w}$ in (A13) and rearrange.

$$\hat{P}_N - \sum_{i \neq N} \frac{\beta_i}{\sum_{i' \neq N} \beta_{i'}} \hat{P}_i + \frac{1}{(1-\beta_N)\sum_{i'} \lambda_{i'} \frac{\sigma_{i'}}{\theta_{i'}}} \hat{L} = \frac{\theta_N}{(1-\beta_N)\sigma_N(1-\theta_N)} \hat{Y}_N^p \tag{A21}$$

Holding labor fixed, this expression shows that production shifts *toward* the nontraded good if the nontraded price increases relative to a weighted average of traded goods prices, with weights based on the industry's size and labor demand elasticity captured in β_i.

Equations (A19) and (A21) relate very closely to the intuition described in Section 2.2 for the case of only one traded good. Consumption shifts away from the nontraded good if its price rises relative to average traded goods prices while production shifts toward it. Since regional consumption and production of the nontraded good must be equal, with a single traded good the nontraded price must exactly track the traded good price. Combining (A19) and (A21) shows that in the many-traded-good case, the nontraded price change is a weighted average of traded goods price changes.

$$\hat{P}_N = \frac{\eta_L - \frac{\sigma_N}{\theta_N} \frac{(1-\theta_N)}{\sum_i \lambda_i \frac{\sigma_i}{\theta_i}}}{\sum_{i' \neq N} \left[\frac{\sigma_N}{\theta_N}(1-\theta_N)\beta_{i'} + \varphi_{i'}\right]} \hat{L} + \sum_{i \neq N} \xi_i \hat{P}_i \tag{A22}$$

[29] Since the amount of labor in the economy may change due to migration, the revenue maximization problem is specified as follows.

$$r(P_1 ... P_N, L) = \max_{Y_i^p, L_i} \sum_i P_i Y_i^p \quad s.t. \quad \sum_i L_i \leq L \quad \text{and} \quad Y_i^p \leq F^i(L_i, T_i)$$

$$\text{where} \qquad \xi_i = \frac{\frac{\sigma_N}{\theta_N}(1-\theta_N)\beta_i + \varphi_i}{\sum_{i' \neq N}\left[\frac{\sigma_N}{\theta_N}(1-\theta_N)\beta_{i'} + \varphi_{i'}\right]} \tag{A23}$$

A.4 Restriction to Drop the Nontraded Sector from Weighted Averages

Under Cobb-Douglas production with equal factor shares across industries, $\theta_i = \theta$, and $\sigma_i = 1\ \forall i$. This implies that $\beta_i = \lambda_i$, the fraction of labor in each industry. Similarly,

$$\varphi_i \equiv \frac{P_i Y_i^p}{\sum_{i'} P_{i'} Y_{i'}^p} = \frac{\lambda_i \eta_L}{1-\theta} = \lambda_i. \tag{A24}$$

since $\eta_L = 1 - \theta$ when $\theta_i = \theta$. When $\beta_i = \varphi_i$, the weights ξ_i in (A22) are

$$\xi_i = \frac{\beta_i}{\sum_{i' \neq N} \beta_{i'}} \tag{A25}$$

Plug these weights into (3) and then plugging that into (1) yields a result equivalent to dropping the nontraded sector from the weighted average in (1) and (2).

$$\hat{w} = \frac{\sum_{i \neq N} \beta_i \hat{P}_i}{\sum_{i' \neq N} \beta_i'} \tag{A26}$$

A.5 Wage Impact of Changes in Regional Labor

This section shows that an increase in regional labor decreases the regional wage. Substituting (A22) into (A13), holding traded goods prices fixed ($\hat{P}_i = 0\ \forall i \neq N$), and rearranging yields

$$\hat{w} = \left[\frac{-\frac{\sigma_N}{\theta_N}(1-\theta_N) - (1-\varphi_N) + \lambda_N \frac{\sigma_N}{\theta_N}\eta_L}{\left(\sum_i \lambda_i \frac{\sigma_i}{\theta_i}\right)\left[\frac{\sigma_N}{\theta_N}(1-\theta_N)(1-\beta_N) + (1-\varphi_N)\right]}\right]\hat{L}. \tag{A27}$$

Since the denominator is strictly positive, the sign of the relationship between $\hat{L}$ and $\hat{w}$ is determined by the numerator. An increase in regional labor will decrease the wage if an only if

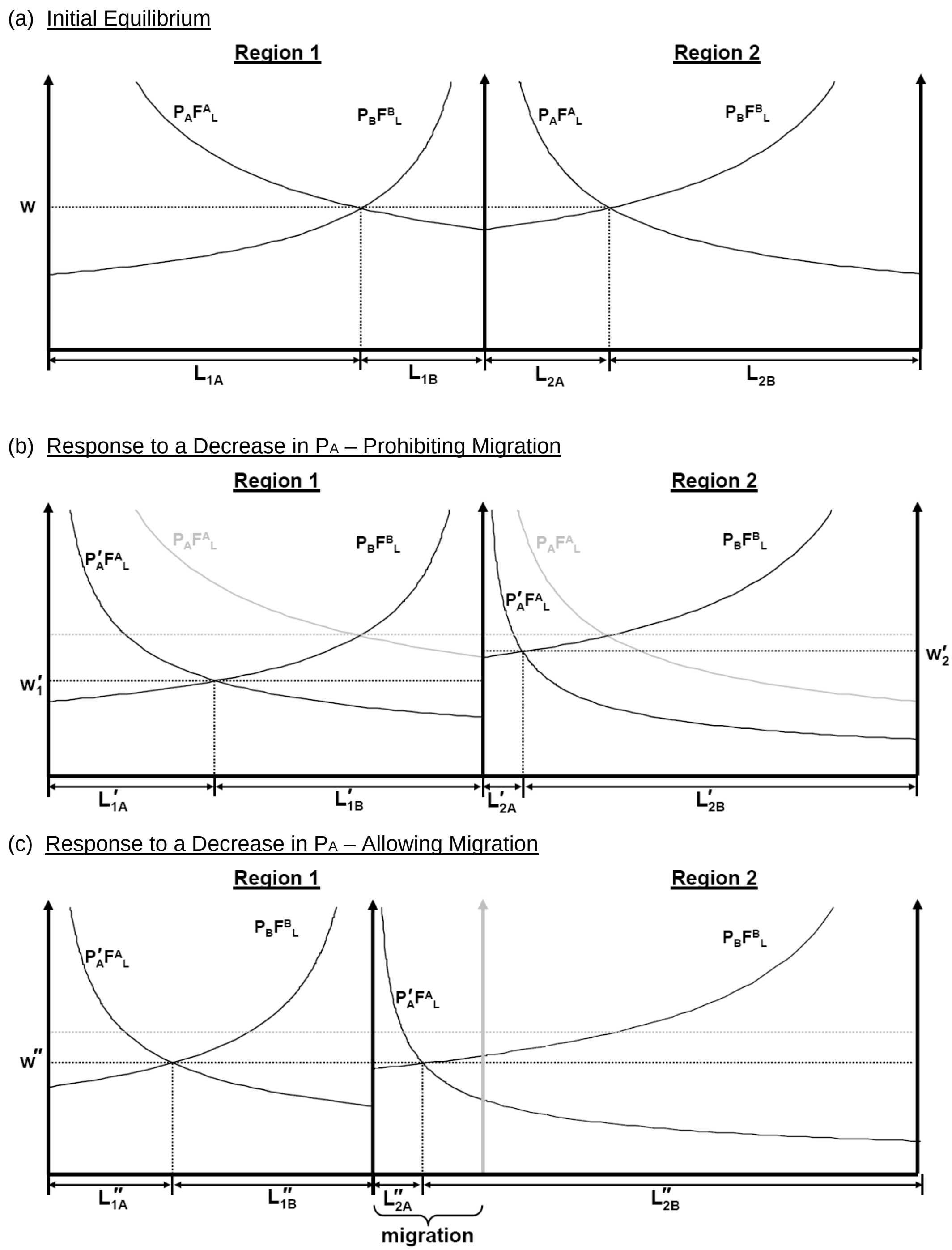
$$(1-\varphi_N) > \frac{\sigma_N}{\theta_N}(\lambda_N \eta_L - (1-\theta_N)). \tag{A28}$$

Using the fact that $\eta_L = (\sum_i \lambda_i (1-\theta_i)^{-1})^{-1}$ the previous expression is equivalent to

$$(1-\varphi_N) > \frac{\sigma_N}{\theta_N}(1-\theta_N)\left(\frac{\lambda_N(1-\theta_N)^{-1}}{\sum_i \lambda_i (1-\theta_i)^{-1}} - 1\right). \tag{A29}$$

The left hand side is positive while the right hand side is negative. Thus, the inequality always holds, and an increase in regional labor will always lower the regional wage.

Figure A1: Graphical Representation of Specific Factors Model of Regional Economies



B Data Appendix

Industry Crosswalk

National accounts data from IBGE and trade policy data from (Kume et al. 2003) are available by industry using the the Nível 50 and Nível 80 classifications. The 1991 Census reports individuals' industry of employment using the *atividade* classification system, while the 2000 Census reports industries based on the newer *Classificação Nacional de Atividades Econômicas - Domiciliar* (*CNAE-Dom*). Table B1 shows the industry definition used in this paper and its concordance with the various other industry definitions in the underlying data sources. The concordance for the 1991 Census is based on a crosswalk between the national accounts and *atividade* industrial codes published by the IBGE (2004). The concordance for the 2000 Census is based on a crosswalk produced by IBGE's Comissão Nacional de Classificação, available at *http://www.ibge.gov.br/concla/*.

Trade Policy Data

Nível 50 trade policy data come from Kume et al. (2003). Depending on the time period, they aggregated tariffs on 8,750 - 13,767 individual goods first using unweighted averages to aggregate individual goods up to the Nível 80 level, and then using value added weights to aggregate from Nível 80 to Nível 50.[30] In order to maintain this weighting scheme, I weight by value added when aggregating from Nível 50 to the final classification listed in Table B1. For reference, Figure B1 and Figure B2 show the evolution of nominal tariffs and effective rates of protection in the ten largest sectors by value added. Note that along with a general reduction in the level of protection, the dispersion in protection was also greatly reduced during liberalization, consistent with the goal of aligning domestic production incentives with world prices. It is clear that the move from a high-level, high-dispersion tariff structure to a low-level low-dispersion tariff distribution generated substantial variation in protection changes across industries; industries with initially high levels of protection experienced the largest cuts, while those with initially lower levels experienced smaller cuts.

Goods Prices

Brazilian goods prices are measured using the series IPA-OG (Indice de Precos por Atacado - Oferta Global), which was abandoned in 2008, and the newer IPA-Origem (Indice de Precos ao Productor Amplo - Origem). Both series are prepared by the Instituto Brasileiro de Economia of the Fundação Getulio Vargas (FGV-IBRE) and available online through the IPEAdata website, *http://www.ipeadata.gov.br/*. I use IPA-OG for most series and the IPA-Origem for a few aggregate indexes not reported for IPA-OG[31] I then assign these indexes to Nível 100 industries using the mapping in Muendler (2003b), aggregate using unweighted averages from Nível 100 to Nível 50, and then weighted by value added from Nível 50 to the final industry classification.

U.S. goods prices by Nível 100 industry come from Muendler (2003a). They were generated from BLS PPI by SIC industry, including all traded industries except for agriculture. I use yearly averages from the monthly data, aggregate using unweighted averages from Nível 100 to Nível 50,

[30] Email correspondence with Honório Kume, March 12, 2008.

[31] The aggregate indexes are 3 - Average Price; 18 - Agricultural Products, Total; 27 - Industrial Products, Total; 28 - Mineral Mining; and 29 - Manufacturing, Total.

and then weighted by imports from Nível 50 to the final industry classification. For agricultural prices, I use the yearly USDA price series for Prices Received - All Farm Products, available online from the National Agricultural Statistics Service[32].

Cross-Sectional Wage Regressions

In order to calculate the regional wage change for each microregion, I estimate standard wage regressions separately in 1991 and 2000. Wages are calculated as an individual's monthly earnings / 4.33 divided by weekly hours in their main job (results using all jobs are nearly identical). I regress the log wage on age, age-squared, a female indicator, an inner-city indicator, four race indicators, a marital status indicator, and fixed effects for each year of education from 0 to 17+, 21 industry fixed effects, and 494 microregion fixed effects. By running these regressions separately by survey year, I control for changes in regional demographic characteristics and for changes in the national returns to those characteristics. The results of these regressions are reported in Table B2. All terms are highly statistically significant and of the expected sign. The regional fixed effects are normalized relative to the national average log wage, and their standard errors are calculated using the process described in Haisken-DeNew and Schmidt (1997). The regional wage change is then calculated as the difference in these normalized regional fixed effects between 1991 and 2000:

$$d\ln w_r = (\ln w_r^{2000} - \ln w_r^{1991}) - (\overline{\ln w_r^{2000}} - \overline{\ln w_r^{1991}}) \tag{B1}$$

The regional wage changes are shown in Figure 4. As mentioned in the main text, some sparsely populated regions have very large measured regional wage changes. To demonstrate that these results are driven by the data and not some artefact of the wage regressions, Figure B3 shows a similar map plotting regional wage changes that were generated without any demographic or industry controls. They represent the change in the mean log wage in each region. The amount of variation across regions is similar, and these unconditional regional wage changes are highly correlated with the conditional versions used in the analysis, with a correlation coefficient of 0.93.

Price Change Regression Elements

The following figures present the individual elements used in the price vs. protection regressions in Section 5.2. The change in log price faced by producers is shown in Figure B4. The change in the log of one plus the effective rate of protection (ERP) is shown in Figure B5. Import penetration, calculated as imports divided by the sum of imports and domestic production based on national accounts data, is shown in Figure B6.

Region-Level Tariff Change Elements

The fixed factor share of input costs is measured as one minus the labor share of value added in national accounts data. The resulting estimates of θ_i are plotted in Figure B7. Theta calculation from national accounts. The distribution of laborers across industries in each region (λ_{ri}) are calculated in the 1991 Census using the sample of employed individuals, not enrolled in school, aged 18-55.

[32] *http://usda.mannlib.cornell.edu/MannUsda/viewDocumentInfo.do?documentID=1003*

Figure B1: Nominal Tariff Timeline

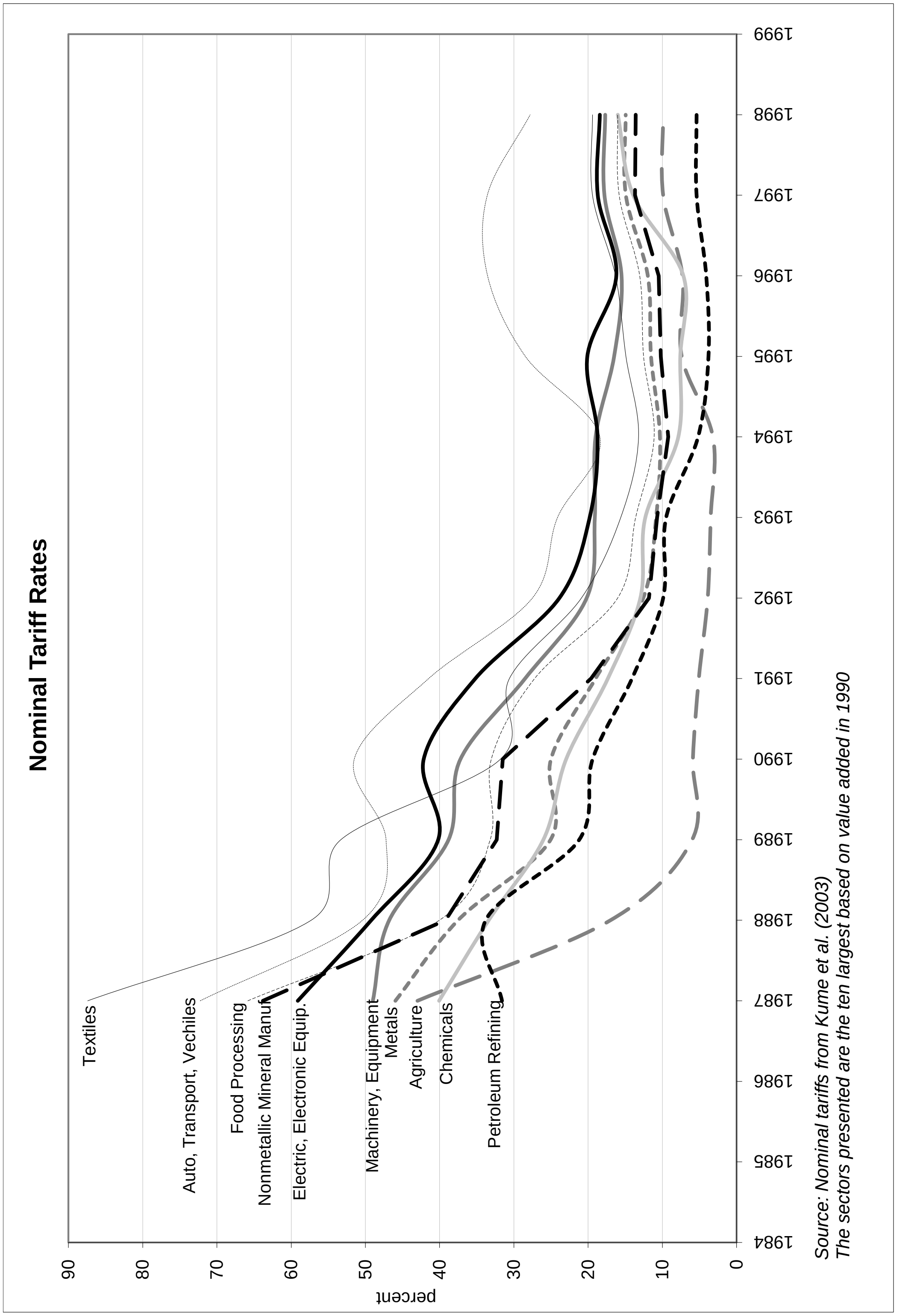
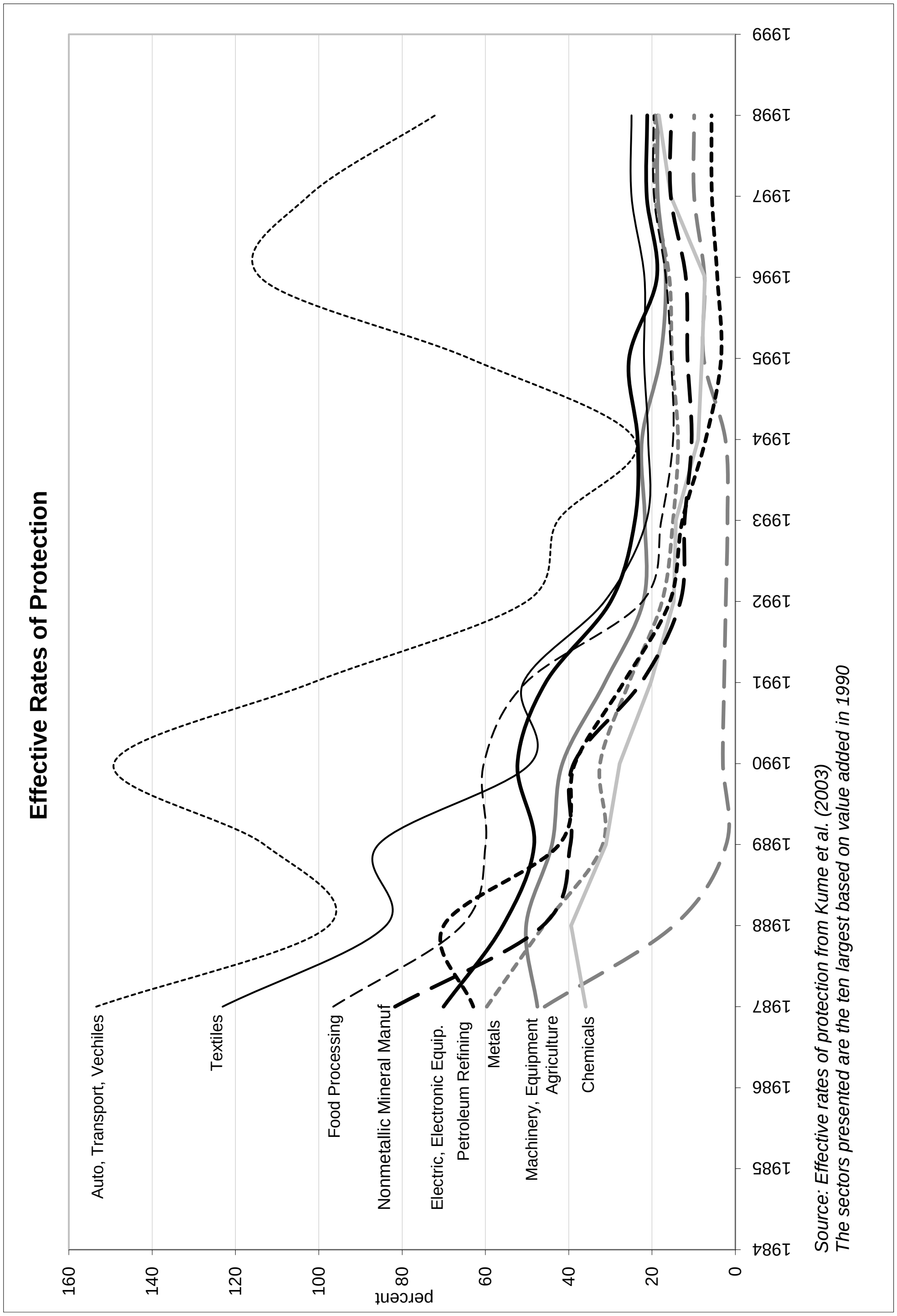
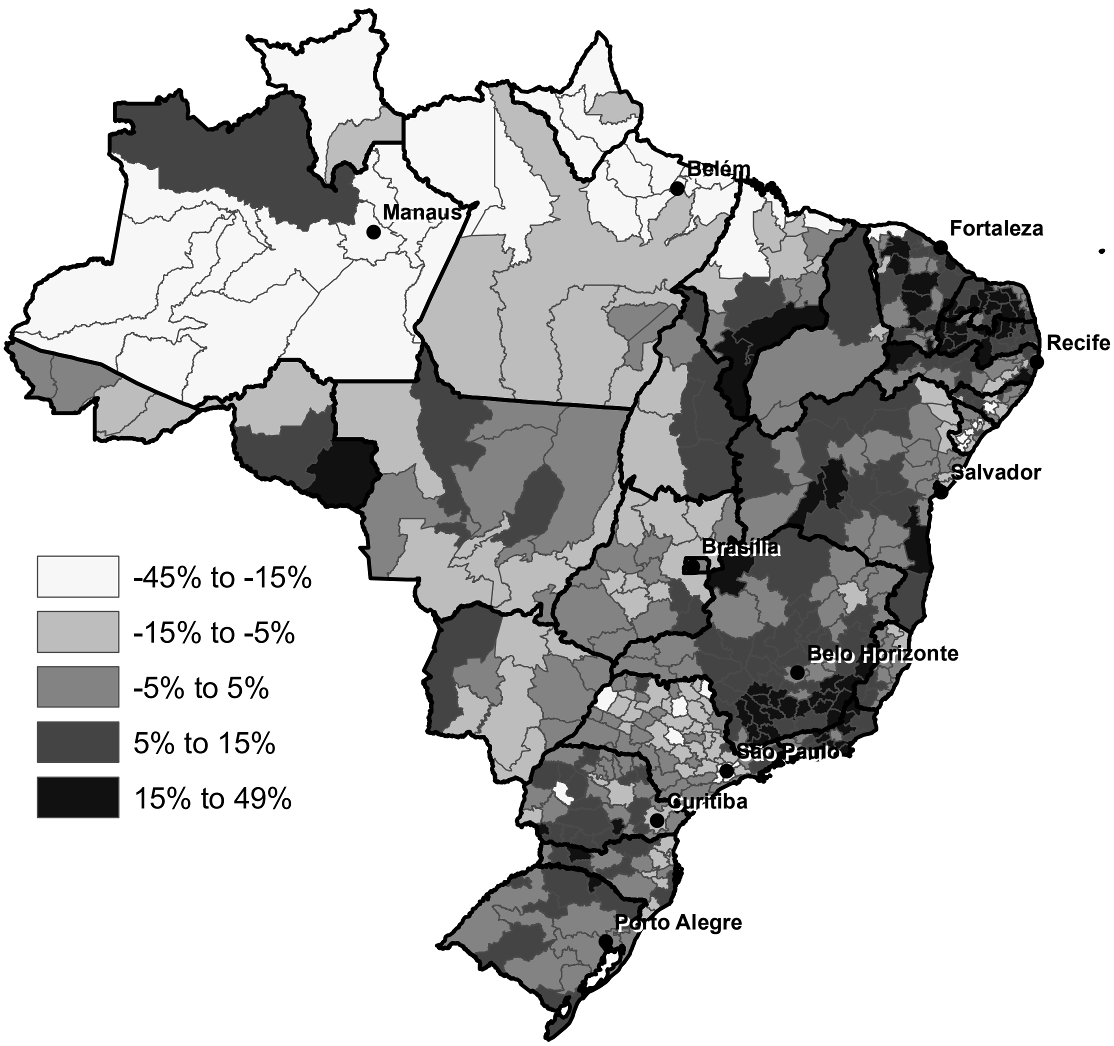


Figure B2: Effective Rate of Protection Timeline



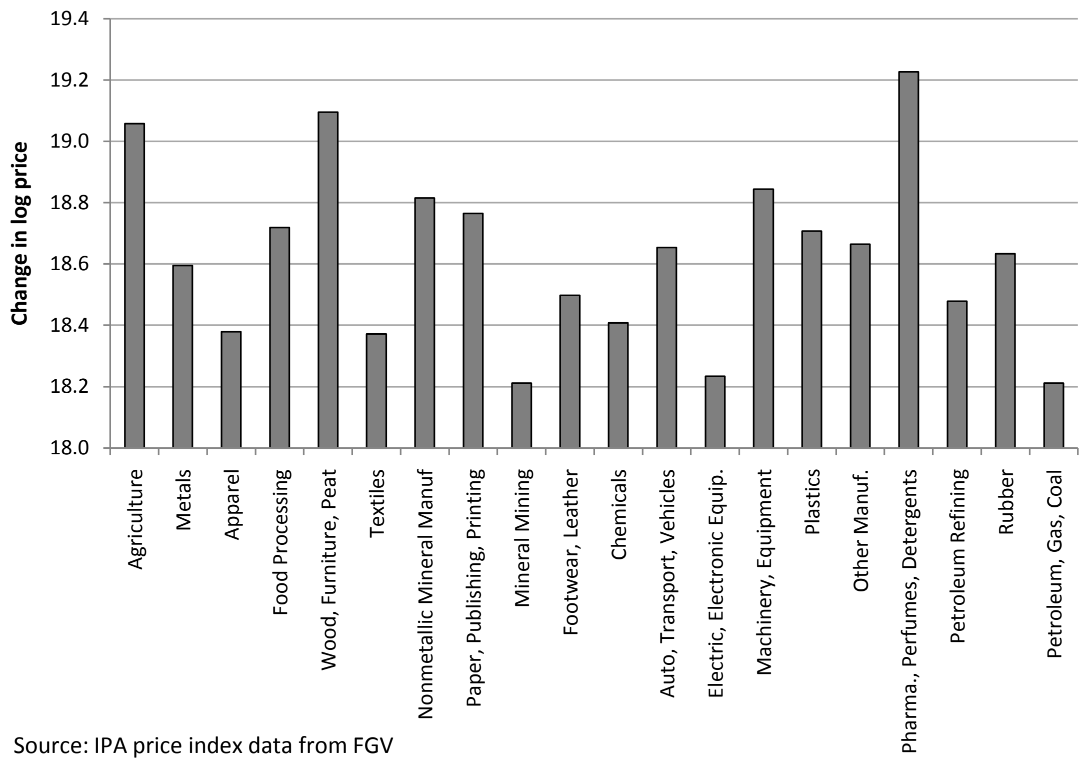
Source: Effective rates of protection from Kume et al. (2003)
The sectors presented are the ten largest based on value added in 1990

Figure B3: Unconditional Regional Wage Changes



Proportional wage change by microregion - normalized change in microregion fixed effects without demographic controls

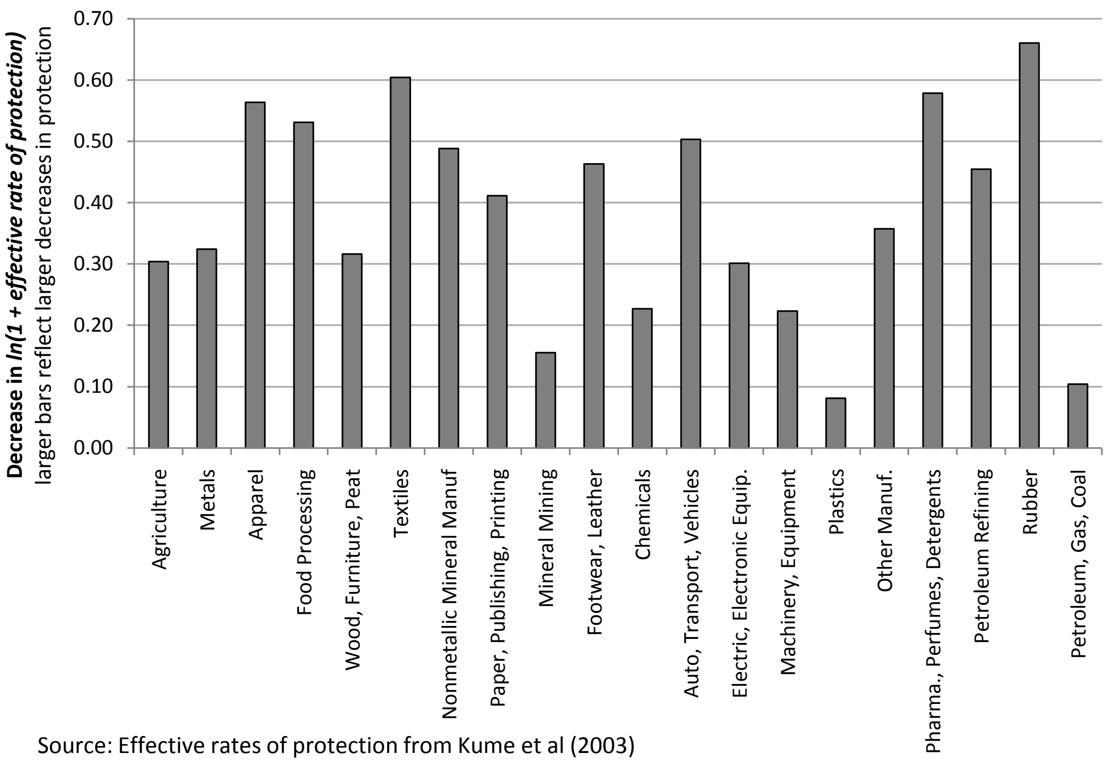
Figure B4: Price Changes



Source: IPA price index data from FGV

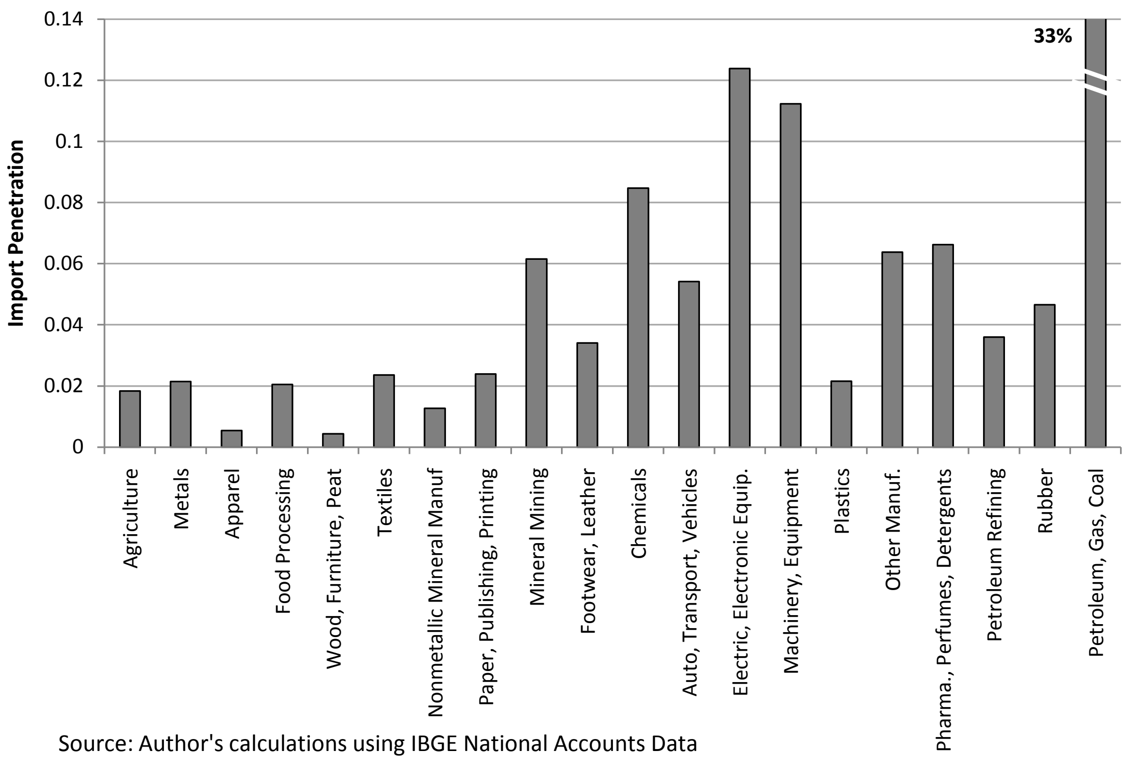
Sorted by national industry employment share in 1990 (largest to smallest)

Figure B5: Change in Effective Rate of Protection



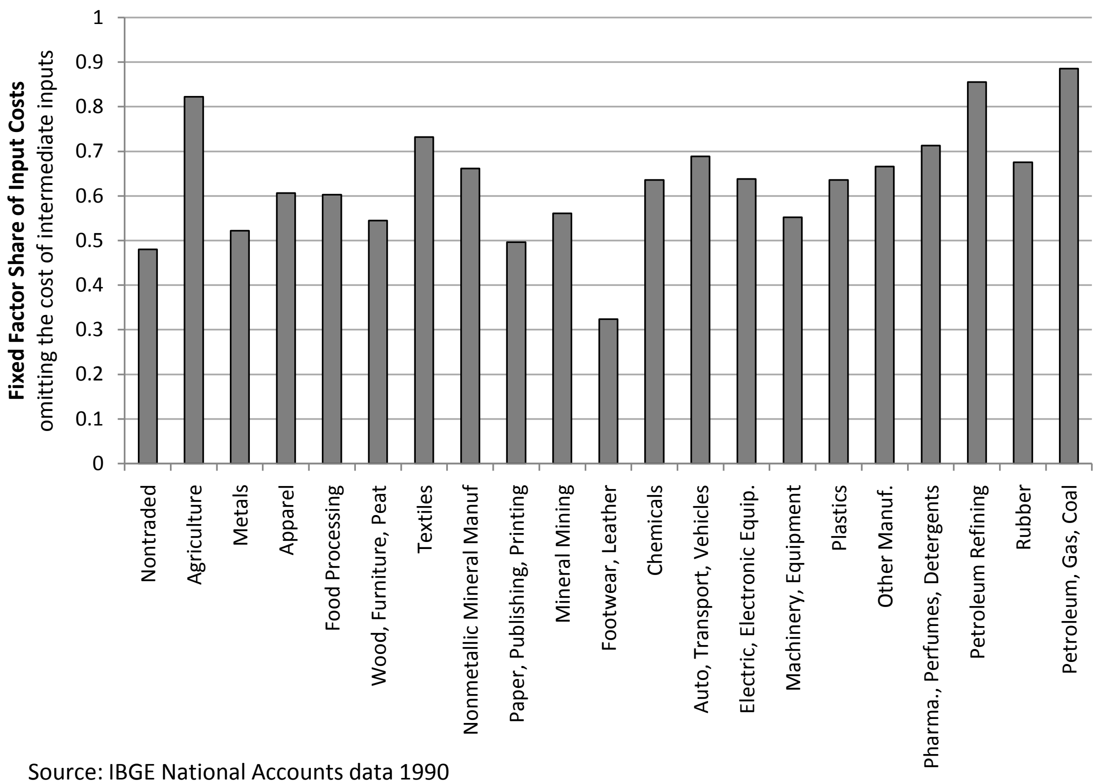
Source: Effective rates of protection from Kume et al (2003)
Sorted by national industry employment share in 1990 (largest to smallest)

Figure B6: Import Penetration



Source: Author's calculations using IBGE National Accounts Data
Sorted by national industry employment share in 1990 (largest to smallest)

Figure B7: Fixed Factor Share of Input Costs



Source: IBGE National Accounts data 1990
Sorted by national industry employment share in 1990 (largest to smallest)

Table B1: Industry Aggregation and Concordance

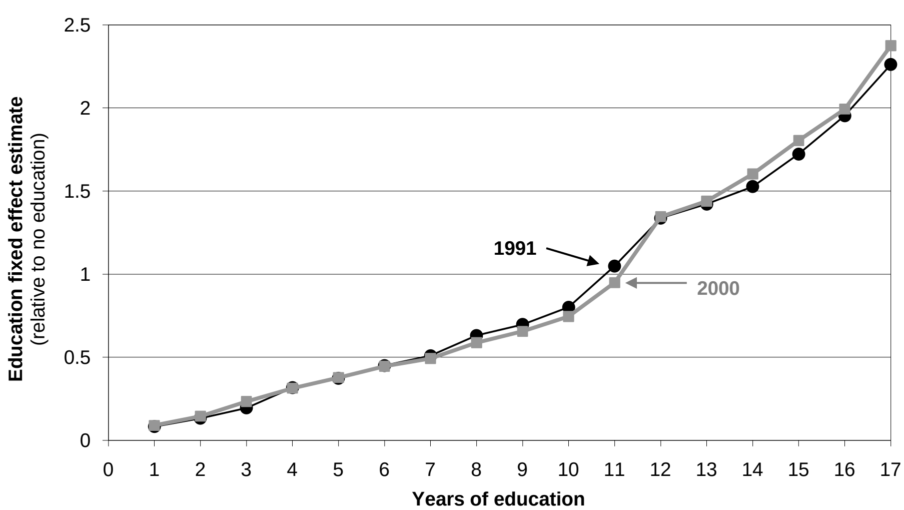
Final Industry	Sector Name	Nível 50	Nível 80	1991 Census (*atividade*)	2000 Census (*CNAE-Dom*)
1	Agriculture	1	101-199	011-037, 041, 042, 581	1101-1118, 1201-1209, 1300, 1401, 1402, 2001, 2002, 5001, 5002
2	Mineral Mining (except combustibles)	2	201-202	050, 053-059	12000, 13001, 13002, 14001-14004
3	Petroleum and Gas Extraction and Coal Mining	3	301-302	051-052	10000, 11000
4	Nonmetallic Mineral Goods Manufacturing	4	401	100	26010, 26091, 26092
5	Iron and Steel, Nonferrous, and Other Metal Production and Processing	5-7	501-701	110	27001-27003, 28001, 28002
8	Machinery, Equipment, Commercial Installation Manufacturing, and Tractor Manufacturing	8	801-802	120	29001
10	Electrical, Electronic, and Communication Equipment and Components Manufacturing	10-11	1001-1101	130	29002, 30000, 31001, 31002, 32000, 33003
12	Automobile, Transportation, and Vehicle Parts Manufacturing	12-13	1201-1301	140	34001-34003, 35010, 35020, 35030, 35090
14	Wood Products, Furniture Manufacturing, and Peat Production	14	1401	150, 151, 160	20000, 36010
15	Paper Manufacturing, Publishing, and Printing	15	1501	170, 290	21001, 21002, 22000
16	Rubber Product Manufacturing	16	1601	180	25010
17	Chemical Product Manufacturing	17,19	1701-1702, 1901-1903	200	23010, 23030, 23400, 24010, 24090
18	Petroleum Refining and Petrochemical Manufacturing	18	1801-1806	201, 202, 352, 477	23020
20	Pharmaceutical Products, Perfumes and Detergents Manufacturing	20	2001	210, 220	24020, 24030
21	Plastics Products Manufacturing	21	2101	230	25020
22	Textiles Manufacturing	22	2201-2205	240, 241	17001, 17002
23	Apparel and Apparel Accessories Manufacturing	23	2301	250,532	18001, 18002
24	Footwear and Leather and Hide Products Manufacturing	24	2401	190, 251	19011, 19012, 19020
25	Food Processing (Coffee, Plant Products, Meat, Dairy, Sugar, Oils, Beverages, and Other)	25-31	2501-3102	260, 261, 270, 280	15010, 15021, 15022, 15030, 15041-15043, 15050, 16000
32	Miscellaneous Other Products Manufacturing	32	3201	300	33001, 33002, 33004, 33005, 36090, 37000
99	Nontraded Goods and Services	33-43	3301-4301	340, 351-354, 410-424, 451-453, 461, 462, 464, 471-476, 481, 482, 511, 512, 521-525, 531, 533, 541-545, 551, 552, 571-578, 582-589, 610-619, 621-624, 631, 632, 711-717, 721-727, 901, 902	1500, 40010, 40020, 41000, 45001-45005, 50010, 50020, 50030, 50040, 50050, 53010, 53020, 53030, 53041, 53042, 53050, 53061-53068, 53070, 53080, 53090, 53101, 53102, 53111-53113, 55010, 55020, 55030. 60010. 60020. 60031. 60032. 60040. 60091. 60092, 61000, 62000, 63010, 63021, 63022, 63030, 64010, 64020, 65000, 66000, 67010, 67020, 70001, 70002, 71010, 71020, 71030, 72010, 72020, 73000, 74011, 74012, 74021, 74022, 74030, 74040, 74050, 74060. 74090. 75011-75017. 75020. 80011. 80012. 80090, 85011-85013, 85020, 85030, 90000, 91010, 91020, 91091, 91092, 92011-92015, 92020, 92030, 92040, 93010, 93020, 93030, 93091, 93092, 95000, 99000

Table B2: Cross-Sectional Wage Regressions - 1991 and 2000 Census

dependent variable: log wage = ln((monthly earnings / 4.33) / weekly hours) at main job

Year	1991	2000
Age	0.057	0.061
	(0.000)**	(0.000)**
Age^2 / 1000	-0.575	-0.601
	(0.004)**	(0.004)**
Female	-0.392	-0.335
	(0.001)**	(0.001)**
Inner City	0.116	0.101
	(0.001)**	(0.001)**
Race		
Brown (parda)	-0.136	-0.131
	(0.001)**	(0.001)**
Black	-0.200	-0.173
	(0.002)**	(0.001)**
Asian	0.154	0.122
	(0.006)**	(0.006)**
Indigenous	-0.176	-0.112
	(0.010)**	(0.006)**
Married	0.186	0.163
	(0.001)**	(0.001)**
Fixed Effects		
Years of Education (18)	X	X
Industry (21)	X	X
Microregion (494)	X	X
Observations	4,721,996	5,135,618
R-squared	0.518	0.500

Robust standard errors in parentheses
+ significant at 10%; * significant at 5%; ** significant at 1%
Omitted category: unmarried white male with zero years of education, outside inner city, working in agriculture



C Industry Aggregation

Start with the relationship between regional wages and industry prices described in (1) and make a slight change in notation. Assume industries i each consist of many goods g. Define $\mathbf{1}(ipc_{ig})$ as an indicator function for whether or not a good g in industry i faces import price competition. P_{ig}^{W} is the world price for that good.

$$\hat{P}_{ig} = \mathbf{1}(ipc_{ig})(1 \hat{+} \tau_i) + \hat{P_{ig}^{W}}. \tag{C1}$$

Plug this into (1), under the new notation including goods within industries.

$$\begin{aligned} \hat{w}_r &= \sum_i \sum_{g \in i} \beta_{rig} (\mathbf{1}(ipc_{ig})(1 \hat{+} \tau_i) + \hat{P}_{ig}^{W}) && \text{(C2)} \\ &= \sum_i (1 \hat{+} \tau_i) \sum_{g \in i} \beta_{rig} \mathbf{1}(ipc_{ig}) + \sum_i \sum_{g \in i} \beta_{rig} \hat{P}_{ig}^{W} && \text{(C3)} \end{aligned}$$

The empirical analysis will impose the additional restriction of Cobb-Douglas production, as it is not feasible to calculate elasticities of factor substitution by industry and region. This restriction along with identical technologies across regions implies that $\sigma_{rig} = 1$ and $\theta_{rig} = \theta_i$. Imposing this restriction implies

$$\begin{aligned} \sum_{g \in i} \beta_{rig} \mathbf{1}(ipc_{ig}) &= \frac{\frac{1}{\theta_i} \sum_{g \in i} L_{rig} \mathbf{1}(ipc_{ig})}{\sum_{i'} \frac{1}{\theta_{i'}} \sum_{g' \in i'} L_{ri'g'}} && \text{(C4)} \\ &= \frac{\frac{L_{ri}}{\theta_i} \frac{\sum_{g \in i} L_{rig} \mathbf{1}(ipc_{ig})}{L_{ri}}}{\sum_{i'} L_{ri'} \frac{1}{\theta_{i'}}} && \text{(C5)} \\ &= \beta_{ri} \phi_{ri} && \text{(C6)} \end{aligned}$$

$$\text{where} \qquad \phi_{ri} \equiv \sum_{g \in i} \frac{L_{rig}}{L_{ri}} \mathbf{1}(ipc_{ig}) \tag{C7}$$

ϕ_{ri} is the fraction of industry i workers producing goods that face import competition. Now consider the second term in (C3).

$$\begin{aligned} \sum_{g \in i} \beta_{rig} \hat{P}_{ig}^{W} &= \frac{\frac{1}{\theta_i} \sum_{g \in i} L_{rig} \hat{P}_{ig}^{W}}{\sum_{i'} \frac{1}{\theta_{i'}} \sum_{g' \in i'} L_{ri'g'}} && \text{(C8)} \\ &= \frac{\frac{L_{ri}}{\theta_i} \frac{\sum_{g \in i} L_{rig} \hat{P}_{ig}^{W}}{L_{ri}}}{\sum_{i'}} && \text{(C9)} \\ &= \beta_{ri} \hat{P}_{i}^{W} && \text{(C10)} \end{aligned}$$

$$\text{where} \qquad \hat{P}_{i}^{W} \equiv \sum_{g \in i} \frac{L_{rig}}{L_{ri}} \hat{P}_{ig}^{W} \tag{C11}$$

$\hat{P}_i^W$ is the average proportional change in prices in industry i, with weights based on the amount of labor producing each good in the industry. Although it is impossible to obtain world prices with this particular weighting scheme, it is likely that industry level world prices calculated with a similar weighted mean structure will closely approximate this expression. Plugging these results back into (C3), yields the result of the aggregation.

$$\hat{w}_r = \sum_i \beta_{ri}(\phi_{ri}(1 \hat{+} \tau_i) + \hat{P}_i^W) \tag{C12}$$

In the empirical analysis, industry import penetration, γ_i, is used as a proxy for the fraction of goods in the industry facing import competition, ϕ_{ri}.

D Location Choice Estimation Equation Derivation

This appendix follows Scanlon et al. (2002) and Cadena (2007) to difference out time invariant unobservable terms from the location choice specification described in (12). The observed share of individuals in source s who choose to live in destination d at time t, S_{sdt}, consists of the true choice probability, π_{sdt}, and mean zero random sampling error, ξ_{sdt}.

$$S_{sdt} = \frac{e^{V_{sdt}}}{D_{st}} + \xi_{sdt} \tag{D1}$$

Taking logs yields

$$\ln S_{sdt} = ln(e^{V_{sdt}} + \xi_{sdt} D_{st}) - \ln D_{st}. \tag{D2}$$

A first-order Taylor series approximation evaluated at $\xi_{sdt} = 0$ yields

$$\ln S_{sdt} \approx V_{sdt} - lnD_{st} + \frac{\xi_{sdt}}{\pi_{sdt}}. \tag{D3}$$

Plugging in the definition of V_{sdt} from (11),

$$\ln S_{sdt} \approx \alpha \ln w_{dt} + \mu_{sdt} + \eta_{sd} - lnD_{st} + \frac{\xi_{sdt}}{\pi_{sdt}}. \tag{D4}$$

The model is still nonlinear in α, due to its presence within D_{st}. This term can be canceled by subtracting the log share of an arbitrary reference destination. For convenience, the reference region is s, the individual's initial region of residence.

$$\ln S_{sdt} - \ln S_{sst} \approx \alpha(\ln w_{dt} - \ln w_{st}) + (\mu_{sdt} - \mu_{sst}) + (\eta_{sd} - \eta_{ss}) + \left(\frac{\xi_{sdt}}{\pi_{sdt}} - \frac{\xi_{sst}}{\pi_{sst}}\right) \tag{D5}$$

Although the preceding expression is linear in α, it still contains unobserved components that may be correlated with log wages. The time invariant unobserved components, η_{sd}, can be canceled out by differencing over time.

$$d\ln S_{sd} - d\ln S_{ss} \approx \alpha(d\ln w_d - d\ln w_s) + \left[(d\mu_{sd} - d\mu_{ss}) + d\left(\frac{\xi_{sd}}{\pi_{sd}}\right) - d\left(\frac{\xi_{ss}}{\pi_{ss}}\right)\right] \tag{D6}$$