

RESEARCH SEMINAR IN INTERNATIONAL ECONOMICS

Gerald R. Ford School of Public Policy  
The University of Michigan  
Ann Arbor, Michigan 48109-3091

Discussion Paper No. 606

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

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January 10, 2010

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# Regional Labor Market Effects of Trade Policy: Evidence from Brazilian Liberalization\*

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Initial Draft: July 10, 2008  
This Draft: January 10, 2010

## Job Market Paper

### Abstract

This paper quantifies the effects of trade liberalization on local labor market outcomes and workers' migration patterns. I extend the classic specific-factors model to examine the impact of national price changes on local labor markets. The model describes how tariff changes across industries affect wages in local labor markets within the liberalizing country. In particular, 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. This result provides a theoretical foundation for a prevalent empirical approach in previous studies of local labor markets and lends economic interpretations to estimates that allow the researcher to evaluate the magnitude of results along with their direction.

I then use these theoretical results to measure how Brazil's 1987-1995 trade liberalization affected wages and interstate migration within the country. I find that regions whose output faced a 10% larger liberalization-induced price decline experienced a 7% larger wage decline. In addition, liberalization resulted in a substantial shift in migration patterns. The most affected Brazilian states gained or lost approximately 2% of their populations as a result of liberalization-induced shifts in migration patterns. These results demonstrate the empirical value of the specific-factors framework developed here and represent the first systematic evaluation of the effects of liberalization on internal migration.

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\*I would like to thank Martha Bailey, Rebecca Blank, Charlie Brown, Brian Cadena, Alan Deardorff, John DiNardo, Juan Carlos Hallak, Benjamin Keys, Osborne Jackson, David Lam, Alexandra Resch, James Sallee, Jeff Smith, and seminar participants at the University of Michigan, the University of Western Ontario, the Brazilian Econometric Society meetings in Salvador de Bahia, and the Pacific Conference for Development Economics at San Francisco State University for helpful comments on this research. Special thanks are due to Honorio Kume at IPEA for providing detailed information on the trade policy data utilized in this study. The author also gratefully acknowledges fellowship support from the Population Studies Center and the Rackham Graduate School at the University of Michigan.

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## 1 Introduction

Between 1988 and 1995, the Brazilian government abandoned a policy of import substitution in favor of drastic reductions in overall trade restrictions and a decrease in the variation of trade restrictions across industries. Along with the removal of non-tariff barriers, between 1987 and 1995 average tariffs fell from 54.9% to 10.8%, and the standard deviation of tariffs across industries fell from 21.3 to 7.4. Since the industrial composition of the labor force is quite varied across Brazilian states, the effects of trade liberalization were likely to have varying effects across different local labor markets in the country. In this paper, I develop a specific-factors model of regional economies to examine the relationships between trade liberalization and regional labor market outcomes. I then use the model's predictions to measure the liberalization's effect on wages in local labor markets and the effect on interstate migration patterns in Brazil.

I find that local labor markets whose workers are 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 7% larger wage decline, relative to other regions. Moreover, I find that workers responded to this change in the geographic returns to work by shifting inter-state migration patterns, with increased migration flows out of states whose labor force faced the largest tariff cuts and into states facing smaller cuts. The most affected Brazilian states gained or lost approximately 2% of their populations as a result of liberalization-induced shifts in migration patterns. Both of these findings support the theoretical predictions of the specific-factors model of regional economies and confirm its value in guiding empirical specifications.

This is, to my knowledge, the first study to systematically evaluate the effects of national trade policy on internal migration.<sup>1</sup> The findings contribute to the empirical trade and local labor markets literatures in a number of ways. First, the results demonstrate a fundamental link between national trade policy and regional employment, housing, transportation, and poverty policy. The

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<sup>1</sup>Although Aguayo-Tellez, Muendler and Poole (2009) do not measure the effect of trade liberalization on internal migration, they demonstrate that globalization in general may influence workers' location choices, finding that Brazilian workers at exporting firms are less likely to migrate and that migrants tend to choose destinations with a high concentration of foreign-owned firms.

theoretical and empirical results imply that trade policy makers can use their knowledge of the pre-liberalization industrial mix of different regions to predict what regions are likely to see the largest wage changes and subsequent migration due to a proposed change in tariff structure. This will allow national governments pursuing large trade reforms to anticipate which regions will experience increased demand for infrastructure and public services, facilitating coordination of regional policies with changes in national trade policy.

Second, the model presented here provides a clear theoretical foundation in which to understand the circumstances under which national trade policies have disparate effects across different regions of a country. Previous empirical studies examining India's trade liberalization utilize the pre-liberalization industry mix of a region's workforce to determine how the region will be affected by a set of tariff changes (Topalova 2005, Edmonds, Pavcnik and Topalova 2007, Hasan, Mitra and Ural 2007, Hasan, Mitra and Ranjan 2009).<sup>2</sup> The model developed here provides a theoretical foundation for the use of pre-liberalization industry mix to infer the effects of subsequent tariff changes. In particular, the model provides guidance on how to treat the nontraded sector and yields predictions both for the sign of liberalization's effects, but also for their magnitudes. This allows for sharper tests of the mechanisms through which liberalization effects local labor markets, and the empirical results support the model's predictions quite closely.

Third, this paper contributes to a growing empirical literature evaluating the effects of Brazilian trade liberalization on labor market outcomes. Since Brazil's liberalization was large, quickly implemented, and well documented, it has been a fruitful ground for research on the relationship between trade policy and inequality.<sup>3</sup> This paper broadens the scope of this previous literature by examining the differential effects of liberalization across geographic regions of Brazil, rather than only considering country-wide impacts of liberalization.

Finally, the results complement the conclusions of previous work examining the effects of national shocks on local labor markets in the U.S. (Bartik 1991, Blanchard and Katz 1992, Bound and

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<sup>2</sup>McCaig (2009) examines the effect of U.S. liberalization on labor market outcomes across Vietnamese regions, using a very similar empirical approach. To the extent that U.S. liberalization caused price changes faced by Vietnamese producers to vary across industries, the model developed here can be applied to that context as well.

<sup>3</sup>Goldberg and Pavcnik (2007) provide a summary, and more recent work includes Ferreira, Leite and Wai-Poi (2007) and Gonzaga, Filho and Terra (2006).

Holzer 2000). These studies examine the effects of changes in national industry mix on local labor markets, assuming that industry employment changes at the national level are exogenous to regional performance. This paper similarly maps national shocks into their regional effects, but contributes an explicit economic mechanism explaining the variation in national industry mix, showing that changes in national industry employment are driven by plausibly exogenous trade policy variation.<sup>4</sup> Since the specific-factors model of regional economies is based upon 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.

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 applies the specific-factors model in the context of trade liberalization and compares the resulting empirical specifications motivated by the model to those in previous work. Section 4 describes the data sets used, and Section 5 describes the specific trade policy changes implemented in Brazil's liberalization along with evidence in favor of the exogeneity of the tariff changes to industry performance. Section 6 presents an empirical analysis of the effects of trade liberalization on wages across local labor markets, and Section 7 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 8 concludes.

## 2 Specific-Factors Model of Regional Economies

This section 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. Each region's endowment of industry-specific factors drives the equilibrium allocation of labor across industries and determines the effect of goods price changes on regional wages. In the baseline model, price changes in industries that use a large amount of regional labor and have highly elastic labor demand will have the greatest impact on regional wages. Adding a nontraded sector to the model

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<sup>4</sup>See Figure 4 and the discussion in Section 5.

shows that local nontradables prices move with tradable prices, informing their empirical treatment. The section concludes by discussing the role of labor migration across regions in smoothing regional wage variation.

## 2.1 Baseline model

The baseline model treats each region within a country as a Jones (1975) specific-factors economy.<sup>5</sup> 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 industries, 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 (considered below). The second input,  $T$ , is specific to each industry in each region, i.e. it 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.<sup>6</sup> 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. Note that all theoretical results are derived in Appendix A (the following expression is (A13) with labor held constant).

$$\hat{w}_r = \sum_i \beta_{ri} \hat{P}_i \quad \forall r, \quad (1)$$

<sup>5</sup>The specific-factors model is generally used to model a country rather than a region. In such a framework, the current model could be applied to a customs union in which all member countries impose identical trade barriers and face identical prices.

<sup>6</sup>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  $T^\alpha$  is isomorphic to variation in the productivity term  $A$ .

$$\text{where } \beta_{ri} = \frac{L_{ri} \frac{\sigma_{ri}}{\theta_{ri}}}{\sum_{i'} L_{ri'} \frac{\sigma_{ri'}}{\theta_{ri'}}}. \quad (2)$$

Hats represent proportional changes,  $\sigma_{ri}$  is the elasticity of substitution between  $T$  and  $L$ , and  $\theta_{ri}$  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 labor demand elasticity equals  $\frac{\sigma}{\theta}$ .<sup>7</sup> 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 a large amount of labor from other industries in order to effect the decrease in the marginal product of labor necessary to restore equilibrium. Thus, price changes in these industries result in larger wage changes after the industrial reallocation of labor.

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 gen-

<sup>7</sup>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.

erally represent a large fraction of the economy under study. As in the baseline model, industries are indexed by  $i = 1 \dots N$ . 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 results of the baseline model, but makes it necessary to describe regional consumers' preferences to determine the nontraded good's equilibrium price. I assume throughout that all individuals have identical homothetic preferences, permitting the use of a representative regional consumer. In particular, assume that each region's representative consumer has CES preferences over all goods and receives as income all wages and specific factor payments earned in the region.

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 (A39) with labor held constant).

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

$$\text{where } \xi_{ri} = \frac{\frac{(1-\theta_{rN})}{\theta_{rN}} \sigma_{rN} \beta_{ri} + \varphi_{ri} + (\sigma - 1) \mu_{ri}}{\sum_{i' \neq N} \frac{(1-\theta_{rN})}{\theta_{rN}} \sigma_{rN} \beta_{ri'} + \varphi_{ri'} + (\sigma - 1) \mu_{ri'}}. \quad (4)$$

$\varphi_{ri}$  is the share of regional production value accounted for by industry  $i$ ,  $\sigma$  is the elasticity of substitution across goods in consumption (not to be confused with  $\sigma_{ri}$ , the elasticity of substitution in production), and  $\mu_{ri}$  is the share of regional consumers' expenditure allocated to good  $i$ . 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.

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 strategies. The first approach assumes no price change for nontraded goods, since trade liberalization has no direct impact on the nontraded sector. This approach is not supported by the theory, which predicts that nontraded prices move with traded prices. Artificially setting the price change to zero in the large nontraded sector would greatly understate the scale of liberalization's impact on regional wages. The second approach removes the nontraded sector from the weighted average in (1). This approach is more consistent with the theoretical findings. If the nontraded price changes



by approximately the same amount as the average traded price, then dropping the nontraded price from (1) will have very little effect upon the overall average. Appendix A describes the conditions under which the nontraded sector will have exactly no effect on the overall average and can be omitted.<sup>8</sup> Ideally, one would simply calculate the terms in (4) using detailed data on production values and consumption shares across industries at the regional level. However, when data on regional production and consumption patterns are limited, the model implies that dropping the nontraded sector is likely to provide a 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. This interregional migration will tend to equalize the impact of the price change across regions.

The mechanisms behind this equalization are demonstrated graphically in Figure 1, which represents a two-region ( $r = 1, 2$ ) and two-industry ( $i = A, B$ ) version of the baseline model.<sup>9</sup> 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. Although not necessary for any of the more general results, the

<sup>8</sup>Omitting the nontraded sector will have no effect on the overall average when  $\xi_{ri} = \frac{\beta_{ri}}{1-\beta_{rN}}$ . Appendix A demonstrates this fact and describes the restrictions under which the condition will hold exactly, though  $\xi_{ri}$  and  $\beta_{ri}$  are likely to be closely related in general, since part of the cross-industry variation in  $\xi_{ri}$  comes directly from  $\beta_{ri}$ , and  $\varphi_{ri}$  is also likely to be highly correlated with  $\beta_{ri}$ .

<sup>9</sup>Figure 1 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,  $P_A = 0.5$ .

figures are generated under the assumption of costless interregional migration for ease of exposition.

Panel (a) of Figure 1 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  $A$  when wages are equalized. Panel (b) shows the effect of a 50% decrease in the price of good  $A$ , so the marginal value product curve in both regions moves down halfway toward the x-axis. As described in (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 to be allocated 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. Migration continues until regional wages are equalized.

The same equalizing effect of regional migration will occur in the more general model. The baseline model with variable labor demonstrates this effect (the following equation is (A13) with prices held fixed).

$$\hat{w}_r = \frac{-1}{\sum_i \lambda_{ri} \frac{\sigma_{ri}}{\theta_{ri}}} \hat{L}_r, \quad (5)$$

where  $\lambda_{ri} = \frac{L_{ri}}{L_r}$  is the fraction of regional labor allocated to industry  $i$ . This expression indicates that the aggregate regional labor demand elasticity is a weighted average of industry labor demand elasticities, with weights based on the allocation of labor across industries. As individuals migrate away from regions that were impacted relatively negatively by price changes and toward regions affected relatively positively, the wage difference between locations will shrink. 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 wage analysis presented in Section 6, which finds evidence of some equalizing migration, but not enough to completely equalize cross-region wage impacts of liberalization.

Migration in the presence of nontraded goods poses two 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, 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.<sup>10</sup> 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.

This section has described a specific-factors model of regional economies including many regions and many industries. The model yields predictions for the effects of goods price changes on regional wages, the prices of nontraded goods, and the incentives to migrate between regions. The framework developed here can be used to measure the local impacts of any event in which a country faces price changes that vary exogenously across industries. I apply the model to the analysis of trade policy and devote the next section to operationalizing the model in the context of trade liberalization.

### 3 Applying the Model to Trade Liberalization

The previous section described a general framework linking national price changes to wage changes in regional labor markets. Here, I apply the model's insights to the question of how trade liberalization impacts local labor markets within the liberalizing country. I first link the model's price-based predictions to trade liberalization by describing the relationship between tariff changes and price

<sup>10</sup>In particular, the proportional change in a region's real wage,  $\omega_r$ , can be expressed as follows:

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

where  $\mu_i$  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.

changes when using industry-level data. Then I compare the resulting empirical framework to the previous literature on the local effects of liberalization. The model's predictions motivate empirical specifications that are similar to those in previous work, but exhibit some important differences regarding functional forms, the treatment of nontradables, and the interpretation of the magnitude of local effects.

### 3.1 Relating Tariff Changes to Price Changes

In order to use the specific-factors model in Section 2 to measure the effects of trade liberalization on local labor markets within the liberalizing country, I first need to determine how tariff cuts affected the prices faced by producers. For simplicity I make the small country assumption that tariff changes do not affect world prices (i.e. no terms of trade effects). In the Brazilian context, the researcher must use industry-level tariff and price data rather than information on tariffs and prices for individual goods (see Section 4 for more details). I address the issue of industry tariff pass-through by modeling industries as aggregations over a number of goods, some of which face import competition while others do not. This simple aggregation strategy yields an estimation framework for measuring the effect of tariff changes on price changes at the industry level.

Starting with the result from the baseline model described in (1), make a slight change of notation. Industries  $i$  now consist of many goods  $g$ . Define  $\mathbf{1}(ipc_{ig})$  as an indicator function for whether or not good  $g$  in industry  $i$  faces import price competition and  $P_{ig}^W$  as the world price. The price faced by producers is then,

$$P_{ig} = (1 + \tau_i)^{\mathbf{1}(ipc_{ig})} P_{ig}^W \quad (6)$$

For particular goods that are exported and thus do not face import price competition,  $\mathbf{1}(ipc_{ig}) = 0$ , and the price faced by producers equals the world price. For imported goods,  $\mathbf{1}(ipc_{ig}) = 1$  and producers face the world price plus the tariff. Taking proportional changes,

$$\hat{P}_{ig} = \mathbf{1}(ipc_{ig})(1 + \hat{\tau}_i) + \hat{P}_{ig}^W. \quad (7)$$

Appendix B plugs this expression into (1) and aggregates from individual goods up to the industry level. The aggregation requires the additional restriction of Cobb-Douglas production (which was necessary for the empirical analysis in any case, since it is not feasible to calculate elasticities of factor substitution by industry and region). The result of the aggregation is

$$\hat{w}_r = \sum_i \beta_{ri}(\phi_{ri}(1 + \hat{\tau}_i) + \hat{P}_i^W), \quad (8)$$

where  $\phi_{ri}$  is the fraction of industry  $i$  workers in region  $r$  producing goods that face import competition. As described below, the empirical analysis uses industry import penetration as a proxy for cross-industry variation in  $\phi_{ri}$ . Import penetration measures are only available at the national level, and hence do not vary by region. Accordingly, I assume constant import competition exposure across regions for a given industry, so  $\phi_{ri} = \phi_i$ . Imposing this restriction in (8), and comparing the result to (1), we have

$$\hat{P}_i = \phi_i(1 + \hat{\tau}_i) + \hat{P}_i^W. \quad (9)$$

Thus, tariff changes will have the largest effect on prices in industries facing large amounts of import competition ( $\phi_i$  close to 1), and small effects on prices in export industries ( $\phi_i$  close to 0).

### 3.2 Summary and Comparison to Previous Work

The specific-factors model of regional economies in Section 2 describes the relationship between the prices of tradable goods and regional wages. To understand the model's predictions for the local effects of trade liberalization, plug the price-tariff relationship from (9) into (1) (setting world price changes to zero), and drop the nontraded sector as discussed in Section 2.2. This yields the following expression describing the effect of tariff changes on regional wages.

$$\hat{w}_r = \sum_{i \neq N} \beta_{ri} \phi_i (1 + \hat{\tau}_i) \quad \forall r, \quad (10)$$

$$\text{where} \quad \beta_{ri} = \frac{L_{ri} \frac{\sigma_{ri}}{\theta_{ri}}}{\sum_{i' \neq N} L_{ri'} \frac{\sigma_{ri'}}{\theta_{ri'}}}. \quad (11)$$

The empirical analysis below uses this relationship to measure the effects of trade liberalization on regional wages and subsequent interregional migration.

The expression in (10) is quite similar to the empirical specifications employed in previous studies of the effect of liberalization on local market outcomes such as poverty, child labor, and unemployment in India (Topalova 2005, Edmonds et al. 2007, Hasan et al. 2007, Hasan et al. 2009), with some important differences. In these papers, changes in “district-level tariffs,”  $\tau_r^D$ , are computed as follows (using present notation).<sup>11</sup>

$$\tau_r^D = \sum_i \delta_{ri} \Delta \tau_i \quad \forall r \quad (12)$$

$$\text{where} \quad \delta_{ri} = \frac{L_{ri}}{\sum_{i'=1}^I L_{ri'}}$$

Expressions (10) and (12) are both weighted averages of tariff changes with weights based (at least partly) on the region’s industrial allocation of labor. However, a number of differences are present as well.

First, in (12) tariff changes are expressed as simple differences rather than proportional changes in  $(1 + \tau_i)$ . For small  $\tau_i$ ,  $\ln(1 + \tau_i) \approx \tau_i$ , so proportional changes may approximate changes in tariff levels.<sup>12</sup> Second, the tariff pass-through adjustment,  $\phi_i$ , is omitted. Although this adjustment is essential when analyzing aggregate industry data in the Brazilian case, disaggregate data were used in the studies of India, so the pass-through adjustment may be less important in that context. Third, the weights omit the labor demand elasticity terms,  $\frac{\sigma_{ri}}{\theta_{ri}}$ , essentially assuming that these terms are equal across all industries and regions. It is well beyond the scope of this paper to estimate elasticities of substitution between labor and other factors that vary across all industries and regions of Brazil, so I assume Cobb-Douglas production with factor shares free to vary across industries. This restriction implies that  $\sigma_{ri} = 1$  and  $\theta_{ri} = \theta_i$ . I can calculate rough estimates of  $\theta_i$  from Brazilian national accounts data and find that including them in the calculation of  $\beta_{ri}$  or

<sup>11</sup>Note that Hasan et al. (2007) and Hasan et al. (2009) also use measures of non-tariff barriers.

<sup>12</sup>Although Brazil’s liberalization involved large tariff cuts, making the approximation quite inaccurate, tariff changes based on tariff levels yield roughly the same ranking of industries as proportional changes in  $(1 + \tau_i)$ , so the choice does not affect the sign of the results.

omitting them does not substantially change the empirical results. Thus, although these differences should be accounted for in future work, none appears to cause economically significant deviations from the model's predictions.

The model also provides guidance on treatment of the nontraded sector. Topalova (2005) and Edmonds et al. (2007) estimate two versions of the weighted average in (12), one with the nontraded price change set to zero, and one dropping the nontraded sector, as in (10). The latter version is then used as an instrument for the former. Hasan et al. (2007) and Hasan et al. (2009) simply drop the nontraded sector and use that measure directly. As discussed in Section 2.2, the analysis presented here strongly favors dropping the nontraded sector. This measure should be used directly, omitting the version with zero nontraded price change entirely. Keep in mind that in cases where detailed production and expenditure data are available by region, the researcher can simply calculate the predicted tariff-induced nontraded price change in each region based on (3).

The theory-motivated approach clarifies the labor demand channel through which liberalization impacts regional labor markets and allows the researcher to carefully evaluate the magnitude of the effects of liberalization in testing the model's predictions. The model relates wage changes with tariff changes, and predicts a one-to-one relationship between proportional regional wage changes and the weighted average of tariff changes in (10). In the empirical analysis of Section 6, I examine this relationship directly, and find slightly smaller effects than the one-to-one relationship, as expected given some regional migration. Without the theoretical predictions, such a test of the sign and *magnitude* of local effects would not be possible. Thus, the theory allows the analysis to move beyond examining only the sign of estimates and provides a sharper test of the empirical model.

Given the many similarities, the model developed here provides a theoretical foundation for the general approach employed by previous empirical work on the local effects of liberalization. However, the differences just discussed provide important guidance on the appropriate implementation of empirical analyses. The remainder of this paper tests the model's predictions regarding the impact of trade policy changes on regional wages and interregional migration patterns in the context of Brazil's 1987-1995 trade liberalization, and finds strong evidence supporting the model.

## 4 Data

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), who aggregated tariffs on 8,750 - 13,767 individual goods, depending on the time period. 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.

Import penetration data, used to proxy for tariff pass-through adjustment in (9), were calculated from Brazilian National Accounts data available from the Brazilian Census Bureau (Instituto Brasileiro de Geografia e Estatística - IBGE). Following Gonzaga et al. (2006), I measure import penetration as imports divided by the sum of imports and domestic production. Ferreira et al. (2007) implement a similar pass-through adjustment using import penetration data from Muendler (2003b), which is calculated using a slightly different formula. The results presented here have also been generated using these alternative import penetration adjustments without any substantive differences. Since Brazil does not calculate a producer price index (Muendler 2003a), I use the wholesale price index, IPA-OG maintained by Fundação Getúlio 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.

Wage data come from the long form Brazilian Demographic Censuses (Censo Demográfico) for 1991 and 2000 from IBGE. In both 1991 and 2000, the long form was applied to a 10% sample of households in municipalities whose estimated population exceeded 15,000 and a 20% sample in smaller municipalities (IBGE 2002). The survey is nationally representative and yielded sample sizes of approximately 4 million households consisting of 17 million individuals in 1991 and 5.3 million households consisting of 20.3 million individuals in 2000. The wage analysis presented in Section 6 uses the microregion as the geographic unit of observation. Each of 558 microregions



is a grouping of economically integrated municipalities with similar geographic and productive characteristics (IBGE 2002). Wages are calculated as monthly earnings at the individual's main job divided by 4.33 times weekly hours at that job. The Census also reports employment status and industry of employment, which permits the calculation of the industrial distribution of labor in each microregion. While it would be ideal to have wage and employment information in 1987, just prior to liberalization, the wage analysis uses the 1991 Census as the baseline period under the assumption that wages and employment shares adjusted slowly to the trade liberalization.

Migration data come from the Pesquisa Nacional por Amostra de Domicílios (PNAD), a survey of Brazilian households conducted by IBGE. The survey has been conducted yearly since 1976 except census years (1980, 1991, 2000) and 1994. The survey is nationally representative, with the exception of the rural Northern region, corresponding to the Amazon rainforest. Since the survey is not representative of the entire Northern region, which accounted for only 6.8% of the national population in 1991, I omit it from the empirical analysis. Figure 10 shows the states included in the migration analysis. Note that I combine Tocantins and the Distrito Federal into the state of Goiás in order to maintain consistent state classifications over time.<sup>13</sup> The PNAD sample size is approximately 100,000 households including roughly 300,000 individuals, covering about 0.2% of the population. The survey includes information on employment status and industry of employment, which permits the calculation of the industrial distribution of labor in each state. Migration data are available in the core survey from 1992 to the present. Questions include the current and previous state of residence and the years since the last interstate migration, topcoded at 10 years. Given that migration questions in the PNAD describe geography at the state level, I define "migration" as moving from one state to another.

In both the wage and migration analyses, I restrict the sample to individuals aged 18-55 in order to focus on people who are most likely to be tied to the labor force. In the migration analysis presented in Section 7, I also generate results that further restrict the sample based on employment and family status in an effort to abstract from issues of tied movers and family size. In order to utilize these disparate data sets in the analysis, it was necessary to construct a common industry

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<sup>13</sup>Tocantins split from Goiás in 1988.

classification that was consistent across data sources. The classification is based upon a crosswalk between the national accounts and PNAD industrial codes published by the IBGE (2004). The final industry classification consists of 21 industries, including agricultural and nontraded goods, shown in Table 1.

## 5 Trade Liberalization in Brazil

Brazil's large, quickly implemented, and well-documented trade liberalization in the early 1990's provides an excellent context in which to study the effects of trade policy changes on other economic outcomes. Brazil's liberalization generated substantial variation in tariff changes across industries by moving from a tariff regime with high tariff levels and high cross-industry tariff dispersion to a low level, low dispersion tariff regime. Qualitative and quantitative evidence supports the exogeneity of cross-industry variation in tariff changes to counterfactual industry performance, allowing causal interpretations of the subsequent empirical results using this variation.

### 5.1 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, import bans on certain products, yearly maximum import levels per firm, assorted surcharges, prior authorization for imports of certain goods, and restricted credit for the purchase of imports (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 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 Política Aduaneira) proposal of a sharp tariff reduction and the removal of many non-tariff barriers.<sup>14</sup> 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.

Figures 2 and 3 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 tariff levels, the dispersion in tariffs was also greatly reduced during liberalization, consistent with the goal of aligning domestic production incentives with world prices. Before liberalization, effective rates of protection were higher than nominal tariffs because of a graduated tariff structure that imposed higher tariffs on final goods than on imported intermediates. As the dispersion in the tariff structure fell during liberalization, the graduated structure was eliminated and effective rates of protection fell to approximately the same level as nominal tariffs.

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<sup>14</sup>See Kume et al. (2003) for a detailed account of Brazil's liberalization, from which this paragraph is drawn.

It is clear in the figures that the move from a high-level, high-dispersion tariff structure to a low-level low-dispersion tariff distribution generated substantial variation in tariff changes across industries; industries with initially high tariffs experienced the largest tariff cuts, while those with initially lower tariff levels experienced smaller cuts. These large differences in tariff cuts across industries provide the identifying variation in the empirical analysis below and make Brazil an ideal context in which to study the differential impact of liberalization across regions with varying industrial distributions.

## 5.2 Exogeneity of Tariff Changes to Industry Performance

The empirical analysis below utilizes variation in tariff changes across industries. Figure 4 shows that industries facing larger tariff cuts shrank in terms of total workers employed, while industries facing smaller tariff cuts expanded their employment (The “tariff-induced price change,” calculated based on (9) is described in detail in the next section).<sup>15</sup> Interpreted causally, this result implies that the cross-industry variation in tariff cuts generated changes in the national industry mix that may have induced workers to move from regions with many shrinking industries to regions with many growing industries. However, in order to make this causal claim, it is essential that the tariff changes were not correlated with counterfactual industry performance in the absence of liberalization. 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.

There are a number of reasons to believe that these general concerns were not realized in the specific case of Brazil’s trade liberalization. As mentioned above, 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

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<sup>15</sup>A figure similar to Figure 4, appearing in Ferreira et al. (2007), provided the initial motivation for undertaking the present study.

(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. This pattern is apparent in Figure 6. 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  $-0.9$ .<sup>16</sup> The pre-liberalization tariff regime was based upon a tariff schedule developed in 1957 (Kume et al. 2003). Since the structure of the liberalization 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.

Finally, one can gain insight into the exogeneity of tariff changes by observing their relationship to industry growth. This relationship is demonstrated in Figure 4. 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<sup>17</sup>. 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 should be a correlation between pre-liberalization industry employment growth and subsequent tariff changes. As shown in Figure 5, 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.

<sup>16</sup>The results for nominal tariffs are essentially identical, with a correlation of  $-0.95$ .

<sup>17</sup>This interpretation is somewhat implausible, since the observed pattern of tariff cuts were precisely the opposite of what one would expect if policy makers were trying to protect declining industries. The observed pattern would imply that policy makers cut tariffs most on declining industries that were most in need of protection

## 6 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 tariff changes on wages as predicted by the model in (10). 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. The results strongly confirm the model's prediction, implying that regions facing a 10% larger tariff decline experience 6.3%-7.6% larger wage declines. This finding is consistent with some equalizing interregional migration, motivating the subsequent migration analysis.

### 6.1 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 wage change in a given region could be due changes in individual characteristics, changes in the returns to those characteristics, or changes in regional labor demand due to liberalization. In order to isolate the last effect, 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. The results of these regressions are reported in Table 2. I then calculate the regional wage change as the change in microregion fixed effects, plus a term reflecting the change in wages for an average 1991 individual. The addition of this average wage change term is purely for interpretation, as it does not vary across regions. It means that the regional wage change is interpreted as the proportional wage change an average 1991 individual would expect to face living in each microregion.

Figure 7 shows the 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 and more yellow experienced the largest wage declines during the 1991-2000 time period, while darker and bluer regions experienced the largest wage increases. As the scale indicates, some observations are quite large in magnitude. Happily, only 8 observations fall outside the  $\pm 0.3$  range, and these

are all in sparsely populated areas, leading to imprecise estimates.

## 6.2 Tariff-Induced Price Changes

The discussion of industry aggregation in Section 3.1 suggests that tariff changes will have a larger impact on prices in industries where many of the goods comprising the industry face import competition. In particular, (9) suggests multiplying the tariff changes by the fraction of industry workers producing import-competing goods. This fraction is unknown, but we can proxy for it with industry import penetration,  $\gamma_i$ , calculated as imports divided by the sum of imports and domestic production.<sup>18</sup> Although I expect this measure to substantially understate the level of import competition in a given industry (i.e.  $\gamma_i < \phi_i$ ), it is likely to capture the relative degree of import competition across industries. As a proxy for world prices in (9), I use U.S. prices. Using these proxies and allowing for random measurement error in prices  $u_i$ , the proportional change in the Brazilian price level  $\pi$ , and the proportional change in the Real-dollar exchange rate  $S$ , equation (9) becomes

$$\hat{P}_i = \pi + \gamma_i(1 + \hat{\tau}_i) + S + \hat{P}_i^{US} + u_i. \quad (13)$$

This relationship is estimated as

$$d \ln(P_i) = \psi_0 + \psi_1 \gamma_i d \ln(1 + \tau_i) + \psi_2 d \ln(P_i^{US}) + u_i, \quad (14)$$

where  $d$  represents the long difference between 1997 and 1995,  $\psi_0$  captures the effect of inflation and exchange rate changes, and  $\psi_1$  is likely to be substantially larger than one given that import penetration understates the level of import price competition in each industry.

The results of estimating (14) are shown in Table 3. Columns (1) and (2) omit the tariff pass through term and find no relationship between tariff changes and price changes. This result is consistent with the findings of Gonzaga et al. (2006), and demonstrates the importance of the import penetration adjustment in capturing variable tariff pass through across industries. Columns (3) and (4) include the import penetration adjustment, finding a positive and statistically significant

<sup>18</sup>Alternative measures of import penetration have been used as well with no qualitative changes in results.

relationship between price changes and tariff changes. The estimate's large size suggests that import penetration does underestimate the scale of import competition, as expected. Letting hats represent estimates (rather than proportional changes as in the theory section), The tariff-induced price change is calculated as

$$d \ln(\hat{P}_i) = \hat{\psi}_1 \gamma_i d \ln(1 + \tau_i). \quad (15)$$

By omitting  $\hat{\psi}_0$  from this expression, tariff-induced price changes are calculated relative to changes in the overall price level. Figure 8 shows the tariff-induced price changes resulting from this calculation. Since these measures are normalized relative to the overall price level, they may be positive or negative in individual industries even though all tariffs were cut. This reflects the inherently cross-sectional nature of the empirical exercise. The goal is to measure the different effects of tariff changes on prices across industries rather than the overall effect of the liberalization on the price level.

### 6.3 Region-Level Tariff Changes

Based on (10), the effect of a given set of tariff changes on a region's wages is determined by a weighted average of tariff-induced price changes. In what follows, I call this weighted average the "region-level tariff change." Calculating the  $\beta_{ri}$  terms in (11) requires information on the allocation of labor across industries and on labor demand elasticities. The industrial allocation of labor is calculated for each microregion from the 1991 Census. As mentioned above, it is not feasible to calculate elasticities of factor substitution across regions and industries, so I restrict production to be Cobb-Douglas. This implies that  $\sigma_{ri} = 1$  and  $\theta_{ri} = \theta_i$ , which is calculated 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.

$$RTC_r = \sum_{i \neq N} \beta_{ri} d \ln(\hat{P}_i) \quad (16)$$

$$\text{where } \beta_{ri} = \frac{L_{ri} \frac{1}{\theta_{ri}}}{\sum_{i' \neq N} L_{ri'} \frac{1}{\theta_{ri'}}}. \quad (17)$$



The results of this calculation appear in Figure 9. Lighter and more yellow microregions faced the most negative tariff-induced price changes, while darker and bluer microregions faced more positive price changes. Recall that the tariff-induced price changes are calculated relative to the overall price level, so although all tariffs were cut, they may be positive or negative. This normalization is reflected here in the region-level tariff changes as well.

#### 6.4 Wage-Tariff Relationship

Given the empirical estimates of the regional wage changes and region-level tariff changes, it is now possible to examine the effect of tariff changes on regional wages predicted by the specific-factors model. I form an estimating equation from (10) as

$$d\ln(w_r) = \zeta_0 + \zeta_1 RTC_r + \epsilon_r, \quad (18)$$

where  $d\ln(w_r)$  is calculated as described in Section 6.1. Since these wage changes are estimates, I weight the regression by the inverse of the standard error of the estimates.  $RTC_r$  is given by (16).  $\zeta_0$  captures the increase in average real wages between 1991 and 2000. In the model without migration, the 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 arbitrated 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  $\epsilon_r$  captures any drivers of wage change that are unrelated to liberalization. In case of changes in state policies that may have influenced wages similarly across microregions within the state, I will also include state-level fixed effects.<sup>19</sup>

Table 4 presents the results of regressing the regional wage changes on the region-level tariff

<sup>19</sup>State-specific minimum wages were not implemented until 2002, so this does not confound the analysis.

changes. As expected, the effect of region-level tariffs on regional wages is positive and statistically significant. This implies that microregions facing the largest tariff declines, as predicted by the model, did experience slower wage growth than regions facing smaller tariff cuts. The point estimates for  $\zeta_1$  are both less than 1, indicating the presence of some equalizing interregional migration.<sup>20</sup> The following section will examine migration patterns directly, corroborating this finding. The addition of state fixed effects lowers the magnitude of the point estimate somewhat, but remains qualitatively similar. In case of remaining covariance in the error term across microregions in a given state beyond a common additive component captured by the state fixed effects, I report standard errors clustered by state. This reduces the significance in column (1), but leaves the fixed effect specification essentially unchanged. Recall that these results are interpreted cross-sectionally - they do not measure the effect of liberalization on national wage growth or contraction, but rather describe the different effects of liberalization across regions of the liberalizing country. Thus, the estimate in column (2) implies that a region facing a 10% larger tariff decline will experience a 6.3% larger wage decline relative to other regions.

These results confirm the model's prediction, particularly in finding an estimate of the expected sign that is significantly different from zero, but below one. This supports the assumption that cross-region differences in the effects of liberalization are correctly measured and can be applied to other labor market outcomes of interest. The next section does this by examining the effects of liberalization on inter-state migration. The wage results also have implications for policy makers considering undertaking a large trade policy change, as they imply a clear link between trade policy decisions at the national level and local policy challenges. Given the predictions of the model, national policy makers could use the pre-liberalization distribution of labor across industries in different regions to determine what regions' workers are most likely to be negatively impacted by a proposed trade policy change. They can then coordinate with local policy makers to respond to the expected local impacts of the national policy change.

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<sup>20</sup>Although the estimate in column (1) is not statistically different from 1 at the 5% level

## 7 The Effect of Liberalization on Interstate Migration

The preceding section showed that trade liberalization caused substantial variation in wage changes across Brazilian microregions, and suggested that workers responded by migrating away from locations facing the most negative wage changes to locations facing the most positive changes. This section directly measures the impact of liberalization on migration patterns utilizing detailed survey data on interstate migration from Brazil's yearly household survey. The results show that migration patterns changed as a result of liberalization, with more individuals moving away from states facing the largest tariff cuts and toward states facing smaller cuts. Counterfactual simulations imply that the most affected Brazilian states gained or lost approximately 2% of their populations as a result of liberalization-induced shifts in migration patterns.

### 7.1 Location Choice Specification

This section derives a framework for estimating the effect of tariff changes on individuals' location choice from a model of individual maximizing behavior. Although wages are an important aspect of location choice, other considerations such as local amenities, proximity to friends and relatives, and costs of moving to a particular location will also be relevant. These various aspects of location choice can be captured in the following additive random utility model.

$$U_{igt} = V_{gt} + \epsilon_{igt} \tag{19}$$

$$V_{gt} \equiv \alpha_g \ln w_{gt} + \mu_{gt} + \eta_{gd} \tag{20}$$

$U_{igt}$  is the utility that individual  $i$  in group  $g$  (described below) receives from living in destination state  $d$  at time  $t$ .  $V_{gt}$  represents the average utility individuals in group  $g$  receive from living in location  $d$  at time  $t$ , while  $\epsilon_{igt}$  represents individual idiosyncratic deviations from the average. The average utility in a given destination depends upon wages,  $w$ , and unobservable characteristics of the destination, some of which vary over time,  $\mu_{gt}$ , and some of which are fixed over time,  $\eta_{gd}$ .

The "group" subscript,  $g$ , determines how the unobservable terms in (20) vary. In this analysis,

groups are always at least based upon state of residence, and potentially upon other demographic characteristics such as age or gender. Grouping by state of residence implies that the unobserved terms,  $\mu_{gdt}$  and  $\eta_{gd}$ , vary by state pairs. Any moving costs associated with the distance between two states are therefore subsumed in these unobserved effects. Location-specific amenities such as natural beauty or urban nightlife will similarly be captured by these terms. Now suppose that groups are defined by state of residence and by age. This allows the value of these location-specific amenities to vary across age groups. 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 the average preferences of one's group, is captured in the error term  $\epsilon_{idgt}$ . By careful group definition, the model can capture many rich and complex considerations that are relevant to location choice. The parameter of interest is  $\alpha_g$ , the importance of wages in location decisions. Note that this parameter may also be assumed to vary across groups, as indicated by the group subscript. The empirical results presented below include specifications in which  $\alpha$  is assumed constant across groups and others in which  $\alpha_g$  may vary across groups.

Individuals compare all states and choose to live in the state that maximizes utility. Assuming that the  $\epsilon_{idgt}$  are independently drawn from a Type I extreme value distribution, the probability  $\pi_{gdt}$  that an individual in group  $g$  chooses location  $d$  at time  $t$  is

$$\pi_{gdt} = \frac{e^{V_{gdt}}}{D_{gt}} \quad \text{where} \quad D_{gt} \equiv \sum_{d'} e^{V_{gd't}}. \quad (21)$$

In the absence of the unobservable  $\eta_{gd}$  and  $\mu_{gd}$  terms in  $V_{gdt}$ , 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  $\alpha_g$ . Thus, an alternative approach is necessary. 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 through the use of a first-order Taylor series approxima-

tion. This process, implemented in Appendix C, yields the following equation.

$$d \ln S_{gd} - d \ln S_{gh} \approx \alpha_g (d \ln w_d - d \ln w_h) + \left[ (d\mu_{gd} - d\mu_{gh}) + d \left( \frac{\xi_{gd}}{\pi_{gd}} \right) - d \left( \frac{\xi_{gh}}{\pi_{gh}} \right) \right] \quad (22)$$

Before describing the notation, replace the wage change terms with liberalization's effect on regional wages, the region-level tariff change ( $RTC$ ), calculated in (16).

$$d \ln S_{gd} - d \ln S_{gh} \approx \alpha_g (RTC_d - RTC_h) + \left[ (d\mu_{gd} - d\mu_{gh}) + d \left( \frac{\xi_{gd}}{\pi_{gd}} \right) - d \left( \frac{\xi_{gh}}{\pi_{gh}} \right) \right] \quad (23)$$

For simplicity, assume for the moment that  $g$  represents only state of residence, without any distinctions between demographic groups.  $S_{gd}$  is the observed share of individuals from state  $g$  choosing to locate in destination state  $d$ . The subscript  $h$  represents the current state of residence, or “home,” so  $S_{gh}$  is the share of people from state  $g$  choosing to stay there rather than relocate. Thus the left hand side of (23) is the change in the share of individuals from  $g$  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,  $\eta_{gd}$ . The independent variable of interest is the liberalization induced wage change in destination  $d$ , again relative to the same expression at home. Having an estimate of the coefficient on this term,  $\alpha_g$ , makes it possible to run counterfactual simulations describing how individuals would have moved under different circumstances. I do this below to measure the impact of liberalization on the distribution of population across Brazilian states. The term in brackets represents the error term, consisting of two parts. The first is the difference in time varying unobservable amenities. The presence of this expression in the error term makes clear the additional identification assumption necessary to estimate (23) 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.  $\xi_{gd}$  is random sampling error in measuring  $S_{gd}$ , generating heteroskedasticity. I therefore weight by the square-root of the number of observations used to calculate  $S_{gd}$ .

## 7.2 Location Choice Results

I calculate region-level tariff changes by state (rather than by microregion as in the wage analysis) in the same manner as described in Section 6.3, the only difference being that employment shares were calculated using the 1987 PNAD rather than the Census. Figure 10 shows the results. The left hand side of (23) is calculated using migration data from the PNAD. Table 5 presents summary statistics regarding inter-state migration in Brazil among different demographic groups. The first column presents the fraction of the total population in each demographic group, while subsequent columns describe the fraction of individuals in each demographic group reporting different migration behaviors. Inter-state 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, inter-state migration in Brazil is much more common than in India. Topalova (2005) reports that only 3-4% of people migrated between Indian districts within a ten-year time period, whereas 9.7% of Brazilians report moving between states during a ten year period. Districts in India are very small compared to Brazilian states (on average each Indian state consists of 16 districts), so the difference in mobility is particularly striking.

The analysis compares individuals' location decisions just preceding trade liberalization (September 1982 - September 1988) to those just after liberalization (September 1996 - September 2002). The final two columns of Table 5 present the fraction of each demographic group that migrated in each of these periods. A number of patterns emerge. Consistent with the early observations of Sjaastad (1962) and nearly every subsequent study of migration, younger individuals are more likely to move. More educated individuals are more mobile, although the effect is not monotone over years of schooling, and those with larger families are far less mobile than individuals or couples. Whites and those of mixed heritage (reporting *Pardo*) are much more mobile than Blacks. Contrary to expectations, married people generally report more mobility than unmarried people, although the sample fractions are nearly equal in the post-migration period. 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 baseline results of estimating (23) are presented in Table 6. In the first row, grouping is by state of residence (source state) only. Thus, each observation represents a source-destination state pair. Since the equation for the share of individuals choosing to stay in the same state has been differenced from each observation, and there are 19 states included in the analysis, the total number of potential observations is  $19 * 18 = 342$ . The analysis drops any state pairs in which the share term,  $S_{gd}$ , was estimated using less than five underlying observations, so the realized number of observations is 168 rather than 342.<sup>21</sup> The estimate of  $\alpha$  in the first row of Table 6 is 1.92. In order to assess the scale of this estimate, note that the estimating equation admits a convenient reduced-form interpretation that can be obtained by differentiating (21) with respect to  $\ln w_{dt}$  for all  $d$ .

$$d\pi_{sd} = \alpha\pi_{sdt} \left( (1 - \pi_{sdt})d \ln w_d - \sum_{d' \neq d} \pi_{sd't} d \ln w_{d'} \right) \quad (24)$$

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  $\alpha$ , the observed pre-liberalization migration fractions, and the tariff-induced wage changes given by (16), 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 state population in 1988 and summing over all sources for a given destination, it is possible to calculate the number of people accounted for by liberalization-induced shifts in the interstate migration pattern. The results of this exercise are shown in Table 7. The first column reports the number of people in each state that are accounted for by liberalization-induced shifts in migration patterns and the final column reports the same number as a fraction of the state's 1988 population. For those states facing the largest and smallest tariff cuts, liberalization accounts for gains or losses of approximately 2% 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 remaining rows in Table 6 differ from the specification in the first row in that grouping

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<sup>21</sup>Since (23) requires taking logs of  $S_{gd}$ , group-destination bins containing zero observations, i.e. when no one in a particular group chooses a given destination, *must* be dropped. Although cells generated with 1-4 observations are technically usable, they are omitted in order to avoid wildly inaccurate estimates. A more stringent rule, dropping observations based on less than 10 underlying observations yield similar results.

is based on state of residence (source) and on demographic characteristics. Given that group-destination pairs containing less than five underlying observations are dropped from the analysis, each demographic characteristic is separated into only two bins in order to avoid creating such fine grouping classifications that many group-destination pairs are dropped. When source state and demographic groups are considered, the number of potential observations is  $19 \times 18 \times 2 = 684$ . Although grouping by demographics increases the potential number of observations, the number of clusters (19) remains constant across all specifications, so demographic grouping does not inappropriately increase statistical power by “inventing” more observations. When grouping by demographic characteristics, two different specifications are considered. The first, labeled “homogeneous effect across groups,” restricts the estimate of  $\alpha$  to be constant across demographic groups, but does not place any restriction on the unobserved effects across demographic groups. Different age groups can value unobserved amenities differently even though  $\alpha$  is constant across groups.

The second specification, labeled “heterogeneous effects across groups,” allows  $\alpha_g$  to vary across demographic groups, along with accounting for differences in unobserved effects across groups. In these specifications, it is expected that younger and more mobile individuals and those who are more connected to the labor market will exhibit stronger effects of location choice on tariff-induced wage changes, since these individuals have more to gain in expectation from choosing a new location. The results for age, gender, and family size all demonstrate the expected pattern - the more mobile group exhibits a stronger relationship (in statistical and economic terms) between tariff-induced wage changes and location choice. Note that the point estimates for the mobile groups are in a few cases much larger than the estimate from the first row considered above, indicating substantially larger liberalization-induced migration responses for these demographic groups. The results for education in Table 6 are more surprising. Although those with fewer years of education are less mobile in general (see Table 5), less educated individuals exhibit a very strong location response to liberalization. The result may indicate that labor markets are segmented between high-skilled and lower-skilled workers, and that employers adjust to tariff changes primarily through changes in lower-skilled labor demand. This is an area for further study in a framework that accounts for worker heterogeneity in production.



These findings provide strong evidence that the disparate effects of trade liberalization on labor market conditions across Brazilian states led individuals to alter their location choices, moving away from states facing the largest tariff-induced price declines and toward states facing smaller cuts. The results also demonstrate the importance of accounting for variation in unobserved components of utility across demographic groups, and the fact that groups that are more mobile and more connected to labor market outcomes are most influenced by the geographic variation in the returns to work. As in the wage analysis, these results have important policy implications in linking trade policy decisions at the national level to local policy challenges. If a country's regions have different industrial compositions, then the adjustment to a large change in trade policy will necessarily involve some movement of workers from regions with many contracting industries toward regions with many growing industries. The results presented here show that the specific factors model of regional economies provides a means of predicting the pattern of interregional migration resulting from liberalization. Given this information, national policy makers can work at the local level to help individuals make the geographic transitions that necessarily come with a large industrial reorganization.

## 8 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 interstate 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 7% larger wage decline. Liberalization also caused a substantial shift in migration patterns. The most affected Brazilian states gained or lost approximately 2% of their populations as a result of liberalization-induced shifts in migration patterns.

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. (2009) 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. Particularly mobile groups of individuals will be able to smooth out regional wage variation by migrating while less mobile individuals will not. 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.

This paper's findings have important implications in linking national policy changes, such as liberalization, to local policy challenges involving migration, transportation, and housing, as individuals migrate to restore geographic equilibrium. National policy makers can use the specific-factors model's predictions to assess what areas are likely to experience an influx of migrants hoping to gain employment in an area with many expanding industries 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 play an important role when pursuing a change in national policy that requires substantial industrial reallocation.

## A Specific Factors Model Solution

### 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. Consider a particular region,  $r$ , suppressing that subscript on all terms. Industries are indexed by  $i = 1 \dots N$ .  $L$  is the total amount of labor and  $T_i$  is the amount of industry-specific factor for industry  $i$  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

$$a_{Ti}Y_i = T_i \quad \forall i, \quad (\text{A1})$$

$$\sum_i a_{Li}Y_i = L. \quad (\text{A2})$$

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

$$a_{Li}w + a_{Ti}R_i = P_i \quad \forall i \quad (\text{A3})$$

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

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

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

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

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

$$\hat{Y}_i = -\hat{a}_{Ti} \quad \forall i \quad (\text{A6})$$

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

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

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

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

Substituting this into (A7) yields

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

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

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

Rewrite this expression as follows for convenience of notation.

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

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

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

Note that the inverse of the diagonal matrix  $\Theta$  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:

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

$$\text{where } \beta_i = \frac{\lambda_i \frac{\sigma_i}{\theta_i}}{\sum_{i'} \lambda_{i'} \frac{\sigma_{i'}}{\theta_{i'}}} \quad (\text{A14})$$

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

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

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

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

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

## A.2 Nontraded goods prices

As in the previous section, consider a particular region, omitting the  $r$  subscript on all terms. Industries are indexed by  $i = 1 \dots N$ . The final industry, indexed  $N$ , is nontraded, while other industries ( $i < N$ ) 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 CES preferences over goods from each industry. This implies the following goods demands.

$$Y_i^c = \left( \frac{\alpha_i}{P_i} \right)^\sigma \frac{m}{\sum_j \alpha_j^\sigma P_j^{1-\sigma}}, \quad (\text{A17})$$

where  $Y_i^c$  is consumer demand,  $m$  is total consumer income,  $\alpha_i$  is the CES share parameter,  $\sigma$  is the elasticity of substitution in consumption (not to be confused with  $\sigma_i$ , which is the elasticity of substitution between factors of production), and  $P_i$  is the good's price. To simplify future expressions, define  $\bar{P}$  as the CES price index,

$$\bar{P} \equiv \left( \sum_i \alpha_i^\sigma P_i^{1-\sigma} \right)^{\frac{1}{1-\sigma}}. \quad (\text{A18})$$

Substituting this into (A17) and calculating the proportional change in  $Y_i^c$  yields

$$\hat{Y}_i^c = \hat{m} - \sigma \hat{P}_i + (\sigma - 1) \hat{P}, \quad (\text{A19})$$

where hats represent proportional changes. The goal of the remaining steps is to express the terms of (A19) in terms of price changes and changes in labor.

**Change in the Price Level.** Given the definition of  $\bar{P}$ ,

$$\hat{P} = \sum_i \mu_i \hat{P}_i \quad (\text{A20})$$

$$\text{where } \mu_i \equiv \frac{\alpha_i^\sigma P_i^{1-\sigma}}{\sum_j \alpha_j^\sigma P_j^{1-\sigma}}. \quad (\text{A21})$$

**Change in Consumer Income.** Consumer income equals total factor payments, so

$$\hat{m} = \eta_L (\hat{w} + \hat{L}) + \sum_i \eta_i \hat{R}_i, \quad (\text{A22})$$

where  $\eta_L$  and  $\eta_i$  are, respectively, the share of labor earnings and industry  $i$  specific factor earnings in total income. Substituting (A13) and (A16) into (A22) and collecting terms yields

$$\begin{aligned} \hat{m} &= \underbrace{\sum_i \eta_L \beta_i \hat{P}_i + \sum_j \eta_j \beta_j \hat{P}_j + \sum_j \eta_j \sum_{k \neq j} \beta_k \hat{P}_k}_{\mathcal{X}} + \underbrace{\sum_j \frac{\eta_j}{\theta_j} (1 - \beta_j) \hat{P}_j - \sum_j \frac{\eta_j}{\theta_j} \sum_{k \neq j} \beta_k \hat{P}_k}_{\mathcal{Y}} \quad (\text{A23}) \\ &\quad + \underbrace{\left( \eta_L \left( \sum_i \lambda_i \frac{\sigma_i}{\theta_i} - 1 \right) + \sum_i \eta_i \frac{(1 - \theta_i)}{\theta_i} \right) \frac{\hat{L}}{\sum_i \lambda_i \frac{\sigma_i}{\theta_i}}}_{\mathcal{Z}} \end{aligned}$$

Examining the group of terms labeled  $\mathcal{X}$ ,

$$\mathcal{X} = \sum_i \eta_L \beta_i \hat{P}_i + \sum_i \eta_i \left( \beta_i \hat{P}_i + \sum_{k \neq i} \beta_k \hat{P}_k \right) \quad (\text{A24})$$

$$= \sum_i \eta_L \beta_i \hat{P}_i + \sum_i \eta_i \sum_j \beta_j \hat{P}_j \quad (\text{A25})$$

$$= \left( \eta_L + \sum_i \eta_i \right) \sum_j \beta_j \hat{P}_j \quad (\text{A26})$$

$$= \sum_i \beta_i \hat{P}_i \quad (\text{A27})$$

where the final equality follows from noting that  $\eta_L + \sum_j \eta_j = 1$  by construction. Now examining the group of terms labeled  $\mathcal{Y}$ , first note that  $\frac{\eta_i}{\theta_i} = \frac{P_i Y_i}{\sum_j P_j Y_j}$ , which is industry  $i$ 's share of total production value; call

this share  $\varphi_i$ .

$$\mathcal{Y} = \sum_j \varphi_j \hat{P}_j - \sum_j \varphi_j \beta_j \hat{P}_j - \sum_j \varphi_j \sum_{k \neq j} \beta_k \hat{P}_k \quad (\text{A28})$$

$$= \sum_j \varphi_j \hat{P}_j - \sum_j \varphi_j \sum_k \beta_k \hat{P}_k \quad (\text{A29})$$

$$= \sum_j \varphi_j \hat{P}_j - \sum_k \beta_k \hat{P}_k, \quad (\text{A30})$$

where the final equality comes from the fact that  $\sum_i \varphi_i = 1$ . Finally, examine the group of terms labeled  $\mathcal{Z}$ . Note that

$$\sum_i \eta_i \frac{(1 - \theta_i)}{\theta_i} = \sum_i \frac{R_i T_i}{m} \frac{w L_i}{P_i Y_i} \frac{P_i Y_i}{R_i T_i} = \frac{w}{m} \sum_i L_i = \eta_L. \quad (\text{A31})$$

Plugging this into the expression for  $\mathcal{Z}$ ,

$$\mathcal{Z} = \left( \eta_L \left( \sum_i \lambda_i \frac{\sigma_i}{\theta_i} - 1 \right) + \eta_L \right) \frac{\hat{L}}{\sum_i \lambda_i \frac{\sigma_i}{\theta_i}} \quad (\text{A32})$$

$$= \eta_L \hat{L} \quad (\text{A33})$$

Combining these results implies

$$\hat{m} = \eta_L \hat{L} + \sum_i \varphi_i \hat{P}_i \quad (\text{A34})$$

**Change in Nontraded Good Production.** For the nontraded good, regional production equals consumption, so  $\hat{Y}_N^c = \hat{Y}_N^p$ . Substitutions using (A4), (A5), (A6), and (A8) yield the following expression for the change in nontraded good output.

$$\hat{Y}_N^p = \frac{(1 - \theta_N)}{\theta_N} \sigma_N (\hat{P}_N - \hat{w}) \quad (\text{A35})$$

$$= \frac{(1 - \theta_N)}{\theta_N} \sigma_N \left( \hat{P}_N + \frac{\hat{L}}{\sum_i \lambda_i \frac{\sigma_i}{\theta_i}} - \sum_i \beta_i \hat{P}_i \right) \quad (\text{A36})$$

**Combining Terms.** Plugging (A20), (A34), and (A36) into (A19) for the nontraded industry  $N$  yields

$$\frac{(1 - \theta_N)}{\theta_N} \sigma_N \left( \hat{P}_N + \frac{\hat{L}}{\sum_i \lambda_i \frac{\sigma_i}{\theta_i}} - \sum_i \beta_i \hat{P}_i \right) = \eta_L \hat{L} + \sum_i \varphi_i \hat{P}_i - \sigma \hat{P}_N + (\sigma - 1) \sum_i \mu_i \hat{P}_i. \quad (\text{A37})$$

Isolate and collect terms including  $\hat{P}_N$

$$\begin{aligned} \left[ \frac{(1 - \theta_N)}{\theta_N} \sigma_N (1 - \beta_N) - \varphi_N + \sigma - (\sigma - 1) \mu_N \right] \hat{P}_N &= \left[ \eta_L - \frac{(1 - \theta_N)}{\theta_N} \frac{\sigma_N}{\sum_i \lambda_i \frac{\sigma_i}{\theta_i}} \right] \hat{L} \\ &+ \sum_{i \neq N} \left[ \frac{(1 - \theta_N)}{\theta_N} \sigma_N \beta_i + \varphi_i + (\sigma - 1) \mu_i \right] \hat{P}_i \end{aligned} \quad (\text{A38})$$

Grouping terms on the left hand side and solving for  $\hat{P}_N$ ,

$$\hat{P}_N = \frac{\eta_L - \frac{(1-\theta_N)}{\theta_N} \frac{\sigma_N}{\sum_i \lambda_i \frac{\sigma_i}{\theta_i}}}{\sum_{i' \neq N} \frac{(1-\theta_N)}{\theta_N} \sigma_N \beta_{i'} + \varphi_{i'} + (\sigma - 1) \mu_{i'}} \hat{L} + \sum_{i \neq N} \xi_i \hat{P}_i \quad (\text{A39})$$

$$\text{where } \xi_i = \frac{\frac{(1-\theta_N)}{\theta_N} \sigma_N \beta_i + \varphi_i + (\sigma - 1) \mu_i}{\sum_{i' \neq N} \frac{(1-\theta_N)}{\theta_N} \sigma_N \beta_{i'} + \varphi_{i'} + (\sigma - 1) \mu_{i'}} \quad (\text{A40})$$

### A.3 Restrictions to Drop the Nontraded Sector from Weighted Averages

Under Cobb-Douglas production with equal factor shares across industries ( $\theta_i = \theta \forall i$ ), the first order conditions imply that, for all  $i$

$$P_i(1 - \theta) \frac{Y_i}{L_i} = w \quad (\text{A41})$$

$$\varphi_i(1 - \theta) = \eta_L \lambda_i \quad (\text{A42})$$

$$\varphi_i = \left( \frac{\theta}{(1 - \theta)} \eta_L \sum_{i'} \lambda_{i'} \frac{\sigma_{i'}}{\theta_{i'}} \right) \beta_i \quad (\text{A43})$$

$$\varphi_i = \kappa \beta_i, \quad (\text{A44})$$

where the final equality comes from defining  $\kappa$  as the coefficient on  $\beta_i$ , which does not vary across industries. Restrict consumer preferences to be Cobb-Douglas ( $\sigma = 1$ ). Under this restriction, and plugging in (A44),  $\xi_i$  is

$$\xi_i = \frac{\left( \frac{(1-\theta_N)}{\theta_N} + \kappa \right) \beta_i}{\sum_{i' \neq N} \left( \frac{(1-\theta_N)}{\theta_N} + \kappa \right) \beta_{i'}} \quad (\text{A45})$$

$$= \frac{\beta_i}{\sum_{i' \neq N} \beta_{i'}} \quad (\text{A46})$$

Plug this result into (3) and (1)

$$\hat{w} = \sum_{i \neq N} \beta_i \hat{P}_i + \beta_N \left( \frac{\sum_{i \neq N} \beta_i \hat{P}_i}{\sum_{i \neq N} \beta_i} \right) \quad (\text{A47})$$

$$= \left( 1 + \frac{\beta_N}{1 - \beta_N} \right) \sum_{i \neq N} \beta_i \hat{P}_i \quad (\text{A48})$$

$$= \frac{\sum_{i \neq N} \beta_i \hat{P}_i}{\sum_{i' \neq N} \beta_{i'}} \quad (\text{A49})$$

This is equivalent to omitting the nontraded industry  $N$  from the sums in (1) and (2).

## B Industry Aggregation

Begin with equation (7).

$$\hat{P}_{ig} = \mathbf{1}(ipc_{ig})(1 + \hat{\tau}_i) + P_{ig}^W. \quad (\text{B1})$$

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

$$\hat{w}_r = \sum_i \sum_{g \in i} \beta_{rig} \mathbf{1}(ipc_{ig}) (1 + \hat{\tau}_i) + \hat{P}_{ig}^W \quad (\text{B2})$$

$$= \sum_i (1 + \hat{\tau}_i) \sum_{g \in i} \beta_{rig} \mathbf{1}(ipc_{ig}) + \sum_i \sum_{g \in i} \beta_{rig} \hat{P}_{ig}^W \quad (\text{B3})$$

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

$$\sum_{g \in i} \beta_{rig} \mathbf{1}(ipc_{ig}) = \frac{\frac{1}{\theta_i} \sum_{g \in i} L_{rig} \mathbf{1}(ipc_{ig})}{\sum_{i'} \frac{1}{\theta_{i'}} \sum_{g' \in i'} L_{ri'g'}} \quad (\text{B4})$$

$$= \frac{\frac{L_{ri} \sum_{g \in i} L_{rig} \mathbf{1}(ipc_{ig})}{L_{ri}}}{\sum_{i'} L_{ri'} \frac{1}{\theta_{i'}}} \quad (\text{B5})$$

$$= \beta_{ri} \phi_{ri} \quad (\text{B6})$$

$$\text{where } \phi_{ri} \equiv \sum_{g \in i} \frac{L_{rig}}{L_{ri}} \mathbf{1}(ipc_{ig}) \quad (\text{B7})$$

$\phi_{ri}$  is the fraction of industry  $i$  workers producing goods that face import competition. Now consider the second term in (B3).

$$\sum_{g \in i} \beta_{rig} \hat{P}_{ig}^W = \frac{\frac{1}{\theta_i} \sum_{g \in i} L_{rig} \hat{P}_{ig}^W}{\sum_{i'} \frac{1}{\theta_{i'}} \sum_{g' \in i'} L_{ri'g'}} \quad (\text{B8})$$

$$= \frac{\frac{L_{ri} \sum_{g \in i} L_{rig} \hat{P}_{ig}^W}{L_{ri}}}{\sum_{i'}} \quad (\text{B9})$$

$$= \beta_{ri} \hat{P}_i^W \quad (\text{B10})$$

$$\text{where } \hat{P}_i^W \equiv \sum_{g \in i} \frac{L_{rig}}{L_{ri}} \hat{P}_{ig}^W \quad (\text{B11})$$

$\hat{P}_i^W$  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 (B3), yields the result of the aggregation.

$$\hat{w}_r = \sum_i \beta_{ri} (\phi_{ri} (1 + \hat{\tau}_i) + \hat{P}_i^W) \quad (\text{B12})$$

## C 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 (21). The observed share of individuals in group  $g$  who choose to live in location  $d$  at time  $t$ ,  $S_{gdt}$ , will consist of the true choice probability,  $\pi_{gdt}$ , and mean zero random sampling error,  $\xi_{gdt}$ .

$$S_{gdt} = \frac{e^{V_{gdt}}}{D_{gt}} + \xi_{gdt} \quad (\text{C1})$$



Taking logs yields

$$\ln S_{gdt} = \ln(e^{V_{gdt}} + \xi_{gdt}D_{gt}) - \ln D_{gt}. \quad (\text{C2})$$

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

$$\ln S_{gdt} \approx V_{gdt} - \ln D_{gt} + \frac{\xi_{gdt}}{\pi_{gdt}}. \quad (\text{C3})$$

Plugging in the definition of  $V_{gdt}$  from (20),

$$\ln S_{gdt} \approx \alpha_g \ln w_{dt} + \mu_{gdt} + \eta_{gd} - \ln D_{gt} + \frac{\xi_{gdt}}{\pi_{gdt}}. \quad (\text{C4})$$

The model is still nonlinear in  $\alpha_g$ , due to its presence within  $D_{gt}$ . This term can be canceled by subtracting the log share of an arbitrary reference destination. For convenience, the reference state is  $h$ , the state of residence of individuals in group  $g$ .

$$\ln S_{gdt} - \ln S_{ght} \approx \alpha_g (\ln w_{dt} - \ln w_{ht}) + (\mu_{gdt} - \mu_{ght}) + (\eta_{gd} - \eta_{gh}) + \left( \frac{\xi_{gdt}}{\pi_{gdt}} - \frac{\xi_{ght}}{\pi_{ght}} \right) \quad (\text{C5})$$

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

$$d \ln S_{gd} - d \ln S_{gh} \approx \alpha_g (d \ln w_d - d \ln w_h) + \left[ (d\mu_{gd} - d\mu_{gh}) + d \left( \frac{\xi_{gd}}{\pi_{gd}} \right) - d \left( \frac{\xi_{gh}}{\pi_{gh}} \right) \right] \quad (\text{C6})$$

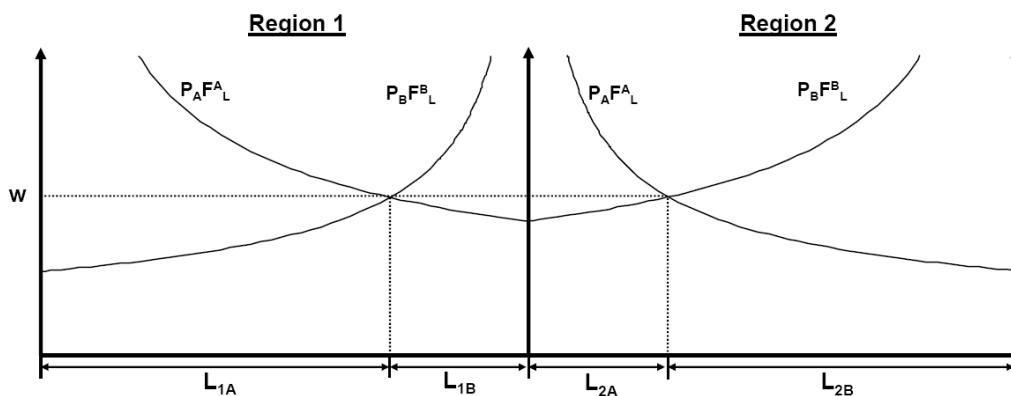
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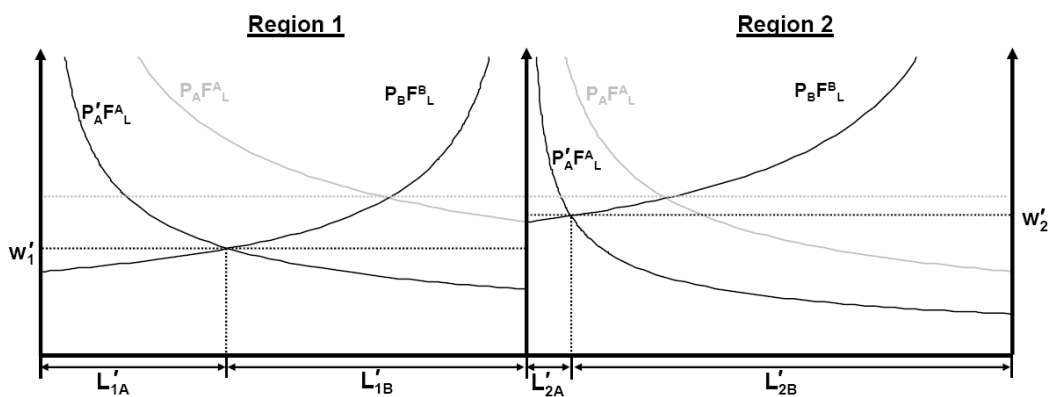
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Figure 1: Graphical Representation of Specific Factors Model of Regional Economies

(a) Initial Equilibrium



(b) Response to a Decrease in  $P_A$  – Prohibiting Migration



(c) Response to a Decrease in  $P_A$  – Allowing Migration

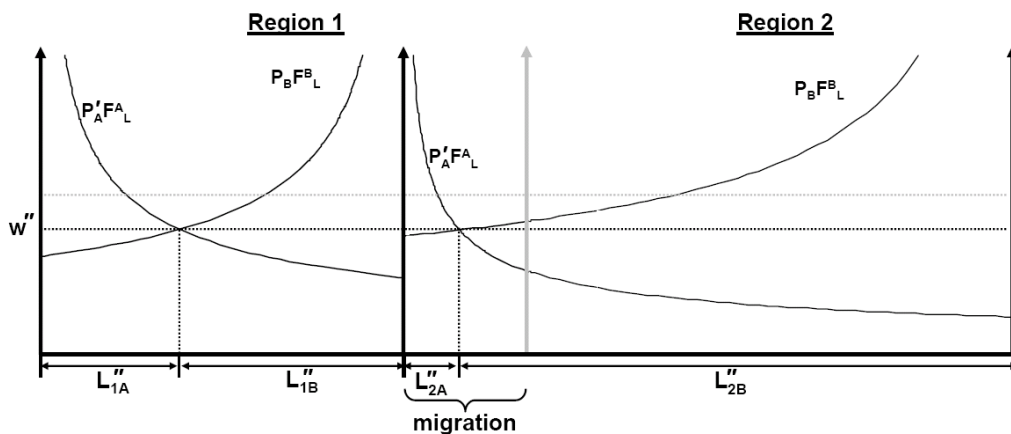


Figure 2: Nominal Tariff Timeline

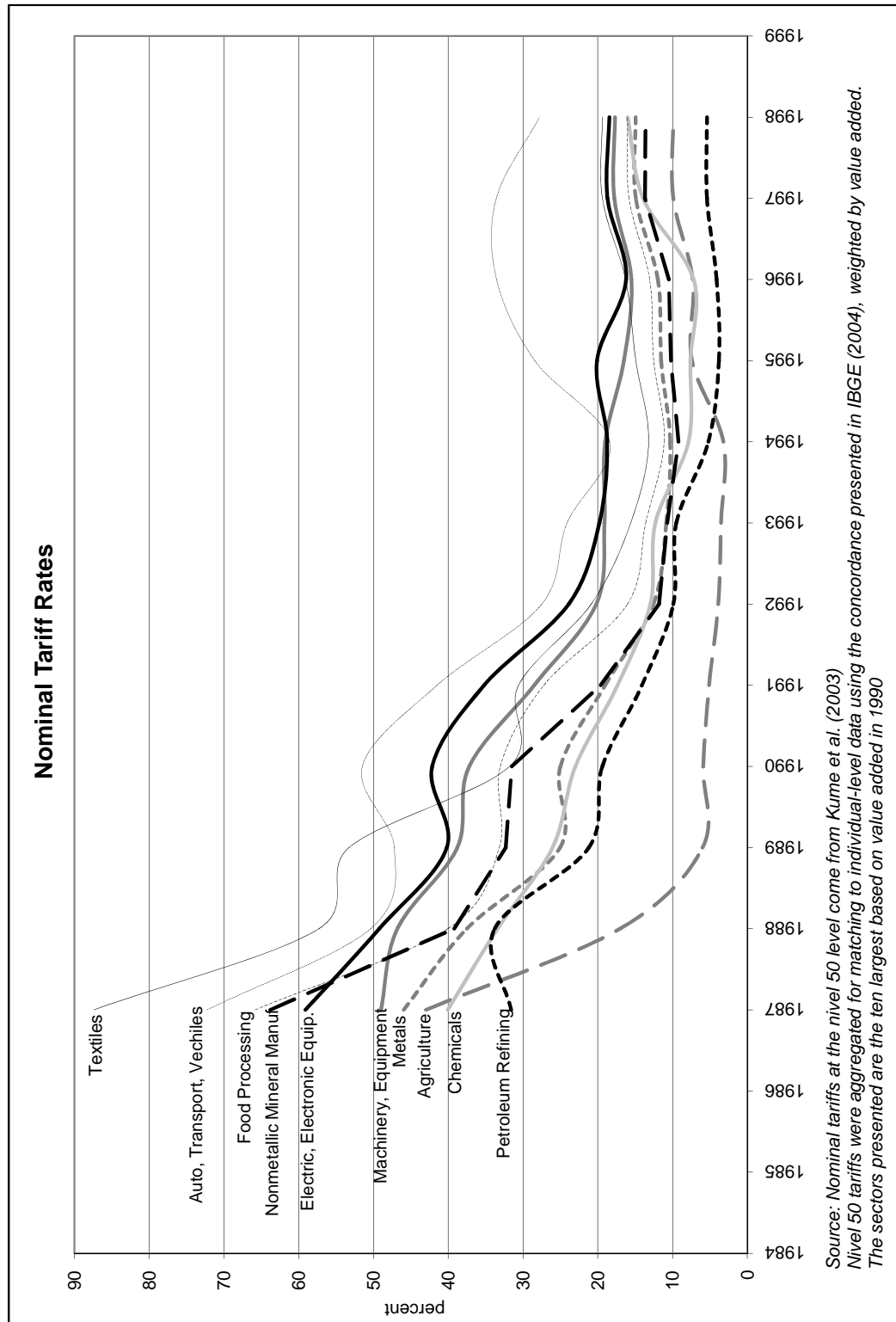


Figure 3: Effective Rate of Protection Timeline

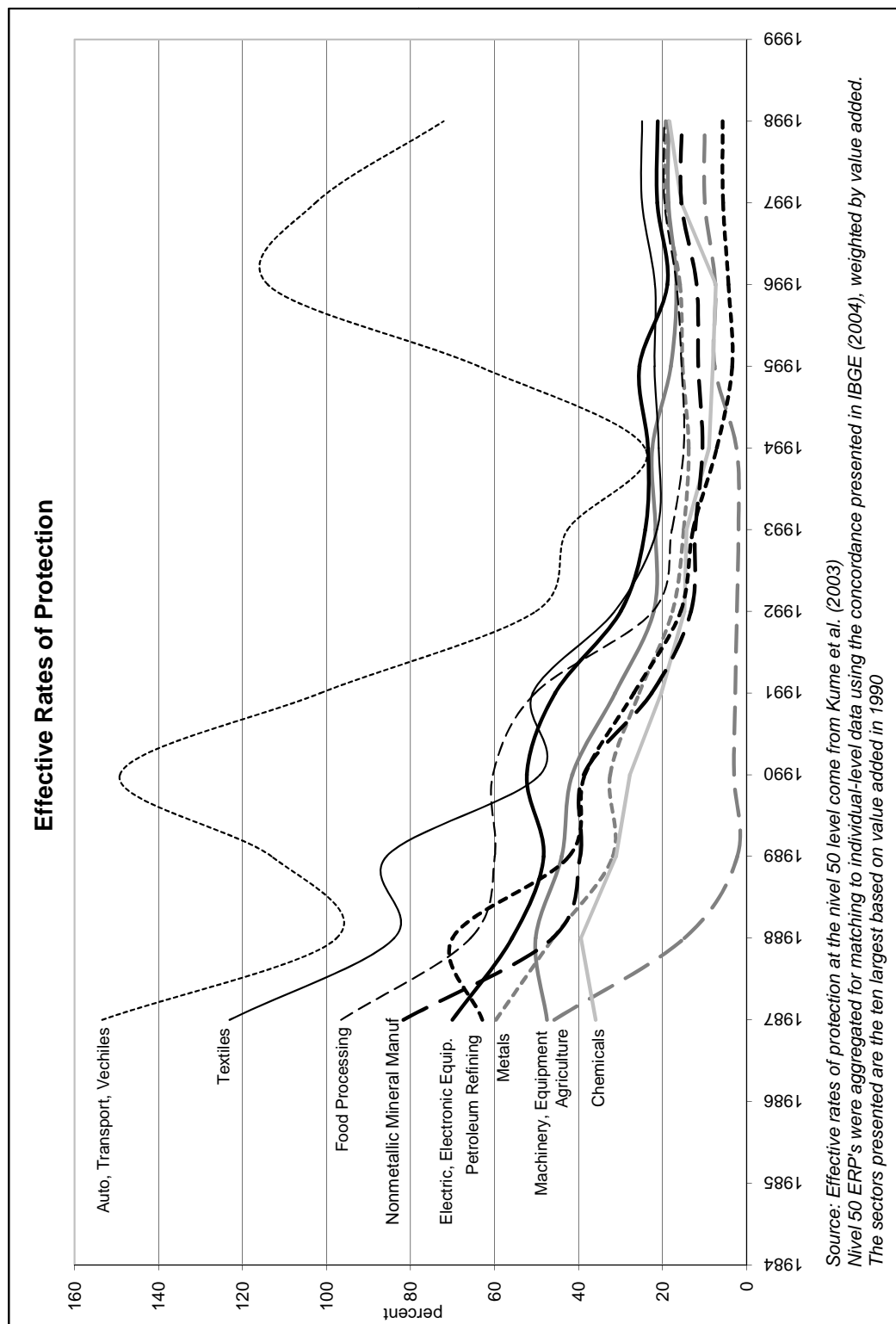


Figure 4: Industry Employment Growth and Tariff Changes

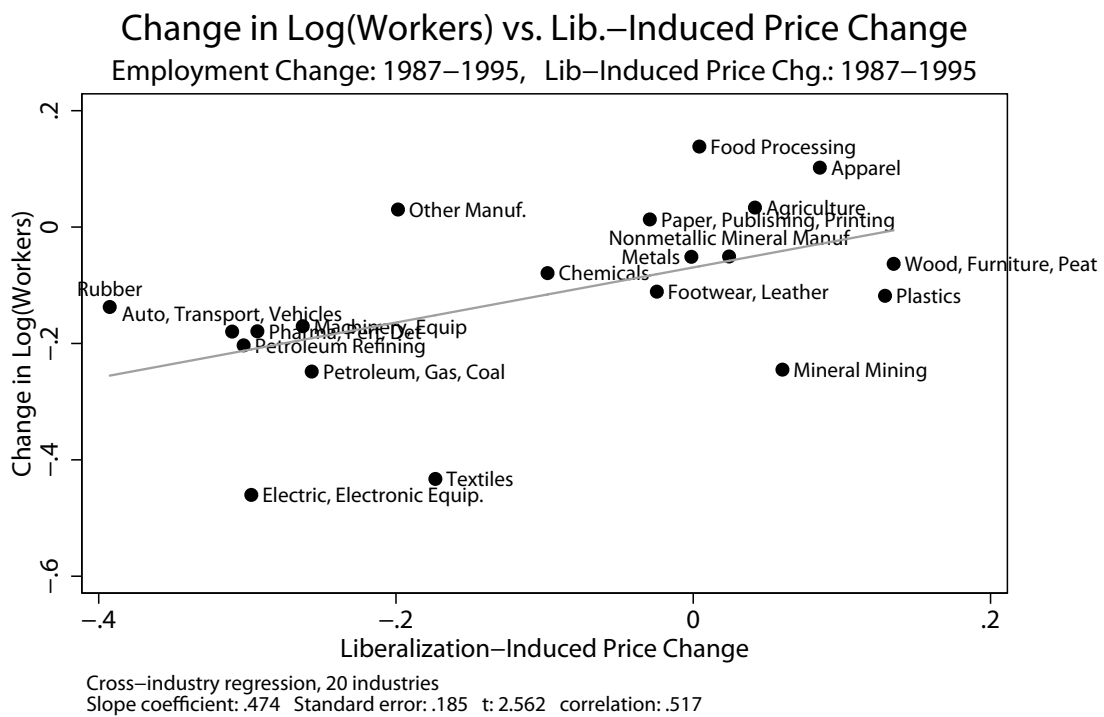


Figure 5: Industry Employment Growth and Tariff Changes - False Experiment

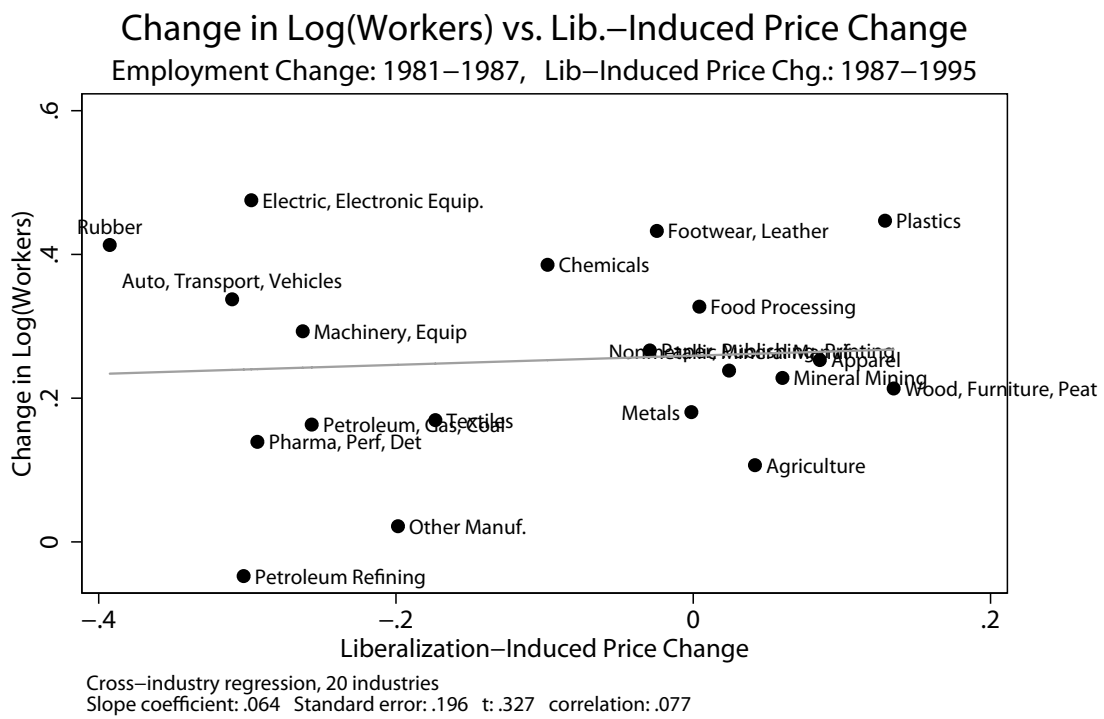




Figure 6: Relationship Between Tariff Changes and Pre-Liberalization Tariff Levels

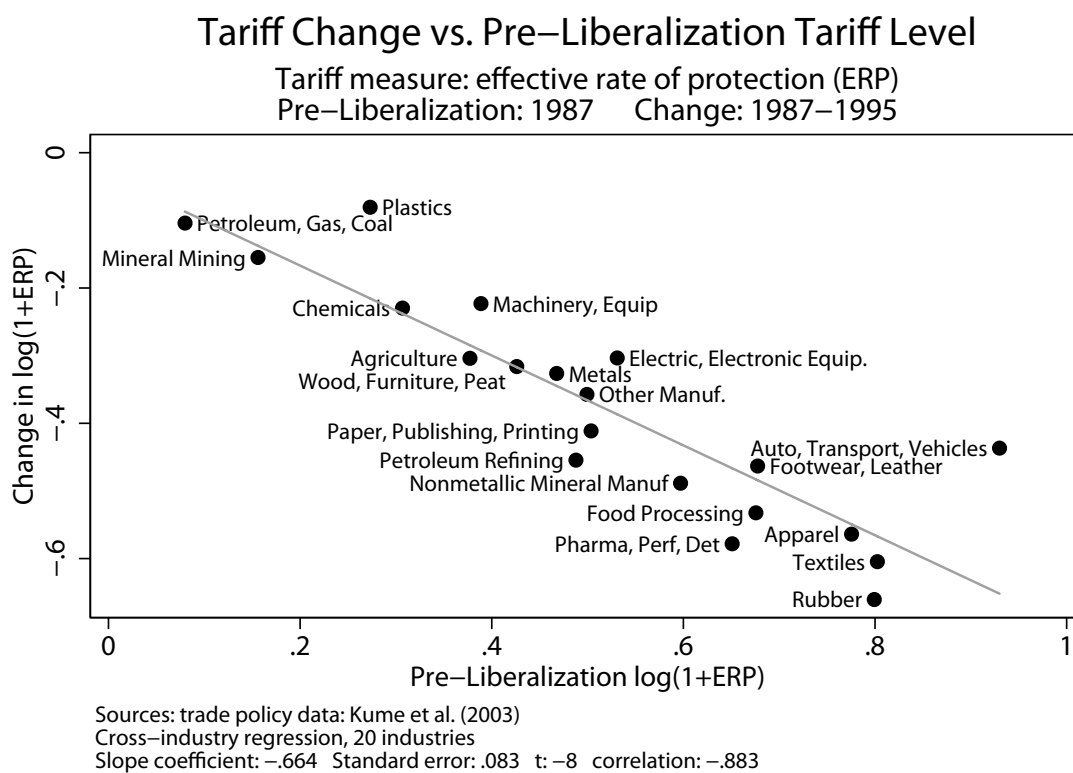
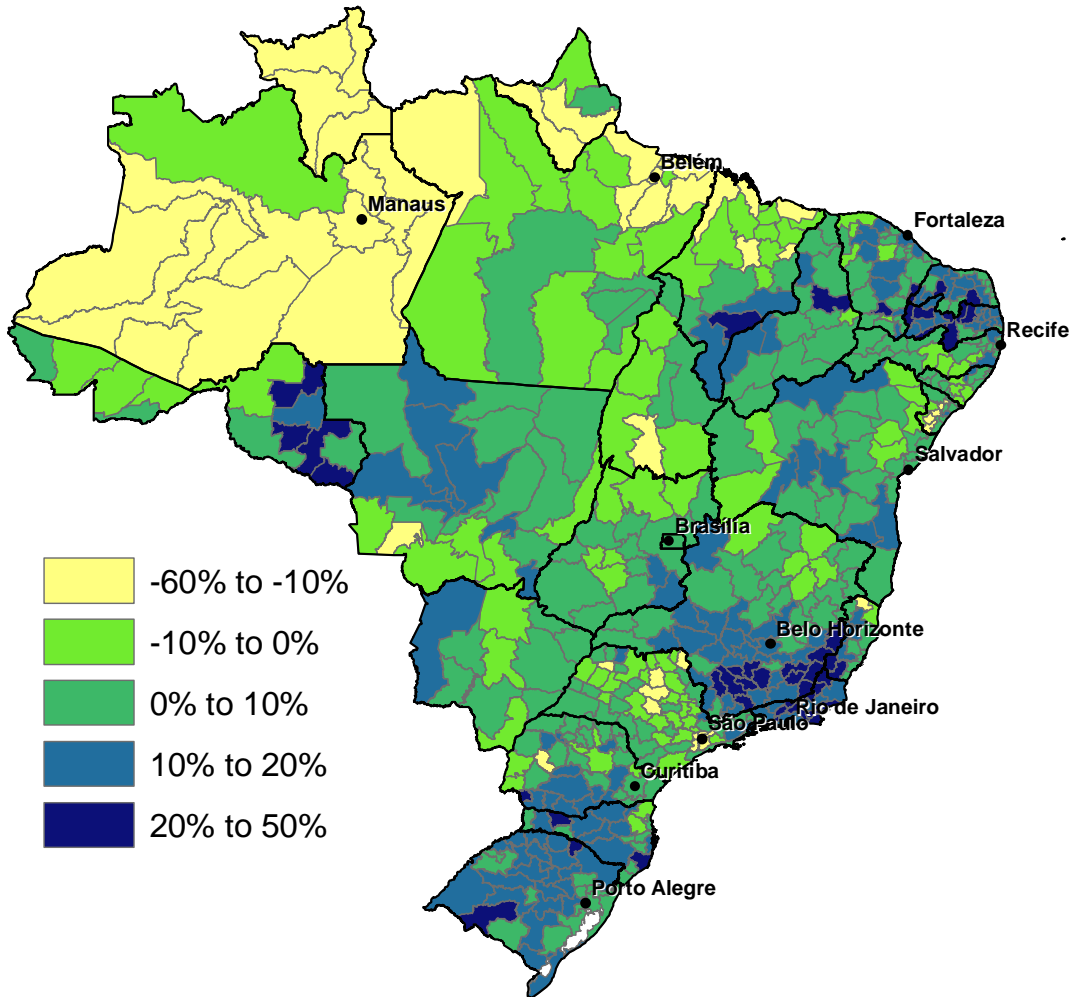
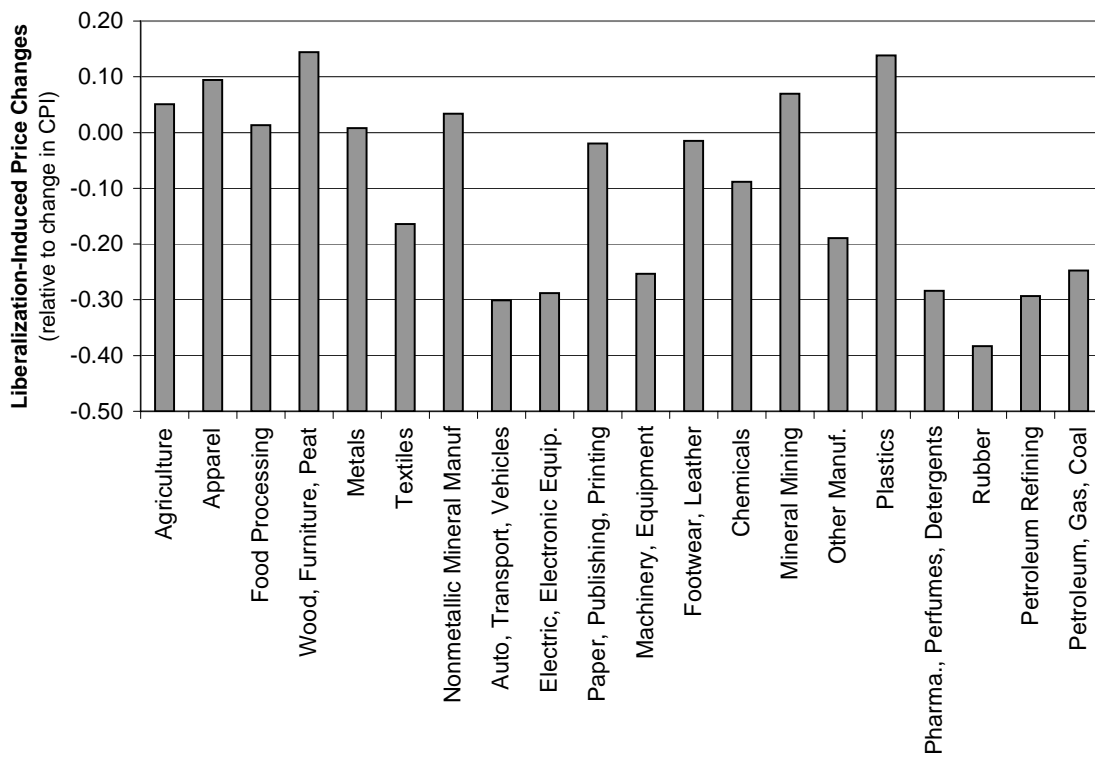


Figure 7: Regional Wage Changes



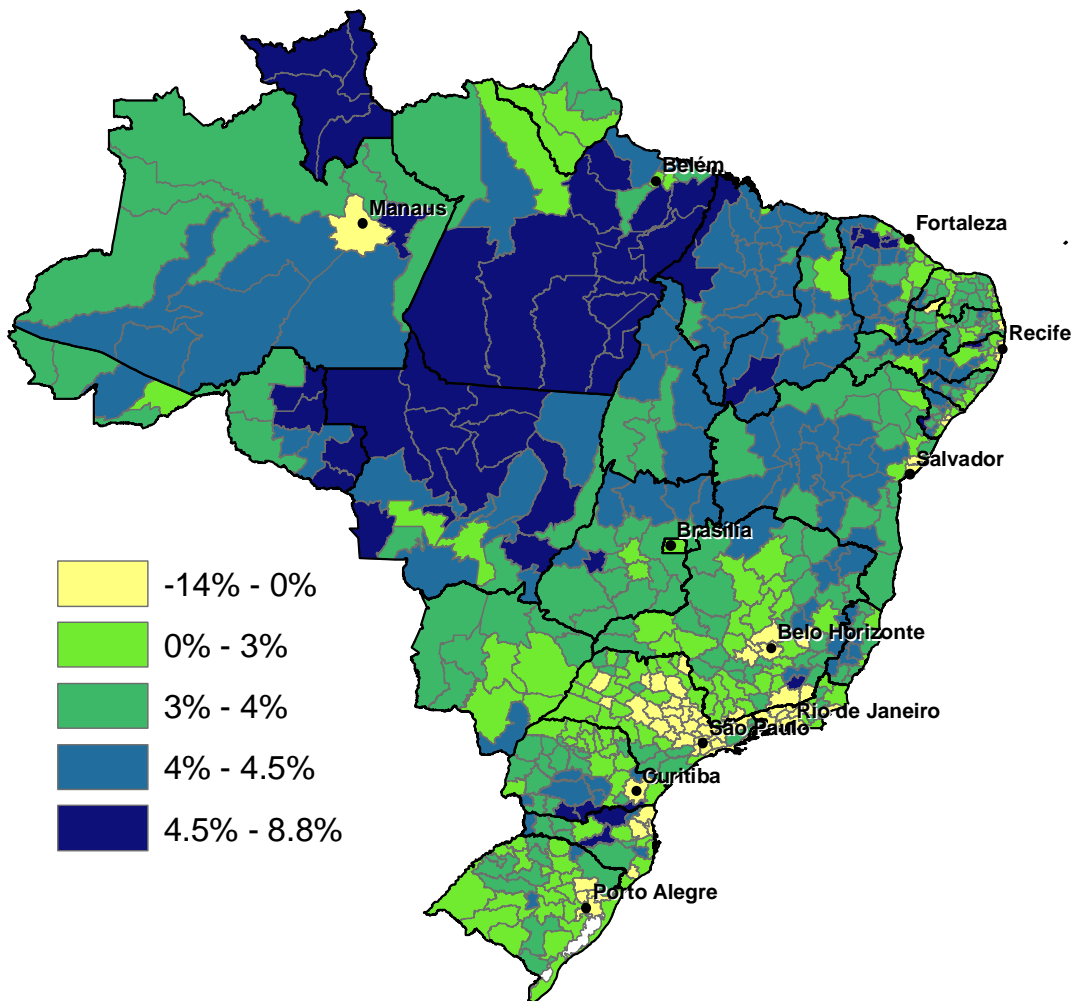
*Proportional wage change by microregion - change in microregion fixed effects from wage regression*

Figure 8: Tariff-Induced Price Changes



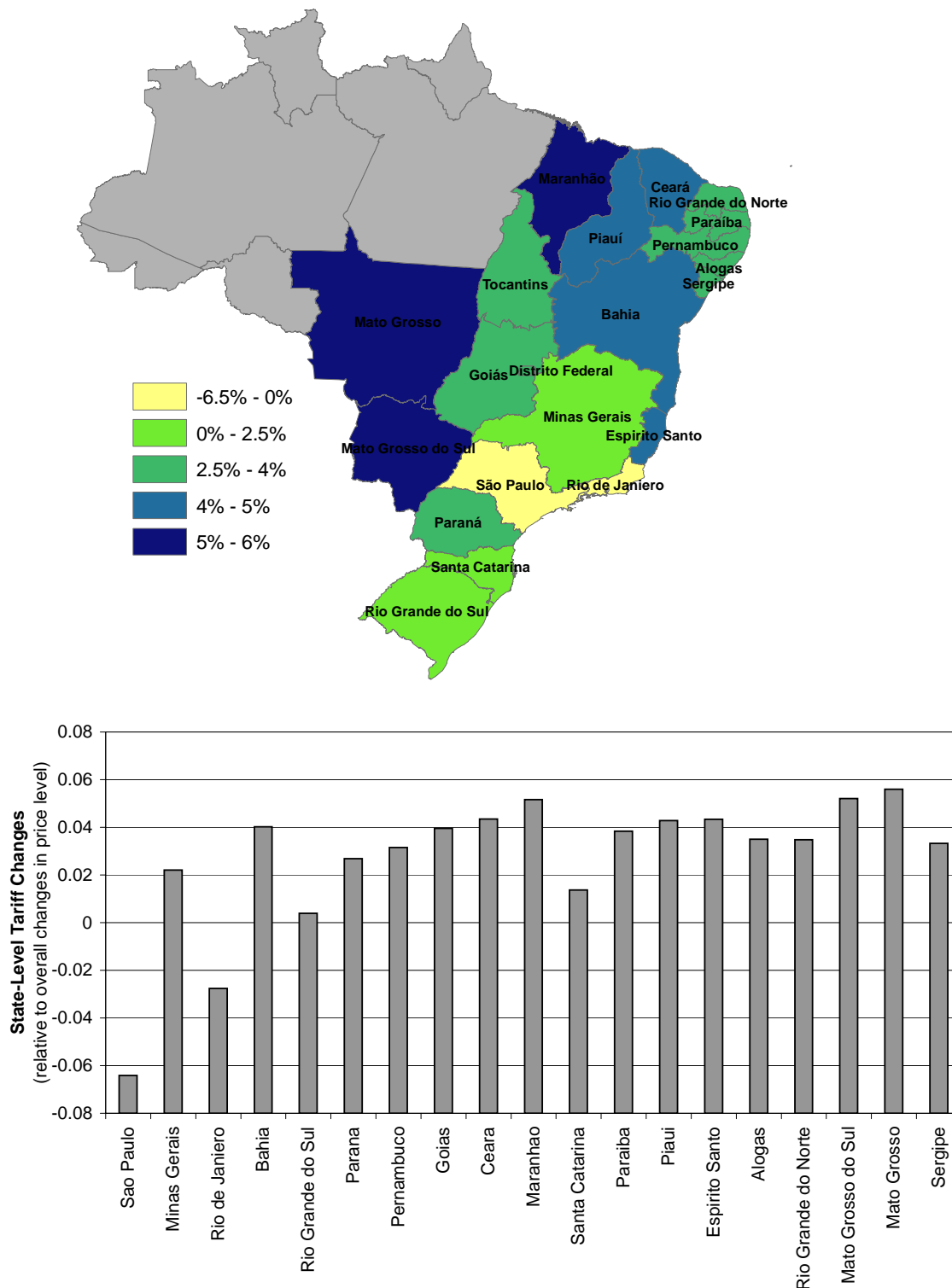
Source: Author's calculations - see text  
 Industries sorted by 1987 employment share (descending order)

Figure 9: Region-Level Tariff Changes



*Weighted average of proportional tariff changes - see text for details.*

Figure 10: State-Level Tariff Changes



*Weighted average of proportional tariff changes - see text for details.*

Table 1: Industry Aggregation and Concordance

Final Industry	Sector Name	Nivel 50	Nivel 80	PNAD / 91 Census
1	Agriculture	1	101-199	011-037, 041, 042, 581
2	Mineral Mining (except combustibles)	2	201-202	050, 053-059
3	Petroleum and Gas Extraction and Coal Mining	3	301-302	051-052
4	Nonmetallic Mineral Goods Manufacturing	4	401	100
5	Iron and Steel, Nonferrous, and Other Metal Production and Processing	5-7	501-701	110
8	Machinery, Equipment, Commercial Installation Manufacturing, and Tractor Manufacturing	8	801-802	120
10	Electrical, Electronic, and Communication Equipment and Components Manufacturing	10-11	1001-1101	130
12	Automobile, Transportation, and Vehicle Parts Manufacturing	12-13	1201-1301	140
14	Wood Products, Furniture Manufacturing, and Peat Production	14	1401	150, 151, 160
15	Paper Manufacturing, Publishing, and Printing	15	1501	170, 290
16	Rubber Product Manufacturing	16	1601	180
17	Chemical Product Manufacturing	17,19	1701-1702, 1901-1903	200
18	Petroleum Refining and Petrochemical Manufacturing	18	1801-1806	201, 202, 352, 477
20	Pharmaceutical Products, Perfumes and Detergents Manufacturing	20	2001	210, 220
21	Plastics Products Manufacturing	21	2101	230
22	Textiles Manufacturing	22	2201-2205	240, 241
23	Apparel and Apparel Accessories Manufacturing	23	2301	250, 532
24	Footwear and Leather and Hide Products Manufacturing	24	2401	190, 251
25	Food Processing (Coffee, Plant Products, Meat, Dairy, Sugar, Oils, Beverages, and Other)	25-31	2501-3102	260, 261, 270, 280
32	Miscellaneous Other Products Manufacturing	32	3201	300
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

Table 2: 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	0.060 (0.000)**	0.067 (0.000)**
Age <sup>2</sup> / 1000	-0.616 (0.004)**	-0.690 (0.003)**
Female	-0.364 (0.001)**	-0.310 (0.001)**
Inner City	0.102 (0.001)**	0.081 (0.001)**
Race		
Brown (parda)	-0.129 (0.001)**	-0.124 (0.001)**
Black	-0.192 (0.002)**	-0.164 (0.001)**
Asian	0.137 (0.006)**	0.111 (0.005)**
Indigenous	-0.158 (0.010)**	-0.102 (0.006)**
Married	0.190 (0.001)**	0.161 (0.001)**
Fixed Effects		
Years of Education (18)	X	X
Industry (21)	X	X
Microregion (558)	X	X
Observations	4962311	5664677
R-squared	0.517	0.503

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

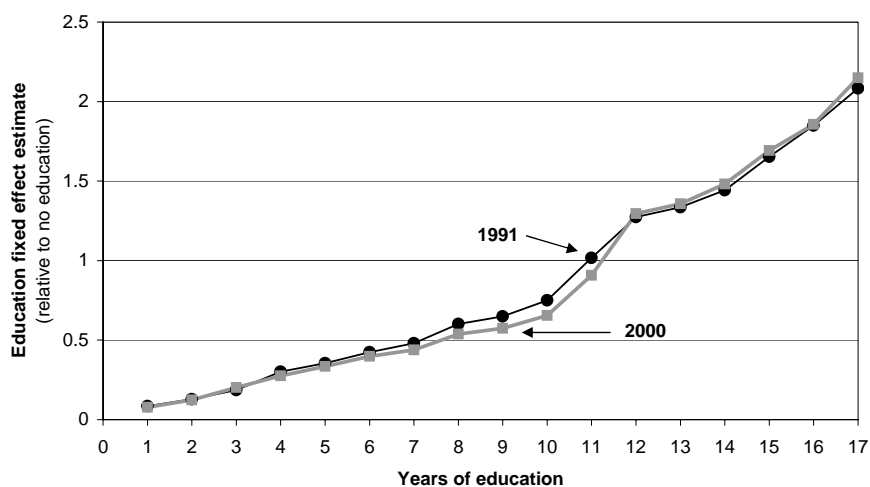


Table 3: The Effect of Tariff Changes on Price Changes

**Effect of Tariff Changes on Price Changes 1987-1995**

*dependent variable: change in log wholesale price in Brazil*

	(1)	(2)	(3)	(4)
$\Delta \ln(1 + \tau_i)$	0.029 (0.512)	0.142 (0.491)		
$\gamma_i \Delta \ln(1 + \tau_i)$			12.587 (5.446)*	12.240 (6.117) <sup>+</sup>
$\Delta \ln P_{US,i}$		0.694 (1.014)		0.397 (0.849)
Constant	18.686 (0.260)**	18.603 (0.327)**	18.933 (0.159)**	18.855 (0.295)**
R-squared	0.000	0.036	0.281	0.293

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

Table 4: The Effect of Liberalization on Local Wages

*dependent variable: change in log microregion wage<sup>a</sup>*

	(1)	(2)
Regional Liberalization Shock	0.764 (0.242)** [0.381] <sup>+</sup>	0.629 (0.171)** [0.158]**
Constant	0.037 (0.007)** [0.022]	
State Fixed Effects (27)		X
R-squared	0.061	0.620

Heteroskedasticity robust standard errors in parentheses ( )  
Standard errors adjusted for 27 clusters by state in brackets [ ]  
+ significant at 10%; \* significant at 5%; \*\* significant at 1%  
558 microregion observations  
Weighted by inverse of standard error of microregion wage premium estimate  
<sup>a</sup> Change in microregion wage premium, calculated from microregion fixed effects  
in cross-sectional wage regressions



Table 5: Migration Summary Statistics

Demographic Group	Fraction of Total Population in Demographic Group	Fraction of Demographic Group that Ever Migrated	Migrated in the 10 Years Preceding the Survey	Migrated during Sept 1982 - Sept 1988 (Pre-Liberalization)	Migrated during Sept 1996 - Sept 2002 (Post-Liberalization)
All	100.00%	29.04%	9.65%	6.22%	5.90%
Gender					
Female	51.54%	28.46%	9.21%	5.94%	5.66%
Male	48.46%	29.65%	10.12%	6.51%	6.15%
Age					
18-24	24.93%	19.83%	10.58%	5.63%	7.19%
25-34	30.27%	28.36%	12.43%	8.09%	7.65%
35-54	44.80%	34.63%	7.26%	5.15%	4.11%
Education					
0	11.72%	30.06%	8.80%	5.39%	5.57%
1-3	14.51%	31.01%	9.82%	6.31%	5.99%
4-7	31.46%	29.92%	9.90%	6.32%	5.91%
8-10	16.30%	27.58%	9.76%	6.28%	5.76%
11-14	20.00%	25.81%	9.06%	6.14%	5.73%
15+	2.22%	36.99%	14.05%	10.02%	9.10%
Race					
White	55.50%	28.85%	9.22%	6.30%	5.51%
Brown ( <i>Pardo</i> )	38.16%	30.09%	10.65%	6.49%	6.51%
Black	5.72%	23.41%	6.95%	3.71%	4.83%
Asian	0.48%	33.81%	12.63%	4.97%	10.39%
Indigenous	0.14%	30.34%	10.55%	3.23%	10.77%
Marital Status					
Married	62.23%	32.46%	10.13%	6.88%	5.89%
Unmarried	37.77%	23.40%	8.87%	5.04%	5.91%
Family Size					
1-2	15.51%	31.31%	11.95%	7.14%	7.87%
3-4	49.84%	28.93%	9.58%	6.45%	5.67%
5-6	25.89%	29.32%	8.98%	5.97%	5.16%
7+	8.76%	24.80%	8.00%	4.93%	4.90%
Baseline Population	81,099,568	81,099,568	81,099,568	72,282,488	92,562,936
Survey Years Utilized	1992-2002	1992-2002	1992-2002	1992	2002
Observations	1,612,368	1,612,368	1,612,368	158,061	208,080

Source: Author's calculations based on 1992-2002 PNAD  
Sample: Individuals age 18-55

Table 6: The Effect of State-Level Tariff Changes on Location Choice

Additional grouping beyond source state	Homogeneous effect across groups	Heterogeneous effect across groups	Observations
<u>None</u>	1.920 (0.983) <sup>+</sup>		168
<u>Age</u>	1.848 (0.872) <sup>+</sup>		253
Age 18-34		2.524 (0.864) <sup>**</sup>	
Age 35-55		0.576 (1.725)	
<u>Gender</u>	1.768 (0.975) <sup>+</sup>		258
Male		2.118 (1.072) <sup>+</sup>	
Female		1.381 (0.898)	
<u>Education</u>	2.733 (1.048) <sup>*</sup>		251
0-7 Years		3.583 (1.006) <sup>**</sup>	
8+ Years		1.048 (1.335)	
<u>Race</u>	1.828 (1.029) <sup>+</sup>		241
White		2.376 (1.348) <sup>+</sup>	
Non-white		1.054 (0.844)	
<u>Family Size</u>	2.291 (1.119) <sup>+</sup>		231
4 or fewer		2.413 (1.043) <sup>*</sup>	
5 or more		2.061 (1.369)	

Standard errors clustered by 19 destination states  
+ statistically significant at 10%, \* at 5%, \*\* at 1%

Observations represent group (including source state) x destination pairs  
Sample: Individuals age 18-55 at time of survey  
Dropping groups with less than 5 observations in either period  
Weighted by the square root of the number of observations in each cell

Source: Author's calculations based upon the following data sets  
Migration and employment: 1987, 1992, and 2002 PNAD  
Trade policy: Kume et al. (2003)  
Import penetration: IBGE Brazil national accounts

Table 7: Liberalization-Induced Population Shifts

<b>State</b>	<b>Liberalization-induced population change</b>	<b>1988 population (aged 18-55)</b>	<b>Proportional population change</b>
Mato Grosso	20,655	780,113	2.65%
Mato Grosso do Sul	19,820	853,602	2.32%
Paraiba	18,420	1,347,519	1.37%
Espirito Santo	13,553	1,161,370	1.17%
Alogas	10,877	987,854	1.10%
Bahia	53,789	4,936,731	1.09%
Pernambuco	33,794	3,209,443	1.05%
Ceara	27,433	2,703,695	1.01%
Parana	40,412	4,375,543	0.92%
Piaui	8,541	1,051,501	0.81%
Minas Gerais	58,643	7,481,558	0.78%
Sergipe	4,141	580,131	0.71%
Rio Grande do Norte	6,453	1,016,421	0.63%
Goiias	17,062	3,244,584	0.53%
Maranhao	9,111	2,007,522	0.45%
Santa Catarina	3,236	2,197,374	0.15%
Rio Grande do Sul	-4,749	4,672,987	-0.10%
Rio de Janiero	-43,956	7,331,464	-0.60%
Sao Paulo	-297,234	16,810,570	-1.77%

*Source: Author's calculations - see text for details*