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Regional Effects of Trade Reform: What is the Correct Measure of Liberalization?†

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Over the last 40 years, trade barriers around the world have fallen to historically low levels. As part of this process, many developing countries abandoned import substituting industrialization policies by sharply lowering trade barriers, motivating a large literature examining the effects of trade liberalization on various national labor market outcomes such as poverty and inequality.[1] The focus on national outcomes follows the approach of classical trade theory, which takes the country as the geographic unit of analysis. A small but growing literature takes a different approach, examining the effects of trade liberalization on labor market outcomes at the subnational level. The papers in this literature measure the local effect of liberalization using a weighted average of changes in trade policy, with weights based on the industrial distribution of labor in each region.[2] In this article, I develop a specific-factors model of regional economies that yields a very similar weighted-average relationship between regional wage changes and liberalization-induced price changes across industries. The model provides a theoretical foundation for this intuitively appealing empirical approach and provides guidance on important choices faced by researchers when constructing regional measures of trade liberalization.

The model shows that liberalization in a particular industry will have a larger effect on local wages when (i) liberalization has a larger effect on the prices faced by producers, (ii) the industry accounts for a larger share of local employment, and (iii) labor demand in the industry is more elastic.[3] The model also incorporates the nontraded sector, showing that nontraded prices move with traded goods prices during liberalization. This finding supports omitting the nontraded sector from the local weighted average and suggests that an alternative approach implicitly assuming that

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

[2] Topalova's (2007) influential paper was followed by Edmonds, Pavcnik, and Topalova (2010), Hasan, Mitra, and Ural (2006), Hasan et al. (2012), McCaig (2011), Topalova (2010), McLaren and Hakobyan (2010), and Autor, Dorn, and Hanson (2012). Note that in contrast to the other papers in the literature, Autor, Dorn, and Hanson (2012) study the effects of increased imports, rather than the effects of changes in trade policy.

[3] See Section I for a definition of the relevant labor demand concept.

nontraded prices are unaffected by liberalization will yield substantially upward-biased estimates when studying liberalization's effect on wages.[4]

I empirically examine the model's predictions in the context of Brazil's trade liberalization in the early 1990s. Brazilian liberalization involved drastic reductions in overall trade restrictions and a decrease in the variation of trade restrictions across industries, implying wide cross-industry variation in tariff cuts. Additionally, the industrial composition of the labor force varies substantially across Brazilian regions. This variation in tariff changes across industries and industrial composition across regions combine to identify the effect of liberalization on local wages.

The empirical results confirm the model's predictions. Local labor markets whose workers were concentrated in industries facing the largest tariff cuts were generally affected more negatively, while markets facing smaller cuts were more positively affected. A region facing a 10 percentage point larger liberalization-induced price decline experienced a 4 percentage point larger wage decline (smaller wage increase) than a comparison region. I also investigate deviations from the weighted-average measure supported by the model and find that treatment of the nontraded sector is quite important in determining the magnitude of liberalization's effect on wages, as predicted by the theoretical analysis. The model supports omitting the nontraded sector from the regional weighted average, and estimates obtained when doing so are of the expected magnitude. In contrast, estimates that implicitly assume that nontraded prices were unaffected by liberalization are more than four times larger.

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

The remainder of the paper is organized as follows. Section I develops the 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 II describes the datasets used, and Section III describes the specific trade policy changes implemented in Brazil's liberalization. Section IV empirically examines the effects of trade liberalization on wages across local labor markets, including an investigation of various alternative regional measures of liberalization. Section V concludes.

[4] Note that much of the prior literature studies nonwage outcomes such as intraregional inequality (Topalova 2007, 2010) and child labor (Edmonds, Pavcnik, and Topalova 2010). The present model does not incorporate these features and, hence, has little to say about treatment of the nontraded sector in those contexts.

I. Specific-Factors Model of Regional Economies

A. *Price Changes' Effects on Regional Wages*

Each region within a country is modeled as a Jones (1975) specific-factors economy.[5] Consider a country with many regions, indexed by r. The economy consists of many industries, indexed by i. Production uses two inputs. Labor, L, is assumed to be mobile between industries, immobile between regions, supplied inelastically, and fully employed. The second input, T, is not mobile between industries or regions. This input represents fixed characteristics of a region that increase the productivity of labor in the relevant industry. Examples include natural resource inputs such as mineral deposits, fertile land for agriculture, regional industry agglomerations that increase productivity (Rodriguez-Clare 2005), or fixed industry-specific capital. 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).

This setup yields the following relationship between regional wages and goods prices. All theoretical results are derived in online Appendix A (the following expression is (A13) with labor held constant).

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

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

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

Equation (1) describes how a particular region's wage will be affected 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, $\frac{\sigma}{\theta}$ represents a labor demand elasticity in which the specific factor is held fixed, while output and the specific factor price

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

may vary.[6] 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 need to absorb large amounts of labor from other industries in order to decrease the marginal product of labor sufficiently to restore equilibrium. Thus, price changes in these industries have more weight in determining equilibrium wage changes. For further intuition, see the graphical treatment in online Appendix A.

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.

B. *Nontraded Sector*

This subsection introduces a nontraded sector in each region, demonstrating that nontraded prices move with traded prices. This finding guides the empirical treatment of nontradables, which generally represent a large fraction of modern economies. As above, 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 prior results but makes it necessary to describe regional consumers' preferences to determine the nontraded good's equilibrium price in each region. I assume that all individuals have identical Cobb-Douglas preferences, permitting the use of a representative regional consumer who receives as income all wages and specific factor payments earned in the region.[7]

This setup yields the following relationship between the regional price of nontradables and tradable goods prices (the following expression is (A22) with labor held constant):

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

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

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

[6] 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.

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

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

This finding is important in guiding the empirical treatment of the nontraded sector. Previous empirical studies of trade liberalizations' effects on regional labor markets pursue two different approaches. The first approach sets the nontraded term in (1) to zero, since trade liberalization has no direct impact on the nontraded sector.[8] In the context of the present model examining wages, this is equivalent to assuming no price change for nontraded goods. This approach is not supported by the model presented here, which predicts that nontraded prices move with traded prices. Setting the price change to zero in the large nontraded sector would lead the weighted average to substantially understate the size of liberalization's effect on regional labor demand and overstate its effect on regional wages.

The second approach removes the nontraded sector from the weighted average in (1) and rescales the weights for the traded industries in (2) such that they sum to one.[9] This approach more closely conforms to the model just described. If the nontraded price changes by approximately the same amount as the average traded price, as described in (3), then dropping the nontraded price from (1) will have very little effect upon the overall average.[10] Ideally, one would simply calculate the terms in (4) using detailed data on production values across industries at the regional level and substitute the result into (1). However, when data on regional output by industry are unavailable, as is the case in the empirical analysis below, the model implies that dropping the nontraded sector is likely to provide a very close approximation to the ideal calculation.

It should be noted that a number of previous papers study nonwage outcomes such as poverty, inequality, and child labor that the present model does not address.[11] Even when examining wages, however, under additional technological and labor market restrictions, setting the nontraded price change to zero is equivalent to multiplying the full weighted average by a positive scalar.[12] This difference will have no effect on sign tests but will affect the size of the estimates. If the additional restrictions hold, conclusions regarding the effects on liberalization across regions remain largely unaffected.

[8] This approach is used in Autor, Dorn, and Hanson (2012), Edmonds, Pavcnik, and Topalova (2010), McCaig (2011), McLaren and Hakobyan (2010), Topalova (2007), and Topalova (2010).

[9] This approach is used in Hasan et al. (2012) and Hasan, Mitra, and Ural (2006), presented as a robustness check in McCaig (2011), and used as an instrumental variable in Edmonds, Pavcnik, and Topalova (2010), Topalova (2007), and Topalova (2010).

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

[11] Edmonds, Pavcnik, and Topalova (2010) focus on child labor, while McCaig (2011), Topalova (2007), and Topalova (2010) examine the effects of liberalization on poverty and inequality.

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

II. Data

The preceding section described a specific-factors model of regional economies that yields predictions for the effects of changes in tradable goods' prices on regional wages and the prices of nontraded goods. This framework can be used to measure the local impacts of any event in which a country faces price changes that vary exogenously across industries. In the remainder of the article, I apply the model to the analysis of the regional impacts of trade liberalization in Brazil, requiring the combination of various industry-level and individual-level data sources.

The model is driven by exogenous changes in prices across tradable industries, which in the context of trade liberalization are driven by tariff changes. Trade policy data at the Nível 50 industrial classification level (similar to two-digit SIC) come from researchers at the Brazilian Applied Economics Research Institute (IPEA). Kume, Piani, and de Souza (2003) report nominal tariffs and effective rates of protection using the Brazilian input-output tables. Nominal tariffs are the preferred measure of protection, but all results were also generated using effective rates of protection without any substantive differences from those presented here.

Wage and employment data come primarily from the long form Brazilian Demographic Censuses (*Censo Demográfico*) for 1991 and 2000 from the Brazilian Institute of Geography and Statistics (IBGE). Throughout the analysis, local labor markets are defined as microregions. Each microregion is a grouping of economically integrated contiguous municipalities with similar geographic and productive characteristics (Instituto Brasileiro de Geografia e Estatística 2002).[13] Wages are calculated as earnings divided by hours. The census also reports employment status and industry of employment, which permits the calculation of the industrial distribution of labor in each microregion. I restrict the sample to individuals aged 18–55 who are not currently enrolled in school in order to focus on people who are most likely to be tied to the labor force. The wage analysis in Section IV further restricts the sample to those receiving nonzero wage income. While it would be ideal to have wage and employment information just before liberalization began in March 1990, I use the 1991 census as the baseline period under the assumption that wages and employment shares adjusted slowly to the trade liberalization.[14] An alternative annual household survey, the *Pesquisa Nacional por Amostra de Domicílios* (PNAD), is available yearly but reports only state-level geographic information, making it impossible to identify local markets. I therefore use the census when analyzing the effects of liberalization on local wages.

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

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

[14] See the following section for a discussion of the timing of liberalization.

III. Trade Liberalization in Brazil

A. *Context and Details of Brazil's Trade Liberalization*

From the 1890s to the mid-1980s 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 and nontariff barriers (Abreu 2004a; Kume, Piani, and de Souza 2003). The average tariff level in 1987 was 54.9 percent, with values ranging from 15.6 percent on oil, natural gas, and coal to 102.7 percent 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, Piani, and de Souza 2003). Along with high nominal tariffs was a list of items whose import was prohibited. This list, known as "Anexo C," was in place from 1975 to March 1990 (Hahn 1991) and had extensive coverage; in January 1987, the list of prohibited imports covered 38 percent of individual tariff lines.[15]

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 1980s that the policy was no longer sustainable (Abreu 2004a). Large amounts of international borrowing in response to the oil shocks of the 1970s, followed by slow economic growth in the early 1980s, 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. These events prompted trade reforms, beginning in late 1987 with a governmental Customs Policy Commission (Comissão de Politica Aduaneira) proposal of sharp tariff reductions and the removal of many nontariff barriers.[16]

In June of 1988 the government adopted a reform that removed a few nontariff barriers and lowered nominal tariffs. However, these initial reforms had very little impact on the actual protection faced by producers in Brazil, due to large levels of tariff redundancy resulting primarily from the presence of a system of "special customs regimes" that granted preferential access to particular types of imports, either waiving or reducing import duties for qualifying imports. The system was very extensive; between 1977 and 1985, 69 percent of imports benefited from one or more special customs regimes (Kume 1990). The 1988 tariff cuts primarily served to eliminate tariff redundancy without affecting the actual level of protection, as measured by the gap between international prices and those faced by Brazilian producers.[17]

In March 1990, substantial customs reforms were implemented that abolished the most important nontariff barriers, including the list of suspended import licenses,

[15] Author's calculations based on the list of suspended import licenses presented in Supplement 2 to the *Bulletin International des Douanes*, number 6, 11th edition.

[16] See Kume (1990) and Kume, Piani, and de Souza (2003) for detailed accounts of Brazil's liberalization.

[17] See Kume (1990) for a detailed summary and analysis of the special customs regime system and the 1988 tariff reforms. He shows that even after the 1988 tariff cuts, nominal tariffs were well above the prereform gaps between internal and external prices, reflecting the actual protection Brazilian producers faced. Kume concludes that "… the result achieved [by the tariff reform of June 1988] merely contributed more transparency to the tariff structure, without inducing a process of import liberalization. In this sense, the tariff reform maintained the structure of protection already in place" [author's translation].

and the majority of the special customs regimes.[18] Simultaneously, tariff levels were adjusted to reflect the prior gap between international prices and those in Brazil. This process, known as "tariffization" (*tarifação*), effectively replaced the nontariff barriers and special customs regimes with tariffs providing equivalent levels of protection (de Carvalho 1992; Kume, Piani, and de Souza 2003). By the second half of 1990, tariffs were the primary instruments of protection in Brazil, and they largely reflected the gaps between international prices and those faced by Brazilian producers. I therefore consider 1990 as the starting point for measuring tariff liberalization.[19]

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, Piani, and de Souza 2003). Following 1994, there was a slight reversal of the previous tariff reductions, but tariffs remained essentially stable following this period. Thus, I measure trade liberalization based on the 1990–1995 proportional change in one plus the tariff rate, which corresponds to the proportional price change in the model.

B. *Exogeneity of Tariff Changes to Industry Performance*

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

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

More striking support for exogeneity comes from an examination of the nature of the tariff cuts during Brazil's liberalization, following the approach of Goldberg

[18] The remaining special customs regimes included (i) the "drawback" in which tariffs were refunded for imports that were eventually reexported, (ii) those required by international agreements, and (iii) the Manaus Free Trade Zone (Kume, Piani, and de Souza 2003). Note that Manaus is omitted from the analysis for this reason.

[19] Consistent with the interpretation that genuine liberalization did not begin until 1990 after the removal of nontariff barriers and the special customs regimes, tariff changes between 1987 and 1990 have no relationship to Brazilian wholesale price changes during that period, while subsequent tariff changes between 1990 and 1995 exhibit a strong positive relationship to prices. Results are available upon request.

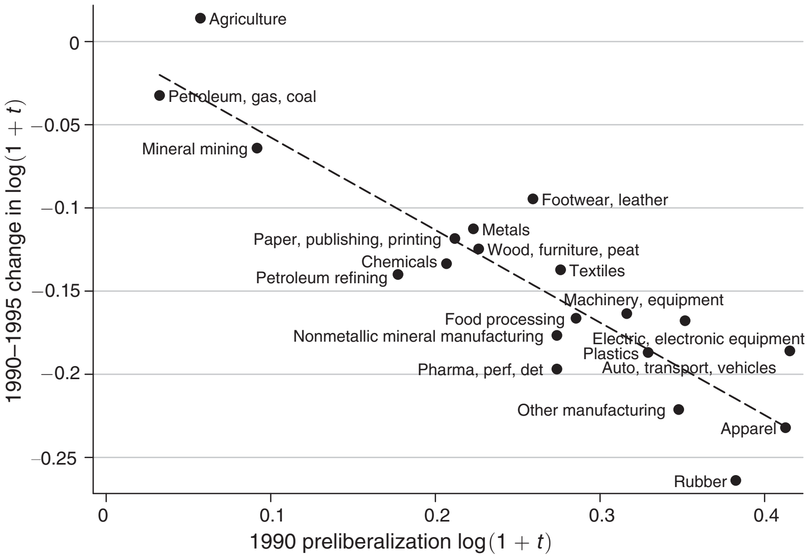

Figure 1. Relationship between Tariff Changes and Preliberalization Tariff Levels

Note: Correlation: −0.899; regression coefficient: −0.556; standard error: 0.064; t: −8.73.
Source: Author's calculations based on data from Kume, Piani, and de Souza (2003).

and Pavcnik (2005). It was a stated goal of policymakers to reduce tariffs in general, and to reduce the cross-industry variation in tariffs to minimize distortions relative to external incentives (Kume, Piani, and de Souza 2003). This equalizing of tariff levels implies that the tariff changes during liberalization were almost entirely determined by the preliberalization tariff levels. Figure 1 shows that industries with high tariffs before liberalization experienced the greatest cuts, with the correlation between the preliberalization tariff level and change in tariff equaling −0.90. Since the liberalization policy imposed cuts based on a protective structure that was set decades earlier (Kume, Piani, and de Souza 2003), it is unlikely that the tariff cuts were manipulated to induce correlation with counterfactual industry performance or with industrial political influence.[20]

IV. The Effect of Liberalization on Regional Wages

A. *Regional Wage Changes*

The model described in Section I considers homogenous labor, in which all workers are equally productive and thus receive identical wages in a particular region.

[20] It should be noted that the 1990–1995 tariff changes are negatively correlated with the preliberalization 1985–1990 growth in industry employment, indicating that industries that were growing more quickly during 1985–1990 subsequently experienced larger tariff cuts during liberalization in 1990–1995. While this correlation is consistent with strategic behavior in which the "strongest" industries were allowed to face increased international competition, under a counterfactual in which the trends would have continued, such a relationship would impart downward bias to the wage results below, going against finding the positive estimates they exhibit.

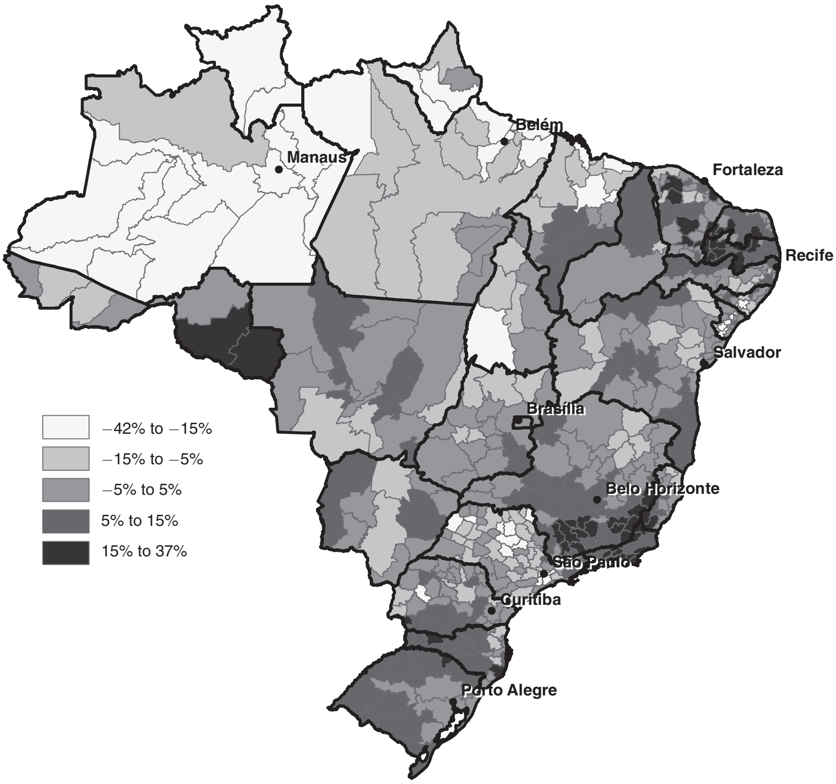

FIGURE 2. REGIONAL WAGE CHANGES

Note: Normalized change in microregion fixed effects from wage regressions.

In reality, wages differ systematically across individuals, and the observed wage change in a given region could be due to changes in individual characteristics or changing returns to those characteristics. In order to net out these effects, I calculate regional wage changes as follows. In 1991 and 2000 I separately estimate a standard wage equation, regressing the log of real wages on demographic and educational controls, industry fixed effects, and microregion fixed effects.[21] I then normalize the microregion fixed effects relative to the average log wage change and calculate the associated standard errors based on Haisken-DeNew and Schmidt (1997).

Figure 2 shows the resulting estimated regional wage changes in each microregion of Brazil. States are outlined in bold, while each smaller area outlined in gray is a microregion. Microregions that are lighter experienced the largest wage declines during the 1991–2000 time period, while darker regions experienced the largest wage increases, relative to the national average. As the scale indicates, some observations are quite large in magnitude, though only seven observations fall outside

[21] The results of these regressions are reported in online Appendix B, Table B2.

the ± 30 percent range, and these are all in sparsely populated areas with imprecise estimates that receive little weight in subsequent analysis.[22]

B. *Region-Level Tariff Changes*

Based on (1), trade liberalization's effect on a region's wages is determined by a weighted average of liberalization-induced price changes. In what follows, I call this weighted average the "region-level tariff change." I measure liberalization-induced price changes as $d\ln(1+\tau_i)$ where τ_i is the tariff rate and d represents the long-difference from 1990 to 1995.[23] Calculating the β_{ri} terms in (1) requires information for each region on the allocation of labor across industries and on labor demand elasticities in each industry. The industrial allocation of labor is calculated for each microregion from the 1991 census. There exist no credible estimates of labor demand elasticities by Brazilian industry and region; in fact, I am unaware of any estimates of industry-specific labor demand elasticities for any country, even restricting the elasticities to be constant across regions. Given this limitation, for the empirical analysis I assume that production in all industries is Cobb-Douglas, and that the factor shares may vary across industries, implying that $\sigma_{ri} = 1$ and $\theta_{ri} = \theta_i$. I calculate θ_i as one minus the wagebill share of industry value added using national accounts data from IBGE.

Given these restrictions I calculate the region-level tariff change (RTC) for each microregion as follows:

$$RTC_r = \sum_{i \neq N} \beta_{ri}\, d\ln(1+\tau_i) \tag{5}$$

$$\text{where} \qquad \beta_{ri} = \frac{\lambda_{ri}\frac{1}{\theta_i}}{\sum_{i' \neq N} \lambda_{ri'}\frac{1}{\theta_{i'}}}. \tag{6}$$

Recall from Section IB that ideally one would directly measure the nontraded prices in each region or model them using the traded goods prices as in (3). Given that neither nontraded prices nor output by industry are available by region in Brazil, these ideal approaches are not feasible in this case. Instead, I drop the nontraded sector from the weighted average in (5) based on the conclusion that nontraded prices move with traded prices, following the discussion in Section IB.

The results of this calculation appear in Figure 3. Lighter microregions faced the largest tariff cuts, while darker microregions faced smaller cuts or small increases. Figure 4 demonstrates the underlying variation driving differences in the region-level tariff changes by comparing the weights, β_{ri}, for the microregion with the most negative region-level tariff change, Rio de Janeiro, to those in the microregion with the most positive region-level tariff change, Traipu in the state of Alagoas. The industries on the x-axis are sorted from the most negative to most positive tariff

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

[23] One can think of this as a reduced-form analysis in which tariff changes act as exogenous instruments for price changes faced by producers.

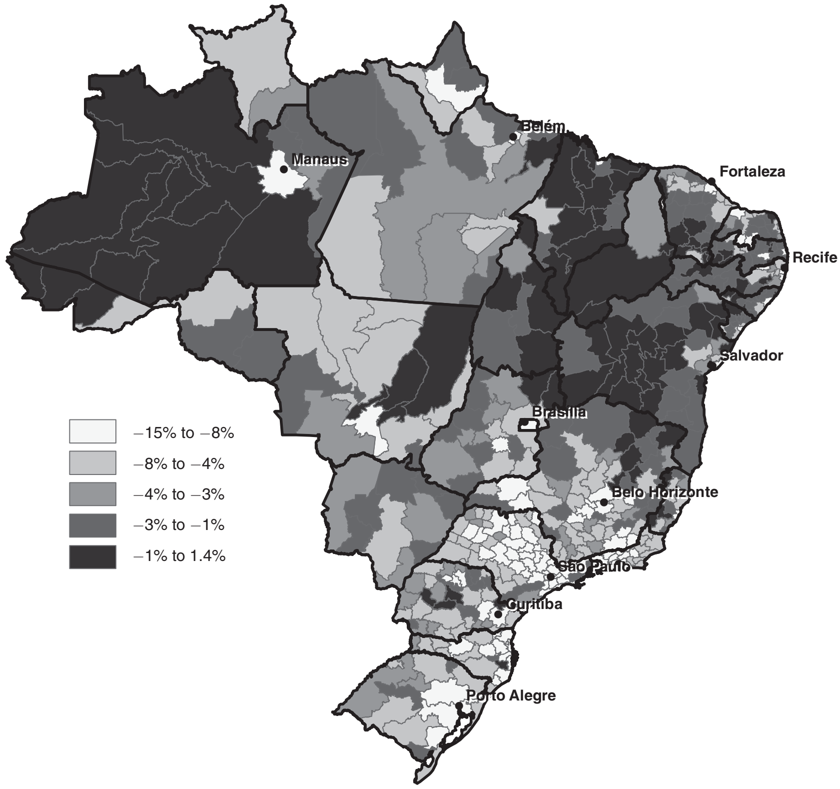

Figure 3. Region-Level Tariff Changes

Notes: Weighted average of tariff changes. See text for details.

change. Rio de Janeiro has more weight in the left side of the diagram, particularly in the apparel and food processing industries. Traipu produces agricultural goods almost exclusively, which faced the most positive tariff changes. Thus, although all regions faced the same set of tariff changes across industries, variation in the weight applied to those industries in each region generates the substantial variation seen in Figure 3.

C. *Wage-Tariff Relationship*

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

$$d\ln(w_r) = \zeta_0 + \zeta_1 RTC_r + \epsilon_r, \tag{7}$$

where $d\ln(w_r)$ is the regional wage change described in Section IVA. Since these wage changes are estimates, I weight the regression by the inverse of the standard error of the estimates based on Haisken-DeNew and Schmidt (1997). ζ_1 captures the regional effect of liberalization on real wages between 1991 and 2000. The model predicts

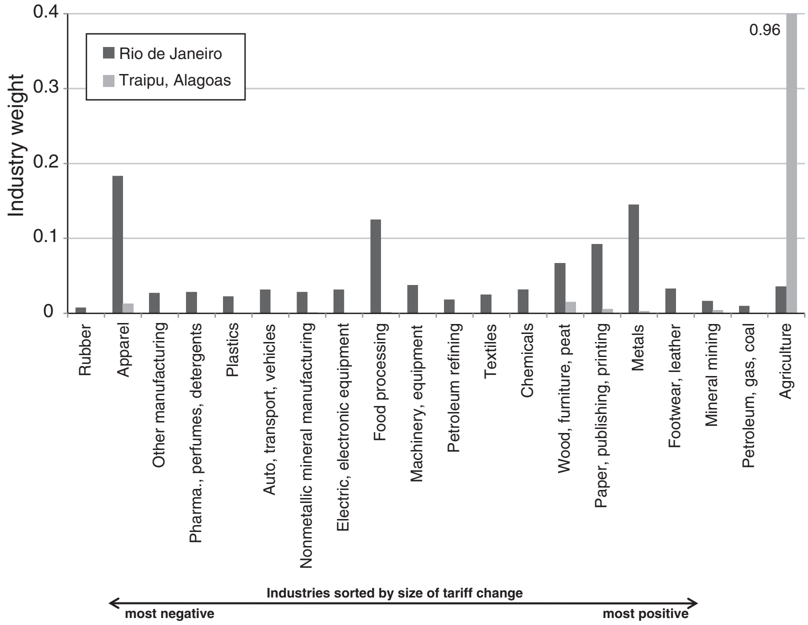

Figure 4. Demonstration of Variation behind Region-Level Tariff Changes

Notes: Author's calculations. See text for details.

that $\zeta_1 = 1$. However, there are a few reasons to expect an estimate between 0 and 1. Any interregional mobility in response to liberalization will smooth out the regional wage variation that would have been observed on impact.[24] In the extreme case of costless, instant worker mobility, all liberalization-induced wage variation would be immediately arbitraged away by worker migration, and there would be no relationship between region-level tariff changes and regional wage changes, i.e., $\zeta_1 = 0$. Since Brazil's population is quite mobile (interstate migration rates are similar to those in the United States), I expect some equalizing migration over the nine-year period being observed. Also, any imperfect pass-through from tariff changes to price changes will be reflected in lower estimated effects, as will random measurement error in the RTC_r measure. Thus, I expect that $0 < \zeta_1 < 1$. Finally, the error term ϵ_r captures any unobserved drivers of wage change that are unrelated to liberalization.

Table 1 presents the results of regressing regional wage changes on region-level tariff changes under various alternate specifications. Each specification is reported with and without state fixed effects, and all standard errors are clustered at the state level, accounting for remaining covariance in the error terms across microregions in the same state.[25] All specifications omit the city of Manaus, which is a free trade area, unaffected by liberalization. Columns 1 and 2 present the main specification as described above. As expected, the relationship between wage changes and

[24] Online Appendix A shows that in the model increasing labor in a region lowers wages, while decreasing labor raises wages. Note that this feature is specific to the model employed here and could be reversed in other models such as those with agglomeration effects.

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

Table 1—The Effect of Liberalization on Local Wages

	Main		No labor share adjustment		Nontraded price change set to zero		Nontraded sector workers' wages	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Regional tariff change	0.404	0.439	0.409	0.439	2.715	1.965	0.417	0.482
Standard error	(0.502)	(0.146)***	(0.475)	(0.136)***	(1.669)	(0.777)**	(0.497)	(0.140)***
State indicators (27)	—	X	—	X	—	X	—	X
Nontraded sector								
Omitted	X	X	X	X	—	—	X	X
Zero price change	—	—	—	—	X	X	—	—
Labor share adjustment	X	X	—	—	X	X	X	X
R^2	0.034	0.707	0.040	0.711	0.112	0.710	0.037	0.763

Notes: 493 microregion observations (Manaus omitted). Standard errors adjusted for 27 state clusters (in parentheses). Weighted by the inverse of the squared standard error of the estimated change in log microregion wage, calculated using the procedure in Haisken-DeNew, and Schmidt (1997).

***Significant at the 1 percent level.

**Significant at the 5 percent level.

*Significant at the 10 percent level.

region-level tariff changes is positive. This implies that microregions facing the largest tariff declines experienced slower wage growth than regions facing smaller tariff cuts, as predicted by the model. The estimate in column 1 of 0.404 implies that a region facing a 10 percentage point larger liberalization-induced price decline experienced a 4 percentage point larger wage decline (or smaller wage increase) relative to other regions. The difference between the region-level tariff change in regions at the 5th and 95th percentile was 12.8 percentage points. Evaluated using the column 1 estimate, a region at the 5th percentile experienced a 5.2 percentage point larger wage decline (or smaller wage increase) than a region at the 95th percentile. The addition of state fixed effects in column 2 has almost no effect on the point estimate but absorbs residual variance such that the estimate is now statistically significantly different from zero at the 1 percent level.

The remaining columns of Table 1 examine the effects of deviations from the preferred specification in columns 1 and 2. Columns 3 and 4 omit the labor share adjustment, which in the context of the model is equivalent to assuming that the labor demand elasticities are identical across industries so that the weights in each region are determined only by the industrial distribution of workers. All of the papers in the previous literature follow this approach. In the Brazilian context, the omission of the labor share adjustment has very little effect on the estimates, as they have little effect on the weights across industries. Taking a region $\times$ industry pair as an observation, the correlation between the weights with and without labor share adjustment is 0.996.

Columns 5 and 6 include the nontraded sector in the regional tariff change calculations, setting the nontraded price change to zero. Footnote 8 lists papers using this approach.[26] This change results in a substantial increase in the point estimates,

[26] The previous literature does not explicitly make assumptions about the price of nontraded goods but rather includes a zero term for the nontraded sector in the weighted averages used in their empirical analyses. In the context of the present model, that is equivalent to assuming zero price change for nontraded goods. However, many

which is precisely in line with the theoretical discussion in Section IB. Setting the nontraded price change to zero understates the size of liberalization's effect on regional labor demand, resulting in an offsetting increase in the estimate of liberalization's effect on regional wages. The relative scale of the estimates with and without the nontraded sector is also of a sensible order of magnitude. Assuming identical Cobb-Douglas technology across industries and constant nontraded sector share of employment (λ_{rN}) across regions, setting the nontraded sector term to zero would inflate the regression coefficient by a factor $\frac{1}{1-\lambda_N}$. Evaluating this at the mean of λ_{rN}, 0.68 (weighted as in Table 1), the coefficients in columns 5 and 6 would be 3.2 times larger than the corresponding coefficients in columns 1 and 2 under the stated restrictions. The actual estimates are 6.7 and 4.5 times larger, respectively, with the difference from 3.2 likely accounted for by the fact that λ_{rN} varies substantially across regions (its standard deviation is 0.13).

One of the benefits of deriving the estimating equation (7) from the theoretical model in Section I is that the model predicts both the sign and magnitude of the coefficient ζ_1. As just mentioned, the coefficient should fall between 0 and 1 depending on the amount of equalizing interregional migration and pass through from tariff changes to price changes faced by producers. Consistent with this prediction, the coefficients in columns 1–4 of Table 1 fall between 0 and 1. In contrast, the coefficients in columns 5 and 6 are much larger than 1, reflecting the upward bias resulting from setting the nontraded sector price change to zero (although doing so also decreases the precision of the estimates such that they are not statistically significantly greater than one).

Columns 7 and 8 of Table 1 present additional evidence that the labor market in the nontraded sector is closely tied to the tariff cuts facing the traded goods sector.[27] These regressions use an alternate version of the dependent variable, which measures the regional change in wages only for workers in the nontraded sector. The calculations generating the independent variable are identical to those used in the main analysis in columns 1 and 2. The model suggests that nontraded sector workers' wages should move with those of traded sector workers even tough nontraded sector workers face the effects of liberalization only indirectly. The results in columns 7 and 8 are very similar to the results for all sectors' wages in columns 1 and 2, further supporting the model's implication that liberalization similarly affects the traded and nontraded sectors.[28]

V. Conclusion

This article develops a specific-factors model of regional economies addressing the local labor market effects of national price changes and applies the model's

previous papers examine outcomes other than wages, about which the present model has little to say. It is quite possible that models allowing for intraregional inequality, for example, would provide a justification for including a zero term for the nontraded sector.

[27] Thanks to an anonymous referee for suggesting this test.

[28] As an additional robustness test, I examined the relationship between regional wage changes in the pre-liberalization intercensal period, 1980–1991, and the regional tariff change during liberalization as in columns 1 and 2 of Table 1 to see whether the main results reflected a change in the pattern of wage growth across regions or whether they were merely part of an ongoing trend. The results, available upon request, show a *negative* relationship, indicating that wages reversed their direction following liberalization. As in footnote 20 this finding suggests that if anything the results in Table 1 are biased downward, against finding the positive estimates shown there.

predictions in measuring the effects of Brazil's trade liberalization on regional wages. The model predicts that local labor markets whose workers are concentrated in industries facing the largest tariff cuts will be more negatively affected, while markets facing smaller cuts will be more positively affected. This relationship takes the form of a weighted average of liberalization-induced price changes across industries, where the weights depend on the industrial distribution of workers in each region and labor demand elasticities in each industry. The empirical findings confirm this prediction. Regions whose output faced a 10 percent larger liberalization-induced price decline experienced a 4 percent larger wage decline than other regions.

The model's weighted-average result provides a theoretical foundation for the similar empirical approach used in an influential literature on the regional effects of trade liberalization, helping clarify the mechanisms through which liberalization affects labor market outcomes. It also provides guidance on how to construct the regional measure of liberalization, particularly regarding the inclusion or exclusion of the nontraded sector from the weighted average. The theoretical results support omitting the nontraded sector from the weighted average and suggest that including a zero term reflecting the lack of tariff change for nontraded goods will inflate the magnitude of liberalization's measured effects on wages. The empirical results confirm this prediction.

Given these results, it seems likely that liberalization has differential 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. (2012) 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. Such a framework would combine the interregional inequality effects of the present study with the inequality across skill groups studied in much of the prior literature on liberalization's effects on inequality.

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