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Trade Reform and Regional Dynamics

Evidence from 25 Years of Brazilian Matched Employer-Employee Data

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Abstract

This paper empirically studies the dynamics of labor market adjustment following the Brazilian trade reform of the 1990s. The paper uses variation in industry-specific tariff cuts interacted with initial regional industry mix to measure trade-induced local labor demand shocks and examines regional and individual labor market responses to those one-time shocks over two decades. Contrary to conventional wisdom, the analysis does not find that the impact of local shocks is dissipated over time through wage-equalizing migration. Instead, it finds steadily growing effects of local shocks on regional formal sector wages and employment for 20 years. This finding can be rationalized in a simple equilibrium model with two complementary factors of production, labor and industry-specific factors such as capital, that adjust slowly and imperfectly to shocks. Next, the paper documents rich margins of adjustment induced by the trade reform at the regional and individual levels. Workers initially employed in harder hit regions face continuously deteriorating formal labor market outcomes relative to workers employed in less affected regions, and this gap persists even 20 years after the beginning of trade liberalization. Negative local trade shocks induce workers to shift out of the formal tradable sector and into the formal nontradable sector. Non-employment strongly increases in harder hit regions in the medium run, but in the longer run, non-employed workers eventually find re-employment in the informal sector. Working age population does not react to these local shocks, but formal sector net migration does, consistent with the relative decline of the formal sector and growth of the informal sector in adversely affected regions.

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Trade Reform and Regional Dynamics: Evidence From 25 Years of Brazilian Matched Employer-Employee Data*

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1 Introduction

The reallocation of resources across economic activities is a key mechanism through which increasing openness leads to welfare gains. Prominent theories of international trade rely on the reallocation of factors across sectors or firms in order to generate production gains from trade. However, academic economists have traditionally paid little attention to the adjustment process, instead focusing on long-run models where reallocation is achieved without frictions. This focus has created a tension between academic economists advocating trade liberalization and policy makers concerned with the labor market outcomes of workers employed in contracting sectors or firms (Salem and Benedetto 2013, Hollweg, Lederman, Rojas and Ruppert Bulmer 2014).

Even though many countries underwent major trade liberalization episodes throughout the 1980s and 1990s (e.g., Brazil, Mexico, and India, among others), we still know very little about the medium- to long-run consequences of these policy reforms on labor markets, particularly for workers who were active when these reforms were implemented. Indeed, empirical studies of the labor market effects of trade liberalization have typically emphasized short- to medium-run effects. Frequently changing designs of cross-sectional household surveys forced researchers to focus on relatively short intervals to guarantee consistency over the periods analyzed (Goldberg and Pavcnik 2007). Furthermore, the lack of long and comprehensive panel data during periods of trade liberalization made it impossible to jointly study the short-, medium- and long-run effects of trade liberalization on individual workers' employment trajectories.

We fill this gap in the literature by studying the empirical dynamics of labor market adjustment in response to a major trade liberalization episode. We use 25 years of matched employer-employee data from Brazil to study the dynamics of local labor market adjustment following the country's trade liberalization in the early 1990s. We exploit variation in the degree of tariff declines across industries and variation in the industry mix of local employment across Brazilian regions to measure changes in local labor demand induced by liberalization. This approach, along with our detailed longitudinal data, allows us to observe labor market dynamics for 20 years following the beginning of the trade policy changes.

The results are striking. We find very large and slowly increasing regional effects on earnings, employment, and other labor market outcomes. Workers whose regions face larger tariff declines experience deteriorating formal labor market outcomes compared to workers in other regions (we discuss informal outcomes below). These effects grow steadily for more than a decade before beginning to level off in the late 2000s. This pattern is not driven by other post-liberalization shocks and is robust to alternative measurement strategies. These findings challenge the conventional wisdom that labor mobility would gradually arbitrage away spatial differences in local labor market outcomes, leading to *declining* rather than *increasing* observed regional effects of liberalization over time (Blanchard and Katz 1992, Bound and Holzer 2000). This suggests substantial barriers to inter-regional mobility, but it also implies that local labor demand in more adversely affected regions keeps falling relative to the rest of the country for years following the end of the liberalization. We provide a simple explanation for the slow growth of the regional effects based on geographical mobility frictions for both labor and other complementary factors of production such as capital, with these frictions mutually reinforcing each other to drive slow adjustment (Dix-Carneiro 2014). Agglomeration economies as in Kline and Moretti (2014) are similar in spirit and consistent with the observed patterns of adjustment, as both mechanisms generate dynamics in local labor demand that persist for long periods of time following a one-time shock. We focus our theoretical interpretation on sluggish factor adjustment because additional evidence on the regional number of establishments, establishment entry and exit rates, and job creation and destruction rates is supportive of the sluggish factor adjustment hypothesis. Furthermore, while both explanations can generate the qualitative patterns we observe, these patterns are quantitatively consistent with the slow adjustment analysis in Dix-Carneiro (2014).

The longitudinal nature of our data allows us to track individual workers over time, observing outcomes for two otherwise identical workers who just before liberalization lived in regions that would subsequently face different local trade shocks. Menezes-Filho and Muendler (2011) pioneered the use of administrative panel data to study trade-induced labor reallocation.[1] Our paper is unique in focusing on the *dynamics* of labor market effects of liberalization, documenting how these effects vary over the medium- to long-run. We find that workers whose initial region faced a larger tariff decline become less and less likely to be formally employed over time, and lose substantial amounts of formal earnings in the years following liberalization. We also observe worker adjustment in the face of negative labor demand shocks. Formal tradable sector workers facing more negative local shocks are more likely to transition into formal nontradable sector employment, but on average cannot offset lost tradable sector employment or earnings.[2]

Using supplementary data from the Decennial Census, we find that regions facing larger tariff declines experience relative increases in informal employment. Informal sector jobs do not provide legally mandated labor protections or other benefits, and often involve lower compensation, fewer opportunities for training and advancement, and generally less favorable working conditions (Goldberg and Pavcnik 2007, Bacchetta, Ernst and Bustamante 2009). In harder-hit regions, non-employment also strongly increases in the medium run, but in the longer run non-employed individuals eventually find employment in the informal sector. These results seem to contradict prior work, which typically finds minimal effects of trade on informality in Brazil (Goldberg and Pavcnik 2003, Menezes-Filho and Muendler 2011, Bosch, Goñi-Pacchioni and Maloney 2012). However, our results can be reconciled based on differences in sectoral coverage, short-run vs. medium- and long-run transitions, and different data structures. In contrast to the prior work on Brazil,

[1] See Krishna, Poole and Senses (2014) for another more recent example.

[2] Menezes-Filho and Muendler (2011) also find evidence for transitions into non-manufacturing employment using industry variation in tariff cuts and using short panels from the *Pesquisa Mensal de Emprego* (PME).

McCaig and Pavcnik (2014) find substantial shifts from household (informal) to enterprise (formal) employment in Vietnam in response to the U.S.-Vietnam Bilateral Trade Agreement, more closely paralleling our findings here.

Overall, geographic migration does not appear to respond to changing local labor demand conditions. However, when restricting attention to formally employed workers in the matched employer-employee data, we find that individuals who migrate tend to avoid adversely affected destinations. Together, these findings suggest that regional adjustment of formal employment occurs primarily through workers transitions into or out of formal employment, rather than by migrating across space.

In contrast to much of the trade literature, we focus on regional dynamics rather than frictions across industries or establishments. A recent literature on the local labor market effects of trade, pioneered by Topalova (2007) in the developing country context and Autor, Dorn and Hanson (2013) in the U.S., points toward significant mobility frictions across regions in both developing and developed countries.[3] When industries are concentrated in different regions, workers must overcome geographical mobility barriers in order to reallocate from contracting to expanding industries or establishments. This reallocation is essential for the economy to realize the production gains from trade. The distinction between geographic and industry mobility frictions is also important for policy, as the optimal policy prescriptions to minimize transitional costs, speed up reallocation, and compensate the losers from trade liberalization will be quite different.

Only recently have researchers begun measuring reallocation costs and the dynamics of labor market adjustment following trade policy reforms. Dix-Carneiro (2014) estimates a structural dynamic equilibrium model of the Brazilian labor market in which workers face various frictions in switching sectors of employment. He then simulates a counterfactual trade liberalization episode to study the quantitative implications of the model, including the dynamics of labor market transition and heterogeneous welfare effects on workers with different characteristics (such as age and education). Other papers follow a similar strategy of calibrating or estimating small open economy models in order to study their quantitative implications for welfare and their implied transitional dynamics when facing hypothetical changes in trade policy (see Kambourov (2009), Artuç, Chaudhuri and McLaren (2010), and Coşar (2013)). Importantly, the labor market dynamics studied in these papers derive from the estimated structural models under consideration rather than reflecting observed responses to liberalization in the data. Our paper is the first to empirically describe transitional dynamics induced by a real-world trade reform.

In related work, Autor, Dorn, Hanson and Song (forthcoming) use U.S. panel data to study the labor market effects of increased Chinese imports across industries. Our individual-level longitu-

[3]This growing literature includes Edmonds, Pavcnik and Topalova (2010), Hasan, Mitra and Ural (2006), Hasan, Mitra, Ranjan and Ahsan (2012), McCaig (2011), Topalova (2010), Topalova (2010), Kovak (2013), Hakobyan and McLaren (2012), Autor et al. (2013), Kondo (2014), Costa, Garred and Pessoa (2014) and others. All of these papers point to the presence of substantial barriers to regional mobility.

dinal analysis is similar in spirit, but differs in a number of ways. Most importantly, we study a discrete policy shock rather than a continuously evolving phenomenon like Chinese export growth. This allows us to study the subsequent dynamics following liberalization without the confounding influence that occurs with a continually evolving shock. We focus on regional rather than industry shocks and study formal sector migration and population responses to regional shocks. Finally, we examine various margins of adjustment, including shifts into informal employment, nonemployment, and transitions between tradable and nontradable sector employment, all of which are salient features of the Brazilian context.

As an additional contribution, we examine administrative employer-employee data alongside more commonly available household survey data from the Brazilian Decennial Census of Population. When possible, we corroborate the findings across the two datasets, and find quite consistent results. Given the growing popularity of matched employer-employee data, it is encouraging that results from more traditional cross-sectional data sources are similar.

To summarize, our paper makes contributions to five strands of the international, labor and development economics literatures. First, this is the first paper empirically describing the transitional labor market dynamics that arise in response to a major trade liberalization episode. Second, we complement a growing literature on the local labor market effects of trade by studying the dynamics of these effects and putting existing static results into context. Third, our findings challenge the conventional wisdom in the labor economics literature that the impact of local shocks will be gradually dissipated through equalizing migration. We provide a theoretical interpretation for these findings and present additional empirical evidence in support of that interpretation. Fourth, we are the first to document the medium- to long-run effects of a major trade policy reform by following workers over time and across sectors and regions. Fifth, we find novel results regarding the effect of trade policy changes on informality in Brazil and reconcile our findings with those of the existing literature.

Our paper proceeds as follows. Section 2 describes the history and institutional context of Brazil's early 1990s trade liberalization. Section 3 describes the data sources, including the matched employer-employee panel and cross-sectional data that we utilize, and describes our definition of local labor markets. Section 4 introduces a model of local labor markets with slow factor adjustment and presents our empirical approach, including a theoretically motivated regional measure of trade liberalization. Section 5 presents the empirical findings for our regional analysis, and Section 6 presents the empirical results in the individual analysis. Section 7 examines outcomes relating to the informal sector, and Section 8 concludes.

2 Trade Liberalization in Brazil

Brazil's trade liberalization in the early 1990s provides an excellent setting in which to study the labor market effects of changes in trade policy. The unilateral trade liberalization involved very large declines in average trade barriers and featured substantial variation in tariff cuts across industries. Many papers have examined the labor market effects of trade liberalization in the Brazilian context to take advantage of this variation.[4]

In the late 1980s and early 1990s, Brazil ended nearly one hundred years with extremely high trade barriers imposed as part of an import substituting industrialization policy.[5] In 1987, nominal tariffs were very high, but the degree of protection actually experienced by a given industry often deviated substantially from the nominal tariff rate due to i) a variety of non-tariff barriers such as suspended import licenses for many goods and ii) a system of "special customs regimes" that lowered or removed tariffs for many transactions (Kume, Piani and de Souza 2003).[6] In 1988 and 1989, in an effort to increase transparency in trade policy, the government reduced tariff redundancy by cutting nominal tariffs and eliminating certain special regimes and trade-related taxes, but there was no effect on the level of protection faced by Brazilian producers (Kume 1990).

Liberalization began in March 1990, when the newly elected administration of President Collor suddenly and unexpectedly abolished the list of suspended import licenses and removed nearly all of the remaining special customs regimes (Kume et al. 2003). These policies were replaced by a set of import tariffs providing the same protective structure, as measured by the gap between prices internal and external to Brazil, in a process known as tariffication (*tarificação*) (de Carvalho, Jr. 1992). In some industries, this process required modest tariff increases to account for the lost protection from abolishing import bans.[7] Although these changes did not substantially affect the protective structure, they left tariffs as the main instrument of trade policy, such that tariff levels in 1990 and later provide an accurate measure of protection.

The main phase of trade liberalization occurred between 1990 and 1995, with a gradual reduction in import tariffs culminating with the introduction of Mercosur. Tariffs fell from an average of 30.5 percent to 12.8 percent, and remained relatively stable thereafter.[8] Along with this large average

[4] Examples include Arbache, Dickerson and Green (2004), Goldberg and Pavcnik (2003), Gonzaga, Filho and Terra (2006), Kovak (2013), Krishna et al. (2014), Menezes-Filho and Muendler (2011), Pavcnik, Blom, Goldberg and Schady (2004), Paz (2014), Schor (2004), and Soares and Hirata (2014) among many others.

[5] Although Brazil was a founding signatory of the General Agreement on Tariffs and Trade (GATT) in 1947, it maintained high trade barriers through an exemption in Article XVIII Section B, granted to developing countries facing balance of payments problems (Abreu 2004). Hence trade policy changes during the period under study were unilateral.

[6] These policies were imposed quite extensively. In January 1987, 38 percent of individual tariff lines were subject to suspended import licenses, which effectively banned imports of the goods in question (Authors' calculations from *Bulletin International des Douanes* no 6 v11 supplement 2). In 1987, 74 percent of imports were subject to a special customs regime (de Carvalho, Jr. 1992).

[7] Figure A1 in Appendix A shows the time series of tariffs. Note the tariff increases in 1990 for the auto and electronic equipment industries.

[8] Simple averages of tariff rates across *Nivel 50* industries, as reported in (Kume et al. 2003).

decline came substantial heterogeneity in tariff cuts across industries, with some industries such as agriculture and mining facing small tariff changes, and others such as apparel and rubber facing declines of more than 30 percentage points. In this paper, we measure liberalization using long-differences in tariffs during the period of liberalization, from 1990 to 1995. Specifically, we use changes in the log of one plus the tariff rate, shown in Figure 1. During this time period, tariffs accurately measure the degree of protection faced by Brazilian producers and reflect the full extent of liberalization faced by each industry. We do not rely on the timing of tariff cuts between 1990 and 1995 because this timing was chosen to maintain support for the liberalization plan, cutting tariffs on intermediate inputs earlier and consumer goods later (Kume et al. 2003).

As discussed below, along with regional differences in industry mix, the cross-industry variation in tariff cuts provides the identifying variation in our analysis. Following the argument in Goldberg and Pavcnik (2005), we note that the tariff cuts were nearly perfectly correlated with the pre-liberalization tariff levels (correlation coefficient = -0.90). These initial tariff levels reflected a protective structure initially imposed in 1957 (Kume et al. 2003), decades before liberalization. This feature left little scope for political economy concerns that might otherwise have driven systematic endogeneity of tariff cuts to counterfactual industry performance. To check for any remaining spurious correlation between tariff cuts and other steadily evolving industry factors, we regress pre-liberalization (1980-1991) changes in industry employment and wage premia on the 1990-1995 tariff changes, with detailed results reported in Appendix A.2. We attempted a variety of alternative specifications and emphasize that the results should be interpreted with care, as they include only 20 tradable industry observations. Most specifications exhibit no statistically significant relationship, but heteroskedasticity-weighted specifications place heavy weight on agriculture and find a negative relationship. Agriculture was initially the least protected industry, and it experienced approximately no tariff change. It also had declining wages and employment before liberalization, driving the negative relationship. Consistent with earlier work, when omitting agriculture, tariff-cuts are unrelated to pre-liberalization earnings trends (Krishna, Poole and Senses 2011). Given these varying results, we include controls for pre-liberalization outcome trends in all of the analyses presented below to account for any potential spurious correlation. Consistent with the notion that the tariff changes were exogenous in practice, these pre-trend controls have little influence on the vast majority of our results.

3 Data

In this paper, we use two main data sources, the *Relação Annual de Informações Sociais* (RAIS), and the Decennial Census of Population. RAIS is a matched employer-employee dataset assembled by the Brazilian Ministry of Labor every year since 1976 and provides a high quality census of the Brazilian formal labor market (De Negri, de Castro, de Souza and Arbache 2001, Saboia and

Tolipan 1985). We utilize RAIS data spanning the period from 1986 to 2010. The Census is a traditional household survey covering the entire population, including informally employed and non-employed workers. We use Census data from 1970-2010.

Originally, RAIS was created as an operational tool for the Brazilian government to i) monitor the entry of foreign workers into the labor market; ii) oversee the records of the FGTS program (a national benefits program consisting of employers' contributions to each of its employees); iii) provide information for administering several government benefits programs such as unemployment insurance; and iv) generate statistics regarding the formal labor market. Today it is the main tool used by the government to enable the payment of the "*abono salarial*" to eligible workers. This is a government program that pays one additional minimum wage at the end of the year to workers whose average monthly wage was not greater than two times the minimum wage, and whose job information was correctly declared in RAIS, among other minor requirements. Thus, workers have an incentive to ensure that their employer is filing the required information. Moreover, firms are required to file, and face fines until they do so. RAIS includes all formally employed workers, meaning those with a signed work card, providing them access to all of the benefits and labor protections afforded them by the legal employment system. It omits those without signed work cards, including interns, the self-employed, elected officials, domestic workers, and other minor employment categories. These data have recently been employed by Menezes-Filho and Muendler (2011), Helpman, Itskhoki, Muendler and Redding (2014), Lopes de Melo (2013), Krishna et al. (2014), and Dix-Carneiro (2014), though these papers utilize a shorter panel.

The data consist of job records identified by both a worker ID number (PIS) and an establishment registration number (CNPJ). These identifiers are unique and do not change over time, allowing us to track workers over time and across establishments. Establishment-level information includes geographic location (municipality), industry (IBGE subsector[9]), and worker-level information includes gender, age, education (9 categories), December earnings, average monthly earnings, tenure, occupation, month of accession into the job (if accession occurred during the current year) and month of separation (if any). Throughout the analysis, we limit our sample to include working-age individuals, aged 18-64, and unless otherwise noted, focus on labor market outcomes reported in December of each year.[10]

These data have various advantages relative to previous work on the effects of trade on local labor markets. First, relative to Kovak (2013) and Autor et al. (2013), we can analyze the dynamics of adjustment to the trade liberalization shock, as RAIS data are available every year and because RAIS is representative at fine geographic levels by including the universe of formally employed

[9]The IBGE subsector classification includes 12 manufacturing industries, 2 primary industries, 11 nontradable industries, and 1 other/ignored.

[10]In the regional analysis, we also omit individuals working in public administration and those reporting other/ignored sectors. We impose additional age restrictions and other sample restrictions in the individual-level analysis, described in Section 6.

workers in every Brazilian municipality.[11] The dynamic patterns we document below would be unobservable using standard household survey or Census data. Second, a very rich set of labor market outcomes can be analyzed with such data, including how liberalization affected i) the duration of non-formal labor market spells; ii) job creation and job destruction rates; iii) the number of active establishments; and iv) the establishment size distribution. Third, the ability to follow workers over time and across establishments and municipalities allows us to analyze the short-, medium- and long-run effects of the reform on individuals' labor market trajectories controlling for observable worker characteristics, including individuals' industry and region of employment just before the policy shock. Fourth, we study a discrete policy shock and observe outcomes for 20 years following the beginning of liberalization, allowing us to study the dynamic response to this well-defined trade policy shock. This contrasts with Autor et al. (forthcoming), who use U.S. panel data to study the effects of growing trade with China. They emphasize that the continuously evolving nature of Chinese trade confounds their ability to study the dynamic response to a trade shock at any given point in time.

As is typically the case in matched employer-employee datasets, the limitation of RAIS is a lack of information on workers who are not formally employed. It is therefore impossible to tell whether a worker is out of the labor force, unemployed, informally employed, or self-employed. This is important in the Brazilian context, with informality rates reaching over 50% of all employed workers during the period under scrutiny.[12] However, we can infer when the worker is not employed in the formal labor market and examine spells outside of formal employment.

To address this limitation, we supplement the RAIS data with five rounds of the Brazilian Demographic Census: 1970, 1980, 1991, 2000, and 2010. While these data provide much smaller samples and do not permit following individual workers over time, they cover all individuals, including the informally employed, unemployed, and those outside the labor force. This allows us to examine trade liberalization's effect on informality, employment, and other outcomes for workers outside formal employment. Moreover, with the increasing availability and popularity of matched employer-employee data, it is useful to compare the empirical relationships in these types of data with those in more traditional cross-sectional surveys. When possible, we corroborate results from RAIS using the Demographic Census, finding very similar results across datasets.

To analyze outcomes by local labor market, we must define the boundaries of each market. We use the "microregion" definition of the Brazilian Statistical Agency (IBGE), which groups together economically integrated contiguous municipalities (counties) with similar geographic and productive characteristics (IBGE 2002), closely paralleling an intuitive notion of a local labor market. When necessary, we combine microregions whose boundaries changed during our sample period to ensure

[11] The National Household Survey (*Pesquisa Nacional por Amostra de Domicílios* - PNAD) would be a natural alternative data source for a yearly analysis, but it only provides geographic information at the state level, does not allow one to follow individual workers over time, and provides a much smaller sample.

[12] Authors' calculations using Brazilian Demographic Census.

that we consistently define local labor markets from 1980-2010.[13] This process leads to a set of 475 consistently identifiable local labor markets.[14]

4 Empirical Framework

4.1 Model of Local Labor Markets with Factor Adjustment

Given our focus on the dynamic regional effects of trade liberalization, we develop a specific-factors model of regional economies that allows for imperfect and slow regional factor adjustment in response to changing local conditions. The model yields a tractable measure of liberalization-induced local labor demand shocks that parallels the empirical approach used throughout the literature on the local effects of trade.[15] By allowing for the possibility of imperfect and slow factor adjustment, the model can also accommodate the dynamic evolution of outcomes that we document below.

The national economy consists of many regions, r, each of which may produce goods in many industries, i. Following Jones (1975), each region is endowed with a vector of industry-specific factors, T_{ri}, and a stock of regional labor, L_r that is costlessly mobile across industries. Goods and factor markets are competitive. Production is Cobb-Douglas, and specific-factor shares, φ_i, may vary across industries. Hats represent proportional changes. Producers in all regions face the same vector of national price changes, $\hat{P}_i$. Kovak (2013) studies a similar model in which local factor supplies are fixed. Here we allow the amounts of labor and specific factors to vary in response to liberalization. We solve this variation of the model in Appendix B, yielding the following equilibrium relationship governing the evolution of wages in a region r.

$$\hat{w}_r = \sum_i \beta_{ri}\hat{P}_i - \delta_r \left(\hat{L}_r - \sum_i \lambda_{ri}\hat{T}_{ri} \right), \tag{1}$$

$$\text{where} \quad \beta_{ri} \equiv \frac{\lambda_{ri}\frac{1}{\varphi_i}}{\sum_j \lambda_{rj}\frac{1}{\varphi_j}}, \qquad \delta_r \equiv \frac{1}{\sum_k \lambda_{rk}\frac{1}{\varphi_k}},$$

[13] This geographic classification is a slightly aggregated version of the one in Kovak (2013), accounting for additional boundary changes during the longer sample period. Related papers define local markets based on commuting patterns (e.g. Autor et al. (2013)). Our local market definition performs well based on this standard as well - only 3.4 and 4.6 percent of individuals lived and worked in different markets in 2000 and 2010, respectively.

[14] The regional definition is shown in Figure 3. The analysis omits 11 microregions, shown with a cross-hatched pattern the figure. These include i) Manaus, which was part of a Free Trade Area and hence not subject to tariff cuts during liberalization, ii) the microregions that constitute the state of Tocantins, which was created in 1988 and hence not consistently identifiable throughout our sample period, and iii) a few other municipalities that are omitted from RAIS in the 1980s. The inclusion or exclusion of these regions when possible has no substantive effect on the results. We also implemented the main analyses using a more aggregate local labor market definition, "mesoregions" defined by IBGE, and results are nearly identical.

[15] See footnote 3 for examples.

and λ_{ri} is the share of regional labor initially allocated to tradable industry i. Although we do not explicitly model the nontradable sector, we follow Kovak (2013) by omitting nontradables from the sums in (1), based on the idea that nontradables prices move with tradables prices, such that dropping them closely approximates the ideal measure. Also, while (1) measures proportional changes in nominal regional wages, if the nontradable goods' share of consumption is constant across regions, and trade balances at the regional level, then local real nominal wage changes are given by a linear transformation of nominal wage changes.[16] In this case, although real wage changes are smaller than nominal wage changes, the explanatory power of regional tariff shocks is unchanged. Real versus nominal wage effects aside, note that our findings below point toward large effects on real local labor market outcomes such as formal and informal employment, unemployment, and spells out of the formal sector.

We apply this model to the formal sector, in which workers have access to legally mandated rights and benefits. Labor supply ($\hat{L}_r$) can therefore adjust either through interregional migration or through shifts out of formal employment. In the Brazilian context, informal employment tends to be an absorbing state, as the costs to enter formal employment are very large (Dix-Carneiro 2014). We think of specific factor reallocation ($\hat{T}_{ri}$) as reflecting capital depreciation in one region and new investment in another region, which occurs slowly over time. We further assume that if an industry was not active in a region at the onset of trade liberalization, fixed costs of building a new industry from scratch are high enough that the industry will not emerge following liberalization.

To fix ideas, define the "regional price change" as $\hat{P}_r \equiv \sum_i \beta_{ri}\hat{P}_i$, and imagine estimating the following reduced-form regression, using data generated by the model described in (1).

$$\hat{w}_r = \alpha + \theta\hat{P}_r + \nu_r, \tag{2}$$

This specification parallels the approach in the literature on the local effects of trade and is a simplified version of the regional estimation strategy we use below in Section 5. Changing factor supplies appearing in the rightmost term in (1) are omitted from (2), and hence are incorporated into the error term. Since (1) represents an equilibrium relationship where changes in factor quantities $\hat{L}_r$ and $\hat{T}_{ri}$ are themselves functions of changes in factor payments, these supply shifts are endogenous.[17] Thus, the observed wage effect, captured in the reduced-form OLS estimate of θ, depends upon the correlation between changing factors supplies in ν_r and the regional price change, $\hat{P}_r$.

With factor supplies fixed ($\hat{L}_r = \hat{T}_{ri} = 0$), there will be no such correlation, and we would expect the estimate of θ to be constant over time and close to one. In this case, the wage change

[16] With constant nontradable consumption shares across regions, and balanced trade region by region, the change in real wage is simply the change in nominal wage scaled down by the tradable goods' share of consumption, plus a term that is constant across regions and drops out of the analysis. See footnote 13 in Kovak (2011) for details.

[17] In principle, one could specify a functional form for the factor supply process to close the model, but the main insights are unchanged when considering general factor adjustment patterns. We have solved a version of the model with a particular functional form for factor supplies, and results are available upon request.

in (1) equals the weighted average of proportional price changes, with weights determined by the region's industry mix, reflected in β_{ri}. Although all regions face the same set of price changes, regions specializing in goods facing larger price declines experience larger wage declines. This weighted average therefore captures the intuitive idea, initially explored by Topalova (2007), that regions experience larger declines in labor demand when their most important industries face larger liberalization-induced price declines. Thus, with factor supplies held fixed, regions facing larger price declines for their most important products would experience relative wage declines, but the model would not include any mechanism to generate wage or employment dynamics.

Now, continue to hold specific factor quantities fixed ($\hat{T}_{ri} = 0$), but allow the quantity of labor to vary. In this case, the model predicts large wage effects just after liberalization, followed by a period of declining regional wage differences as labor reallocates, arbitraging away wage differences across markets. Specifically, assume that employment falls in regions facing more negative labor demand shocks, such that $\text{cov}(\hat{P}_r, \hat{L}_r) > 0$. Since $\delta_r > 0$, this reallocation partly equalizes the wage changes across regions, driving down the estimated value of θ. Now allow $\hat{L}_r$ to vary over time, always measuring changes relative to a fixed pre-liberalization base year. If reallocation continues over time, then $\text{cov}(\hat{P}_r, \hat{L}_r)$ becomes steadily more positive, and one would observe a steadily declining regional wage effect of liberalization. This pattern captures the conventional wisdom in the labor literature, which focuses on worker mobility's role in arbitraging spatial differences in labor market outcomes (Blanchard and Katz 1992, Bound and Holzer 2000).

In contrast to the declining wage effects predicted by labor adjustment alone, when one allows both labor and specific factor supplies to vary, complex patterns can emerge. In fact, in sharp contrast to the conventional wisdom, the empirical analyses in Sections 5 and 6 find steadily *growing* regional wage and employment effects of liberalization. This pattern can be rationalized by the model if $\text{cov}(\hat{P}_r, \hat{L}_r - \sum_i \lambda_{ri}\hat{T}_{ri})$ declines over time. Changes in this covariance are determined by the relative speed of labor and specific factor adjustment. Consistent with the empirical results below, we restrict attention to situations where both labor and specific factors reallocate toward regions with higher factor returns, such that $\text{cov}(\hat{P}_r, \hat{L}_r) > 0$ and $\text{cov}(\hat{P}_r, \sum_i \lambda_{ri}\hat{T}_{ri}) > 0$.[18] If both labor and specific factors reallocate slowly, their respective covariances with $\hat{P}_r$ will increase over time. As an example, if labor reallocates a bit more quickly at first and capital reallocates more quickly later, then $\text{cov}(\hat{P}_r, \hat{L}_r - \sum_i \lambda_{ri}\hat{T}_{ri})$ declines over time, and the estimated effect of liberalization grows over time. Intuitively, when labor and complementary factors adjust slowly, they create a mutually reinforcing incentive for factors to flow away from less favorably affected regions; declining local employment lowers the return to capital, and declining local capital lowers the return to labor. Dix-Carneiro (2014) Section 7.4 studies a similar mechanism in the context of inter-industry reallocation, showing quantitatively that slow capital adjustment can drive increasing

[18] Appendix B shows that the average specific factor return falls more (increases less) in regions with more negative values of $\hat{P}_r$, such that labor and specific-factors have the incentive to reallocate in the same direction.

industry-specific wage effects over long periods of time.

Slow factor adjustment can therefore rationalize a steady growth in regional wage and employment effects of liberalization. It is worth noting that agglomeration economies in manufacturing are similar in spirit, and could yield similar patterns (see, for example, Ellison and Glaeser (1997), Rosenthal and Strange (2004), Greenstone, Hornbeck and Moretti (2010), and Kline and Moretti (2014)). The important common element of the two mechanisms is that they can both generate dynamics in local labor demand that persist for long periods of time following a one-time shock. We focus our theoretical interpretation on sluggish factor adjustment because, as we discuss below, evidence on the regional number of establishments, establishment entry and exit rates, and job creation and destruction rates is supportive of the sluggish factor adjustment hypothesis, and Dix-Carneiro (2014) shows that it can quantitatively explain the patterns we observe.

4.2 Empirical Approach

Following the model just described, we define the "regional tariff change," or RTC_r, as our empirical measure of liberalization's effect on local labor demand. This measure corresponds to the weighted-average regional price change in (1), where we utilize only the variation in prices that is driven by trade liberalization.

$$RTC_r = \sum_i \beta_{ri} d \ln(1 + \tau_i), \tag{3}$$

where i indexes tradable goods industries. To calculate the β_{ri}, we measure λ_{ri} as industry i's initial share of region r formal employment and φ_i as one minus the initial wagebill share of industry value added in industry i.[19] τ_i is the tariff rate in industry i, and d represents the long difference from 1990-1995, the period of Brazilian trade liberalization.

Because Brazilian local labor markets differ substantially in the industry distribution of their employment, the weights β_{ri} vary across regions. Figure 2 demonstrates how variation in industry mix leads to variation in RTC_r. The figure shows the initial industry distribution of employment for the region with the most negative value, Colatina, the second largest city in Espírito Santo state, and the most positive value, Paranatinga, in central Mato Grosso state. The industries on the x-axis are sorted from the most negative to the most positive tariff change. Colatina has more weight on the left side of the diagram, particularly in the apparel and food processing industries. Paranatinga produces agricultural goods and wood products almost exclusively, both of which faced more positive tariff changes. Thus, although all regions faced the same set of tariff changes across

[19] We calculate formal employment shares using the 1991 Census, as it provides a more detailed industry classification than that available in RAIS. We initially use *formal* shares because of our focus on formal sector outcomes and because workers can easily leave formal employment, but returning appears to be quite difficult (Dix-Carneiro 2014). As we show in section 5.3, results are very similar when using overall employment shares, including both formal and informal employment. We also use overall shares when studying outcomes outside the formal sector in Section 7. The φ_i are calculated using IBGE national accounts data.

industries, variation in the industry distribution of employment in each region generates substantial variation in RTC_r. Appendix A.3 shows similar figures for a variety of other regions throughout the country, showing how differences in industry mix drive geographic variation in RTC_r.

This variation is presented in Figure 3. Regions facing larger tariff cuts are presented as darker and bluer, while regions facing smaller cuts are shown as lighter and yellower. The region at the 10th percentile faced a tariff decline of 13.8 percentage points, while the region at the 90th percentile faced a 4.2 percentage point decline. Hence, in interpreting the regression estimates below, we compare regions whose values of RTC_r differ by 10 percentage points, closely approximating the 90-10 gap of 9.6 percentage points. Note that there is substantial variation in the tariff shocks even among local labor markets within the same state. As we include state fixed effects in our analyses, these within-state differences provide the identifying variation in our study.[20]

Note that our theoretical framework rationalizing the use of RTC_r as a valid measure of trade-induced local labor demand shocks assumes that workers are homogeneous, perfect substitutes in production, and perfectly mobile across industries within regions. These assumptions limit our ability to interpret heterogeneous effects of regional tariff changes on wages and employment of workers with different skills and initially employed in different industries. Therefore, we refrain from reporting heterogeneous effects of RTC_r on the various labor market outcomes studied in this paper and postpone such an investigation to future work extending our framework to accommodate heterogeneous workers and complementarities in production.[21]

In the following two sections, we implement empirical analyses at the regional and individual levels. In both cases, we examine how post-liberalization (1995-2010) labor market outcomes evolved for regions facing larger tariff cuts vs. regions facing smaller tariff cuts. Thus, we essentially use a regional difference-in-differences approach with a continuous treatment given by RTC_r. This approach allows us to generate credible estimates of the causal regional effects of liberalization, but cannot capture overall national effects of liberalization that apply across regions. This feature is common to all studies utilizing cross-industry or cross-region variation for identifying the effects of liberalization.

Our approach parallels prior studies examining the local effects of trade liberalization. However, the RAIS data allow us to calculate changes in regional outcomes in each year following liberalization, so rather than observe liberalization's local effect in one post-shock period, we trace out the dynamic regional response to liberalization as it evolves over time. We also include controls for pre-liberalization trends that might otherwise confound this research design, with little effect on the estimates. The following two sections describe our detailed estimation strategies and findings.

[20] A regression of RTC_r on state fixed effects yields an R^2 of 0.23; i.e. 77% of the variation in RTC_r is not explained by state effects. Our main conclusions are unaffected by the inclusion or exclusion of state fixed effects.

[21] In this case, there will no longer be a unique local labor demand shock, but rather different local labor demand shocks facing heterogeneous workers. Because of complementarities, a local labor demand shock to skilled workers will also affect unskilled workers' outcomes in such a way that local labor demand shocks across skill levels will all be intertwined, as shown in Dix-Carneiro and Kovak (2015).

5 Regional Analysis

5.1 Empirical Specification

Our regional analysis compares the evolution of outcomes in markets facing larger tariff cuts to those in markets facing smaller cuts. For each year t following the beginning of liberalization (1992 to 2010), we estimate an equation of the following form:

$$y_{rt} - y_{r,1991} = \theta_t RTC_r + \alpha_{st} + \gamma_t (y_{r,1990} - y_{r,1986}) + \epsilon_{rt}, \tag{4}$$

where y_{rt} is the value of a regional outcome such as earnings or employment, θ_t is the effect of liberalization on outcomes by year t, α_{st} are state fixed effects, and $(y_{r,1990} - y_{r,1986})$ is a pre-liberalization trend in the outcome variable. While the change in outcome varies with the year t under consideration, the liberalization shock, RTC_r, does not. Instead, it always reflects the regional measure of tariff changes *during liberalization*, from 1990 to 1995.[22] Using this strategy, each year's θ_t estimates one point on an impulse response function describing the local effects of liberalization in each post-liberalization year.

We use 1991 as the base year for outcome changes to make results comparable between RAIS and the decennial census. We include state fixed effects to account for any state-specific policies that might commonly affect outcomes for all regions in the same state, such as state-specific minimum wages, introduced in 2002 (Neri and Moura 2006). We include pre-liberalization trends in outcomes $(y_{r,1990} - y_{r,1986})$ to address the possibility of confounding ongoing trends. For our main outcomes, we present results with and without state fixed effects and pre-trends, with little effect on the coefficients of interest. Since many of our dependent variables are themselves estimates, we weight regressions based on the inverse of their standard error to account for heteroskedasticity. We also cluster standard errors at the mesoregion level to account for potential spatial correlation in outcomes across neighboring regions.

To consistently estimate θ_t, ϵ_{rt} must be uncorrelated with RTC_r, conditional on the state fixed effects and outcome pre-trend. For this identification assumption to be violated, there would need to be an omitted variable that i) drives wage or employment growth across regions within a state and ii) is correlated with RTC_r but iii) is *not* captured by pre-liberalization outcome trends. While such a feature is unlikely to exist, in Section 5.3 we confirm that our results are robust to a variety of potential confounders and alternative specifications.

5.2 Regional Earnings and Employment

We begin by examining the local labor market effects of liberalization on formal sector earnings and employment in each year following liberalization. Because hours information is not available

[22] Recall from Section 2 that tariffs declined between 1990 and 1995, after which they remained relatively stable.

in RAIS before 1995, we focus on monthly earnings rather than hourly wages in the RAIS analysis. We supplement these results with monthly earnings and hourly wages calculated using the Census.

First, we calculate regional earnings premia, adjusted for any compositional changes in the regional workforce. For each year, we regress workers' log December earnings on demographic and educational controls, industry fixed effects, and region fixed effects. These regional fixed effect coefficient estimates then reflect the earnings premium a worker receives in the relevant region, beyond what would be expected given their other observable characteristics. By separately estimating these regressions in each year, we allow for changes in the regional composition of workers and changes in the returns to worker characteristics over time.[23] This approach ensures that our results are not driven by secular growth or decline of industries, changing discrimination, changes in the returns to schooling, or any other changes in the returns to observable characteristics that operate at the national level. Our dependent variable is then the change in log regional earnings premium from 1991 to each subsequent year, 1992 to 2010.

Table 1 shows the results of estimating (4) for regional formal sector earnings and wages. All estimates for the coefficient on RTC_r are positive, indicating that regions facing larger tariff declines experience relative earnings or wage declines. Consider Panel A, which presents earnings estimates using the RAIS data. Columns (1) to (3) examine changes in earnings from 1991 to 2000, while columns (4) to (6) examine changes from 1991 to 2010. Columns (2) and (4) add state fixed effects, and columns (3) and (6) add pre-trend controls for the change in the regional formal earnings premium from 1986 to 1990. The coefficient estimate of 0.358 in column (3) indicates that a region facing a 10 percentage point larger tariff decline (approximately the 90-10 gap in RTC_r) experienced a 3.58 percentage point larger proportional decline (smaller increase) in the regional formal earnings premium.

A striking feature of Table 1 is the sharp increase in liberalization's effect on earnings from 2000 to 2010. This indicates that the divergent earnings growth in regions facing different tariff declines continued well beyond the liberalization period. Figure 4 confirms this pattern by plotting the evolution of θ_t in (4). Each point on the figure corresponds to a regression coefficient like those in Table 1, where the dependent variable is the change in earnings premium from 1991 to the year listed on the x-axis. The vertical dashed line indicates that liberalization was complete in 1995. We present coefficient estimates for 1992-94, but these should be interpreted with care, as liberalization was still ongoing. However, the majority of the tariff cuts were implemented by 1993, so these early coefficients are still informative regarding liberalization's short run effects.[24] The local earnings effects of liberalization steadily grow for more than a decade, before leveling off in the late 2000s.

This increasing pattern is very robust to details of the specification, including trimming earnings outliers, restricting the sample based on worker skill, focusing on tradable or nontradable

[23] Alternatively, we can control for *constant, unobservable* worker heterogeneity by pooling across years and including individual fixed effects. As shown in Section 5.3, the results are very similar.

[24] When regressing RTC_r on an alternate version measuring tariff changes from 1990-93, the R^2 is 0.93.

sectors, and restricting the sample based on regional characteristics. Appendix Figure A5 shows the underlying scatterplots, confirming our choice of linear estimating equation and showing that the results are not driven by outliers. Panels B and C of Table 1 further confirm the increasing pattern for formal earnings and formal hourly wages using the Decennial Census. We restrict the Census sample to formally employed workers (those with a signed work card) for comparability with RAIS. The similarity of the earnings and hourly wage results rules out the possibility of substantial offsetting movement in hours, which are unavailable in RAIS. The negative coefficient estimates on the pre-liberalization formal earnings trend controls suggests that, if anything, the growing effects that we see after liberalization represent a reversal of ongoing trends. Figure 5 confirms this point by plotting coefficient estimates from regressing formal earnings growth from 1986 to each pre-liberalization year on RTC_r. It is clear that the growing effects after liberalization represent a reversal of the existing trend.

Having seen that regions facing larger tariff declines experienced relatively deteriorating earnings and wages, Table 2 examines liberalization's effects on regional log formal employment in 2000 and 2010. As with earnings, the patterns are quite consistent across specifications, though the Census point estimates are somewhat smaller.[25] The estimate of 2.762 in Panel A column (3) shows that a region facing a 10 percentage point larger tariff decline experienced a 31.8 percentage point larger proportional decline (smaller increase) in formal employment from 1991 to 2000. Once again, the effect grows substantially from 2000 to 2010, indicating that employment growth continued to diverge for regions facing different regional tariff changes. The Census estimates in Panel B similarly increase over time, and Figure 6 shows that the increase was quite steady for the first decade following liberalization before leveling off in the mid 2000s. Figure 7 shows that there may have been a very modest positive pre-liberalization trend in employment, but it is statistically insignificant and of a far smaller scale than the large post-liberalization growth in Figure 6.

To better understand how regional employment adjusted to local liberalization shocks, Figure 8 separates overall employment into tradable sector and nontradable sector employment. The effects of liberalization are clearly concentrated in the tradable sector, with large and increasing effects on employment during and after liberalization. In the years just after liberalization, harder hit regions experience small relative increases in nontradable sector employment, as the nontradable sector absorbs a small share of the workers displaced from the tradable sector (we will confirm this interpretation below using panel data). However, this reallocation is short-lived; a decade following liberalization, the local nontradable sector has shrunk compared to that in more favorably affected regions.

[25] The magnitude of the Census estimates should be viewed with caution. The employment questions changed from a yearly definition in 1991 to a weekly definition in 2000, so changes in the level of employment are partly confounded by this change in definition. However, since the definition change applied to all local labor markets, we are still able to examine liberalization's effect on employment, by comparing outcomes *across* regions.

5.3 Alternative Hypotheses and Robustness

The preceding subsection documented large regional effects of liberalization on formal earnings and employment. The scale of these effects grew steadily for more than a decade, indicating that outcomes continued to degrade in harder hit regions relative to more favorably affected regions for many years following the end of liberalization. As discussed in Section 4.1, this pattern is consistent with a model of regional labor markets with imperfect and slow adjustment of both labor and complementary factors such as capital. Here, we rule out a number of alternative explanations for the growing effects of liberalization to validate focusing on adjustment frictions and to demonstrate the robustness of our findings.

The effects of liberalization could appear to grow over time because of correlated shocks occurring after trade liberalization. To explain the smooth growth of the effects in Figures 4 and 6, such shocks would need to affect industries or regions similarly to liberalization and would need to grow steadily over time or occur quite regularly. Although these circumstances are unlikely, we construct controls for a number of salient economic shocks in the post-liberalization period, demonstrating that they are not driving our results.

If tariff changes after 1995 were correlated with those occurring during liberalization (1990-95), they might drive apparently increasing effects of liberalization. We calculate post-liberalization regional tariff changes as in (3), but use tariff changes between 1995 and year $t > 1995$ as additional controls.[26] Table 3 Panel A replicates the earnings analysis in columns (3) and (6) of Table 1, including the post-liberalization tariff change up to year t, along with RTC_r. The post-liberalization control has the expected positive coefficient, but its inclusion has very little effect on the liberalization coefficients.

Another possible explanation for the steadily increasing regional effects of liberalization is slow adjustment in import or export markets. Although trade liberalization was implemented quickly and was complete by 1995, it is possible that importers and exporters responded slowly to trade policy changes, perhaps because of difficulty in forming trading links with firms abroad. If import and export quantities responded slowly, it could explain the slowly evolving labor market effects that we observe. To address this possibility, we follow Autor et al. (2013) (ADH) by constructing changes in imports and exports per worker (in \$100,000 units) for each industry and year, and then form regional weighted averages analogous to (3).[27] Figure 9 plots the coefficients from regressing these ADH-style changes in imports and exports on regional tariff changes. If slow import and export adjustment were driving the slow evolution of liberalization's effect on regional earnings, the export coefficients would exhibit a similar shape to the earnings coefficients in Figure 4, and/or the import coefficients would exhibit their mirror image. This is clearly not the case; between 1995

[26] We construct post-liberalization tariff changes using tariff data from UNCTAD TRAINS, as the Kume et al. (2003) tariff series ends in 1998.

[27] Imports and exports data are from Comtrade.

and 2003, the response of earnings to RTC_r increased by a factor of 11.6, while the ADH export response stayed roughly constant, and the ADH import response declined. The import and export coefficients grow only in the last few years of the sample, precisely when the earnings effects level off, but the export standard errors also increase substantially during 2008-10, such that the point estimates are not significantly different from zero. Table 3 Panel B confirms this point by directly controlling for import and export growth, with minimal effect on the liberalization coefficients. Note that these controls also rule out the potential confounding effects of the commodity price boom of the 2000s (Costa et al. 2014) or correlated tariff cuts in Brazil's export destinations (McCaig 2011). Both of these phenomena would be reflected in a positive correlation with RTC_r that grows over time, yet we find no evidence of this pattern.

Other potential confounders are the Brazilian Real devaluations that occurred in 1999 and 2002. To the extent that industries were differentially affected by the exchange rate movements, the depreciations could induce correlation with tariff changes across industries. We construct industry-specific real exchange rates as trade-weighted averages of real exchange rates between Brazil and its trading partners.[28] We then take the change in log real exchange rate from 1990 to year t for each industry, and calculate regional shocks using weighted averages as in (3). Panel C includes this control in the earnings analysis, again with little effect on the scale or time pattern of the liberalization coefficients. There was also a substantial wave of privatization during our sample period. Beginning in 1995, the RAIS data allow us to identify as state-owned any firm at least partly owned by the government.[29] In Panel D of Table 3 we control for the change in the (regional) share of workers employed in state owned firms between 1995 and year t, again with little effect on the liberalization coefficients.

Along with these controls for post-liberalization shocks, we implemented a wide variety of robustness tests. Table 3 Panel E utilizes the panel nature of RAIS to account for unobservable worker characteristics by controlling for worker fixed effects when estimating regional wage premia. This approach is an alternative to our main specifications, which control for observable characteristics and allow the returns to these characteristics to vary over time. As shown in Panel E, both approaches yield similar results, indicating that our findings are not driven by workers sorting on unobservable characteristics. Panel F controls for formal earnings pre-trends over the 1980-1991 time period using Census data rather than the 1986-1990 pre-trends from RAIS in our main specifications. The coefficients on RTC_r are minimally affected.

Panels G and H examine changes in formal earnings premia in the manufacturing and nontradable sectors, respectively. Each individual sector exhibits a very similar pattern to overall earnings across sectors, supporting the assumptions of sectoral factor mobility and equilibrium adjustment

[28] Real exchange rates by country pair and year come from the Penn World Tables, while trade weights, measured for 1990, are based on Comtrade data.

[29] The Brazilian privatization program started in 1991 with the administration of President Collor, but was significantly boosted during President Cardoso's administration (1995-2002).

of nontradable sector prices. Moreover, note that the manufacturing sector exhibits substantial effects immediately after liberalization, in 1995, while other sectors did not experience large positive effects until later. This pattern suggests that liberalization's effects were immediately incident upon the manufacturing sector, and that the earnings effects only appeared in other sectors after workers had time to move between sectors to arbitrage away earnings differences. Panels I and J restrict the sample of regions to include those above median urban share and those above median formal share of employment, respectively. In both cases, the sharply increasing pattern remains. The urban restriction ensures that our results are not driven by the developments in genetically modified crop technology studied by Bustos, Caprettini and Ponticelli (2013). As genetically modified crops were outlawed in Brazil before 2003 and only permanently authorized in 2005, the timing of our effects rules out this channel in any case.

Panel K uses an alternative version of the regional tariff change that is based on changes in effective rates of protection rather than changes in nominal tariffs. These effective rates of protection are calculated by Kume et al. (2003) based on Brazilian input-output relationships, accounting for the effects of tariff changes on an industry's inputs as well as its output.[30] The increasing pattern remains, but the point estimates are somewhat smaller than in the main specification, using regional tariff changes based on nominal tariffs. However, this difference simply reflects the fact that the scale of changes is larger for effective rates of protection; evaluated at the 90-10 difference in regional tariff changes, both measures yield nearly identically sized effects. Panel L uses alternate values for the regional employment weights λ_{ri} in (3), including all workers - not just those in the formal sector. Once again, the results are very similar to our baseline specifications, with steady increases in liberalization's effect on earnings.

The results in Table 3 demonstrate that the striking patterns we observe in liberalization's effects on regional earnings are robust to a variety of changes in specification and were not driven by other major economic shocks during the post-liberalization period.[31] Instead, it appears that the regional labor market effects of liberalization genuinely grew for a decade after the tariff cuts were complete, consistent with our model of slow factor adjustment.

5.4 Factor Adjustment

Having ruled out alternative hypotheses for the steady divergence in wage and employment growth between regions facing more positive and more negative regional tariff changes, we more closely consider the argument that the growing effects were driven by slow factor adjustment. Recall from Section 4.1 that if labor adjusts quickly at first but other complementary factors such as capital gradually adjust later, wage growth will steadily diverge in more positively and more negatively

[30] We do not control for output and input tariffs separately, as they are highly correlated at our level of industry aggregation and would be strongly collinear.

[31] Unreported employment results are similarly robust.

affected regions, precisely the pattern of effects in Figure 4.

Seen through this lens, we can explain more detailed features of Figure 4, following Dix-Carneiro (2014) Section 7.4, which simulates a model of inter-industry reallocation with imperfect and slow adjustment of labor and capital. On impact, wages in more negatively affected markets decline, yielding a positive point estimate at the beginning of liberalization, in 1993. Some laborers reallocate relatively quickly, partly equalizing outcomes across markets and lowering the measured effects for a few years. Then, as specific factors begin reallocating away from more negatively affected markets, workers' marginal products and wages decline in those markets. Continued specific factor reallocation amplifies those wage effects over the course of many years before leveling off after the incentive for specific factors to reallocate has been exhausted.

Figure 6 shows direct evidence of slow labor adjustment. Formal employment fell in regions facing more negative shocks relative to more positively affected regions. This adjustment occurred quickly in the years following liberalization and then stopped around 2004, consistent with initially rapid labor reallocation that tapered off by the mid 2000s. To measure regional capital adjustment in response to trade liberalization, we examine changes in the number and size of formal establishments in a given region. These measures are certain to be closely related to changes in regional stocks of capital and other factors complementary with labor.[32] Figure 10 Panel A shows liberalization's effects on the number of formal establishments, demonstrating that regions facing larger tariff declines experienced larger contractions (smaller increases) in the number of local establishments. This reallocation occurred slowly, with an increase in the early 2000s and leveling out later in the sample period. Figure 10 Panel B shows that the decline in the number of establishments in harder hit locations is not offset by growth in the average size of each firm, which is generally unrelated to RTC_r. This result supports the use of changes in the local number of establishments as proxies for changes in local capital stocks.

We reinforce the idea that sluggish capital adjustment drove regional labor market dynamics by studying how cumulative log job creation and log job destruction rates reacted to regional tariff changes. We define job creation and job destruction as in Davis and Haltiwanger (1990), measuring job growth and contraction starting from 1991 until each year in 1992-2010. Figure 11 shows that markets facing larger tariff declines experienced an immediate and persistent decline in job creation rates relative to less affected local labor markets, suggesting a sharp relative decline in investment rates in more affected local labor markets. In contrast, job destruction rates were essentially unaffected through the late 1990s, after which they began to increase faster in relatively

[32]It is not possible to construct regional capital stocks in Brazil during our sample period. Manufacturing firm-level capital stocks could in principle be constructed from the Annual Manufacturing Survey (PIA) in 1996 and later, but establishment-level investment is not collected, making it impossible to regionally allocate capital for multi-establishment firms. IBGE does not permit regional imputation in this case. Luis Carlos F. Pinto at IBGE, the official in charge of clearing the use of PIA, confirmed that capital stocks cannot be constructed at the regional level. Other data sources covering the remaining sectors of the Brazilian economy do not cover our sample period. The Agricultural Census is available only in 1996 and 2006, and the Annual Service Survey (PAS) only started in 2002.

depressed markets. Figure 12 examines establishment entry and exit rates, showing that the job creation and destruction patterns in Figure 11 were driven largely by establishment entry and exit. Markets facing larger tariff declines immediately exhibit lower entry rates following liberalization, and this difference persists throughout the sample period. Just after liberalization, firm exit rates are somewhat higher in regions facing larger tariff declines, but this effect grows steadily over time. This observation is particularly supportive of the slow capital adjustment hypothesis, as the capital stock associated with exiting establishments is very likely to depreciate or leave the region.

Together, these findings are consistent with the idea that slow factor adjustment accounts for the persistent and growing labor market effects of trade liberalization. Both labor and capital appear to adjust slowly, with relative declines in more negatively affected regions and relative increases in more positively affected regions. While it is difficult to judge the precise relative timing and speed of labor and capital reallocation, the results are consistent with the pattern of reallocation needed to rationalize the observed labor market effects, and Dix-Carneiro (2014) has shown that this mechanism can explain the quantitative size of the response as well.

5.5 Summary of Regional Analysis

In this section, we found large and growing regional effects of liberalization on formal sector earnings and employment. Workers whose regions faced larger tariff declines experienced reductions in earnings and employment compared to workers in more favorably affected regions, and these effects grew steadily for more than a decade before leveling off in the late 2000s. These patterns were not driven by pre-liberalization trends or post-liberalization economic shocks, and are robust to a wide variety of alternative specifications and sample restrictions. The dynamics in labor market outcomes appear to reflect mutually reinforcing slow adjustment of labor and other complementary factors such as capital, as evidenced by results on regional establishment counts and job creation and destruction.

Our cross-regional research design identifies liberalization's *relative* effects on regions facing larger and smaller tariff declines, but cannot identify common effects at the national level. To put our regional results in a national context, we present an alternative version of the earnings and employment analysis. Table 4 breaks the 475 Brazilian microregions into quartiles based on the value of each region's regional tariff change (RTC_r), with quartile 1 facing the largest tariff declines and quartile 4 facing the most positive tariff changes. The top panel presents average log real formal earnings levels for regions in each quartile, controlling for regional worker composition as above. Earnings are deflated to 2010 Reais using the national consumer price index. Earnings levels decline monotonically across quartiles, indicating that regions facing more negative shocks had higher initial earnings. The middle panel reports the change in log real formal earnings following 1991. Brazil was growing quickly during our sample period, and real earnings grew in all four quartiles. However, earnings and employment growth were much faster for regions facing more positive regional tariff changes and slower in regions facing more negative shocks, consistent with our main results in Tables 1 and 2. Liberalization drove a convergence in earnings across regions by accelerating earnings growth in low-earnings locations relative to those in high-earnings regions. Thus, liberalization seems to have spurred growth in formerly marginal cities and towns as compared to the country's more traditional economic engines in large high-wage cities.

6 Individual Analysis

6.1 Empirical Specification

The RAIS data make it possible to follow individual workers over time, so we can observe the medium- to long-run effects of liberalization on individual workers based on their initial region of employment. Our main analysis focuses on a panel of workers who were initially employed in the tradables sector in December 1989, just before trade liberalization began.[33] Our population of interest includes all workers aged 25-44 in December 1989, whose highest paying job was in the tradable sector. For computational tractability, we take a 15% sample of individuals meeting these criteria in regions with more than 2,000 tradable sector workers in 1989 and include all relevant workers from smaller regions, weighting appropriately in subsequent analyses. This process yields 587,575 individuals in our main tradable sector sample. In Section 6.6, we also consider an alternate population of workers initially employed in the nontradables sector, in order to investigate the transmission of the trade shock into this sector. All other restrictions and sampling procedures are the same, yielding a sample of 973,048 nontradable sector workers. Table 5 provides summary statistics for the tradables sector and nontradables sector samples. Finally, we use alternate samples to study in-migration and the length of non-formal spells, described in Sections 6.2 and 6.5, respectively.

The individual-level analysis estimates equations of the following form, separately for each year in 1990-2009:

$$y_{ijrt} = \alpha_{st} + \theta_t RTC_r + X_{ijr,1989}\Phi_t + \epsilon_{ijrt}, \tag{5}$$

where i indexes individuals, j is the worker's initial industry of employment in December 1989, r is the worker's initial region of employment in December 1989, and t indexes years following the start of liberalization (1990 to 2009).[34] y_{ijrt} represents various worker-level post-liberalization outcomes, which we define below. $X_{ijr,1989}$ is a rich set of worker-level controls including demographics (9 education category indicators, age, age-squared), initial job characteristics for the highest-paying job in December 1989 (84 occupation category indicators, 14 tradable industry indicators, 12 non-

[33] In the regional analysis above, we use 1991 as the base period to maintain comparability with the Census. Regional results are unchanged when using 1989 as the base year.

[34] The 2010 RAIS data use a different individual identifier than the previous years, so we are only able to implement the individual analysis through 2009.

tradable industry indicators, tenure at the plant), initial employer characteristics (log employment, exporting indicator, log exports, importing indicator, log imports), and initial region characteristics (state indicators, pre-liberalization (1986-89) earnings and formal employment growth, a control for the pre-liberalization (1986-89) level of each outcome of interest, and controls for other pre-liberalization conditions related to the particular outcome).[35] This specification compares subsequent labor market outcomes for two otherwise observationally equivalent workers who in 1989 happened to live in regions facing different local trade shocks.

This analysis takes advantage of the panel dimension of the RAIS data, and is similar in spirit to the panel-data analysis in Autor et al. (forthcoming). However the studies differ in a number of important ways. We i) focus on regional rather than industry shocks, ii) study a discrete shock and hence focus our analysis on subsequent labor market dynamics, iii) study transitions into the nontradable sector and informal employment, which are salient features of the Brazilian context, and iv) study formal sector migration in response to regional shocks.

6.2 Employment

We begin by examining how the regional tariff change in a worker's initial location affected their subsequent formal employment status. We first calculate the average number of months formally employed per year from 1990 to year t:

$$\frac{1}{t-1989}\sum_{s=1990}^{t} Months_{is}, \tag{6}$$

where $Months_{is}$ is the number of months individual i was formally employed in year s.[36]

Figure 13 shows that the average number of months formally employed per year is lower for workers whose initial region faced a larger tariff decline. The effects increase steadily over time, beginning just after liberalization was complete in 1995. This means that workers whose initial regions faced larger tariff cuts worked fewer and fewer months of the year in the formal sector, compared to otherwise similar workers who started in more favorably affected regions. The 2009

[35] Firm-level imports and exports for 1990 come from customs data assembled by the *Secretaria de Comércio Exterior* (SECEX). Region-specific formal earnings and employment growth between 1986 and 1989 are computed as in Section 5. The pre-liberalization outcome controls are calculated as follows. We draw a sample of workers in December 1986, paralleling the main sample, and estimate a version of (5) replacing RTC_r with region indicators. These first step region indicator coefficients enter as controls in equation (5). For example, when the outcome of interest is given by equation (6), we first relate the average number of months formally employed per year between January 1987 and December 1989 to region indicators. When the outcome of interest is given by equation (7), we first relate the fraction of 1989 spent in the formal sector to region indicators. Note that for accumulated earnings, we are unable to normalize by pre-1986 earnings, so we instead include the pre-liberalization control related to months formally employed. For migration-related outcomes, we additionally include 1986-1991 out-migration probability controls obtained from the Census.

[36] RAIS reports the month of accession and separation (if any) for each job, so that we can observe formal employment at the monthly level.

point estimate is 6.0, implying that a worker whose initial region faced a 10 percentage point larger tariff decline (approximately the 90-10 gap in RTC_r) on average worked in the formal sector for 12 fewer total months during 1990 to 2009. This is a large effect, given that the unconditional average number of total months worked in the formal sector during this time period for workers in our sample is 120.

A closely related measure of liberalization's effect on an individual's formal employment experience is the fraction of year t spent in the formal sector:

$$\frac{Months_{it}}{12}, \tag{7}$$

This measure is similar to (6), but focuses on an individual year rather than the cumulative experience from 1990 onward. Figure 14 shows that workers whose initial regions faced larger tariff cuts spent less of each year in formal employment, and that this effect grew for a decade following liberalization before falling somewhat in the second half of the 2000s. The largest effect, 0.77, appears in 2004. On average, a worker whose initial region faced a 10 percentage point larger tariff decline spent 0.92 fewer months in formal employment, compared to an average of 4.8 for the entire sample in 2004.

Along with transitions out of formal employment, workers also adjust between tradable and nontradable sector employment. Recall that all of the workers in our main sample were initially employed in the tradable sector just prior to liberalization. We examine the average number of months formally employed per year, as in (6), but separate months into those worked in tradable and nontradable sector employment. Figure 15 shows that the overall formal employment response is driven by changes in tradable sector employment. In fact, nontradable employment offsets a fraction of the employment losses in the tradable sector, indicating that some tradable sector workers facing larger regional tariff cuts transitioned into nontradable employment. However, the relative increase in months of nontradable employment only offsets a small share of the losses in tradable employment, such that overall months decline in the hardest hit locations, as seen in Figure 13.

6.3 Mobility

Workers can also adjust by moving to a more favorable location. We investigate this phenomenon by studying both out-migration and in-migration choices for formally employed workers. We begin with out-migration, considering the probability that in year t a worker in our sample is formally employed in a region other than their 1989 location. Figure 16 shows that there is no significant relationship between the local tariff decline in a worker's initial region and the subsequent probability of working in a different region. Note that if migrants leave the formal sector, they leave the sample and their migration will not be observed. To lessen potential bias due to differential attrition from formal employment, we calculate an alternate measure, the share of formally employed months spent away from the initial region:

$$\frac{MonthsAway_{it}}{Months_{it}} \tag{8}$$

This measure mitigates selection concerns because the vast majority of individuals in our sample spend at least one month in the formal sector between 1990 and 2009. Figure 17 once again shows no significant relationship. In fact, if anything, the positive point estimates suggest that more negative local shocks lower the probability of obtaining formal employment outside the initial region.

To study in-migration, we construct separate samples of all formally employed individuals in 1989, 1995, 2000, 2005, and 2009 using the same sampling rates described above. For each of these samples, we record each worker's location one year earlier, five years earlier, and in 1989, just before liberalization began. We control for pre-existing migration patterns using three-year migration rates from 1986-89 and other initial regional characteristics such as earnings and employment growth. The results appear in Table 6. All point estimates are positive, indicating larger declines in in-migration to harder hit regions. For example, the coefficient 0.317 in column (2) for 1995 indicates that a region facing a 10 percentage point larger tariff decline experienced a 3.17 percentage point larger decrease (smaller increase) in the share of in-migrants between 1990 and 1995. Longer timeframes permit larger migration responses, as indicated by decreasing magnitude coefficients moving from left to right across columns and increasing coefficients moving down column (1), where 1989 is always the base year. These in-migration responses coupled with the lack of out-migration responses are consistent with the relative contraction in formal employment in regions facing larger tariff declines.

Note that these RAIS migration analyses can only capture migration events occurring for workers who are formally employed both before and after migrating. We supplement these results by examining changes in working-age population using the Census, which covers all individuals regardless of employment status.[37] Columns (1) and (2) in Table 7 counterintuitively find that population increased in harder-hit regions. However, this result disappears in all specifications that control for pre-liberalization trends in population growth. Column (3) controls for 1980-1991 population growth, while (4) controls for 1970-1980 population growth. Because the 1980-1991 pre-trend and the dependent variable both contain the 1991 log population, in columns (5) and (6) we instrument for the pre-trend using the 1980 population level and the 1970-1980 pre-trend, respectively. In all cases with pre-trend controls, the population response is insignificantly different from zero, and in columns (4)-(6), the point estimate is very small as well.

Thus, even though 11 percent of workers in the formal tradable sector in December 1989 are observed with formal employment in a different region in 2000 (see Table 5), their out-migration decisions are not influenced by local labor demand shocks. On the other hand, conditional on

[37] Note that we use a slightly aggregated region definition in Table 7. See Section 7 for details

migrating, formal sector workers tend to avoid adversely affected regions. These findings parallel those of two recent papers studying workers' location choices in the United States during the Great Recession. Yagan (2014) finds that out-migration does not respond to local labor demand shocks, while Monras (2014) finds that in-migration declines in locations facing labor demand declines. Interestingly, even though 5-year gross migration rates for working-age individuals are large (7.3% between 1995 and 2000), net migration rates are not affected by the trade-induced local shocks. Therefore, it seems that migrants keep moving to harder-hit locations, but they avoid the formal sector in these locations. This pattern results in formal sector net migration (as measured by RAIS) responding to the shocks, whereas overall net migration does not (as measured by Census population). These findings imply that the substantial regional labor adjustment we documented earlier occurred primarily through transitions between formal and informal employment rather than through interregional migration.[38] Recall that informal employment tends to be an absorbing state in the Brazilian context, with very large barriers to shifting from informal to formal employment (Dix-Carneiro 2014).

6.4 Earnings

Along with employment, we consider the effect of the regional trade shock in a worker's initial location on subsequent formal earnings. Keep in mind that formal earnings effects are likely to be upper bounds on the overall earnings effects, as workers losing formal earnings may partially offset these losses through earnings in the informal sector (which cannot be observed in RAIS). However, we present these formal earnings results for completeness.

For each year t from 1990-2009, we calculate a worker's average earnings from 1990 to t as a multiple of the worker's average pre-liberalization (1986-89) annualized earnings:

$$\frac{\frac{1}{t-1989}\sum_{s=1990}^{t} Earnings_{is}}{MeanEarnings_{i,1986-89}}, \tag{9}$$

$$\text{where } MeanEarnings_{i,1986-89} \equiv \frac{\sum_{s=1986}^{1989} Earnings_{is}}{\sum_{s=1986}^{1989} Months_{is}} \times 12$$

The numerator is the worker's average post-liberalization earnings from 1990 to t, and the denominator is the worker's average pre-liberalization annualized formal sector earnings from 1986 to 1989. Note that formal earnings may decline due to lower wages or less time worked in the formal sector. We use this measure because it accounts for worker heterogeneity in initial earnings while still being well defined for workers with zero earnings after 1989, avoiding sample selection issues.

[38] Morten and Oliveira (2014) argue that road construction increased internal migration in Brazil by lowering mobility costs. Such a policy response would be unlikely to increase the mobility response to liberalization, however, because more positive regional tariff shocks tended to occur in traditional migration sending (low wage) locations rather than receiving (high wage) locations.

We then regress this earnings measure for each year t on the regional tariff change (RTC_r) and the extensive set of controls described above. Figure 18 shows the results. The point estimate in 2009 is 0.86, implying that over the course of 20 years a worker whose initial region faced a 10 percentage point larger tariff decline lost 1.72 times their average pre-liberalization wage in the formal sector in relative terms.

Figure 19 shows the results for a closely related earnings measure, yearly earnings as a share of pre-liberalization mean earnings.

$$\frac{Earnings_{it}}{MeanEarnings_{i,1986-89}} \tag{10}$$

Rather than averaging over the sample period, this measure expresses earnings in each year. Once again we see positive and growing coefficients following liberalization, indicating that earnings of workers whose initial regions faced larger tariff declines steadily diverged from those initially in more favorable regions for more than a decade following liberalization.

6.5 Non-Formal Spells

Above, we found that workers facing larger tariff declines spent more time out of formal employment. We complement that analysis by examining the length of non-formal spells. To implement this analysis, we need a representative sample of all job terminations that occur in a given region and year. We draw a 5% sample of workers who appear in the RAIS data between 1986 and 2005. We then record the length of all non-formal spells experienced by workers aged 25-55 and employed in the tradable sector at the time the spell began. Since our data end at December 2009, non-formal spells are right-censored at that point. We use a Tobit specification to relate the log non-formal spell length to the regional tariff change in the worker's region of formal employment when the non-formal spell began and the full set of controls.[39]

The results are shown in Figure 20. Workers losing formal employment in regions that faced larger tariff declines experienced relatively longer non-formal spells. The largest point estimate of -4.57 appears in 1998, implying that among spells beginning in 1998, those in regions facing 10 percentage point larger tariff declines are 45.7% longer, on average.[40]

6.6 Nontradable Sector Workers

Recall that the earlier results in this section examine workers who initially worked in tradable goods industries. We also implemented all of the above individual-level analyses using an alternate group

[39]To calculate the pre-liberalization trend control, we estimate a parallel specification for non-formal spells beginning in 1986 and right-censoring spells extending beyond 1989, in order to avoid contaminating the pre-liberalization non-formal spells with the effects of trade reform.

[40]For reference, the average non-formal spell length among those starting in 1998 and not truncated (finalized before December 2009) is 21.4 months. The median length is 12 months.

of workers who were initially employed in the nontradable sector. Our objective is to see whether workers outside tradable sectors are insulated from the local effects trade liberalization, or whether the tradable and nontradable labor markets are sufficiently integrated that regional trade shocks affect both sectors' workers similarly.

We find similar magnitudes and time patterns of effects for the tradable and nontradable sector workers. Figure 21 analyzes the average months formally employed per year for the nontradable workers. The pattern is very similar to that of the tradable workers (compare to Figure 13), with effects that are about 20-25 percent smaller than those for tradable sector workers. Figure 22 analyzes average earnings. Once again, the results are similar to those for the tradable sector workers (compare to Figure 18), with effects that are about 10-15 percent smaller. Thus, the two portions of the labor market appear to be sufficiently integrated that the nontradable sector is only slightly insulated from shocks in the tradable sector, on average.

The integration of nontradable and tradable sector labor markets is further reinforced by Figure 23, which breaks the employment analysis of Figure 21 into months spent in tradable and nontradable employment. The results are quite different from those for tradable sector workers in Figure 15. The biggest losses for workers initially in the *nontradable* sector occur in the *tradable* sector. Only in the last years of our sample do nontradable sector employment losses become significantly different from zero, while tradable sector losses are large and significant throughout the post-liberalization period. This means that nontradable sector workers regularly transition to tradable employment, but that this adjustment margin is steadily curtailed in markets facing larger tariff declines, driving much the employment losses faced by nontradable sector workers.

6.7 Summary of Individual Analysis

The preceding longitudinal analysis demonstrates that trade liberalization had large and long-lasting effects on workers' formal sector employment and earnings outcomes during the 15 years following liberalization. Workers initially located in more adversely affected places experience large relative losses that grow over time and never recover. Adversely affected workers spend less time formally employed, experience relative declines in formal earnings, and experience longer spells outside formal employment.

We also examine various adjustment margins within formal employment. Workers initially in the tradable sector are more likely to transition into nontradable employment when facing more negative shocks. However, these sectoral transitions are too small on average to compensate for losses in the tradable sector. We find minimal effect of regional shocks on worker mobility, particularly for formal out-migration and overall working-age population. However, those that did migrate in the formal sector tended to avoid the hardest hit destinations, lowering in-migration rates of formally-employed workers to those regions.

Finally, the evidence strongly supports the conclusion that formal tradable and nontradable sectors are strongly integrated. Workers initially employed in the nontradable sector experienced very similar employment and earnings effects to those initially employed in the tradable sector, with only a slightly smaller magnitude. Employment losses for initially tradable sector workers were partly offset by transitions into nontradable employment. More strikingly, employment losses for initially nontradable sector workers occurred primarily through reduced subsequent transitions into tradable employment, highlighting the close integration of the two sectors.

7 Informal Employment and Non-Employment

In the preceding analyses, we focus on outcomes for formally employed workers. The formal sector is of particular interest for a variety of reasons. It is more capital intensive, dynamic, and productive than the informal sector, and formal jobs are generally seen as being of much higher quality than informal jobs (LaPorta and Schleifer 2008, Fajnzylber, Maloney and Montes-Rojas 2011, LaPorta and Schleifer 2014). Formal employment gives workers access to all of the benefits and labor protections afforded them by the legal employment system, while informal jobs generally provide minimal benefits and fail to comply with various labor regulations. Hence, transitions out of formal employment generally involve important declines in worker wellbeing even if displaced workers later find informal employment. Finally, focusing on formal employment allows us to carefully observe worker adjustment by taking advantage of the RAIS administrative panel data.

In this section we supplement the main formal sector results with additional analyses examining the roles of informal employment and non-employment in the regional labor market adjustment patterns documented above.[41] Trade policy's effects on informality are also of independent interest, as evidenced by a large and growing academic literature.[42] Import competition may increase pressure on firms to cut costs by failing to comply with labor regulations, and informal jobs are often characterized as providing lower compensation, fewer opportunities for training and advancement, and generally less favorable working conditions (Goldberg and Pavcnik 2007, Bacchetta et al. 2009). Together, these concerns have made informality a prominent issue in public debates over globalization in the developing world (Bacchetta et al. 2009).

We examine informal employment and non-employment using the Decennial Census, which covers all individuals, including those outside formal employment. Because the Census consists of repeated cross-sections rather than panel data as in RAIS, our conclusions in this section will necessarily be somewhat more speculative than those for the formal sector. The 1991-2010 Censuses report whether a worker has a work card signed by their employer, which determines formal

[41] We focus on non-employment, which includes both unemployment and out of the labor force. This approach allows us to avoid changing labor force definitions over time and captures transitions into unemployment and out of the labor force, both of which may be affected by trade reform.

[42] See Goldberg and Pavcnik (2007) and Paz (2014) for literature reviews with relevant citations.

employment status.[43] Table 8 reports the informal share of employment in each Census year overall and by major sector. Informality increased during the 1990s from 58 percent to 64 percent before falling sharply during the 2000s to 49 percent in 2010. Informality rates vary widely across sectors, with very high levels in agriculture and much lower levels for manufacturing and for mining in the later periods. Informality rates fell steadily over time in agriculture and mining, with particularly sharp declines in mining, while manufacturing and nontradable informality shares followed the overall pattern of rising and then falling. This heterogeneity by sector will be important in interpreting our subsequent results.

Because the Census only provides repeated cross sections rather than panel data, throughout this section we use the regional estimation strategy introduced in Section 5. We construct pre-liberalization trends using the 1980 and, when possible, 1970 Censuses. The 1970 Census contains no information on formal vs. informal employment status, so we use 1970-1980 pre-trends only when they are relevant. Since 1980-1991 pre-trends overlap with the dependent variables in 1991, we instrument for them using the 1980 outcome level. We slightly aggregate regions to be consistent from 1970 to 2010, yielding 411 consistently identifiable regions. As above, we cluster all standard errors by mesoregion to account for potential spatial correlation in outcomes across neighboring regions. Because we are now focusing on labor demand shocks faced by workers in both formal and informal employment, we calculate regional tariff changes using regional industry employment shares, λ_{ri}, for all workers, as in Panel L of Table 3.

7.1 Informal and Non-Employment Shares

In both the regional and individual analyses presented above, it is clear that workers facing larger tariff cuts spend less time in formal employment, but in the RAIS data we are unable to distinguish between informal employment and non-employment. To address these adjustment margins, we use Census data to examine the share of regional working age population that is informally employed or non-employed. Following our wage analyses and Goldberg and Pavcnik (2003), we calculate regional employment shares controlling for worker composition.[44] We regress indicators for informal employment or non-employment on workers' observable characteristics and region indicators, and then use the change in the coefficients on those indicators as our dependent variable.

The results appear in Table 9. Panel A shows that regions facing larger tariff declines experience relative increases in the share informally employed. The estimate of -0.171 in column (3) indicates that by 2000 a region facing a 10 percentage point larger regional tariff decline exhibited a 1.71 percentage point larger increase in its informal employment share. Panel B shows that harder hit regions experienced even larger increases in non-employment (including both unemployed and out of

[43] Since the 1980 Census does not include the work card question, when calculating pre-trends we utilize an alternative question on whether contributions were made to the worker's public pension account. In 1991, the correlation between this measure and the more accurate work card measure is 0.85.

[44] Results are similar for raw shares, unadjusted for worker composition.

the labor force). By 2010, however, the situation is different. Informality shares increase even more in harder hit locations, while the non-employment effect is now zero. In the absence of substantial interregional migration, as documented above, these results imply that many workers whose regions faced larger tariff declines were non-employed in the years following liberalization, but that many of these individuals later found employment in the informal sector. Hence, transitions to informal employment often occured following a lengthy spell of non-employment. Meghir, Narita and Robin (forthcoming) support this interpretation, showing (in their Table 1) very frequent transitions of unemployed workers to informal employment.[45]

This strong effect of liberalization on the local informal employment rate is a novel finding, which may appear to contradict other results in the literature studying the response of Brazilian informality to trade policy changes. The apparent conflict is resolved by noting differences in methodology and margins of adjustment. For example, Goldberg and Pavcnik (2003) do not find an effect of trade policy on informality, a finding corroborated by Bosch et al. (2012). These papers examine changes in informality within manufacturing sectors, so their analyses omit changes in informality within agriculture and mining. As shown in Table 8, during the 1990s, informal shares increased in manufacturing industries, which faced larger tariff cuts, and informal shares declined in agriculture and mining, which faced more positive tariff changes.[46] Moreover, these studies omit transitions to informal employment that occur after a spell of non-employment, which we just found to be an important pattern in the Brazilian context. Menezes-Filho and Muendler (2011) examine yearly employment transitions for individual workers initially employed in manufacturing. They find no significant relationship between tariff changes and the likelihood of transitioning into informal employment, but do find that output tariff declines lead to increased transitions into non-employment. These findings are consistent with our results if, as suggested by Table 9, many displaced formal sector workers spend a year or more in non-employment before eventually obtaining informal employment. Therefore, the differences between our findings and those of prior studies on Brazilian informality result from our inclusion of workers outside manufacturing and our focus on medium- and long-run transitions, even if they involve a spell of non-employment. Our findings more closely parallel those of McCaig and Pavcnik (2014), who find substantial shifts from household (informal) to enterprise (formal) employment in Vietnam in response to the U.S.-Vietnam Bilateral Trade Agreement.[47]

To complete the picture of regional labor market adjustment to liberalization-induced labor demand shocks, we examine changes in the shares of regional employment falling in the following four categories: formal tradable, formal nontradable, informal tradable, and informal nontradable,

[45]Transitions from unemployment to informal employment are 4 to 5 times more frequent than transitions from unemployment to formal employment.

[46]Appendix Figure A6 provides a breakdown of informality changes by more detailed industry.

[47]Paz (2014) and Cruces, Porto and Viollaz (2014) provide two other recent examples that find significant effects of tariff changes on informality, using different methodologies.

with results in Table 10.[48] The results instrumenting for pre-trends in columns (4) and (8) are much less stable than in other analyses, and in Panel D the instruments are very weak, so we focus on columns (1)-(3) and (5)-(7). Formal tradable employment is clearly the category hardest hit when facing negative regional tariff shocks. The offsetting growth in informal employment that we saw in Table 9 occurs primarily in the tradable sector, while nontradables employment experiences much smaller and insignificant effects. Putting these results in context, in Figure 15 we found that formal tradable sector workers were more likely to transition into formal nontradable sector employment when the initial region faced a more negative labor demand shock. Yet here we generally find small positive or insignificant coefficients for the regional formal nontradable employment share, indicating that this portion of the labor market does not expand to absorb the tradable sector workers transitioning into nontradable employment. What, then, happened to nontradable sector workers? Recall from Figure 23 that the biggest employment losses for workers initially in the nontradable sector occurred in the tradable sector. This means that nontradable workers often transition to tradable employment, but these transitions occur much less frequently in markets facing larger tariff declines. It is likely that these nontradable sector workers who are no longer able to find formal tradable or nontradable employment drive a large portion of the growth in informal tradable employment seen in Table 10.

7.2 Wages

In Sections 5 and 6, we found compelling evidence that earnings and wages for formal sector workers declined in regions facing larger tariff cuts, compared to regions facing smaller declines. In Table 11 we perform a similar analysis for informally employed workers, and find quite different results. In 2000 there is no significant relationship between regional wages for informal workers and the regional tariff change. This applies when controlling for 1980-1991 pre-trends, 1970-1980 pre-trends, or when instrumenting the 1980-1991 pre-trend using the 1980 wage premium.[49] The long run results for 2010 are even more imprecisely estimated and vary substantially across specifications. These null results are somewhat surprising, because we expected informal wages to fall along with formal sector wages, both because of declining local labor demand and because displaced formal sector workers flood into informal employment. A potential explanation for the lack of effect on informal wages is that consumers in harder hit regions experience declining incomes and shift toward lower-priced, lower-quality goods produced in the informal sector, offsetting wage declines for informally employed workers.[50]

[48] Note that although these categories partition all employed workers, the coefficients do not precisely add to zero because of differences in weighting and pre-trends across outcomes. Also, given the lack of information on formality in the 1970 Census, we cannot calculate 1970-1980 pre-trends.

[49] We exclude specifications using the 1970-1980 pre-trends as instruments for the 1980-1991 pre-trends, because instrument is very weak.

[50] Burstein, Eichenbaum and Rebelo (2005) show that lower quality goods gain market share in recessions, while McKenzie and Schargrodsky (2011) make a similar argument in the context of the 2002 economic crisis in Argentina.

Other potential explanations are less plausible. First, following the logic of our model in Section 4.1, capital could be slowly reallocating from firms in the formal sector to firms in the informal sector. This mechanism is unlikely, as firms in the informal sector are much less capital intensive than firms in the formal sector (LaPorta and Schleifer 2008, Fajnzylber et al. 2011). Second, displaced formal sector workers who move into informal employment may have more favorable unobserved characteristics than average informal workers, even after controlling for education and other demographics we use when calculating regional wage premia. This mechanism would predict that overall wage effects mirror those for the formal sector, since compositional changes net out when combining informal and formal employment. We show that this is not the case by examining liberalization's effects on overall wages for formal and informal workers together.

Table 12 shows the effects of regional tariff changes on overall wages for all workers irrespective of formal status. In 2000, we find that wages for all workers fell more in regions facing larger tariff declines. The coefficients are positive and significant in many specifications, consistent with the formal sector results in Table 1. However, in contrast to the formal sector, there is little evidence of growing wage effects by 2010. These effects are imprecisely estimated and vary substantially across specifications. More generally, the findings for all workers in Table 12 make it clear that liberalization had very different effects on the formal and informal sectors. These important differences would be missed if one only examined overall wage effects without regard to formality.

8 Conclusion

In this paper, we have examined regional labor market dynamics following the Brazilian trade liberalization of the early 1990s. Using matched employer-employee data as our primary data source, we find large effects of trade liberalization on regional formal labor market outcomes such as earnings and employment. Contrary to conventional wisdom, which assumes equalizing labor adjustment, the regional effects of liberalization grow for more than a decade before leveling off. This pattern is not driven by post-liberalization economic shocks and is robust to a wide variety of alternative specifications. Rather, we argue that it is driven by mutually reinforcing, imperfect, and slow adjustment of labor and complementary factors, driving dynamics in local labor demand. We present a specific-factors model allowing for imperfect and slow factor adjustment, which yields a theoretically consistent measure of liberalization's effects on local labor demand and rationalizes the labor market dynamics we observe. Dix-Carneiro (2014) shows that adjustment frictions across industries in both labor and capital can lead to quantitatively similar effects to those we observe

While there is little direct evidence on the relative quality of goods produced by formal and informal firms, it is well known that informal firms are significantly smaller than formal firms (LaPorta and Schleifer 2014, Meghir et al. forthcoming, Ulyssea 2014), and Kugler and Verhoogen (2011) show that larger firms produce higher quality goods than small firms, on average. Moreover, LaPorta and Schleifer (2008) show that informal firms use lower quality inputs and speculate that they produce lower quality outputs as a result.

across regions. We observe slow changes in regional employment, the regional number of firms, establishment entry and exit rates, and job destruction rates that are consistent with the hypothesis of slow adjustment of complementary factors.

We also study liberalization's effects at the individual level by examining post-liberalization outcomes for otherwise identical workers whose initial regions faced different regional tariff shocks. We find that tradable sector workers whose initial regions faced larger tariff declines subsequently spent less time in formal employment and that this effect grew steadily in the years following liberalization. Tradable sector workers partly smoothed their formal sector outcomes by transitioning into nontradable sector employment, but these transitions were too small on average to fully compensate for losses in the tradable sector. On the other hand, nontradable sector workers were largely driven into informal employment, because they were no longer able to transition into jobs in the shrinking local formal tradable job market. These transitions into informal employment occurred slowly over time and likely involved long periods of non-employment before workers obtained informal employment.

Together, these results describe a rich pattern of dynamic regional adjustment to trade reform in Brazil, in which favorably and unfavorably affected regions experienced very different trajectories for labor market outcomes. It is important to keep in mind that our analysis is cross-sectional, in that it does not identify common national effects of liberalization applying to all regions, but rather identifies differential effects across regions. This feature is common to all papers exploring variation in industry-specific or regional trade or liberalization shocks.

Our results have important implications for our thinking about the labor market effects of trade liberalization. A growing literature has shown in a variety of contexts that trade and trade policy have heterogeneous effects across regions in the short-run. However, most researchers, ourselves included, generally assumed that these effects would be upper bounds on the long-run effects, as labor reallocation would arbitrage away regional differences. This paper finds precisely the opposite. Short-run effects vastly underestimate the long-run effects, indicating that the costs and benefits of liberalization remain unevenly distributed across geography, even twenty years after the policy began.

These novel empirical findings suggest new avenues for future work. We find evidence consistent with the hypothesis that the growing effects of liberalization were driven by slow adjustment of labor and complementary factors. These empirical results reinforce the message of Dix-Carneiro (2014) that jointly quantifying mobility frictions for labor and capital is key to understanding trade adjustment. Artuç, Bet, Brambilla and Porto (2014) take an initial step in this direction. We also find large differences in local labor market effects in the formal and informal sectors, making clear the importance of understanding how the two sectors interact in driving labor market outcomes, and suggesting another avenue for future research. Finally, contrary to previous research on the topic, we find that the informal sector acts as a fallback sector for workers in the aftermath of trade liberalization. It is therefore important to understand i) to what extent adversely affected workers were able to smooth their labor market outcomes by transitioning to informal employment, and ii) how labor market policies should respond to this margin of adjustment, if at all, given that informal jobs are generally of lower quality and on average appear in less productive firms.

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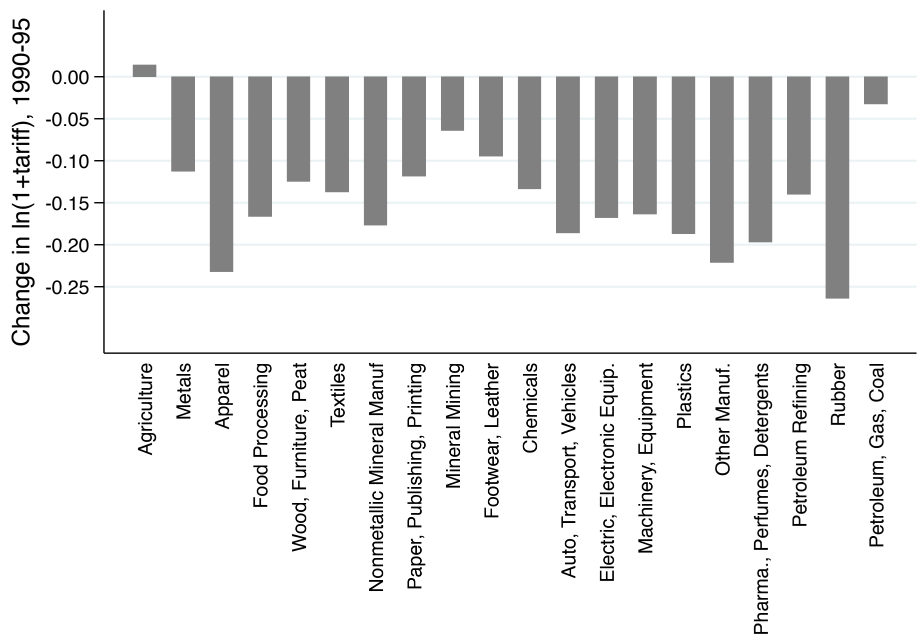
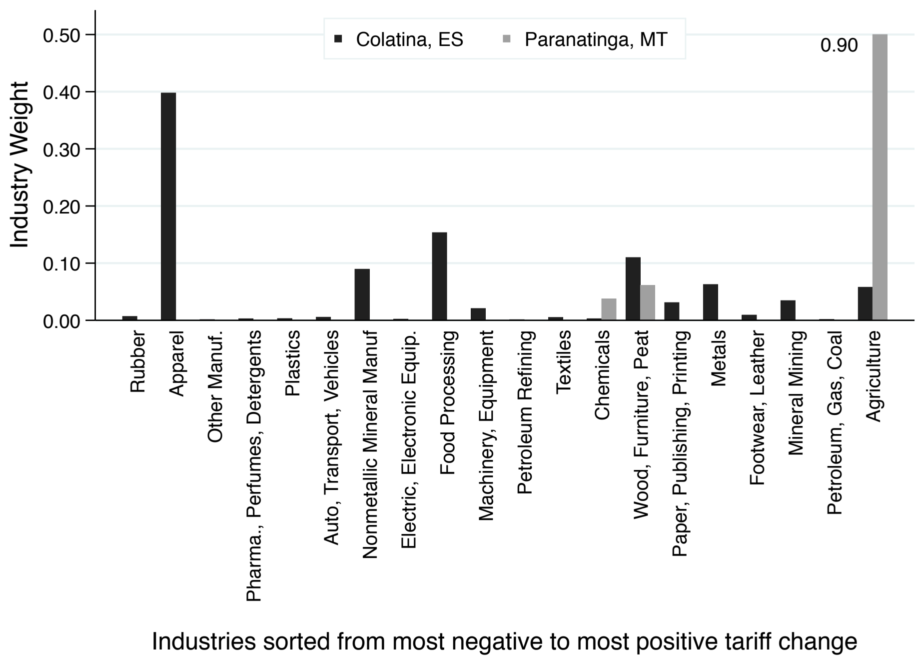
Figure 1: Tariff Changes



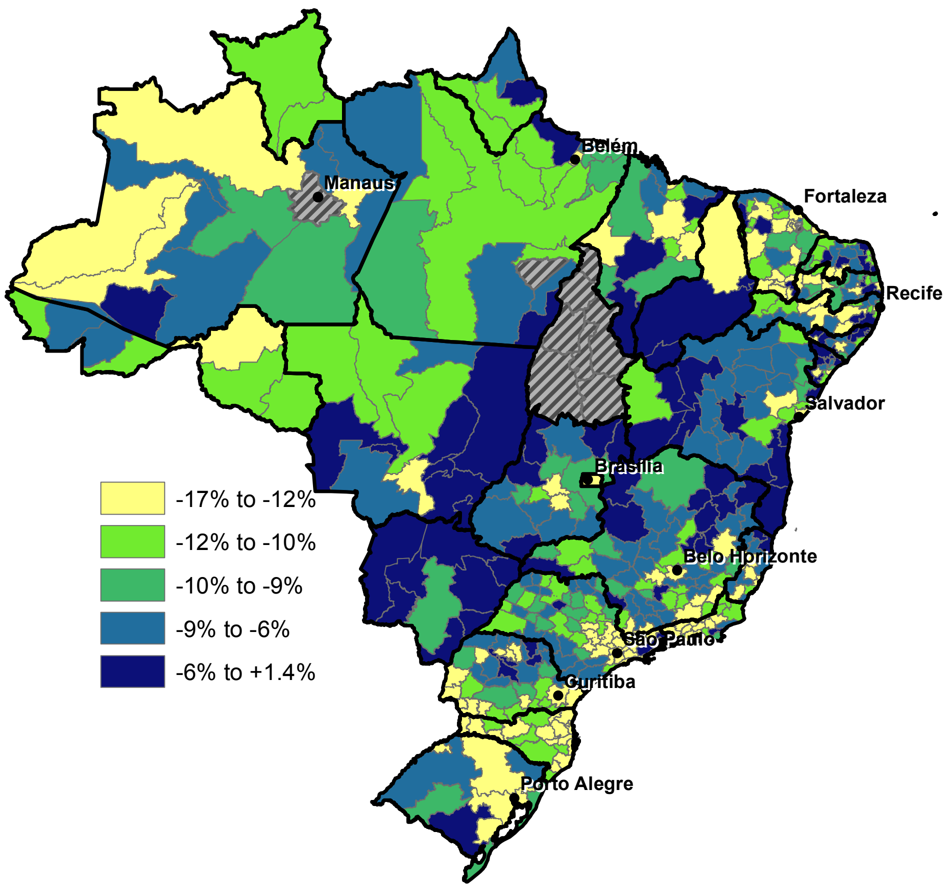
Tariff data from Kume et al. (2003), aggregated to allow consistent industry definitions across data sources. See Appendix Table A1 for details of the industry classification.

Figure 2: Variation Underlying Regional Tariff Change



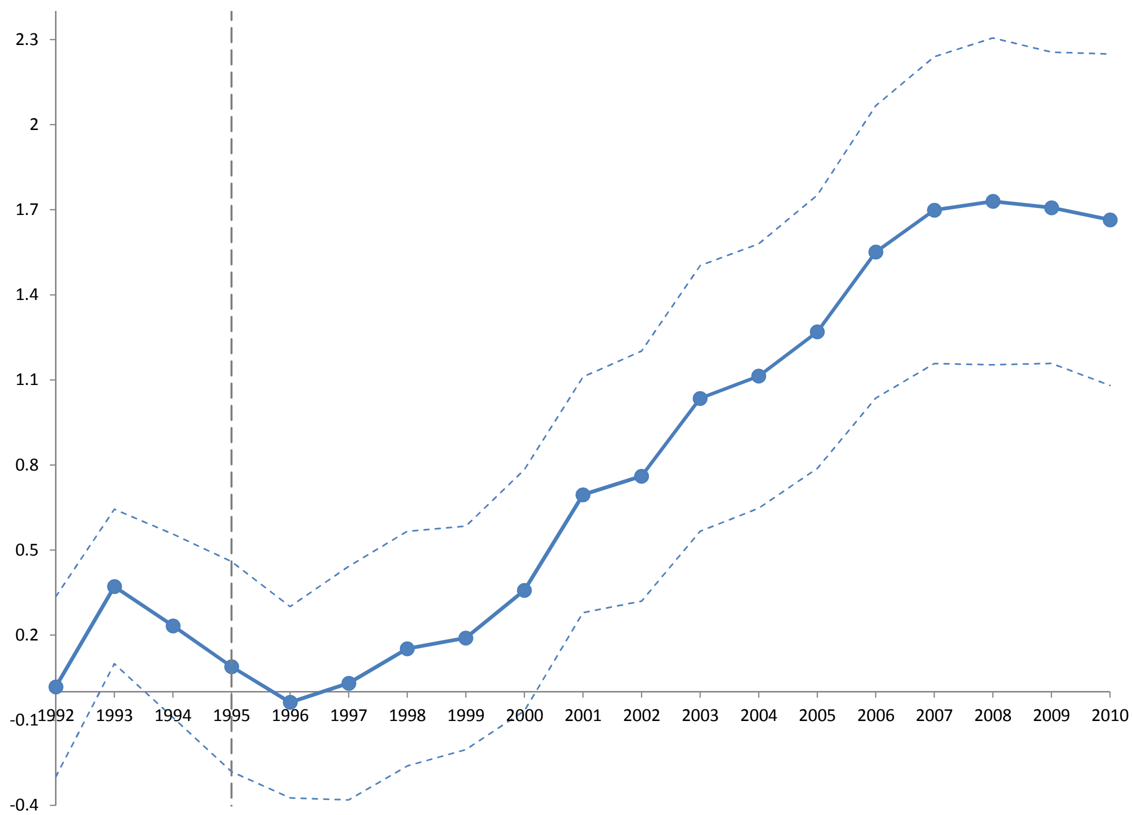
Industry distribution of 1991 employment in the two regions facing the most negative (Colatina, ES) and most positive (Paranatinga, MT) regional tariff changes. Industries sorted by the tariff change, shown in Figure 1. More weight on the left side of the figure leads to a more negative regional tariff change, and more weight on the right side leads to a more positive regional tariff change.

Figure 3: Regional Tariff Changes



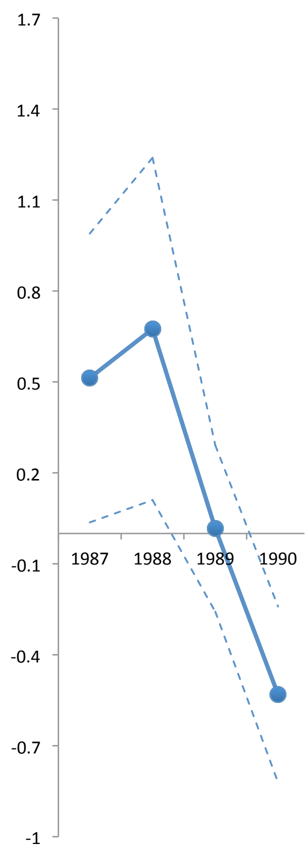
Local labor markets reflect microregions defined by IBGE, aggregated slightly to account for border changes between 1986 and 2010. Regions are colored based on the regional tariff change measure defined in (3). Dark lines represent state borders, gray lines represent consistent microregion borders, and cross-hatched migroregions are omitted from the analysis. These microregions were either i) part of a Free Trade Area ii) part of the state of Tocantins and not consistently identifiable over time, or iii) not included in the sample before 1990.

Figure 4: Regional log Formal Earnings Premia - 1992-2010



Each point reflects an individual regression coefficient where the dependent variable is the change in regional log formal earnings premium from 1991 to the year listed on the x-axis, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with state fixed effects and pre-trend control. Positive estimates imply larger earnings declines in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 5: Regional log Formal Earnings Premia - Pre-Trends - 1987-1990

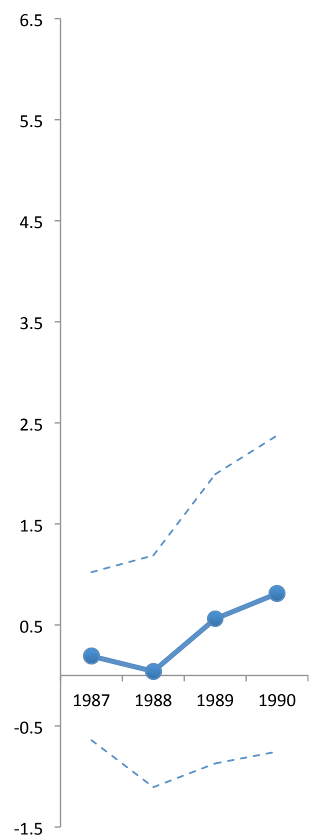


Each point reflects an individual regression coefficient where the dependent variable is the change in regional log formal earnings premium from 1986 to the year listed on the x-axis, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with state fixed effects. Positive estimates imply larger earnings declines in regions facing larger tariff declines. Standard errors adjusted for 112 mesoregion clusters.

Figure 6: Regional log Formal Employment - 1992-2010

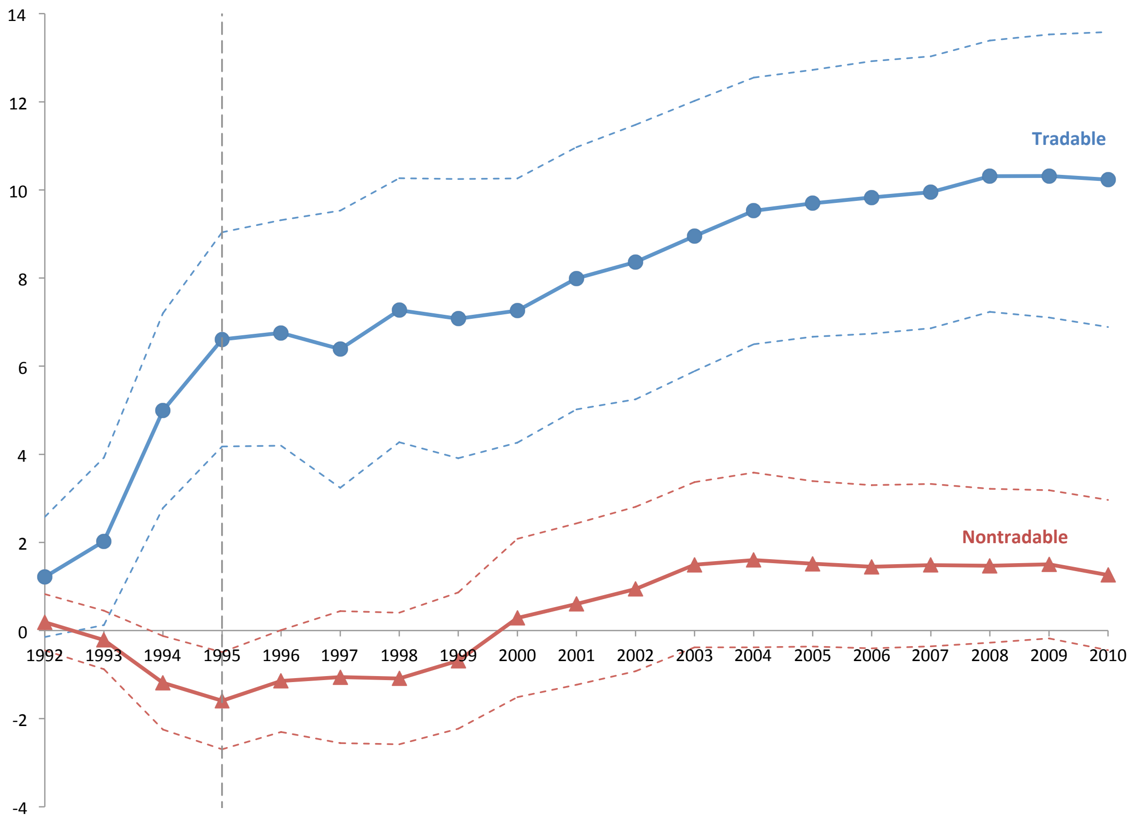
Each point reflects an individual regression coefficient where the dependent variable is the change in regional log formal employment from 1991 to the year listed on the x-axis, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with state fixed effects and pre-trend control. Positive estimates imply larger formal employment declines in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 7: Regional log Formal Employment - Pre-Trends - 1987-1990



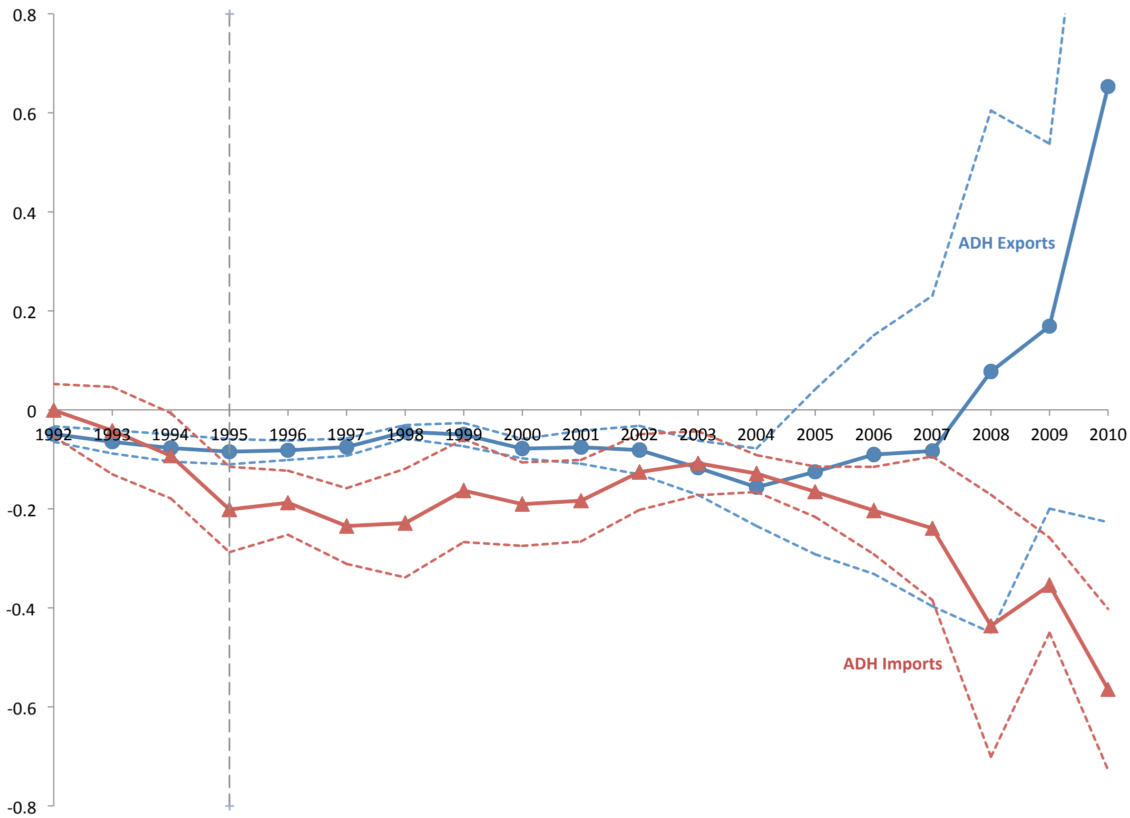
Each point reflects an individual regression coefficient where the dependent variable is the change in regional log formal employment from 1986 to the year listed on the x-axis, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with state fixed effects. Positive estimates imply larger formal employment declines in regions facing larger tariff declines. Standard errors adjusted for 112 mesoregion clusters.

Figure 8: Regional log Formal Employment - Tradable and Nontradable Sectors- 1992-2010



Each point reflects an individual regression coefficient where the dependent variable is the change in regional log formal employment in the tradable or nontradable sector from 1991 to the year listed on the x-axis, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with state fixed effects and pre-trend control. Positive estimates imply larger formal employment declines in the relevant sector in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

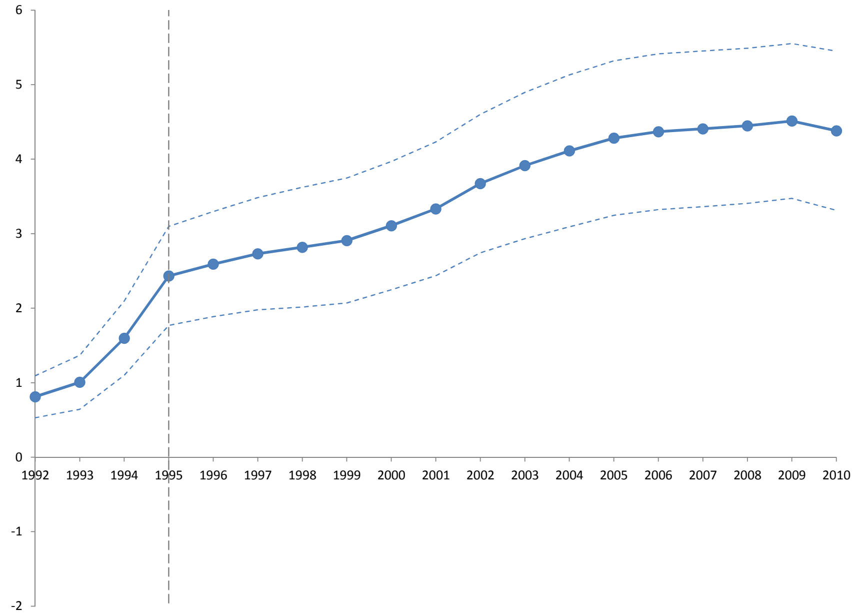
Figure 9: Regional Imports per Worker, and Exports per Worker - 1992-2010



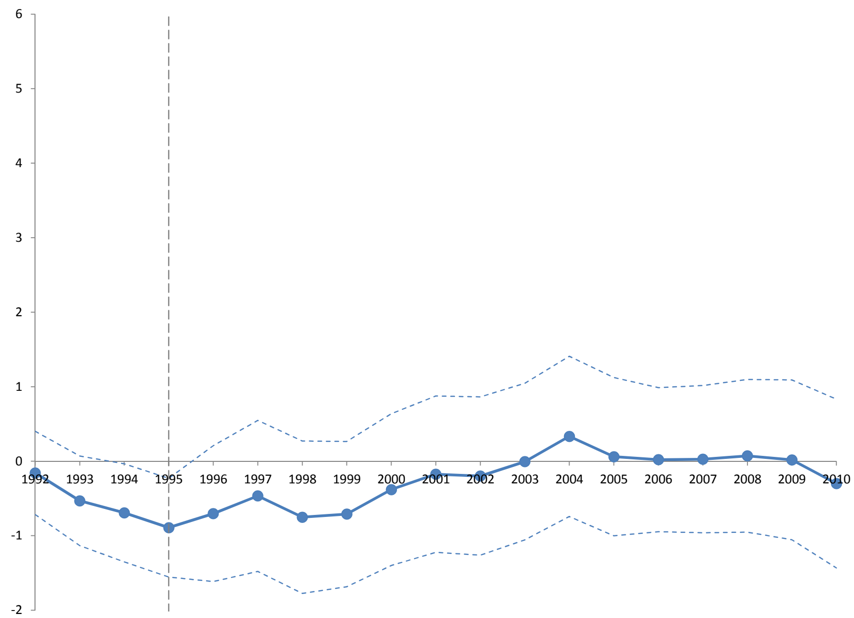
Each point reflects an individual regression coefficient where the dependent variable is the change in log employment (from RAIS), change in imports per worker, or change in exports per worker, calculated from 1991 to the year listed on the x-axis (see text for details). For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with state fixed effects and pre-trend control. Positive (negative) estimates imply that regions facing larger tariff declines experienced larger declines (increases) in exports per worker, or imports per worker. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters. Regional imports and exports per worker calculated as in Autor et al. (2013)

Figure 10: Capital Adjustment

Panel A: Regional log Number of Formal Establishments - 1992-2010

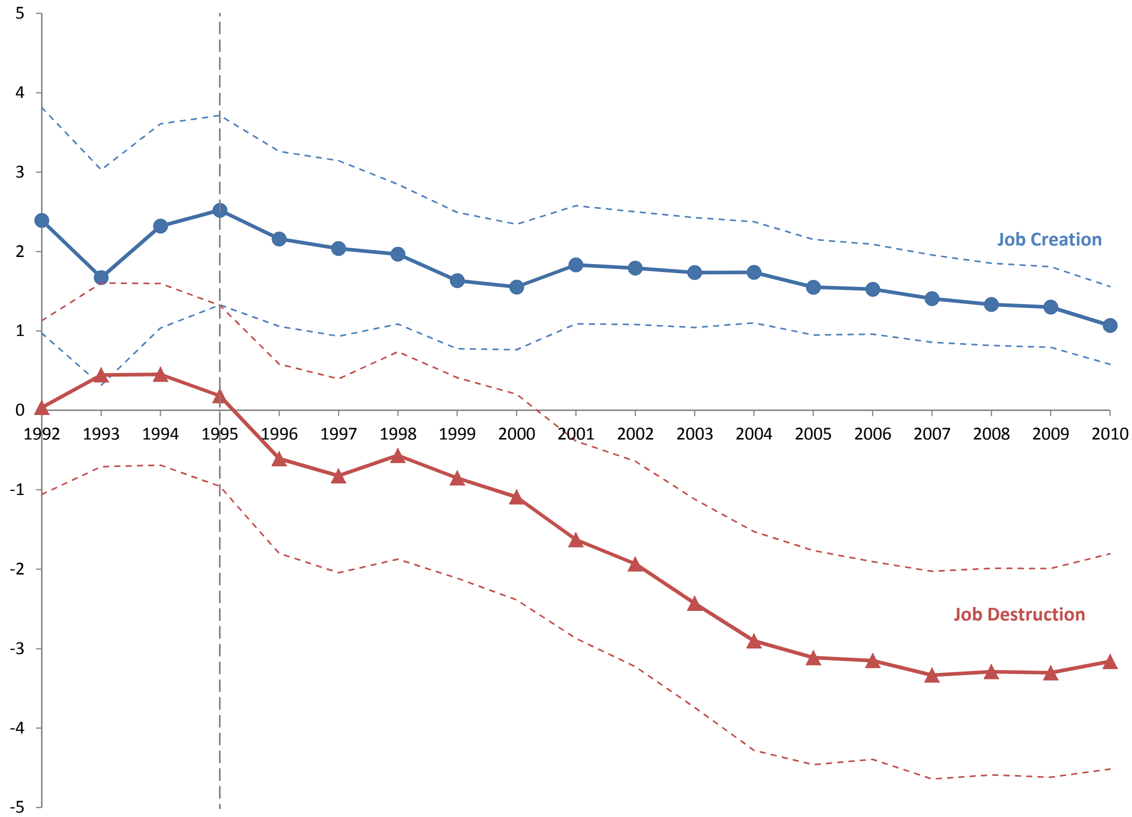


Panel B: Regional log Average Formal Establishment Size (Number of Workers) - 1992-2010



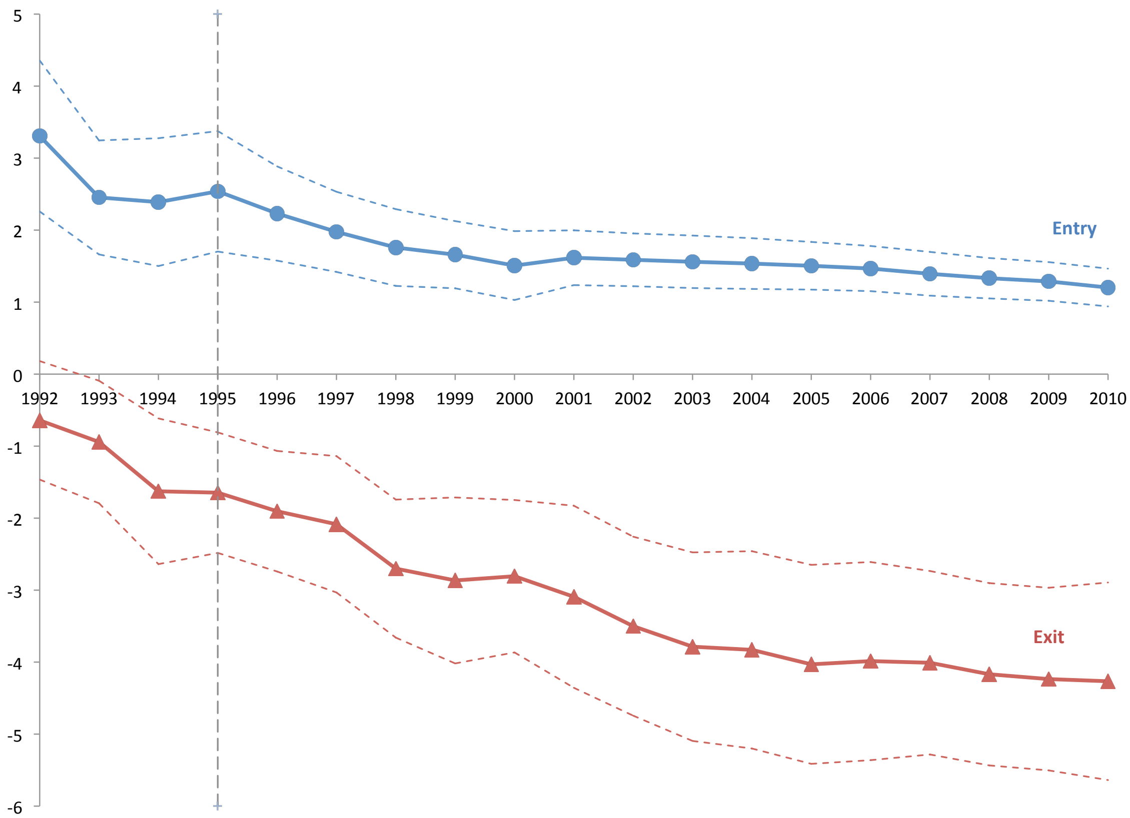
Each point reflects an individual regression coefficient where the dependent variable is the change in log number of establishments or log average establishment size, calculated from 1991 to the year listed on the x-axis using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with state fixed effects and pre-trend control. Positive estimates imply larger declines in the number or average size of formal establishments in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 11: Cumulative Regional log Job Creation and Job Destruction - 1992-2010



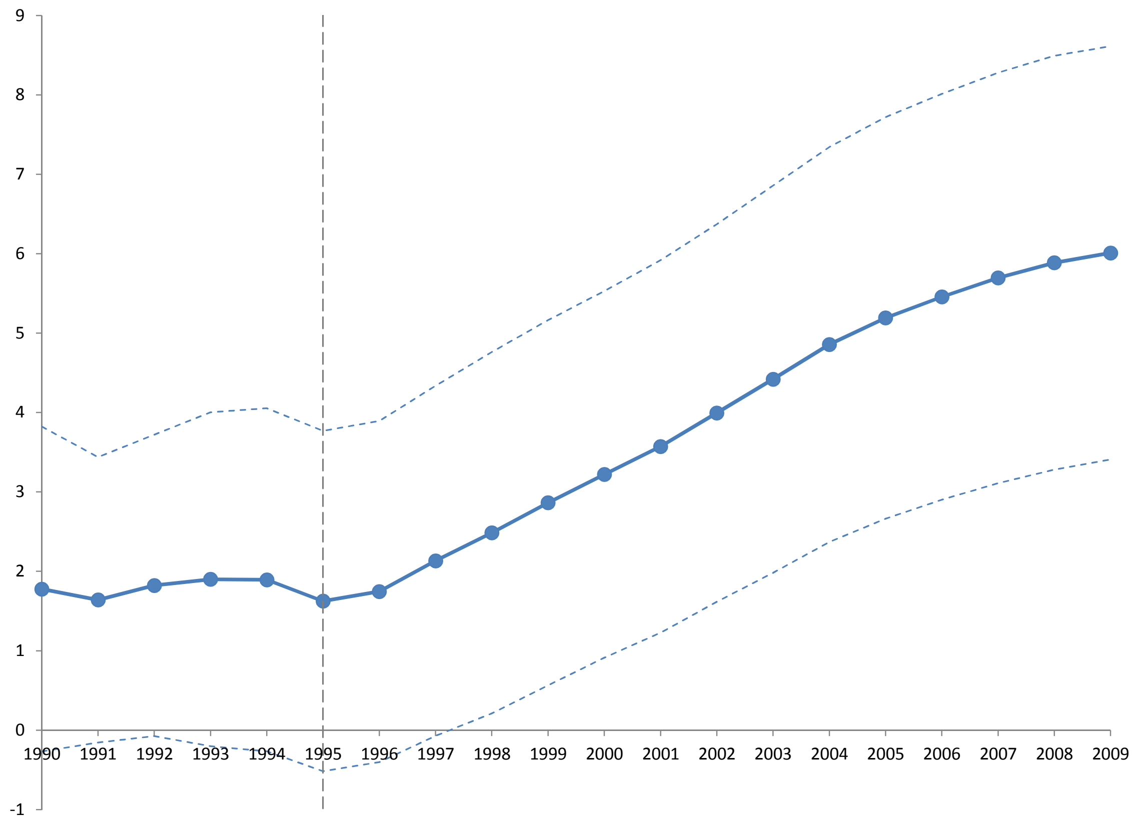
Each point reflects an individual regression coefficient where the dependent variable is the log job creation or log job destruction from 1991 to the year listed on the x-axis, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with state fixed effects and pre-trend control. Positive job creation estimates imply larger declines in cumulative job creation in regions facing larger tariff declines. Negative job destruction estimates imply larger increases in cumulative job destruction in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters. Job creation and destruction measured as in Davis and Haltiwanger (1990).

Figure 12: Regional Formal Establishment Entry and Exit - 1992-2010



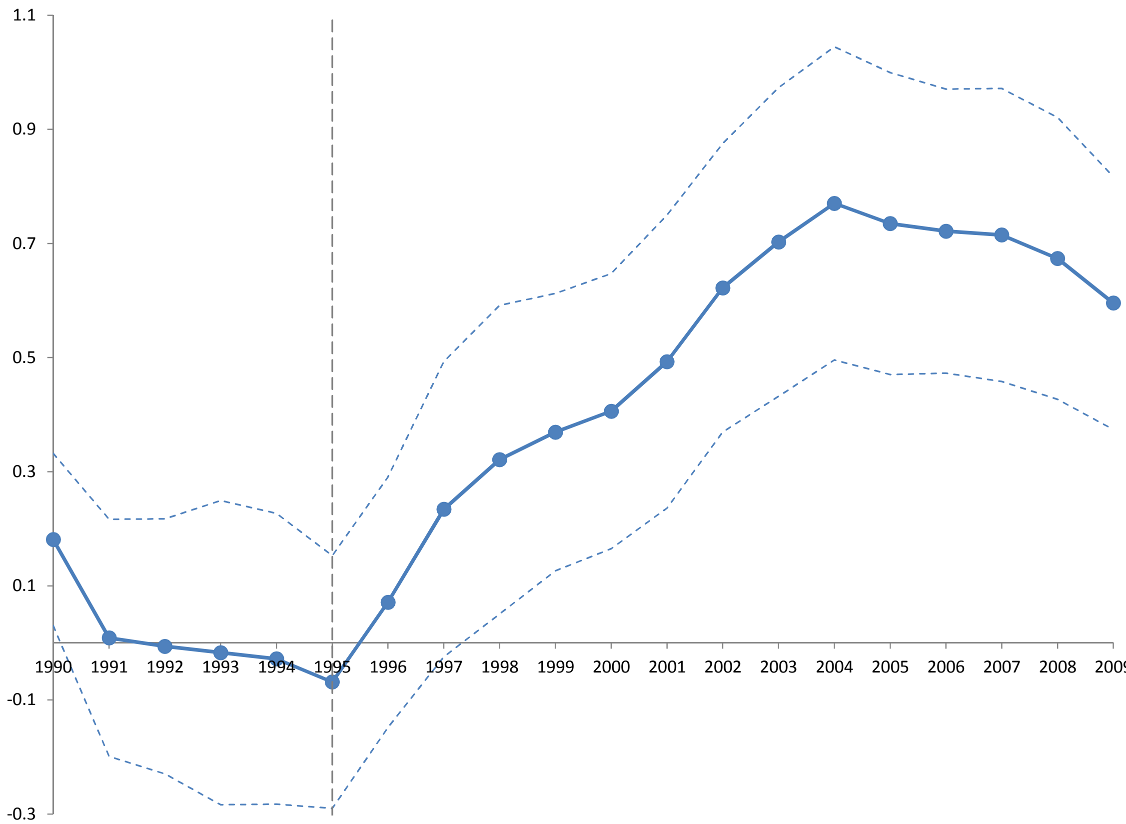
Each point reflects an individual regression coefficient where the dependent variable is the rate of establishment entry or exit from 1991 to the year listed on the x-axis, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with state fixed effects and pre-trend control. Positive entry estimates imply larger declines in firm entry rates in regions facing larger tariff declines. Negative exit estimates imply larger increases in firm exit rates in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 13: Individual Average Months Formally Employed - Tradable Sector Worker Sample - 1990-2009



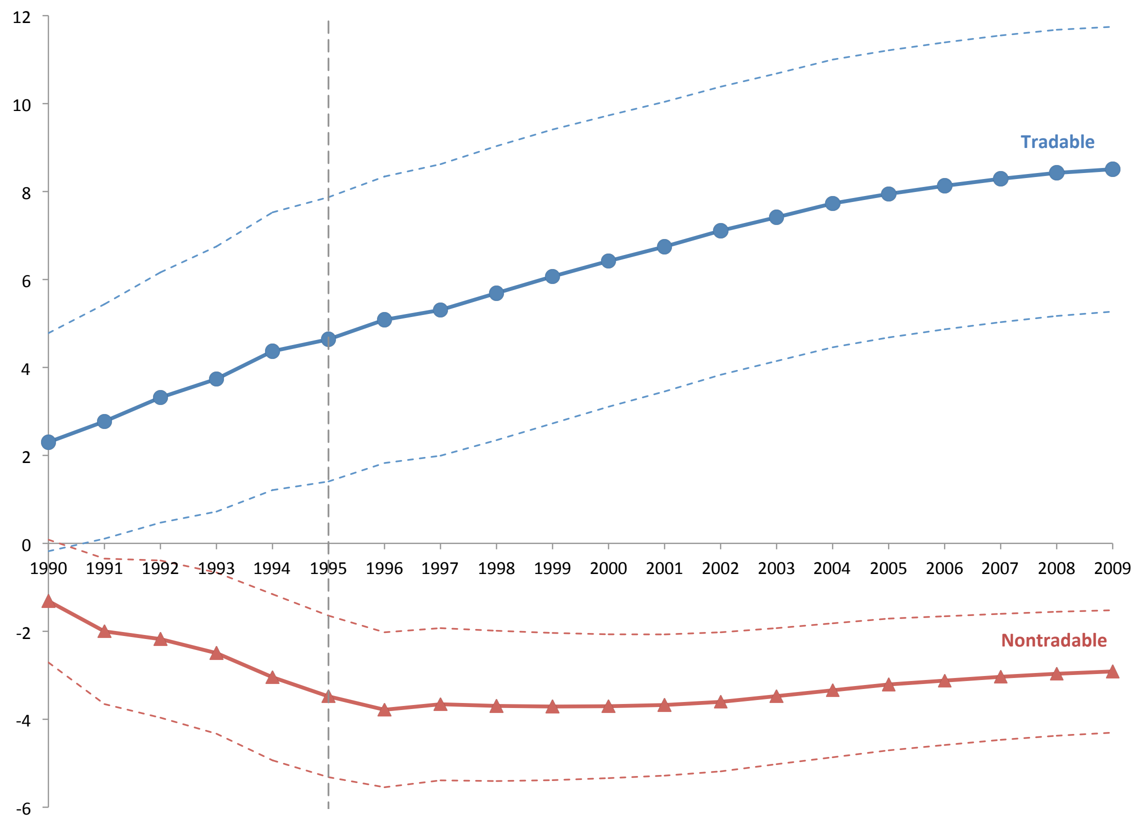
Each point reflects an individual regression coefficient where the dependent variable is the average number of months formally employed per year from 1990 to the year listed on the x-axis, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). Positive estimates imply fewer average months formally employed for workers initially employed in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 14: Fraction of the Year in Formal Employment - Tradable Sector Worker Sample - 1990-2009



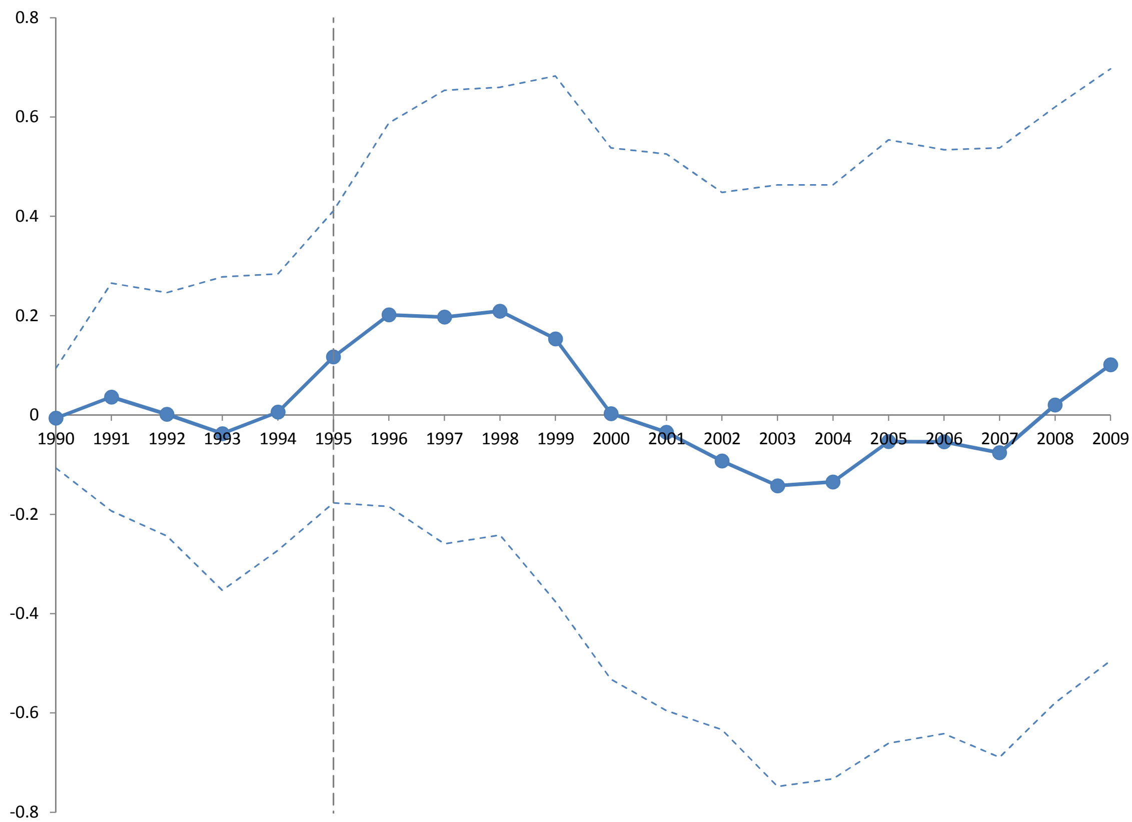
Each point reflects an individual regression coefficient where the dependent variable is the fraction of the year spent in formal employment for the year listed on the x-axis, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). Positive estimates imply smaller fractions of the year in formal employment for workers initially employed in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 15: Individual Average Months Formally Employed, by Sector - Tradable Sector Worker Sample - 1990-2009



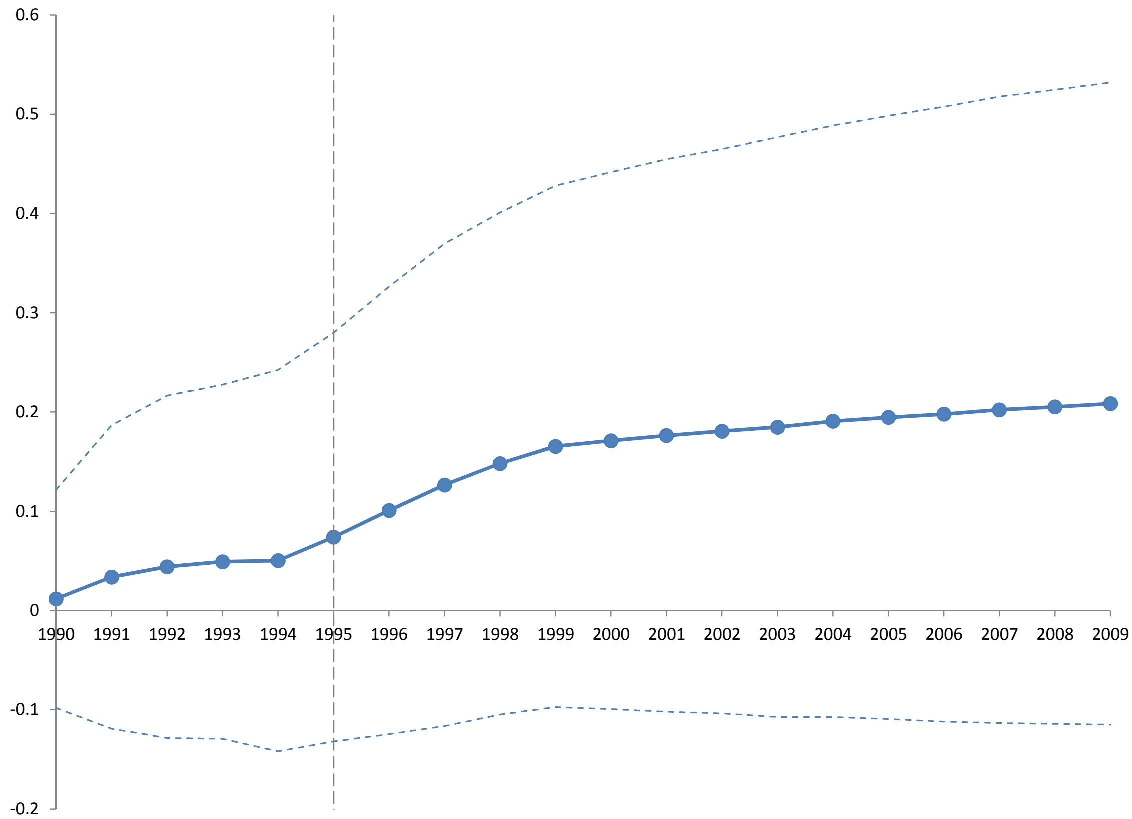
Each point reflects an individual regression coefficient where the dependent variable is the average number of months per year formally employed in the tradable or nontradable sector from 1990 to the year listed on the x-axis, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). Positive (negative) estimates imply fewer (more) average months formally employed in the relevant sector for workers initially employed in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 16: Individual Probability of Formal Employment Away from the Worker's Initial Region - Tradable Sector Worker Sample - 1990-2009



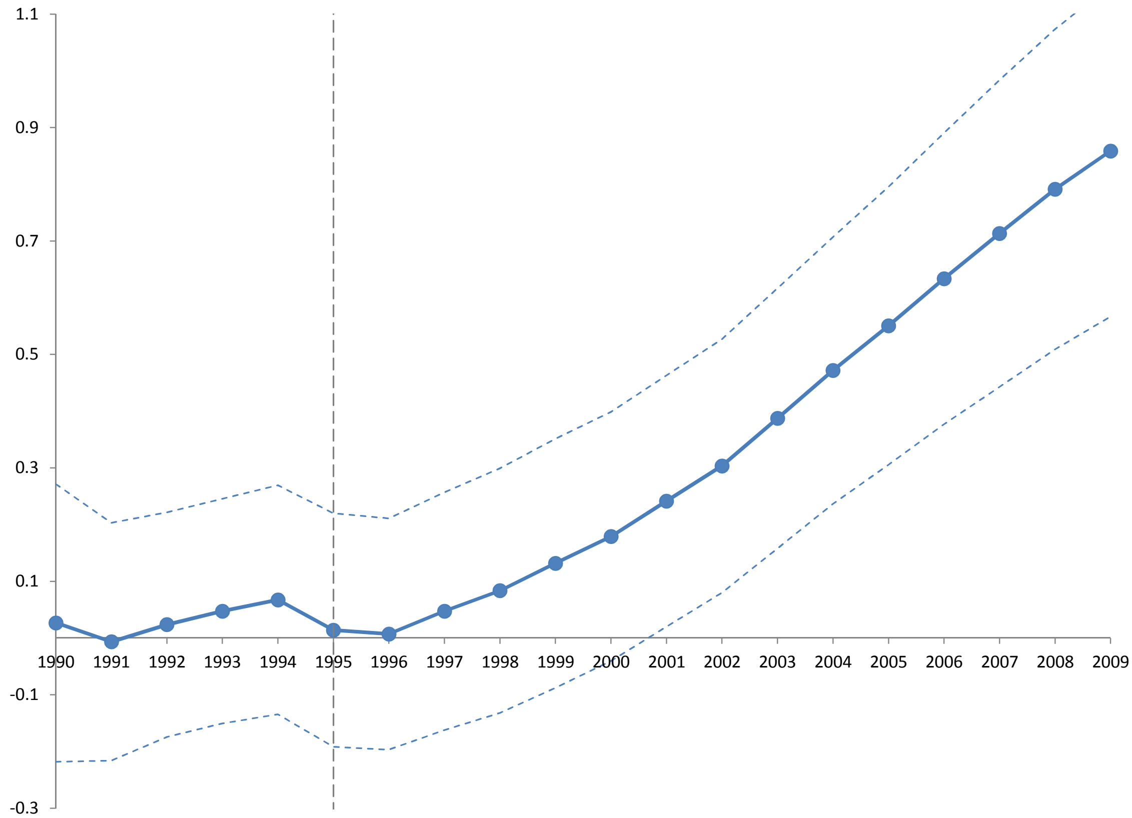
Each point reflects an individual regression coefficient where the dependent variable is an indicator for being observed in a different region in the year listed on the x-axis than in 1989, when the sample was drawn, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). Positive estimates imply a lower probability of being employed away from the initial region for workers initially employed in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 17: Individual Share of Formally Employed Months Away from Initial Region - Tradable Sector Worker Sample - 1990-2009



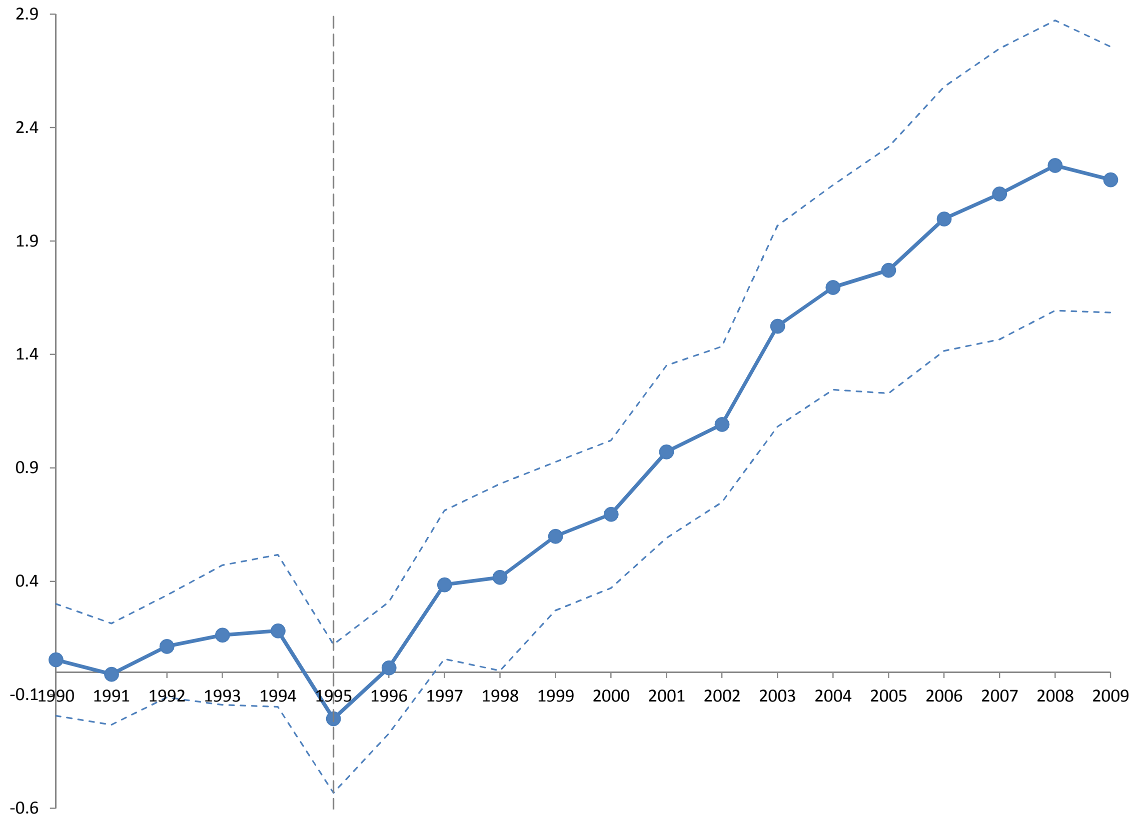
Each point reflects an individual regression coefficient where the dependent variable is the share of formally employed months in the year listed on the x-axis that the individual spent working in a region other than their initial region, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). Positive estimates imply a lower share of formally employed months away from the initial region for workers initially employed in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 18: Individual Average Earnings - Tradable Sector Worker Sample - 1990-2009



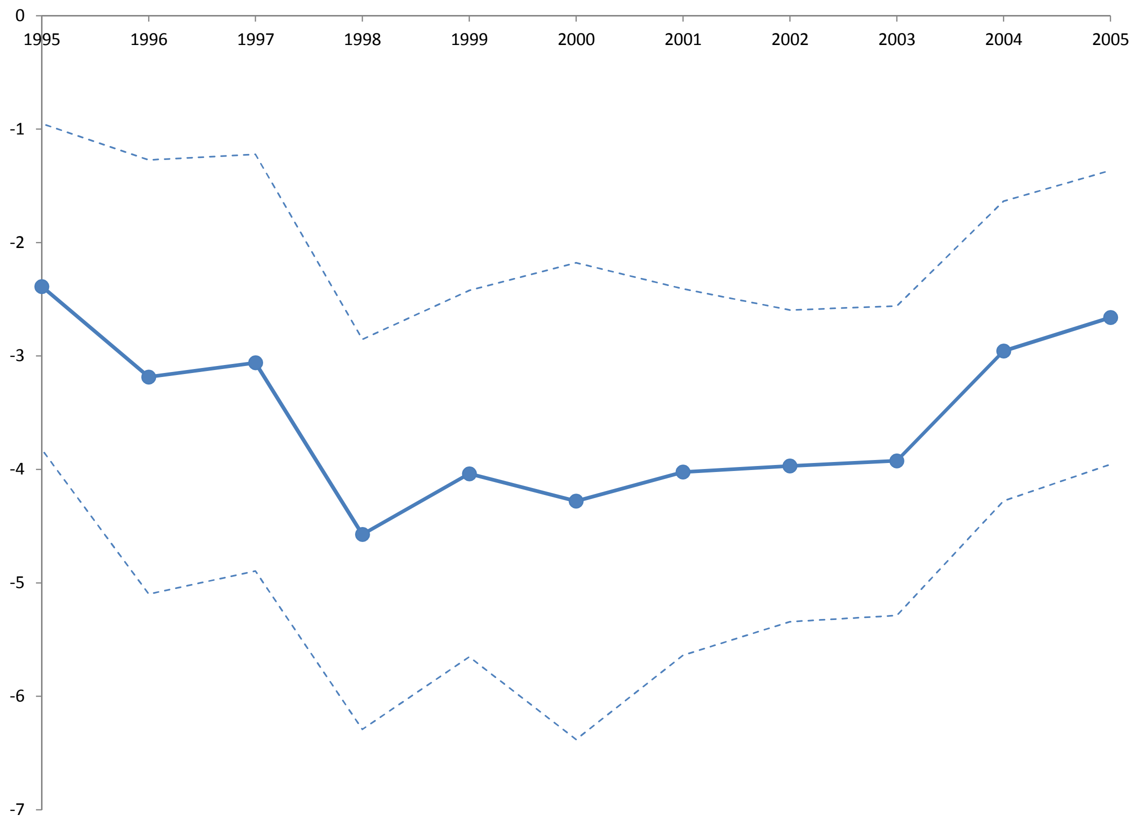
Each point reflects an individual regression coefficient where the dependent variable is the average earnings from 1990 to the year listed on the x-axis as a multiple of the worker's average pre-liberalization (1986-89) earnings, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). Positive estimates imply larger declines in average formal earnings for workers initially employed in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 19: Individual Yearly Earnings - Tradable Sector Worker Sample - 1990-2009



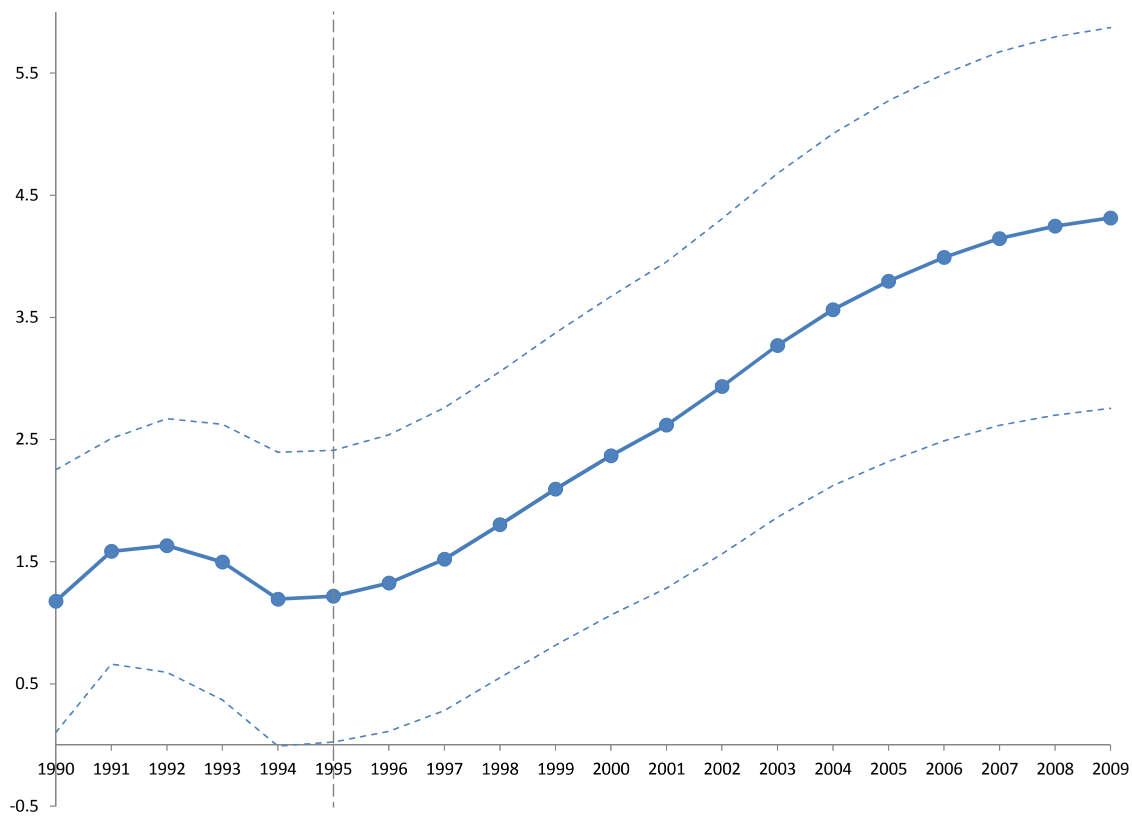
Each point reflects an individual regression coefficient where the dependent variable is the yearly earnings in the year listed on the x-axis as a multiple of the worker's average pre-liberalization (1986-89) earnings, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). Positive estimates imply larger declines in yearly formal earnings for workers initially employed in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 20: log Non-Formal Spell Length (Tobit) - Job Separations Originating in Tradable Sector - 1995-2005



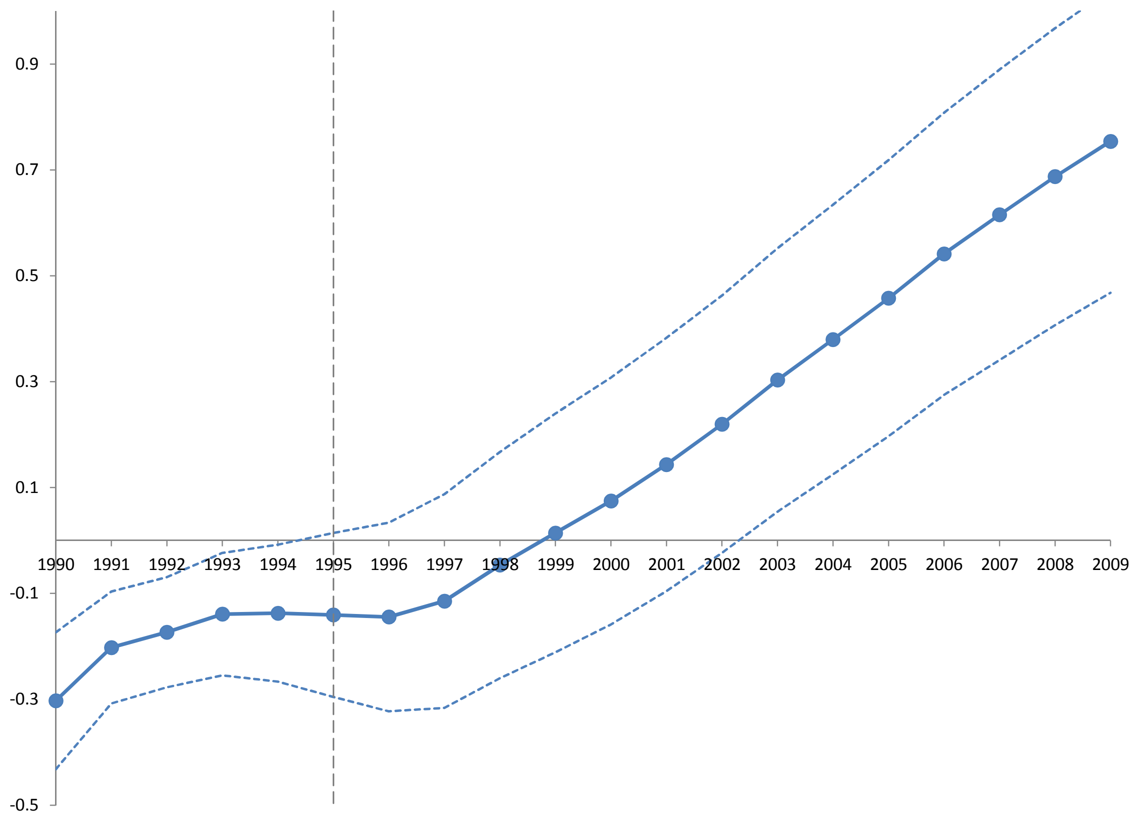
Each point reflects an individual Tobit regression coefficient where observations represent individual spells, and the dependent variable is the log spell length in years, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text) in the worker's region at the beginning of the nonformal spell, with various worker-level controls and pre-trend control (described in the text). Negative estimates imply longer non-formal spell lengths for workers losing formal employment in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 21: Individual Average Months Formally Employed - Nontradable Sector Worker Sample - 1990-2009



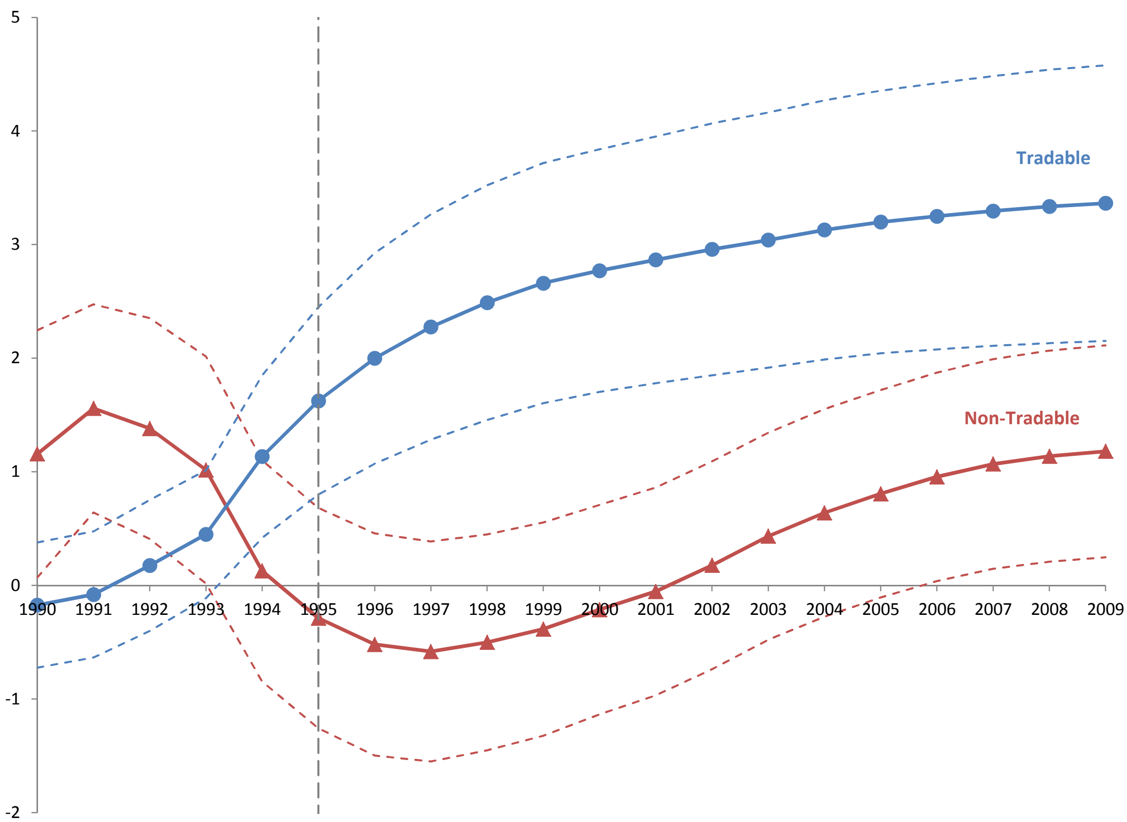
Each point reflects an individual regression coefficient where the dependent variable is the average number of months formally employed per year from 1990 to the year listed on the x-axis, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). Positive estimates imply fewer average months formally employed for workers initially employed in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 22: Individual Average Earnings - Nontradable Sector Worker Sample - 1990-2009



Each point reflects an individual regression coefficient where the dependent variable is the average earnings from 1990 to the year listed on the x-axis as a multiple of the worker's average pre-liberalization (1986-89) earnings, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). Positive estimates imply larger declines in average formal earnings for workers initially employed in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Figure 23: Individual Average Months Formally Employed, by Sector - Nontradable Sector Worker Sample 1990-2009



Each point reflects an individual regression coefficient where the dependent variable is the average number of months formally employed in the relevant sector per year from 1990 to the year listed on the x-axis, calculated using RAIS. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). Positive (negative) estimates imply fewer (more) average months formally employed in the relevant sector for workers initially employed in regions facing larger tariff declines. Vertical bar indicates that liberalization was complete in 1995. Dashed lines show 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

Table 1: Regional log Formal Earnings Premia - 2000, 2010

Change in log formal earnings premium:		1991-2000			1991-2010	
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: RAIS						
Regional tariff change	0.472*	0.536**	0.358	1.891***	1.898***	1.665***
	(0.256)	(0.237)	(0.217)	(0.478)	(0.327)	(0.298)
Formal earnings pre-trend (86-90)			-0.344**			-0.467***
			(0.150)			(0.144)
State fixed effects (26)		✓	✓		✓	✓
R-squared	0.019	0.183	0.236	0.142	0.409	0.454
Panel B: Census Earnings						
Regional tariff change	0.379	0.399**	0.353*	1.342*	1.178***	1.126***
	(0.459)	(0.198)	(0.195)	(0.735)	(0.292)	(0.273)
Formal earnings pre-trend (86-90)			-0.098*			-0.113
			(0.054)			(0.075)
State fixed effects (26)		✓	✓		✓	✓
R-squared	0.013	0.583	0.587	0.071	0.707	0.710
Panel C: Census Hourly Wages						
Regional tariff change	0.584	0.631***	0.572***	1.229*	0.957***	0.894***
	(0.477)	(0.199)	(0.191)	(0.670)	(0.257)	(0.241)
Formal earnings pre-trend (86-90)			-0.126**			-0.139**
			(0.056)			(0.066)
State fixed effects (26)		✓	✓		✓	✓
R-squared	0.029	0.595	0.601	0.057	0.704	0.708

Positive coefficient estimates for the regional tariff change imply larger declines in formal earnings/wages in regions facing larger tariff declines. 475 microregion observations. Standard errors (in parentheses) adjusted for 112 mesoregion clusters. Weighted by the inverse of the squared standard error of the estimated change in log formal earnings or wage premium. Pre-trends computed for 1986-1990 period. *** Significant at the 1 percent, ** 5 percent, * 10 percent level.

Table 2: Regional log Formal Employment - 2000, 2010

Change in log formal employment:		1991-2000			1991-2010	
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: RAIS						
Regional tariff change	2.934***	2.659***	2.762***	4.983***	4.326***	4.380***
	(0.842)	(0.825)	(0.760)	(0.866)	(0.812)	(0.795)
Formal employment pre-trend (86-90)			-0.413**			-0.216
			(0.199)			(0.187)
State fixed effects (26)		✓	✓		✓	✓
R-squared	0.035	0.273	0.300	0.081	0.400	0.405
Panel B: Census						
Regional tariff change	2.319**	1.667***	1.430***	3.330**	2.340***	2.074**
	(0.939)	(0.563)	(0.537)	(1.424)	(0.886)	(0.863)
Formal employment pre-trend (86-90)			0.224***			0.262***
			(0.071)			(0.099)
State fixed effects (26)		✓	✓		✓	✓
R-squared	0.113	0.542	0.563	0.097	0.555	0.568

Positive coefficient estimates for the regional tariff change imply larger declines in formal employment in regions facing larger tariff declines. 475 microregion observations. Standard errors (in parentheses) adjusted for 112 mesoregion clusters. Census results weighted by the inverse of the squared standard error of the estimated change in log formal employment. Pre-trends computed for 1986-1990 period. *** Significant at the 1 percent, ** 5 percent, * 10 percent level.

Table 3: Robustness: Regional log Formal Earnings Premia - 1995, 2000, 2005, 2010

Change in log formal earnings premium:	1991-1995 (1)	1991-2000 (2)	1991-2005 (3)	1991-2010 (4)
Panel A: Post-liberalization tariff changes				
Regional tariff change	0.089	0.319	0.970***	1.847***
	(0.189)	(0.224)	(0.299)	(0.304)
Post-liberalization tariff change	n/a	2.305	10.451**	1.708
		(2.921)	(4.284)	(1.042)
Panel B: Imports and exports (ADH)				
Regional tariff change	0.083	0.363	1.183***	1.542***
	(0.193)	(0.253)	(0.319)	(0.386)
Imports (ADH)	0.003	-0.459**	-0.181	-0.127
	(0.124)	(0.231)	(0.493)	(0.186)
Exports (ADH)	-0.359	1.986	-0.128	-0.051
	(0.950)	(1.219)	(0.375)	(0.067)
Panel C: Real exchange rate				
Regional tariff change	0.133	0.358	1.244***	1.674***
	(0.190)	(0.227)	(0.313)	(0.293)
Real exchange rate change - trade weighted	-0.226	-0.001	-0.092	0.832**
	(0.304)	(0.362)	(0.590)	(0.415)
Panel D: Privatization				
Regional tariff change	0.089	0.328	1.187***	1.585***
	(0.189)	(0.218)	(0.243)	(0.303)
Change in state-owned share of empl.	n/a	0.1785	0.4509**	0.3504
		(0.181)	(0.223)	(0.234)
Panel E: Individual Fixed Effects				
Regional tariff change	0.175	0.265	1.011***	1.201***
	(0.174)	(0.199)	(0.211)	(0.254)
Panel F: 1980-1991 Census Pre-Trends				
Regional tariff change	0.315	0.500*	1.364***	1.832***
	(0.207)	(0.260)	(0.281)	(0.361)
Panel G: Manufacturing Sector Earnings				
Regional tariff change	0.626**	0.630**	1.700***	2.160***
	(0.262)	(0.316)	(0.379)	(0.462)
Panel H: Nontradable Sector Earnings				
Regional tariff change	0.078	0.284	1.221***	1.637***
	(0.238)	(0.259)	(0.268)	(0.299)
Panel I: Sample restriction - above median urban share - 238 obs				
Regional tariff change	0.072	0.272	1.297***	1.939***
	(0.205)	(0.232)	(0.277)	(0.343)
Panel J: Sample restriction - above median formal share - 238 obs				
Regional tariff change	0.182	0.412*	1.417***	1.907***
	(0.205)	(0.223)	(0.244)	(0.311)
Panel K: Alternate tariff measure - effective rate of protection (ERP)				
Regional ERP change	0.009	0.165	0.669***	0.886***
	(0.106)	(0.133)	(0.151)	(0.178)
Panel L: Alternate RTC measure - employment shares including formal and informal labor				
Alternate regional tariff change	-0.000	0.483***	1.367***	1.833***
	(0.150)	(0.167)	(0.167)	(0.215)
Formal earnings pre-trend (86-90)	✓	✓	✓	✓
State fixed effects (26)	✓	✓	✓	✓

Positive coefficient estimates for the regional tariff change imply larger declines in formal earnings in regions facing larger tariff declines. 475 microregion observations, unless otherwise noted. Outcomes calculated using RAIS data. Standard errors (in parentheses) adjusted for 112 mesoregion clusters. Weighted by the inverse of the squared standard error of the estimated change in log formal earnings premium. Pre-trends computed for 1986-1990 period, unless otherwise noted. *** Significant at the 1 percent, ** 5 percent, * 10 percent level.

Table 4: Evolution of Formal Earnings and Employment by RTC Quartile

Regional Tariff Change quartile:	1 most negative	2	3	4 most positive
1991 log real formal earnings level	5.79	5.60	5.53	5.44
change in log real formal earnings				
1991-1995	0.25	0.27	0.25	0.25
1991-2000	0.24	0.25	0.27	0.29
1991-2005	0.23	0.29	0.34	0.38
1991-2010	0.49	0.59	0.64	0.68
change in log formal employment				
1991-1995	0.05	0.09	0.11	0.11
1991-2000	0.10	0.16	0.20	0.23
1991-2005	0.19	0.30	0.34	0.37
1991-2010	0.32	0.43	0.47	0.49

Formal earnings and employment from RAIS data. Regional Tariff Changes calculated based on equation 3. See text for details. 475 microregion observations. Earnings deflated using the national consumer price index. Regional earnings calculated as average regional residuals after controlling for regional worker composition with demographic characteristics and industry fixed effects.

Table 5: Individual Analysis Summary Statsitics

	Tradable Sector Sample		Nontradable Sector Sample	
	mean	std. dev.	mean	std. dev.
Education				
Illiterate	0.02	0.14	0.01	0.11
4th grade incomplete	0.13	0.33	0.10	0.30
4th grade complete	0.25	0.43	0.18	0.38
8th grade incomplete	0.19	0.39	0.14	0.34
8th grade complete	0.15	0.35	0.14	0.35
High School incomplete	0.05	0.21	0.06	0.24
High School complete	0.13	0.33	0.21	0.40
College incomplete	0.02	0.15	0.04	0.19
College complete	0.07	0.26	0.12	0.33
Female	0.24	0.43	0.32	0.46
Age	32.8	5.4	32.8	5.5
December 1989 Earnings (in 2010 R$)	1,911	2,459	1,838	2,670
1989 Yearly Earnings (in 2010 R$)	19,172	23,750	18,701	26,087
Average Annualized Earnings 1986-1989 (in 2010 R$)	19,010	21,043	18,077	21,653
Months formally employed per year				
1990	10.2	3.4	9.9	3.8
1990-1995	8.2	3.8	8.2	3.9
1990-2000	7.1	3.7	7.2	3.9
1990-2005	6.4	3.7	6.6	3.9
1990-2009	6.0	3.7	6.2	3.9
Employed in a different region in 1994 than in 1989	0.09	0.29	0.11	0.31
Employed in a different region in 2000 than in 1989	0.11	0.31	0.12	0.32
Observations	587,575		973,048	

RAIS data. Weighted to account for 15% sample of individuals in regions with more than 2000 traded sector workers in 1989 and 100% sample in other regions. All monetary values reported in 2010 R$. In Dec 31, 2010, a US dollar was worth 1.66 Brazilian Reais.

Table 6: Individual Probability of In-Migration, 1995, 2000, 2005, 2009

	(1)	(2)	(3)
In-Migration Indicator:	1989 to *t*	*t-5* to *t*	*t-1* to *t*
Panel A: 1995 sample			
Regional tariff change	0.633	0.317**	0.476
	(0.393)	(0.157)	(0.335)
Observations	529,940	532,384	613,067
R-squared	0.098	0.094	0.047
Panel B: 2000 Sample			
Regional tariff change	0.637***	0.364**	0.150**
	(0.179)	(0.176)	(0.075)
Observations	399,197	437,615	517,185
R-squared	0.126	0.106	0.050
Panel C: 2005 Sample			
Regional tariff change	0.997***	0.488***	0.170***
	(0.171)	(0.149)	(0.039)
Observations	351,495	387,428	478,112
R-squared	0.128	0.092	0.049
Panel D: 2009 Sample			
Regional tariff change	1.082***	0.268**	0.124***
	(0.158)	(0.104)	(0.034)
Observations	320,840	387,620	460,446
R-squared	0.125	0.090	0.054

Individual-level panel data from RAIS. Each panel represents a separate sample of individuals present in RAIS in the year listed. The dependent variable varies by column. In column (1) it is an indicator for having lived in a different region in 1989 than in the reference year listed in the panel title. In columns (2) and (3) is is an indicator for having lived in a different region five years or one year before the reference year, respectively. Positive coefficient estimates for the regional tariff change imply larger declines in in-migration to regions facing larger tariff declines. All specifications include controls for 1986-89 in-migration probability premia after controlling for individual characteristics. Standard errors adjusted for 112 mesoregion clusters. *** Significant at the 1 percent, ** 5 percent, * 10 percent level.

Table 7: Regional log Working-Age Population

Change in log working-age population:	OLS				IV	
					1980 pop.	70-80 Δpop.
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Census - 1991-2000						
Regional tariff change	-0.176	-0.835***	-0.331	0.055	-0.022	-0.085
	(0.276)	(0.159)	(0.243)	(0.330)	(0.092)	(0.166)
Population pre-trend (80-91)			0.407**		0.656***	0.605***
			(0.164)		(0.038)	(0.106)
Population pre-trend (70-80)				0.295***		
				(0.073)		
State fixed effects (26)		✓	✓	✓	✓	✓
R-squared	0.009	0.399	0.657	0.557		
First stage partial F-Statistic					212.6	32.31
Panel B: Census - 1991-2010						
Regional tariff change	-0.174	-1.175***	-0.389	0.162	0.078	-0.030
	(0.432)	(0.285)	(0.320)	(0.474)	(0.163)	(0.233)
Population pre-trend (80-91)			0.634***		1.010***	0.923***
			(0.225)		(0.059)	(0.147)
Population pre-trend (70-80)				0.441***		
				(0.087)		
State fixed effects (26)		✓	✓	✓	✓	✓
R-squared	0.003	0.422	0.669	0.554		
First stage partial F-Statistic					197.9	31.54

Decennial Census data. Positive (negative) coefficient estimates for the regional tariff change imply larger decreases (increases) in working-age population in regions facing larger tariff declines. 411 microregion observations. Standard errors (in parentheses) adjusted for 112 mesoregion clusters. Weighted by the inverse of the squared standard error of the estimated change in the relevant employment share. Pre-trends computed for 1980-1991 and 1970-1980 periods. To address potential endogeneity of 1980-1991 pre-trend due to overlap with dependent variable in 1991, 1980-1991 pre-trends are instrumented using the 1980 population level in column (5) and the 1970-1980 trend in column (6). First-stage partial F-statistics reported for instrumental variables specifications. *** Significant at the 1 percent, ** 5 percent, * 10 percent level.

Table 8: Informal Share of Employment

	1991	2000	2010
Overall	0.58	0.64	0.49
Agriculture	0.89	0.86	0.83
Mining	0.61	0.45	0.21
Manufacturing	0.28	0.39	0.29
Nontradable	0.55	0.64	0.48

Decennial Census data.

Table 9: Informally Employed and Not-Employed Share of Regional Working-Age Population - 2000, 2010

Change in share:	1991-2000				1991-2010			
	OLS			IV 1980 share	OLS			IV 1980 share
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A: Informally employed								
Regional tariff change	-0.182***	-0.174***	-0.171***	-0.083*	-0.458***	-0.461***	-0.485***	-0.339***
	(0.051)	(0.045)	(0.051)	(0.046)	(0.085)	(0.063)	(0.066)	(0.073)
Dependent variable pre-trend (80-91)			0.009	0.305***			-0.079	0.422***
			(0.041)	(0.071)			(0.060)	(0.123)
State fixed effects (26)		✓	✓	✓		✓	✓	✓
R-squared	0.114	0.334	0.334		0.293	0.569	0.574	
First stage partial F-Statistic				50.20				50.04
Panel B: Not employed								
Regional tariff change	-0.312***	-0.314***	-0.303***	-0.301***	0.126*	0.049	0.026	0.085
	(0.034)	(0.038)	(0.043)	(0.046)	(0.071)	(0.053)	(0.057)	(0.085)
Dependent variable pre-trend (80-91)			0.030	0.036			-0.066	0.107
			(0.045)	(0.070)			(0.056)	(0.149)
State fixed effects (26)		✓	✓	✓		✓	✓	✓
R-squared	0.270	0.474	0.475		0.026	0.579	0.582	
First stage partial F-Statistic				80.47				82.88

Decennial Census data. Negative (positive) coefficient estimates for the regional tariff change imply larger increases (decreases) in informal share or non-employed share in regions facing larger tariff declines. 411 microregion observations. Standard errors (in parentheses) adjusted for 112 mesoregion clusters. Weighted by the inverse of the squared standard error of the estimated change in the relevant employment share. Pre-trends computed for 1980-1991 period. To address endogeneity of pre-trend due to overlap with dependent variable in 1991, pre-trends are instrumented using the corresponding 1980 employment category share. First-stage partial F-statistics reported for instrumental variables specifications. *** Significant at the 1 percent, ** 5 percent, * 10 percent level.

Table 10: Formal/Informal x Tradable/Nontradable Share of Regional Employment - 2000, 2010

Change in share of employment	2000				2010			
	OLS			IV 1980 share	OLS			IV 1980 share
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A: Formal tradable share								
Regional tariff change	0.461***	0.433***	0.386***	0.162***	0.521***	0.560***	0.452***	0.205**
	(0.082)	(0.044)	(0.040)	(0.051)	(0.117)	(0.077)	(0.051)	(0.085)
Dependent var. pre-trend (80-91)			0.180*	1.033***			0.396***	1.302***
			(0.099)	(0.163)			(0.145)	(0.205)
State fixed effects (26)		✓	✓	✓		✓	✓	✓
R-squared	0.454	0.683	0.698		0.303	0.575	0.620	
First stage partial F-Statistic				32.81				30.36
Panel B: Formal nontradable share								
Regional tariff change	0.009	0.121***	0.070	-0.158***	-0.057	0.113*	0.102	-0.080
	(0.041)	(0.042)	(0.060)	(0.044)	(0.053)	(0.063)	(0.094)	(0.084)
Dependent var. pre-trend (80-91)			0.105	0.579***			0.022	0.401***
			(0.075)	(0.073)			(0.115)	(0.118)
State fixed effects (26)		✓	✓	✓		✓	✓	✓
R-squared	0.000	0.376	0.390		0.006	0.590	0.590	
First stage partial F-Statistic				166.1				178.9
Panel C: Informal tradable share								
Regional tariff change	-0.622***	-0.571***	-0.627***	-0.101	-0.803***	-0.771***	-0.825***	0.041
	(0.053)	(0.044)	(0.051)	(0.206)	(0.079)	(0.078)	(0.096)	(0.277)
Dependent var. pre-trend (80-91)			-0.069*	0.573**			-0.065	0.981***
			(0.042)	(0.244)			-0.058	(0.320)
State fixed effects (26)		✓	✓	✓		✓	✓	✓
R-squared	0.535	0.704	0.708		0.486	0.682	0.684	
First stage partial F-Statistic				15.08				16.01
Panel D: Informal nontradable share								
Regional tariff change	0.011	-0.045	-0.060	-0.440**	0.073	-0.023	-0.135*	-0.879**
	(0.052)	(0.047)	(0.048)	(0.224)	(0.103)	(0.073)	(0.073)	(0.368)
Dependent var. pre-trend (80-91)			-0.076	-1.969*			-0.542***	-4.142**
			(0.090)	(1.059)			(0.086)	(1.722)
State fixed effects (26)		✓	✓	✓		✓	✓	✓
R-squared	0.000	0.329	0.332		0.008	0.454	0.534	
First stage partial F-Statistic				4.351				4.985

Decennial Census data. Positive (negative) coefficient estimates for the regional tariff change imply larger decreases (increases) in the relevant share in regions facing larger tariff declines. 411 microregion observations. Standard errors (in parentheses) adjusted for 112 mesoregion clusters. Weighted by the inverse of the squared standard error of the estimated change in the relevant employment share. Pre-trends computed for 1980-1991 period. To address endogeneity of pre-trend due to overlap with dependent variable in 1991, pre-trends are instrumented using the corresponding 1980 employment category share. First-stage partial F-statistics reported for instrumental variables specifications. *** Significant at the 1 percent, ** 5 percent, * 10 percent level.

Table 11: Regional log Wage Premia - Informal Workers - 2000, 2010

Change in log wage premium:	OLS				IV 1980 earn
	(1)	(2)	(3)	(4)	(5)
Panel A: Census - 1991-2000					
Regional tariff change	0.010	0.264	-0.000	0.250	0.442
	(0.408)	(0.167)	(0.158)	(0.163)	(0.313)
Earnings pre-trend (80-91)			-0.235***		0.159
			(0.055)		(0.198)
Earnings pre-trend, all workers (70-80)				0.020	
				(0.071)	
State fixed effects (26)		✓	✓	✓	✓
R-squared	0.000	0.646	0.676	0.646	
First stage partial F-Statistic					24.82
Panel B: Census - 1991-2010					
Regional tariff change	-0.312	-0.334	-0.769***	-0.322	0.057
	(0.590)	(0.329)	(0.261)	(0.324)	(0.508)
Earnings pre-trend (80-91)			-0.380***		0.341
			(0.081)		(0.219)
Earnings pre-trend, all workers (70-80)				-0.019	
				(0.101)	
State fixed effects (26)		✓	✓	✓	✓
R-squared	0.007	0.636	0.680	0.636	
First stage partial F-Statistic					27.19

Decennial Census data. Positive (negative) coefficient estimates for the regional tariff change imply larger declines (increases) in informal workers' wages in regions facing larger tariff declines. 411 microregion observations. Standard errors (in parentheses) adjusted for 112 mesoregion clusters. Weighted by the inverse of the squared standard error of the estimated change in the relevant regional wage premium share. Due to a lack of hours information in 1980, pre-trends for *earnings* computed for 1980-1991 and 1970-1980 periods. To address potential endogeneity of 1980-1991 pre-trend due to overlap with dependent variable in 1991, 1980-1991 pre-trends are instrumented using the 1980 earnings premium level in column (5). The 1970-1980 trend is weakly related to the 1980-1991 trend, so this potential instrument is not utilized. First-stage partial F-statistics reported for instrumental variables specification. *** Significant at the 1 percent, ** 5 percent, * 10 percent level.

Table 12: Regional log Wage Premia - All Workers - 2000, 2010

Change in log wage premium:	OLS				IV 1980 earn
	(1)	(2)	(3)	(4)	(5)
Panel A: Census - 1991-2000					
Regional tariff change	0.413	0.432***	0.174	0.487***	0.843***
	(0.342)	(0.141)	(0.137)	(0.148)	(0.326)
Earnings pre-trend (80-91)			-0.274***		0.437*
			(0.056)		(0.260)
Earnings pre-trend (70-80)				-0.088	
				(0.063)	
State fixed effects (26)		✓	✓	✓	✓
R-squared	0.031	0.648	0.691	0.652	
First stage partial F-Statistic					21.17
Panel B: Census - 1991-2010					
Regional tariff change	0.583	0.165	-0.258	0.251	0.829
	(0.510)	(0.292)	(0.233)	(0.294)	(0.531)
Earnings pre-trend (80-91)			-0.438***		0.688**
			(0.077)		(0.302)
Earnings pre-trend (70-80)				-0.151	
				(0.095)	
State fixed effects (26)		✓	✓	✓	✓
R-squared	0.032	0.601	0.666	0.608	
First stage partial F-Statistic					22.42

Decennial Census data. Positive coefficient estimates for the regional tariff change imply larger declines wages in regions facing larger tariff declines. 411 microregion observations. Standard errors (in parentheses) adjusted for 112 mesoregion clusters. Weighted by the inverse of the squared standard error of the estimated change in the relevant regional wage premium share. Due to a lack of hours information in 1980, pre-trends for *earnings* computed for 1980-1991 and 1970-1980 periods. To address potential endogeneity of 1980-1991 pre-trend due to overlap with dependent variable in 1991, 1980-1991 pre-trends are instrumented using the 1980 earnings premium level in column (5). The 1970-1980 trend is weakly related to the 1980-1991 trend, so this potential instrument is not utilized. First-stage partial F-statistics reported for instrumental variables specification. *** Significant at the 1 percent, ** 5 percent, * 10 percent level.

A Data Appendix and Supplemental Results

A.1 Tariffs

Tariff data come from Kume et al. (2003), who report nominal tariffs and effective rates of protection from 1987 to 1998 using the Brazilian industry classification *Nível* 50. We aggregate these tariffs slightly to an industry classification that is consistent with the Demographic Census data used to construct local tariff shock measures. The classification is presented in Table A1. In aggregating, we weight each *Nível* 50 industry with its 1990 industry value added, as reported in IBGE National Accounts data. Figure A1 shows the evolution of nominal tariffs from 1987 to 1998 for the ten largest industries. The phases of Brazilian liberalization are visible (see Section 2 for a discussion and citations). Large nominal tariff cuts from 1987-1989 had little effect on protection, due to the presence of substantial nontariff barriers and tariff exemptions. In 1990, the majority of nontariff barriers and tariff exemptions were abolished, being replaced by tariffs providing equivalent protection; note the increase in tariffs in some industries in 1990. During liberalization, from 1990 to 1994, tariffs fell in all industries, then were relatively stable from 1995 onward.

A.2 Outcome Pre-Trends vs. Tariff Changes

Along with regional variation in the industrial composition of employment, our analysis relies on variation in tariff cuts across industries. Here we analyze the relationship between tariff cuts during liberalization (1990-1995) and trends in industry wages and employment before liberalization, 1980-1991. We choose these two years in measuring pre trends because this is when Demographic Census data are available. We implemented a similar analysis using another smaller household survey, the PNAD, which is fielded yearly between Censuses. This allowed us to consider pre-trends from 1981-1990 and 1986-1990, avoiding overlap with liberalization in 1991. The results were very similar to the Census results presented here.

We implemented a variety of specifications, with results reported in Table A2. In all specifications, the independent variable is the proportional change in one plus the tariff rate ($\Delta_{1990-95}\ln(1+\tau_i)$), shown in Figure 1. In panels A-C the dependent variable is the change in log industry earnings. Panel A uses unconditional average industry earnings; Panel B uses industry earnings residuals controlling for individual age, sex, education, and formal status; and Panel C uses average log earnings residuals controlling for these individual characteristics and region fixed effects. In panel D, the dependent variable is the change in log industry employment. Column (1) weights industries equally, and presents standard errors based on pairwise bootstrap of the t-statistic, to improve small sample properties with only 20 tradable industry observations. Column (2) uses the same estimator, but drops agriculture. Column (3) uses heteroskedasticity weights and presents heteroskedasticity-robust standard errors, which are likely understated in this small sample (MacKinnon 2011). Column (4) uses the same estimator, but drops agriculture. In all cases, the results should be seen primarily as suggestive, because the analysis only uses 20 observations.

Nearly all of the earnings estimates are negative, indicating that tariffs fell more for industries experiencing more positive wage growth prior to liberalization. The majority of the estimates are insignificantly different from zero, with the exception of weighted results in Panels A and B. These specifications heavily weight agriculture, which exhibited negative wage growth prior to liberalization and experienced essentially no tariff decline during liberalization, driving the strong negative relationship. By dropping agriculture, Column (4) confirms that the significant relationship is driven by agriculture. The employment estimates are larger, and change sign across columns. Given the diversity of findings across wage and earnings specifications, this exercise is somewhat inconclusive. Tariff cuts may or may not have been substantially correlated with pre-liberalization outcome trends. These findings motivate us to control for pre-liberalization outcome trends in the regional analyses presented throughout the paper. This ensures that our results are robust to potential spurious correlation between liberalization-induced labor demand shocks and ongoing trends.

A.3 Regional Tariff Changes

As a supplement to Section 4.2, here we present versions of Figure 2 for a number of additional local labor markets. In Figures A2 - A4, we select states in each of Brazil's major geographic regions, and plot the 1991 distribution of employment across industries for the labor markets with the most positive (light gray), median (medium gray), and most negative (black) value of RTC_r in the relevant state.

Amazonas state in the northwest of Brazil is sparsely populated and primarily agricultural. However, the region facing the most negative tariff change in the state is Rio Negro, which mainly does food processing and produces auto and electronics parts for the major assemblers in the nearby free trade area of Manaus (omitted from the analysis). The median region is in the southeast of the state, while the region facing the most positive regional tariff change is Boca do Acre, which produces paper and agricultural products almost exclusively.

Pernambuco is a densely populated northeastern state. Caruaru faced the largest tariff decline in the state, by virtue of specializing mainly in apparel and food processing. Ipojuca is a tourist destination south of Recife, and faced the median tariff decline as a result of having a relatively even distribution of employment across industries. Agrestina is in the rural interior of the state, and faced the most positive tariff change as a result of its heavy reliance on agriculture.

São Paulo is a southeastern state, home to the metropolis of the same name. Aircraft manufacturer Embraer is located in São José dos Campos, which faced the most negative tariff change in the state because vehicle manufacturing faced one of the larger tariff declines. Votuporanga has a large furniture sector, which faced a middling tariff decline, so the city experienced the median tariff change. Bananal, an ecotourism destination, faced the largest tariff decline due to a large agricultural sector.

Santa Catarina is in the south of Brazil. Joinville is the largest city in the state, and faced the most negative tariff change as a result of its diverse manufacturing base, including apparel, plastics, and electronics. São Miguel do Oeste has large meat and milk processing industries and faced the median tariff change in the state. Lages faced the most positive tariff change due to concentrated employment in wood and paper industries.

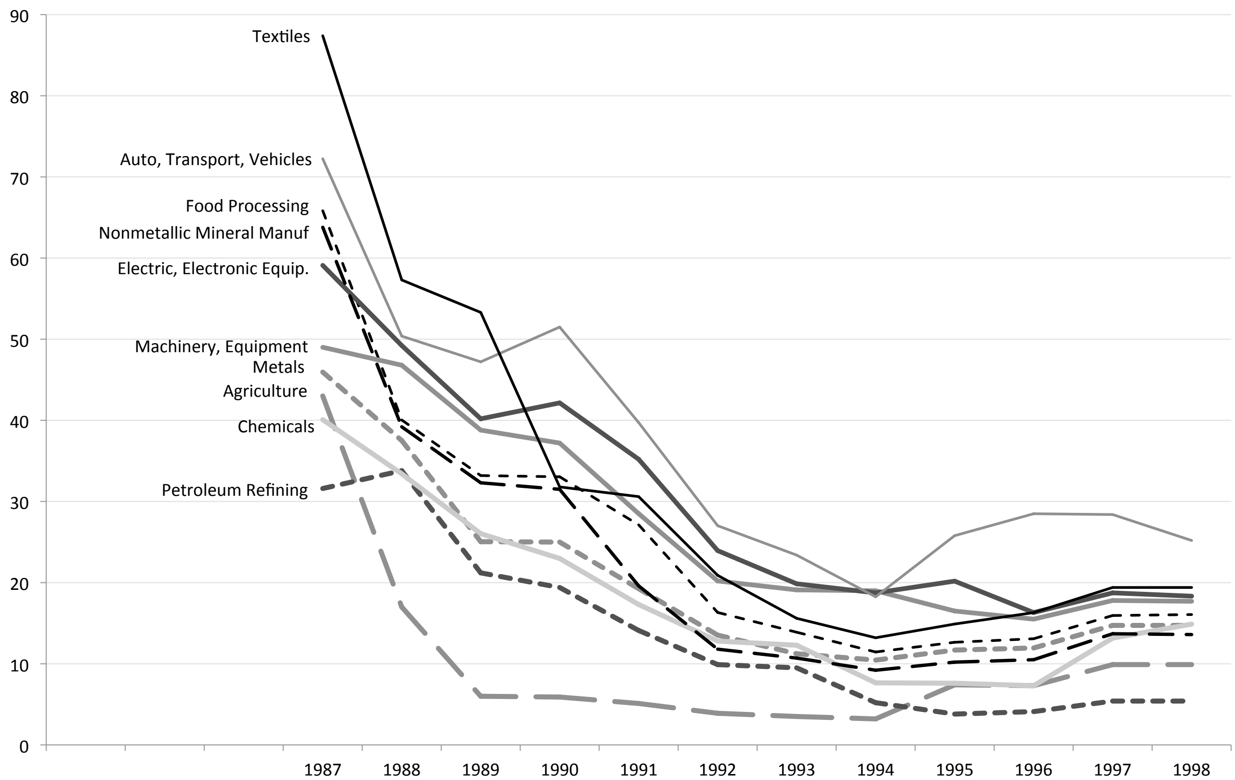
Mato Grosso is a largely rural state in the center-west of Brazil. Cuiabá, the state capital, has some heavy manufacturing and a number of major food processing operations, leading it to have the state's most negative regional tariff change. Matupá produces wood products and agricultural goods almost exclusively, and faced the median tariff change. Paranatinga faced the most positive tariff change by specializing in agricultural output, particularly cereals and oilseeds.

A.4 Formal Earnings Regressions

Figure A5 shows the scatter plots underlying the formal earnings regression estimates in Figure 4 for 1995, 2000, 2005, and 2010. Each marker represents a microregion, and microregions in each major region are shown with a separate type of marker. The size of each marker is proportional to the weight the relevant microregion receives in the estimation. The mean value of the dependent variable is normalized to zero in each year to focus attention on the slope.

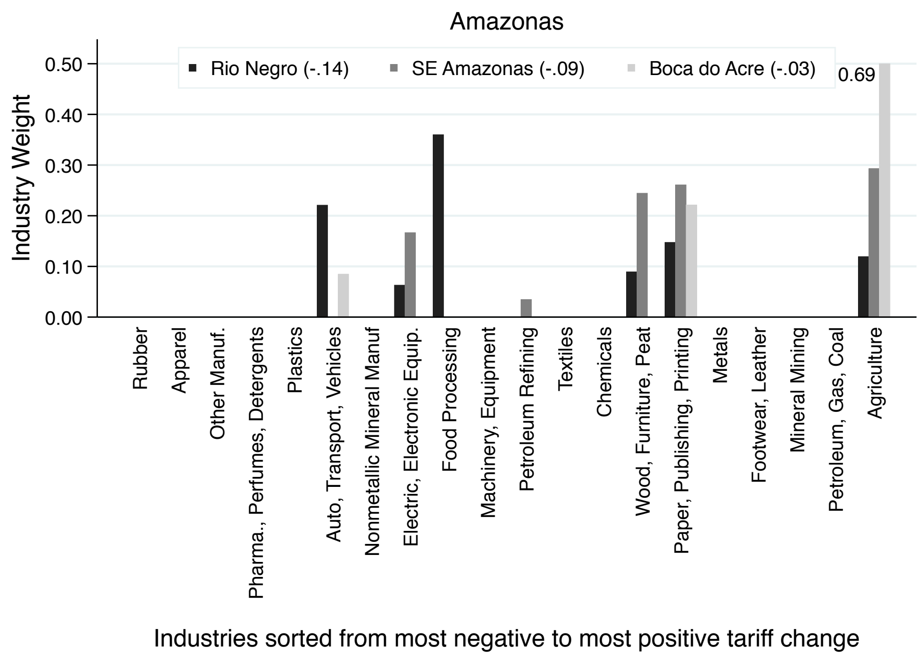
These scatter plots make clear three important points about the earnings estimates. First, as shown in Figure 4, the slope increases substantially and steadily over time, following liberalization. Second, the relationship between changes in formal earnings premia and regional tariff changes is approximately linear in all time periods, justifying our choice of functional form. Third, the increasing slope is driven by shifts in earnings across large numbers of microregions in various parts of the country rather than by a few outliers.

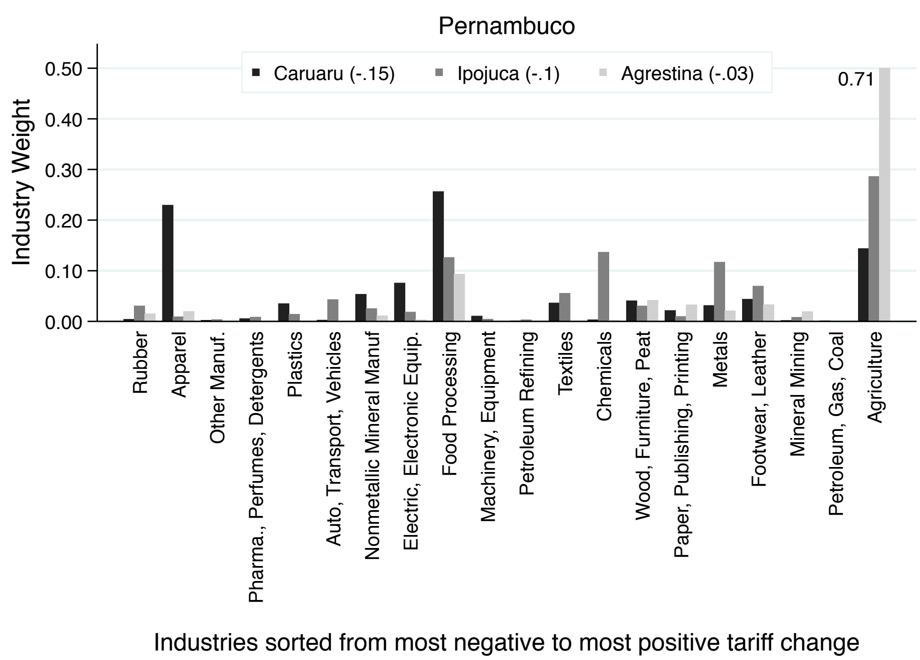
Figure A1: Tariffs - 1987-1998



Nominal tariffs from Kume et al. (2003), aggregated to the industry classification presented in Table A1. Ten largest industries by 1990 value added shown.

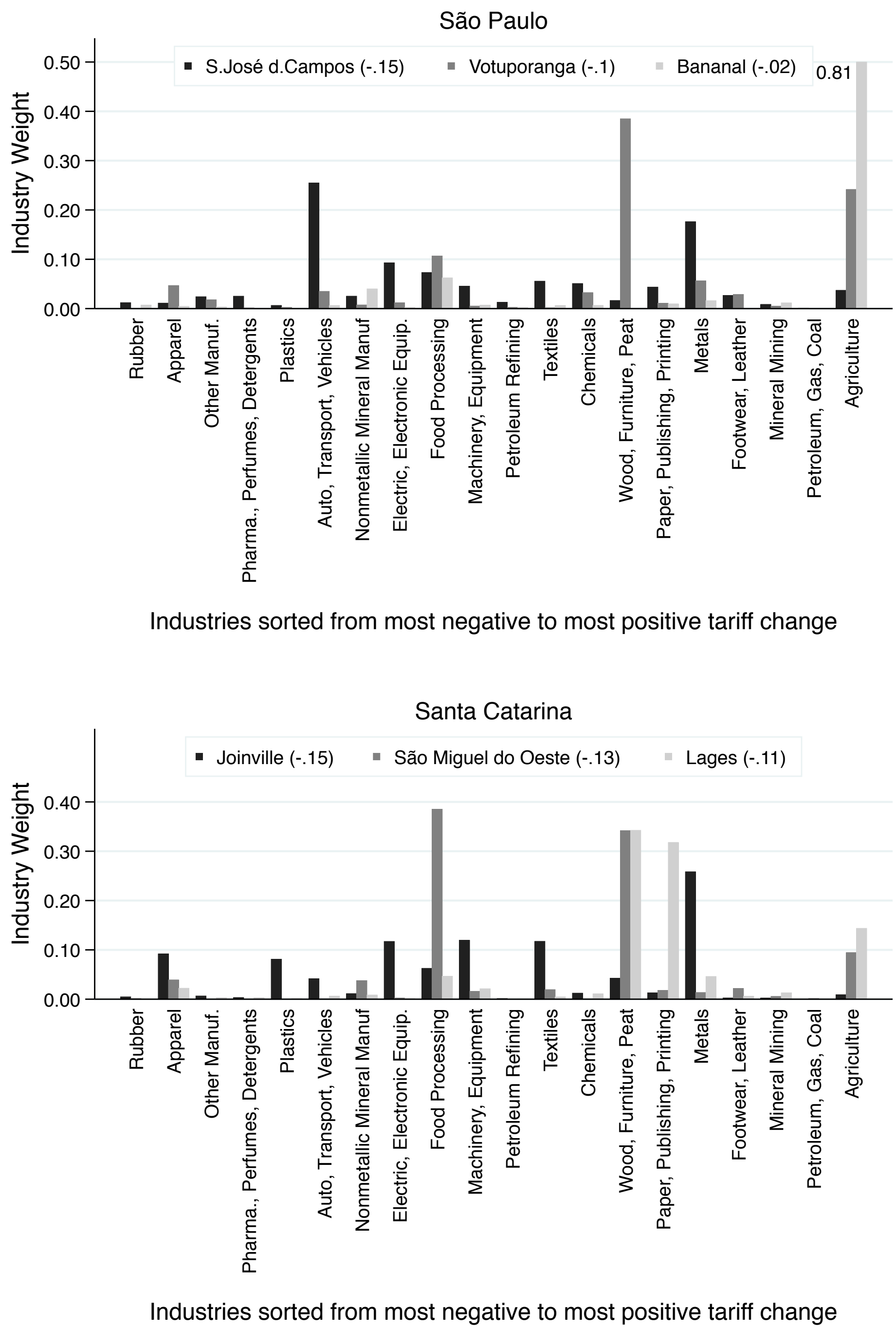
Figure A2: Variation Underlying Regional Tariff Change - Additional Examples (1)





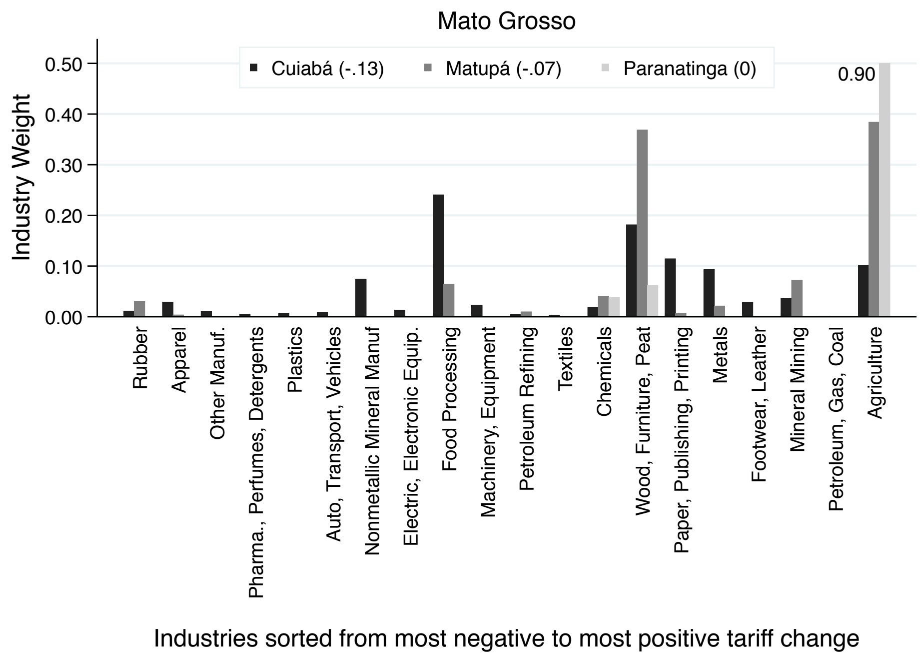
Industry distribution of 1991 employment in the regions facing the most negative, median, and most positive regional tariff changes in the relevant state. Industries sorted by the tariff change, shown in Figure 1. More weight on the left side of the figure leads to a more negative regional tariff change, and more weight on the right side leads to a more positive regional tariff change.

Figure A3: Variation Underlying Regional Tariff Change - Additional Examples (2)



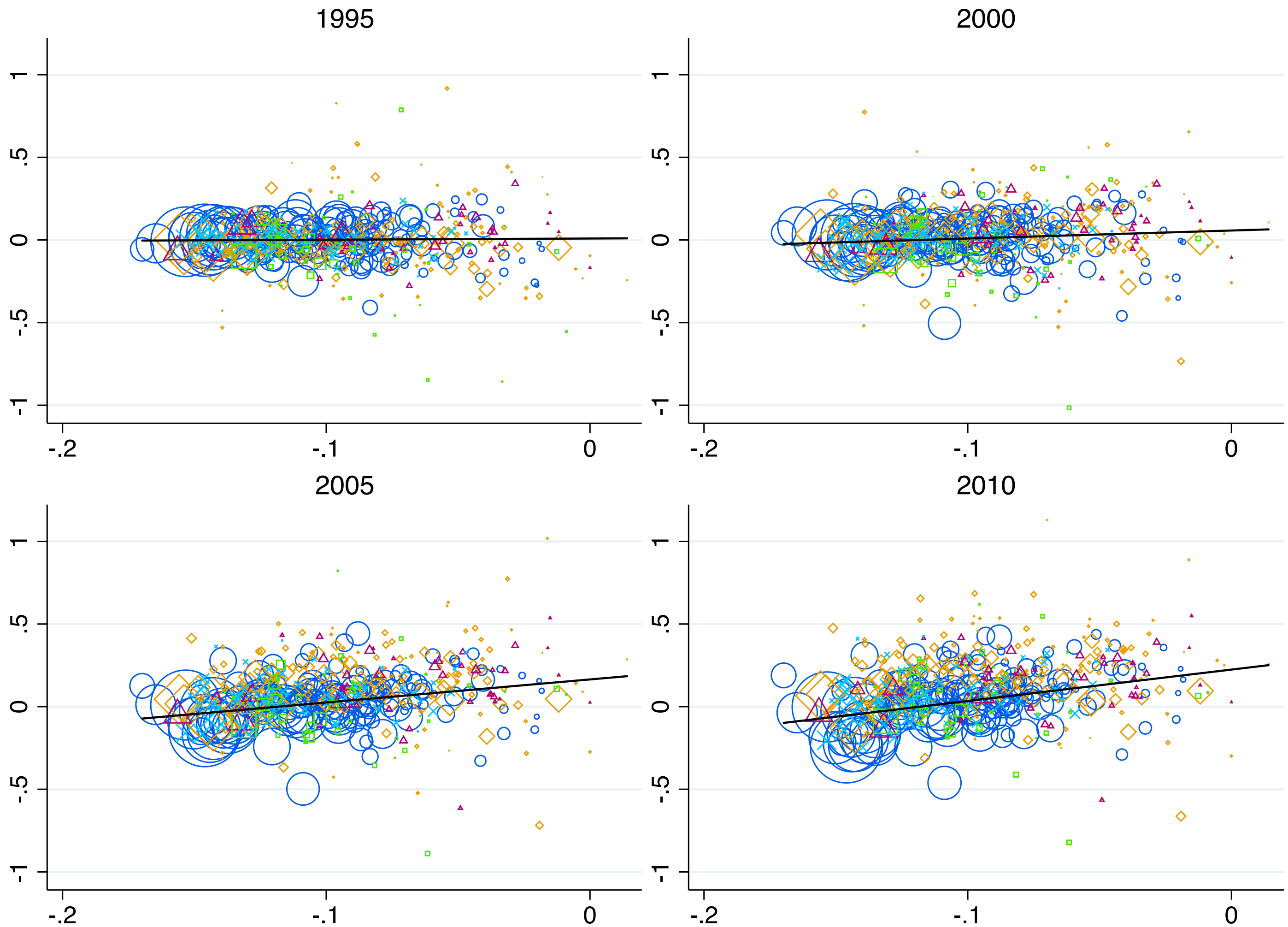
Industry distribution of 1991 employment in the regions facing the most negative, median, and most positive regional tariff changes in the relevant state. Industries sorted by the tariff change, shown in Figure 1. More weight on the left side of the figure leads to a more negative regional tariff change, and more weight on the right side leads to a more positive regional tariff change.

Figure A4: Variation Underlying Regional Tariff Change - Additional Examples (3)



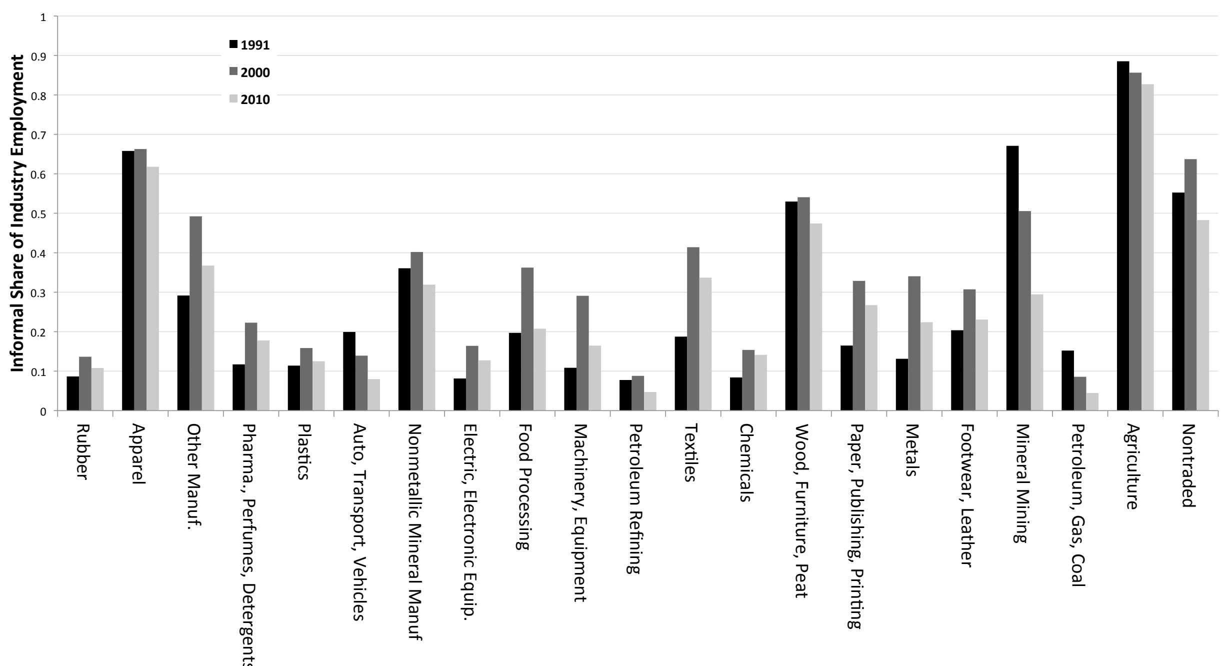
Industry distribution of 1991 employment in the regions facing the most negative, median, and most positive regional tariff changes in the relevant state. Industries sorted by the tariff change, shown in Figure 1. More weight on the left side of the figure leads to a more negative regional tariff change, and more weight on the right side leads to a more positive regional tariff change.

Figure A5: Formal Earnings Premia vs. Regional Tariff Changes - 1995, 2000, 2005, 2010



Scatter plots of regressions behind Figure 4 estimates. 475 microregion observations. Y-axis is the change in log formal earnings premium in a given region from 1991 to the year listed in the title. Dependent variable demeaned in each year. X-axis is the regional tariff change, RTC_r. The size of each marker represents the inverse of the variance of the dependent variable estimate, used to weight the regressions. Major regions shown with different symbols. Southeast: dark blue circles, Northeast: orange diamonds, Center-West: red triangles, South: light blue Xs, North: green squares.

Figure A6: Informal Share of Industry Employment - 1991, 2000, 2010



Decennial Census data. Industries sorted from most negative to most positive tariff change over 1990-1995 (see Figure 1). Nontradable sector faces no tariff change, and is placed at the far right.

Table A1: Industry Classification

	Industry	Industry Name	Nível 50	1980, 1991 Census (*atividade*)	2000, 2010 Census (*CNAE-Dom*)
Tradable	1	Agriculture	1	011-037, 041, 042, 581	1101-1118, 1201-1209, 1300, 1401, 1402, 2001, 2002, 5001, 5002
Tradable	2	Mineral Mining (except combustibles)	2	050, 053-059	12000, 13001, 13002, 14001-14004
Tradable	3	Petroleum and Gas Extraction and Coal Mining	3	051-052	10000, 11000
Tradable	4	Nonmetallic Mineral Goods Manufacturing	4	100	26010, 26091, 26092
Tradable	5	Iron and Steel, Nonferrous, and Other Metal Production and Processing	5-7	110	27001-27003, 28001, 28002
Tradable	8	Machinery, Equipment, Commercial Installation Manufacturing, and Tractor Manufacturing	8	120	29001
Tradable	10	Electrical, Electronic, and Communication Equipment and Components Manufacturing	10-11	130	29002, 30000, 31001, 31002, 32000, 33003
Tradable	12	Automobile, Transportation, and Vehicle Parts Manufacturing	12-13	140	34001-34003, 35010, 35020, 35030, 35090
Tradable	14	Wood Products, Furniture Manufacturing, and Peat Production	14	150, 151, 160	20000, 36010
Tradable	15	Paper Manufacturing, Publishing, and Printing	15	170, 290	21001, 21002, 22000
Tradable	16	Rubber Product Manufacturing	16	180	25010
Tradable	17	Chemical Product Manufacturing	17,19	200	23010, 23030, 23400, 24010, 24090
Tradable	18	Petroleum Refining and Petrochemical Manufacturing	18	201, 202, 352, 477	23020
Tradable	20	Pharmaceutical Products, Perfumes and Detergents Manufacturing	20	210, 220	24020, 24030
Tradable	21	Plastics Products Manufacturing	21	230	25020
Tradable	22	Textiles Manufacturing	22	240, 241	17001, 17002
Tradable	23	Apparel and Apparel Accessories Manufacturing	23	250,532	18001, 18002
Tradable	24	Footwear and Leather and Hide Products Manufacturing	24	190, 251	19011, 19012, 19020
Tradable	25	Food Processing (Coffee, Plant Products, Meat, Dairy, Sugar, Oils, Beverages, and Other)	25-31	260, 261, 270, 280	15010, 15021, 15022, 15030, 15041-15043, 15050, 16000
Tradable	32	Miscellaneous Other Products Manufacturing	32	300	33001, 33002, 33004, 33005, 36090, 37000
Nontradable	91	Utilities	33	351, 353	40010, 40020, 41000
Nontradable	92	Construction	34	340, 524	45001-45005
Nontradable	93	Wholesale and Retail Trade	35	410-424, 582, 583	50010, 50030, 50040, 50050, 53010 ,53020, 53030, 53041, 53042, 53050, 53061-53068, 53070, 53080, 53090, 53101, 53102, 55020
Nontradable	94	Financial Institutions	38	451-453, 585, 612	65000, 66000, 67010, 67020
Nontradable	95	Real Estate and Corporate Services	40, 41	461-464, 543, 552, 571-578, 584, 589	63022, 70001, 71020, 72010, 74011, 74012, 74021, 74022, 74030, 74040, 74050, 74090, 92013, 92014, 92015, 92020
Nontradable	96	Transportation and Communications	36, 37	471-476, 481, 482, 588	60010, 60020, 60031, 60032, 60040, 60091, 60092, 61000, 62000, 63010, 63021 ,64010 ,64020, 91010
Nontradable	97	Private Services	39, 43	511, 512, 521-523, 525, 531, 533, 541, 542. 544, 545, 551, 577, 586, 587, 613-619, 622-624, 632, 901, 902	1500, 50020, 53111, 53112, 53113, 55010, 55030, 63030, 70002, 71010, 71030, 72020, 73000, 74060, 80011, 80012, 80090, 85011, 85012, 85013, 85020, 85030, 90000, 91020, 91091, 91092, 92011, 92012, 92030, 92040, 93010, 93020, 93030, 93091, 93092, 95000
	98	Public Administration	42	354, 610, 611, 621, 631, 711-717, 721-727	75011-75017, 75020

Consistent industry classification used in generating local tariff shocks from *Nível* 50 tariff data in Kume et al. (2003) and Decennial Census data.

Table A2: Pre-Liberalization Industry Trends - 1980-1991

	unweighted, bootstrapped	unweighted, bootstrapped, omitting agriculture	weighted	weighted, omitting agriculture
1980-1991 change in log:	(1)	(2)	(3)	(4)
Panel A: average earnings				
Regional tariff change	-0.345	-0.111	-1.029***	-0.510
	(0.322)	(0.354)	(0.139)	(0.582)
Panel B: earnings premia (with individual controls)				
Regional tariff change	-0.203	0.017	-0.610***	0.235
	(0.273)	(0.311)	(0.157)	(0.350)
Panel C: earnings premia (with individual and region controls)				
Regional tariff change	-0.135	-0.044	-0.184	-0.018
	(0.177)	(0.209)	(0.158)	(0.222)
Panel D: employment				
Regional tariff change	1.564	2.669**	-0.799*	1.638
	(1.268)	(1.331)	(0.408)	(1.868)
Observations	20	19	20	19

Decennial Census data. 20 industry observations (19 omitting agriculture). See text for details of dependent and independent variable construction. Column (1) weights industries equally, and presents standard errors based on pairwise bootstrap of the t-statistic. Column (2) uses the same estimator as Column (1), but drops agriculture. Column (3) uses heteroskedasticity weights and presents heteroskedasticity-robust standard errors. Column (4) uses the same estimator as Column (3), but drops agriculture. *** Significant at the 1 percent, ** 5 percent, * 10 percent level.

B Model

B.1 Setup and Equilibrium Restriction

This section introduces a specific-factors model allowing for imperfect and slow factor adjustment. The economy consists of many regions, indexed by $r = 1...R$, which may produce goods in many industries, indexed by $i = 1...I$. Production in each industry exhibits constant returns to scale using two inputs (below, we will additionally restrict production to be Cobb-Douglas). Labor L_r is assumed to be perfectly mobile between industries within a region and an industry-specific factor T_{ri} is usable only in its respective industry i. Goods and factor markets are perfectly competitive, and producers face exogenous prices P_i, common across regions and fixed by world prices and tariffs.

Consider a particular region r and suppress the region subscript. Let a_{Li} and a_{Ti} be the respective amounts of labor and specific factor used in producing one unit of industry i output. Letting Y_i be output in industry i, factor market clearing implies

$$a_{Ti}Y_i = T_i \quad \forall i, \tag{B1}$$

$$\sum_i a_{Li}Yi = L. \tag{B2}$$

Perfect competition implies that the price equals factor payments,

$$a_{Li}w + a_{Ti}R_i = P_i \quad \forall i, \tag{B3}$$

where w is the wage and R_i is the price of industry i specific factor.

We will consider the effects of price changes. Define $\hat{x}$ as the proportional change in x, and let φ_i be the cost share of the specific factor in industry i. Differentiating (B3),

$$(1-\varphi_i)\hat{w} + \varphi_i\hat{R}_i = \hat{P}_i \quad \forall i, \tag{B4}$$

which follows from the envelope theorem for cost minimization,

$$(1-\varphi_i)\hat{a}_{Li} + \varphi_i\hat{a}_{Ti} = 0 \quad \forall i. \tag{B5}$$

Now we differentiate the factor market clearing conditions, and allow for the possibility that $\hat{L} \neq 0$ and $\hat{T}_i \neq 0$.

$$\hat{Y}_i = \hat{T}_i - \hat{a}_{Ti}. \tag{B6}$$

$$\sum_i \lambda_i(\hat{a}_{Li} + \hat{Y}_i) = \hat{L}, \tag{B7}$$

where $\lambda_i \equiv \frac{L_i}{L}$ is the share of regional labor allocated to industry i. By definition, the elasticity of substitution in production between T_i and L, satisfies

$$\hat{a}_{Li} - \hat{a}_{Ti} = \sigma_i(\hat{R}_i - \hat{w}). \tag{B8}$$

Combining (B6) - (B8) yields

$$\sum_i \lambda_i\sigma_i(\hat{R}_i - \hat{w}) = \hat{L} - \sum_i \lambda_i\hat{T}_i \tag{B9}$$

Equations (B4) and (B9) can be expressed in matrix form as

$$\left[\begin{array}{cccc|c} \varphi_1 & 0 & \dots & 0 & 1-\varphi_1 \\ 0 & \varphi_2 & \dots & 0 & 1-\varphi_2 \\ \vdots & & \ddots & \vdots & \vdots \\ 0 & 0 & \dots & \varphi_N & 1-\varphi_N \\ \hline \lambda_1\sigma_1 & \lambda_2\sigma_2 & \dots & \lambda_N\sigma_N & -\sum_i \lambda_i\sigma_i \end{array}\right] \left[\begin{array}{c} \hat{R}_1 \\ \hat{R}_2 \\ \vdots \\ \hat{R}_N \\ \hline \hat{w} \end{array}\right] = \left[\begin{array}{c} \hat{P}_1 \\ \hat{P}_2 \\ \vdots \\ \hat{P}_N \\ \hline \hat{L} - \sum_i \lambda_i \hat{T}_i \end{array}\right]. \tag{B10}$$

For convenience of notation, rewrite this expression as

$$\left[\begin{array}{c|c} \mathbf{\Phi} & \boldsymbol{\varphi}_{\mathbf{L}} \\ \hline \boldsymbol{\lambda}' & -\sum_i \lambda_i\sigma_i \end{array}\right] \left[\begin{array}{c} \hat{\mathbf{R}} \\ \hline \hat{w} \end{array}\right] = \left[\begin{array}{c} \hat{\mathbf{P}} \\ \hline \hat{L} - \sum_i \lambda_i \hat{T}_i \end{array}\right] \tag{B11}$$

Using Cramer's rule, which in this case involves taking the determinant of a partitioned matrix, we can solve for the proportional change in the regional wage,

$$\hat{w} = \frac{\left(\hat{L} - \sum_i \lambda_i \hat{T}_i\right) - \boldsymbol{\lambda}'\mathbf{\Phi}^{-1}\hat{\mathbf{P}}}{-\sum_i \lambda_i\sigma_i - \boldsymbol{\lambda}'\mathbf{\Phi}^{-1}\boldsymbol{\varphi}_{\mathbf{L}}} \tag{B12}$$

Since the inverse of the diagonal matrix $\mathbf{\Phi}$ is a diagonal matrix of $\frac{1}{\varphi_i}$'s, we can simplify this expression. We also reintroduce regional subscripts and impose the restriction that all production functions are Cobb-Douglas, so $\varphi_{ri} = \varphi_i$ and $\sigma_{ri} = 1$. This yields

$$\hat{w}_r = \sum_i \beta_{ri}\hat{P}_i - \delta_r\left(\hat{L}_r - \sum_i \lambda_{ri}\hat{T}_{ri}\right), \tag{B13}$$

$$\text{where} \quad \beta_{ri} \equiv \frac{\lambda_{ri}\frac{1}{\varphi_i}}{\sum_j \lambda_{rj}\frac{1}{\varphi_j}} \qquad \text{and} \quad \delta_r \equiv \frac{1}{\sum_k \lambda_{rk}\frac{1}{\varphi_k}}.$$

Plug this into (B4) to solve for the change in specific factor returns.

$$\hat{R}_{ri} = \frac{1-\varphi_i}{\varphi_i}\delta_r\left(\hat{L}_r - \sum_i \lambda_{ri}\hat{T}_{ri}\right) - \frac{1-\varphi_i}{\varphi_i}\sum_j \beta_{rj}\hat{P}_j + \frac{1}{\varphi_i}\hat{P}_i. \tag{B14}$$

Equations (B13) and (B14) represent equilibrium restrictions implied by the model. Changes in factor supplies are themselves functions of changes in wages and specific factor returns, and hence are endogenous. By holding factor supplies constant, we observe the changes in factor returns that occur on impact and that generate the incentives for factor reallocation. Setting $\hat{L}_r = \hat{T}_{ri} = 0$ in (B13), it is clear that wages decline more in regions facing larger average price declines, so labor has an incentive to reallocate away from those regions to more favorably affected regions. For the specific factors, consider a change in the average regional specific-factor return: $\sum_i \lambda_{ri} \hat{R}_{ri}$. Plugging in (B14) with factors held fixed yields

$$\begin{aligned}
\sum_i \lambda_{ri} \hat{R}_{ri} &= \sum_i \lambda_{ri} \left(-\frac{1-\varphi_i}{\varphi_i} \sum_j \beta_{rj} \hat{P}_j + \frac{1}{\varphi_i} \hat{P}_i \right) \\
&= \left(\sum_j \beta_{rj} \hat{P}_j \right) \sum_i \lambda_{ri} \left(-\frac{1-\varphi_i}{\varphi_i} \right) + \sum_i \lambda_{ri} \frac{1}{\varphi_i} \hat{P}_i \\
&= \left(\sum_j \beta_{rj} \hat{P}_j \right) \sum_i \lambda_{ri} \left(-\frac{1-\varphi_i}{\varphi_i} \right) + RTC_r \sum_i \lambda_{ri} \frac{1}{\varphi_i} \\
&= \left(\sum_j \beta_{rj} \hat{P}_j \right),
\end{aligned} \tag{B15}$$

where the third equality uses the definition of β_{rj}. Thus, before factor supplies adjust, the change in average regional specific-factor return precisely equals the change in wage, such that both types of factors have an incentive to reallocate away from places facing larger weighted-average price declines.