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

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Abstract

We study how local labor demand shocks induced by the Brazilian trade reform of the 1990s affected regional and individual labor market outcomes for over two decades. Contrary to conventional wisdom, we do not find that the impact of local shocks is dissipated over time through wage equalizing migration. Instead, we find steadily growing effects of local shocks on regional wages and employment. We rationalize this finding with a simple equilibrium model with two complementary factors of production that adjust imperfectly to shocks: labor and industry-specific factors, such as capital. We also find that workers initially employed in regions that were more affected by trade reform face continuously deteriorating formal labor market outcomes relative to workers employed in less affected regions, and this persists even 20 years after the beginning of trade liberalization. Negative local trade shocks induce workers to shift out of the formal sector and to shift out of the formal tradable sector and into the formal nontradable sector. The latter response crowds out workers initially employed in the nontradable sector, given that it also contracts. Interestingly, we find that workers' formal sector migration decisions do not react to these local shocks. In summary, our findings point towards extremely slow and persistent regional adjustment to trade reform at both the aggregate and individual levels.

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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 the investigation of short-, medium- and long-run effects of trade liberalization on individual workers' employment trajectories impossible.

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 developing regional effects on formal earnings, employment, and other labor market outcomes. Workers whose regions face larger tariff declines experience deteriorating labor market outcomes compared to workers in other regions. 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). Another explanation that is consistent with the observed patterns of adjustment involves agglomeration economies as in Kline and Moretti (2014), but we do not attempt to distinguish between these two mechanisms, as both generate similar patterns of adjustment for the variables we can observe.

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. This represents one of the first evaluations of the labor market effects of liberalization to utilize panel data, and the first to analyze medium- to long-run effects of trade liberalization on individual workers' labor market outcomes.[1] 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 nontradable sector employment, but on average cannot offset lost tradable sector earnings.[2] This integration between tradable and nontradable sectors leads to similar magnitude effects for workers initially in the nontradable sector.

Using supplementary data from the Decennial Census, we find that workers facing more negative local shocks are more likely to transition into informal employment, without access to labor protections and other benefits. This result is apparently in contrast to prior work, which typically finds minimal effects of trade on informality (Goldberg and Pavcnik 2003, Bosch, Goñi-Pacchioni and Maloney 2012). However, our results can be reconciled if formal tradable sector workers transition to informal employment in the nontradable sector or if many move to informality following nontrivial periods of unemployment. These margins of adjustment are omitted from prior work. Finally, we find little evidence for geographic migration in response to local demand conditions in spite of substantial unconditional mobility rates. This finding is similar to Yagan (2014) for the U.S. during the Great Recession, and suggests that many migration decisions involve non-labor-market considerations. Together, these findings suggest that formal sector regional adjustment occurs primarily through workers transitioning out of formal employment rather than migrating across space.

In contrast to much of the trade literature, we focus on regional dynamics rather than frictions

[1] Prior examples of papers investigating the labor market effects of trade liberalization using panel data include Menezes-Filho and Muendler (2011) and Krishna, Poole and Senses (forthcoming).

[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.

across industries or establishments. A recent literature on the local labor market effects of trade – pioneered by Topalova (2007) – points toward significant mobility frictions across regions in both developing and developed countries.[3] In light of this recent evidence, understanding geographical mobility barriers is a priority for international trade economists. 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 speed 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 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 (2013b) use U.S. panel data to study the labor market effects of increased Chinese imports across industries. Our individual-level longitudinal analysis is similar in spirit, but differs in a number of important ways. Most importantly, we study a discrete policy shock rather than a continuously evolving phenomenon such as 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 in response to regional shocks. Finally, we examine transitions into the nontradable sector and into informal employment, both of which are salient features in 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.

[3]Topalova's (2007) seminal paper was followed by Edmonds, Pavcnik and Topalova (2010), Hasan, Mitra and Ural (2006), Hasan, Mitra, Ranjan and Ahsan (2012), McCaig (2011), Topalova (2010), Kovak (2013), Hakobyan and McLaren (2012), Autor, Dorn and Hanson (2013a), Kondo (2014), Costa, Garred and Pessoa (2014) and others. All of these papers point to the presence of substantial barriers to regional mobility.

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.

In all, our findings suggest a role for policies that help the economy adjust more quickly and smoothly to trade reform. Some of these policies are already featured in the U.S. Trade Adjustment Assistance (TAA) program, such as worker retraining in order to overcome mobility barriers across sectors or moving allowances in order to overcome barriers of mobility across space. In addition, our results suggest that providing incentives to locate plants in areas with large pools of unemployed or informally employed labor or investing in infrastructure projects can reduce adjustment costs following trade reform. According to Morten and Oliveira (2014), reducing the marginal costs of traveling through infrastructure projects such as building roads has important effects on migration. This suggests that infrastructure projects facilitating movement of people across space can be an effective complementary policy to trade policy reform. In terms of trade theory and quantitative analysis, our results imply that most quantitative trade models abstract from important adjustment frictions when treating large and diverse countries like Brazil as frictionless, integrated economic units.

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, we challenge the conventional wisdom in the labor economics literature that the impact of local shocks will be gradually dissipated through equalizing migration and provide an interpretation for these findings. 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, and argues that our results reflect the causal effects of liberalization. Section 3 describes the data sources, including the matched employer-employee panel and cross-sectional data that we utilize. Section 4 describes our definition of local labor markets, defines our regional measure of trade liberalization, and describes our empirical approach. Section 5 presents the empirical findings for our regional analysis, and Section 6 presents the empirical results in the individual analysis. Section 7 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. This unilateral trade liberalization involved very large declines in trade barriers and featured substantial variation in tariff cuts across industries. Importantly, this cross-industry variation is plausibly exogenous to counterfactual industry performance, which lends our results a causal interpretation. Here we briefly summarize the related discussion in Kovak (2013).

In the late 1980s and early 1990's, Brazil implemented a large unilateral trade liberalization that ended nearly one hundred years with extremely high trade barriers imposed as part of an import substituting industrialization policy. In March 1990, the country's extensive nontariff barriers were abolished and replaced with tariffs providing approximately equivalent degrees of protection. Tariffs then fell substantially in the following five years, from an average of 30.5 percent in 1990 to 12.8 percent in 1995, remaining relatively stable thereafter.[4] Along with this large average 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. Figure 1 shows the cross-industry variation in tariff changes from 1990-1995.[5]

As discussed below, along with regional variation in the industry mix, the cross-industry variation in tariff cuts provides the identifying variation in our analysis. Thus, for our results to reflect the causal effect of trade liberalization, it is necessary for this variation to be uncorrelated with counterfactual industry performance. Following the argument in Goldberg and Pavcnik (2005), the tariff cuts were nearly perfectly correlated with the pre-liberalization tariff levels, which themselves were based on a protective structure that was set decades before liberalization (Kume et al. 2003). Thus, the tariff cuts were unlikely to be related to contemporaneous industry performance during liberalization. Moreover, the tariff cuts largely anticipated the eventual common external tariff negotiated with other members of Mercosur, making it more difficult for Brazilian policy makers to strategically manipulate tariff changes (Kume et al. 2003).

These institutional features leave little scope for endogeneity of tariff cuts to counterfactual industry performance. In case of any remaining spurious correlation between tariff cuts and other steadily evolving industry factors, we include controls for pre-liberalization outcome trends in all analyses. Consistent with the argument that our tariff changes were exogenous, these pre-trend controls have little influence on the vast majority of our results.

[4] Simple averages of tariff rates across *Nivel 50* industries, as reported in (Kume, Piani and de Souza 2003).

[5] We use long-differences in tariff changes from 1990 to 1995 because tariff cuts prior to March 1990 were non-binding, and the timing of tariff cuts within this window was likely endogenous to industry performance. See Kovak (2013) and Bazzi, Menezes-Filho and Muendler (2013) for details.

3 Data

The data used in this paper primarily come from the *Relação Annual de Informações Sociais* (RAIS), a matched employer-employee dataset assembled by the Brazilian Ministry of Labor every year since 1976. RAIS is 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 data spanning the period from 1986 to 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. These data have recently been employed by Menezes-Filho and Muendler (2011), Lopes de Melo (2013), Krishna et al. (forthcoming), 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) and industry (IBGE subsector[6]), 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.[7]

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. (2013a), 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

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

[7]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.

employed workers in every Brazilian municipality.[8] 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 15 years following liberalization, allowing us to study the dynamic response to this well-defined trade policy shock. This contrasts with Autor et al. (2013b), 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.[9] However, we can observe 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 four rounds of the Brazilian Demographic Census: 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 formality, employment, and labor force participation. 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 the RAIS using the Demographic Census.

[8] 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.

[9] Authors' calculations using Brazilian Demographic Census.

4 Empirical Approach

4.1 Local Labor Markets

We define local labor markets based on the "microregion" definition of the Brazilian Statistical Agency (IBGE). Microregions are groupings of 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 that we consistently define local labor markets from 1986-2010.[10] This process leads to a set of 475 consistently identifiable local labor markets.[11]

4.2 Local Measure of Trade Liberalization

A growing literature measures the local effects of national industry-level shocks using weights reflecting the initial industrial distribution of employment in each region. In seminal work, Bartik (1991) examined the local effects of national growth and contraction in employment across industries. Topalova (2007) brought this local markets approach into the field of international trade by examining the local effects of trade liberalization in India. This foundational work has spawned a rapidly growing literature using similar empirical techniques to study the local effects of changes in trade policy or increased imports from developing countries.[12]

We follow the approach of Kovak (2013), which derives a local measure of trade liberalization from a simple model of local labor markets. The model assumes that each region is endowed with a fixed vector of industry-specific factors, as in Jones (1975), and that workers are homogenous and are mobile across industries and immobile across regions. Production is Cobb-Douglas with potentially varying factor shares across industries, and producers in all regions face the same vector of national trade-policy induced price changes. Under these assumptions, Kovak (2013) derives the

[10] This geographic classification is a slightly aggregated version of the one in Kovak (2013), accounting for additional boundary changes during the 2000 to 2010 period. Related papers define local markets based on commuting patterns (e.g. Autor et al. (2013a)). 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.

[11] The regional definition is shown in Figure 3. The analysis omits 11 microregions, shown in gray in 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 the RAIS in the 1980s. The inclusion or exclusion of these regions when possible has no substantive effect on the results. We also implemented the analysis using a more aggregate local labor market definition, "mesoregions" defined by IBGE, and results are nearly identical.

[12] See footnote 3 for examples.

"regional tariff change" or RTC_r, given by

$$RTC_r = \sum_i \beta_{ri} d\ln(1+\tau_i) \quad \text{where} \quad \beta_{ri} \equiv \frac{\frac{1}{\theta_i}\lambda_{ri}}{\sum_j \frac{1}{\theta_j}\lambda_{rj}}, \tag{1}$$

r indexes regions, i and j index tradable goods industries, λ_{ri} is the initial share of region r formal employment in industry i, θ_i is one minus the initial wagebill share of industry value added in industry i, and d represents the long difference from 1990-1995, the period of Brazilian trade liberalization.[13]

Because Brazilian local labor markets differ substantially in the industry distribution of their employment, the weights β_{ri} vary across regions. The weighted average in (1) therefore reflects the intuitive idea, explored by Topalova (2007) and others, that regions experience larger declines in labor demand when their most important industries face larger tariff declines during liberalization. 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 industries, variation in the industry distribution of employment in each region generates substantial 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.[14]

While we calculate RTC_r following the model in Kovak (2013), the measure remains informative

[13] We calculate formal employment shares using the 1991 Census, as it provides a more detailed industry classification than that available in RAIS. We use *formal* shares because the RAIS data cover only formal employment, and hence we focus on formal sector outcomes. Workers appear to be able to transition from formal to informal employment relatively easily, but the reverse transition is more difficult. Dix-Carneiro (2014) estimates very large costs of moving into formal sector employment. As we show in section 5.3, results are very similar when using overall employment shares, including both formal and informal employment. The θ_i are calculated using IBGE national accounts data.

[14] 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.

even under other circumstances. For example, if workers are imperfectly mobile across industries, the weighted average still reflects the average tariff shock initially faced by workers in the region. If workers are mobile across regions, RTC_r remains informative regarding the liberalization-induced labor *demand* shock faced by a given region, which in turn determines the incentive for workers to relocate. We explicitly consider this generalization in Section 5.4.

The 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. 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.[15]

4.3 Research Design

We implement empirical analyses at the regional and individual levels. In both cases, we examine how post-liberalization (1995-2010) labor market outcomes diverged 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. While this approach allows us to generate credible estimates of the causal regional effects of liberalization, it has the limitation that we cannot estimate the national effects of liberalization that apply across regions (which is shared by any investigation exploring regional- or industry-specific variation in tariffs or other trade-related variables).

Our approach parallels the 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.

Our regional analysis, presented in Section 5, studies the effects of local trade shocks on the evolution of region-level outcomes such as regional earnings or employment. The RAIS data also allow us to track individual workers over time, to help understand how trade shocks in a worker's initial region of employment affected their subsequent outcomes. We implement an analysis that is similar in spirit to the industry-level analysis in Autor et al. (2013b), but differs in that we i) focus

[15] 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.

on regional rather than industry shocks, ii) study a discrete shock and hence focus our analysis on subsequent labor market dynamics (see above), iii) study transitions into nontradable sector and informal employment, which are salient in the Brazilian context, and iv) study formal sector migration in response to regional shocks.

The following two sections describe our detailed estimation strategies and findings.

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:

$$y_{rt} - y_{r,1991} = \theta_t RTC_r + \alpha_{st} + \gamma_t(y_{r,1990} - y_{r,1986}) + \epsilon_{rt}, \tag{2}$$

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.[16] 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. This also allows us to compare our results when $t = 2000$ to the parallel specifications in Kovak (2013). 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. 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 spatial correlation in outcomes across neighboring regions.

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

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. Conventional wisdom would predict large wage effects just after liberalization, followed by a period of declining effects as labor migrates, arbitraging away differences across markets (Blanchard and Katz 1992, Bound and Holzer 2000). By estimating (2) in each year following liberalization, we can observe this adjustment process directly.

Because hours information is not available 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.[17] We then calculate the change in this log earnings premium for each region between 1991 and each subsequent year, from 1992 to 2010.

The results of estimating (2) for regional earnings and wages are shown in Table 1. 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. The scale of the coefficients in 2000 are very similar to those in Kovak (2013), which studied wage results in the overall labor market including formal and informal sector workers. Also consistent with the findings in Kovak (2013), the pre-liberalization trend is negatively related to subsequent earnings growth, but the inclusion of this control drives only a modest decline in the measured effect of tariff changes.

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. Each point on the figure corresponds to a regression coefficient like those in Table 1, where the dependent variable is

[17] Alternately, we can i) control for *constant, unobservable* worker heterogeneity by pooling across years and including individual fixed effects or ii) use average outcomes unadjusted for composition. Doing so has no substantive effect on the results.

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.[18] 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 sectors, and restricting the sample based on regional characteristics. Panel B of Table 1 further confirms this increasing pattern for earnings in the Decennial Census. While the point estimates are smaller than those in Panel A, they also exhibit an increasing pattern over time. Panel C shows results for hourly wages in the Census. The point estimates are similar to those for earnings, ruling out the possibility of offsetting movement in hours, which are unavailable in RAIS.

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 for earnings, the results are quite consistent across specifications. Panel A column (3) uses RAIS to show that a region facing a 10 percentage point larger tariff decline experienced a 27.6 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 are smaller, but also increasing over time.[19] Figure 5 shows that the employment effect increased quite steadily over time before leveling off in the mid 2000s.

To better understand liberalization's effects on changing regional employment, we investigate how 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 from 1991 to each year in 1992-2010. In Figure 6, it is clear that the dynamics of employment are primarily driven by dynamics in job creation in the short run and by dynamics in job destruction in the medium- to long-run. New entry and hiring decisions of incumbents respond immediately to the tariff shock and remain relatively steady over time. However, exit or downsizing decisions do not respond immediately. Job destruction is unrelated to RTC_r in the years immediately following liberalization, but then steadily increases for nearly a decade in harder hit locations before leveling off in the mid 2000s. These patterns closely mirror the employment effects in Figure 5, and suggests a scenario in which employers in depressed local markets maintain

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

[19] 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 *across* regions.

employment for a few years before beginning to exit or shed employees.

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. In what follows, we will argue that these patterns reflect slow adjustment of factor markets. But first we rule out a number of alternative explanations for the growing effects of liberalization to validate our focus on adjustment frictions and demonstrate the robustness of our findings.

The effects of liberalization may 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 5, such shocks would need to affect industries 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 to ensure that they are not driving our results.

If tariff changes following 1995 were correlated with those occurring during liberalization (1990-95), this could account for a portion of the growing labor market effects that we observe. We calculate post-liberalization regional tariff changes as in (1), but use tariff changes between 1995 and year $t > 1995$ as additional controls.[20] 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, and 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 to 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. (2013a) (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 (1).[21]

Figure 7 plots the earnings coefficients from Figure 4 alongside similar coefficients in which ADH-style changes in imports and exports are the dependent variables. If slow import and export adjustment were driving the slow evolution of liberalization's effect on regional earnings, the export

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

[21] Imports and exports data are from Comtrade.

coefficients would exhibit a similar shape to the earnings coefficients, and/or the import coefficients would exhibit their mirror image. This is clearly not the case; between 1995 and 2002, the response of earnings to RTC_r increased by a factor of 8.5, while the ADH export response stayed 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.[22] Table 3 Panel B confirms this point by directly controlling for import and export growth, with minimal effect on the liberalization coefficients.[23]

Other potential confounders are the Brazilian 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 therefore construct industry-specific real exchange rates as trade-weighted averages of real exchange rates between Brazil and its trading partners.[24] 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 (1). Panel C includes this control in the earnings analysis, again with little effect on the scale or time pattern of the liberalization coefficients.

Along with these controls for post-liberalization shocks, we implemented a variety of robustness tests. Table 3 Panels D and E restrict the sample to include, respectively, only the most urban regions and those with the largest share formally employed, based on information from the 1991 Census. These restrictions ensure that our results are not driven solely by comparing the most urban and most rural areas or comparing areas with very different formal shares of employment. In both cases, these restrictions have little effect on the liberalization coefficients. We also considered two alternative measures of RTC_r. The first, in Panel F, uses an alternative tariff measure called the effective rate of protection (ERP), which takes into account the effect of tariff changes on firms' output *and inputs*.[25] The coefficient estimates are somewhat smaller in this case, but since the ERP measure has a larger scale than nominal tariffs in the Brazilian context, they imply similar effects to nominal tariffs.[26] The second, in Panel G, uses alternate values for the regional employment weights λ_{ri} in (1), 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

[22]The export standard errors also increase substantially during 2008-10, such that the point estimates are not significantly different from zero.

[23]These controls also rule out the potential confounding effect of the commodity price boom of the 2000s, which is reflected in growth in export values for commodity producing regions (Costa et al. 2014)

[24]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.

[25]Kume et al. (2003) provide ERP measures along with nominal tariffs. See Corden (1974) for details on ERP definition and construction.

[26]This feature reflects the graduated tariff structure Brazil employed as part of its import-substituting industrialization strategy prior to liberalization.

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.[27] Instead, it appears that the regional labor market effects of liberalization genuinely grew for a decade after the tariff cuts were complete. We argue in the following subsection that the steady divergence in outcome growth for positively and negatively affected regions is the result of frictions in the regional adjustment of labor and complementary factors.

5.4 Factor Adjustment

We have documented large regional effects of liberalization that grow over time, such that outcomes steadily deteriorate for a decade in regions facing larger tariff cuts, compared to those in regions facing smaller cuts. Following that period of divergence, the effects remain steady until the end of the sample period, with little sign of convergence. These results are at odds with conventional wisdom, in which equalizing labor migration partly arbitrages away differences in wage growth, particularly over the course of a decade (Blanchard and Katz 1992, Bound and Holzer 2000). Dix-Carneiro (2014) suggests an alternative mechanism in which both labor and complementary factors such as capital face important adjustment barriers. These frictions are mutually reinforcing, and can lead to precisely the pattern of wage adjustment documented in Figure 4.

Another potential explanation for the patterns documented in Figures 4 and 5 involves the existence of agglomeration economies in manufacturing (see, for example, Ellison and Glaeser (1997), Rosenthal and Strange (2004), Greenstone, Hornbeck and Moretti (2010), and Kline and Moretti (2014)). We do not attempt to distinguish between these two mechanisms as they have similar implications for the behavior of variables we can observe. The important common element of these 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. Nevertheless, we focus our theoretical interpretation of the results on sluggish capital adjustment because it entails a simple extension of the model in Kovak (2013) and because Dix-Carneiro (2014) has shown that this mechanism can indeed generate adjustment patterns that are quantitatively similar to those in Figure 4 – at least in the context of inter-industry adjustment.

As outlined above, the model in Kovak (2013) assumes each region is endowed with a vector of industry-specific factors and a stock of regional labor that is costlessly mobile across industries. Goods and factor markets are competitive, production is Cobb-Douglas with potentially varying factor shares across industries, and producers in all regions face the same vector of national price changes. The extension we introduce here is to allow the amounts of labor and specific factors to vary in response to liberalization. We solve this variation of the model in Appendix A, yielding the

[27] Unreported employment results are similarly robust.

following equilibrium relationship governing the evolution of wages in a region r.

$$\hat{w}_r = \sum_i \beta_{ri} \hat{P}_i - \delta_r \left(\hat{L}_r - \sum_i \lambda_{ri} \hat{T}_{ri} \right), \tag{3}$$

$$\text{where} \quad \beta_{ri} \equiv \frac{\lambda_{ri} \frac{1}{\theta_i}}{\sum_j \lambda_{rj} \frac{1}{\theta_j}}, \qquad \delta_r \equiv \frac{1}{\sum_k \lambda_{rk} \frac{1}{\theta_k}},$$

L_r represents regional employment, T_{ri} is the amount of industry i specific factor in region r, and P_i is the price faced by producers in industry i.

It is important to emphasize that (3) represents an equilibrium relationship where changes in factor quantities $\hat{L}_r$ and $\hat{T}_{ri}$ are themselves functions of changes in factor payments, and hence are endogenous. In principle, one could specify a functional form for the factor supply process to close the model.[28] Instead we compare the structural equation in (3) to its reduced form estimating equation in (2). The changing factor supplies appearing in the rightmost term in (3) are omitted from the estimation equation, and hence appear in the error term. The reduced form estimate of θ_t therefore depends upon the correlations between changing factor supplies, RTC_r, and the outcome of interest. By imposing general restrictions on these relationships, we determine what patterns of factor adjustment can rationalize our empirical results.

Consider the case where factor quantities are held fixed, i.e $\hat{L}_r = \hat{T}_{ri} = 0$. The proportional wage change then equals the weighted average of proportional price changes, just as in the definition of RTC_r in (1). In other words, RTC_r reflects the shift in regional labor demand resulting from the industry price changes induced by liberalization. Thus, in the absence of factor adjustment, one would observe a constant relationship between wages and RTC_r in all years following liberalization.

If we allow the quantity of labor to vary but continue to hold specific factors fixed, the model yields the conventional wisdom. Assume that either due to migration or shifts out of the formal sector, formal sector employment declines in regions facing more negative wage changes, such that $\text{cov}(RTC_r, \hat{L}_r) > 0$. Since $\delta_r > 0$, this reallocation will partly equalize the wage changes across regions, driving down the observed regional effect of liberalization. Now allow $\hat{L}_r$ to vary over time, always measuring changes relative to a fixed base year. If labor adjusts slowly, then $\text{cov}(RTC_r, \hat{L}_r)$ becomes steadily more positive over time, and one would observe steadily declining regional wage effects following liberalization.

In order to rationalize *growing* effects over time, we must allow both labor and specific factors to vary in response to liberalization. Once again, labor may adjust ether through migration or shifts into or out of formal employment. Dix-Carneiro (2014) finds that workers face relatively small costs of leaving formal employment, but that the costs to reenter formal employment are very large. Thus, even without migration, the size of the local formal labor pool can adjust as workers

[28] We have done so, and results are available upon request.

transition into informal employment or leave the labor force. We think of specific factor reallocation as reflecting capital depreciation in one region and new investment in another region – to simplify the argument we take the view that installed capital is not mobile across regions. Mechanically, the observed wage effect will grow when $\text{cov}(RTC_r, \hat{L}_r - \sum_i \lambda_{ri}\hat{T}_{ri})$ grows over time. Changes in this covariance are determined by the relative speed of labor and specific factor adjustment. We restrict attention to situations where both labor and specific factors reallocate toward regions with higher returns, such that $\text{cov}(RTC_r, \hat{L}_r) > 0$ and $\text{cov}(RTC_r, \sum_i \lambda_{ri}\hat{T}_{ri}) > 0$.[29] If both labor and specific factors reallocate slowly, their respective covariances will increase over time. As an example, if labor reallocates a bit more quickly at first and capital reallocates more quickly later, then $\text{cov}(RTC_r, \hat{L}_r - \sum_i \lambda_{ri}\hat{T}_{ri})$ grows over time, and one will observe growing wage effects as in Figure 4.

Seen through this lens, we can explain various features of Figure 4, following Dix-Carneiro (2014) Section 7.4, which simulates a model of inter-industry reallocation with 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.

How does this mechanism compare to the data? Figure 5 shows that formal employment fell in regions facing more negative shocks relative to more positively affected regions. Also, this adjustment stopped around 2004, reflected in constant coefficients thereafter. This pattern is consistent with initially rapid labor reallocation, which tapered off by the mid 2000s. Below, in Section 6.2, we will show that the reallocation occurred primarily through movement of workers into and out of formal sector employment rather than through interregional migration.

Reallocation of specific factors is more difficult to measure, particularly at the local level. Even if we restrict attention to capital, there exist no measures of capital stocks or investment at the level of local labor markets. The closest available proxy is the number and size of firms in a given region. Figure 8 Panel A examines the effects of liberalization on the number of formal establishments. Regions facing larger tariff declines experience larger contractions (smaller increases) in the number of local establishments, likely reflecting adjustment of the capital associated with those establishments. This reallocation occurs slowly, with an increase in the rate of reallocation in the early 2000s before leveling out later in the sample period. Figure 8 Panel B shows that the decline in the number of establishments in harder hit locations is not offset by growth in the average size

[29] Appendix A shows that the average specific factor return falls more (increases less) in regions with more negative values of RTC_r, such that labor and specific-factors have the incentive to reallocate in the same direction.

of each firm, which is generally unrelated to RTC_r. As a final proxy for capital adjustment, if regional production reflects a Cobb-Douglas production function, changes in the regional wagebill reflect changes in capital expenditure. Figure 9 shows a similar pattern, with an even stronger acceleration in the reallocation in the early 2000s.

Figure 6 also shows suggesting evidence that the sluggish capital adjustment mechanism we propose may indeed be driving the pattern in Figure 4. More affected local labor markets face a sharp and permanent decline in job creation rates relative to less affected local labor markets. This suggests a relative decline in investment rates in more affected local labor markets that reacts sharply and permanently. On the other hand, job destruction rates are essentially unaffected until 1996 when they start to increase faster in relatively depressed markets. This pattern suggests that the lack of investment in harder hit areas leads to substantial reductions in capital stocks around 1996, which coincides with the inflection point of 4.

Overall, 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. Given our rough measures of capital adjustment, it is difficult to judge the relative timing and speed of labor and capital reallocation, but the results are consistent with the pattern of reallocation needed to rationalize the observed labor market effects.

6 Individual Level 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.[30] Our population of interest includes all workers aged 25-55 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 2000 tradable sector workers in 1989 and include all relevant workers from smaller regions, weighting appropriately in subsequent analyses. This process yields 682,210 individuals in our main tradable sector sample.

In Section 6.5, 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, leaving 1,134,900 workers in the

[30] 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 in the regional analysis.

nontradable sector worker sample.[31] Table 4 provides summary statistics for the tradables sector and nontradables sector samples.

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

$$y_{ijrt} = \alpha_{st} + \theta_t RTC_r + X_{ijrt}\Phi_t + \epsilon_{ijrt}, \tag{4}$$

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).[32] y_{ijrt} represents various worker-level post-liberalization outcomes, which we define below. X_{ijrt} is a rich set of worker-level controls including worker 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 nontradable industry indicators, tenure at the plant), initial employer characteristics (log employment, exporting indicator, log exports, importing indicator, log imports[33]), and initial region characteristics (state indicators, pre-liberalization (1986-89) earnings growth 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).[34] This specification compares subsequent labor market outcomes for two otherwise observationally equivalent workers in 1989 who happen to live in regions facing different local trade shocks.

6.2 Employment Status

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:

$$\frac{1}{t-1989}\sum_{s=1990}^{t} Months_{is}, \tag{5}$$

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

Figure 10 shows that the average number of months formally employed per year is lower for

[31] We use yet another population to study the length of non-formal spells, described in Section 6.4

[32] 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. We expect to remedy this issue in a subsequent revision.

[33] Firm-level imports and exports come from customs data assembled by the *Secretaria de Comércio Exterior* (SECEX).

[34] The pre-liberalization outcome level and trend controls are calculated as follows. We draw a sample of workers in December 1986, paralleling the main sample, and estimate a version of (4) replacing RTC_r with region indicators. We then control for growth in these indicators for log earnings and log employment, and the levels of the indicators for each outcome considered below. 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.

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 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 formally 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 outside the formal sector (which may include informal employment, self-employment, unemployment, and out of the labor force):

$$\frac{12 - Months_{it}}{12}, \tag{6}$$

This measure is similar to (5), but focuses on an individual year rather than the cumulative experience from 1990 onward. Figure 11 shows that workers whose initial regions faced larger tariff cuts spent more of each year outside 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 more months outside formal employment, compared to an average of 7.2 for the entire sample in 2004.

Workers facing larger tariff cuts spend less time in formal employment, but in the RAIS data we cannot determine whether they are not employed at all or employed in the informal sector. To address this question, we use the Decennial Census, which covers all individuals, including those outside formal employment. As the Census provides only repeated cross-sections, we use the regional estimation strategy in (2) to examine the share of the local population that is informally employed, and not employed. Formal employment is defined as having a signed work-card, indicating that the worker is registered with the federal employment authorities and has access to formal labor protections and benefits. The dependent variable is the change in the share of regional working-age individuals falling in the relevant employment category, which we regress on RTC_r as in (2). Following Goldberg and Pavcnik (2003), we calculate these shares by regressing indicators for category membership on workers' observable characteristics and region indicators in order to calculate regional premia. Then, we use the change in regional indicators as our dependent variable. We present specifications with and without state fixed effects and pre-trends for the dependent variables. Since RAIS pre-trends are not available for these variables, we calculate them using the 1980 and 1991 Census. This requires us to use a slightly aggregated regional definition, with 411 consistently identifiable microregions. It also introduces the complication that both the

pre-trends and the dependent variables contain the 1991 share. To address this endogeneity, we instrument the pre-trends using the 1980 value for the relevant share. All instruments are very strong, with first-stage F statistics greater than or equal to 65. It is reassuring that the results are very similar across specifications.

The results appear in Table 5. Panel A shows that regions facing larger tariff declines experience relative increases in the share informally employed. The estimate of -0.248 in column (3) indicates that a region facing a 10 percentage point larger regional tariff decline exhibited a 2.48 percentage point larger increase in its informal employment share. Panel B shows that by 2000, regions facing larger tariff declines also experienced increases in non-employment (including unemployed and out of the labor force). However, by 2010, the sign reverses. This surprising effect can be explained by large increases in self-employment in the hardest hit regions during the 2000s, as seen in the right hand side of Panel D. It appears that many unemployed workers in depressed markets sought self-employment as a means to obtain some income when faced with a lack of more traditional employment opportunities. When grouping together non-employed and self-employed individuals, Panel E finds the expected negative coefficients that grow over time, reflecting the declining opportunities for employees in the hardest hit regions.

These results suggest that trade liberalization drove shifts out of formal employment and into informal employment, not-employment and self-employment. This is a novel finding, which apparently contradicts 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 whether informality increases within manufacturing sectors that face the largest tariff declines, so their analyses do not capture transitions from formal manufacturing employment into informal non-manufacturing employment. Also, their methodology omits transitions into informality that occur through transitory spells in unemployment. Our approach is able to capture both margins. More recently, Menezes-Filho and Muendler (2011) allow for transitions into non-manufacturing sectors and into unemployment. They still do not find any effect of trade policy changes on transitions into informality, but they find significant effects on transitions into unemployment. A key distinction of our methodologies is that Menezes-Filho and Muendler (2011) only observe short-run (yearly) transitions, whereas we focus on transitions at various horizons using RAIS and the Census. Our results are compatible if displaced workers from contracting sectors spend time unemployed before transitioning into informal or self-employment.

Along with adjustment between formal and informal employment, workers also adjust between tradable and nontradable sector employment. We examine the average number of months formally employed per year, as in (5), but separate months into those worked in 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. Figure 12 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 10.

Workers can also adjust by moving to a more favorable location. We split the average number of months formally employed per year into those worked in the individual's initial local labor market and those worked elsewhere. Figure 13 shows the results. Workers initially employed in more adversely affected regions faced a reduction in formal employment on both margins: at their initial location and elsewhere, though these effects are imprecisely estimated. To better understand whether workers initially employed in adversely affected regions were relatively more likely to migrate away we examine the share of time spent in formal employment that occurred away from the initial region:

$$\frac{MonthsAway_{it}}{Months_{it}} \tag{7}$$

This measure approximates the probability of migrating to obtain a formal sector job, conditional on holding a formal sector job anywhere. Figure 14 shows that workers in more negatively affected regions were not more likely to migrate to a formal job in more positively shocked regions. However, keep in mind that we only observe workers' job locations in RAIS if they hold a formal sector job. Therefore, we cannot strictly rule out that aggregate migration flows from negatively shocked places to more positively shocked places. The RAIS analysis above is silent regarding i) the migration decisions of workers who were not employed in the formal sector in December of 1989 and ii) the migration decisions of formal sector workers who migrate to other regions but no longer hold a formal sector job. However, these margins of adjustment are unlikely given that total local population in negatively shocked places does not decrease relatively to more positively affected places.

Even though formal sector migration rates across microregions between 1989 and 1994 are relatively large at 8.6 percentage points (see Table 4), we find that these migration decisions are not influenced by local labor demand shocks, a finding that resembles those of Yagan (2014) for the United Sates. Even though migration rates across counties in the United States are very large (20% change counties over the course of 5 years), these migration rates do not seem to be systematically responding to changes in local labor market conditions following the Great Recession.

Taken together, these findings show that workers whose initial locations faced larger shocks spent substantially more time out of formal employment in the subsequent 15 years. Negatively affected workers were more likely to transition into the nontradables sector or the informal sector,

but neither of these adjustments appears sufficient to offset the formal employment losses.

6.3 Earnings

Along with employment, we consider the effect of the regional trade shock in a worker's initial location on subsequent earnings. 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:

$$\frac{\frac{1}{t-1989}\sum_{s=1990}^{t} Earnings_{is}}{MeanEarnings_{i,1986-89}}, \tag{8}$$

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

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. 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 15 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.

Figure 16 shows the results for a closely related earnings measure, yearly earnings as a share of pre-liberalization mean earnings.

$$\frac{Earnings_{it}}{MeanEarnings_{i,1986-89}} \tag{9}$$

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.

6.4 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 use a somewhat different sample. 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 at the time the spell began. Since our data end at December 2009, non-formal spells are right-censored at that point. Thus 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.[35]

The results are shown in Figure 17. Workers losing formal employment in regions that faced larger tariff declines experienced 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.[36]

6.5 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 of initially nontradable sector workers. 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 all workers similarly.

We find similar magnitudes and time patterns of effects for the tradable and nontradable sector workers. Figure 18 analyzes the average months formally employed per year for the nontradable workers. The results are very similar to those for the tradable workers (compare to Figure 10), with effects that are about 20-25 percent smaller than those for tradable workers. Figure 19 analyzes average earnings. Once again, the results are similar to those for the tradable sector workers (compare to Figure 15), with effects that are about 10-15 percent smaller.

These findings are consistent with prior work finding that nontradable sector workers face similar local effects to tradable sector workers following trade liberalization (Kovak 2013). This result is consistent with the fact that many tradable sector workers in harder hit locations appear to transition into the nontradable sector, as seen in Figure 12. 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.

6.6 Summary of Individual Analysis

In the preceding longitudinal analysis we have demonstrated that trade liberalization had large and long lasting effects on workers' employment and earnings outcomes during the 15 years following liberalization. Workers initially located in more adversely affected places experience large relative losses that never recover. Workers facing the most negative demand shocks are more likely to leave formal employment or to transition into jobs in the nontradable sector. While these adjustments

[35] To calculate the pre-liberalization trend control, we estimate a parallel specification for non-formal spells in 1986-89, allowing for right censoring in 1989, in order to avoid contaminating the pre-liberalization non-formal spells with the effects of trade reform.

[36] The average non-formal spell started in 1998 that is not truncated (finalized before December 2009) is of 21.4 months. The median is of 12 months.

appear to help workers recover some lost formal earnings in the hardest hit regions, they are too small to compensate for losses in the tradable sector.

Since our longitudinal data only cover the formal sector it remains possible that workers displaced from formal employment switched to informal employment. Supplementary regional analysis using the Census supports this hypothesis, finding relative increases in informal employment shares in more negatively affected regions. However, it is unlikely that these transitions allowed workers to recover their lost formal wages, as informal sector jobs usually pay lower wages and have lower wage growth (Goldberg and Pavcnik 2007). We find little evidence of workers moving to obtain formal employment in stronger labor markets, in spite of substantial unconditional mobility rates in Brazil.[37]

Finally, we find that these effects were quite consistent across different groups of workers. Those initially working in the nontradable sector experienced similar outcomes to those initially in the tradable sector, indicating substantial integration of the two sectors. In unreported results, we restrict attention to a smaller group of highly attached workers, those with 36 months or more formal employment during 1986-1989. The responses in this more restricted population were similar to those in the overall population, indicating that both highly connected and more casual workers were similarly affected by liberalization.

7 Conclusion

In this paper, we have documented the regional dynamics following the Brazilian trade liberalization of the early 1990s. Using matched employer-employee records 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, these effects grow for more than a decade before leveling off. This pattern suggests the presence of large mobility frictions for both labor and other complementary factors of production such as capital (see (Dix-Carneiro 2014)). We argued how this mechanism operates within a simple equilibrium model with local labor markets by extending Kovak (2013) and allowing for dynamic changes in equilibrium local quantities of labor and industry-specific factors in response to the trade-induced local shocks. We next showed some evidence in favor of our interpretation of the uncovered empirical pattern.

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 experienced relatively larger formal earnings losses and more time out of formal employment, with each of these effects growing over time. Using data from four rounds of the Brazilian Demographic Census we establish that informality and self-employment increase in regions that are harder hit relative to

[37] This result parallels that of Yagan (2014) for the U.S.

less negatively affected regions, suggesting that workers who drop out of the formal sector due to the trade induced negative local shock transition into informality and self-employment. Tradable sector workers in negatively affected regions are partly compensated for the adverse local shock by shifting into nontradable sectors, but do not respond to the shock by migrating away from their original regions. Workers initially employed in nontradable industries experienced similar patterns of effects, suggesting that local tradable and nontradable labor markets are well integrated.

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, we challenge the conventional wisdom in the labor economics literature that the impact of local shocks will be gradually dissipated through equalizing migration and provide an interpretation for these findings. Fourth, we are the first to document the medium- to long-run effects of a major trade policy reform by following workers over time, across sectors and across regions. Fifth, we find novel results regarding the effect of trade policy changes on informality and reconcile our findings with those of the existing literature.

We believe that a natural next step in this research agenda is to incorporate worker heterogeneity in our theoretical framework and to explore how trade-induced local shocks affected intra- and inter-regional inequality.

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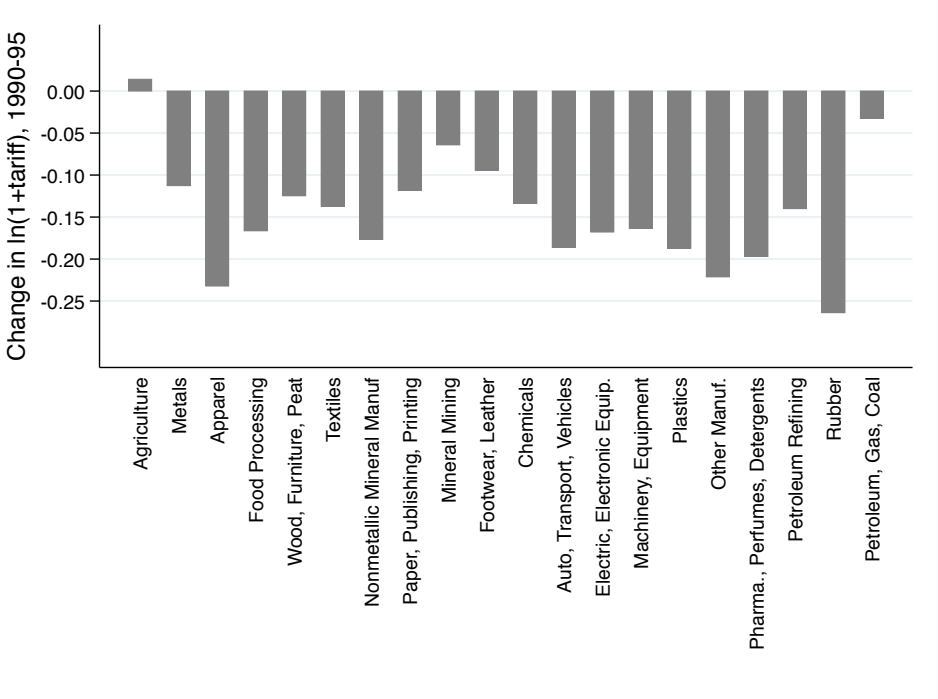
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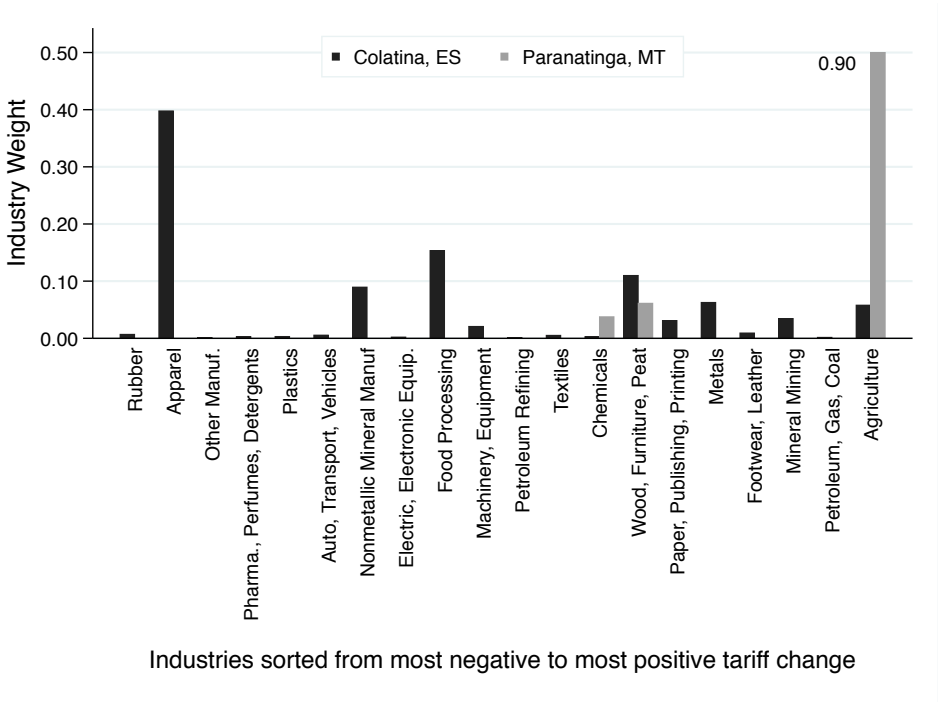
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Figure 1: Tariff Changes



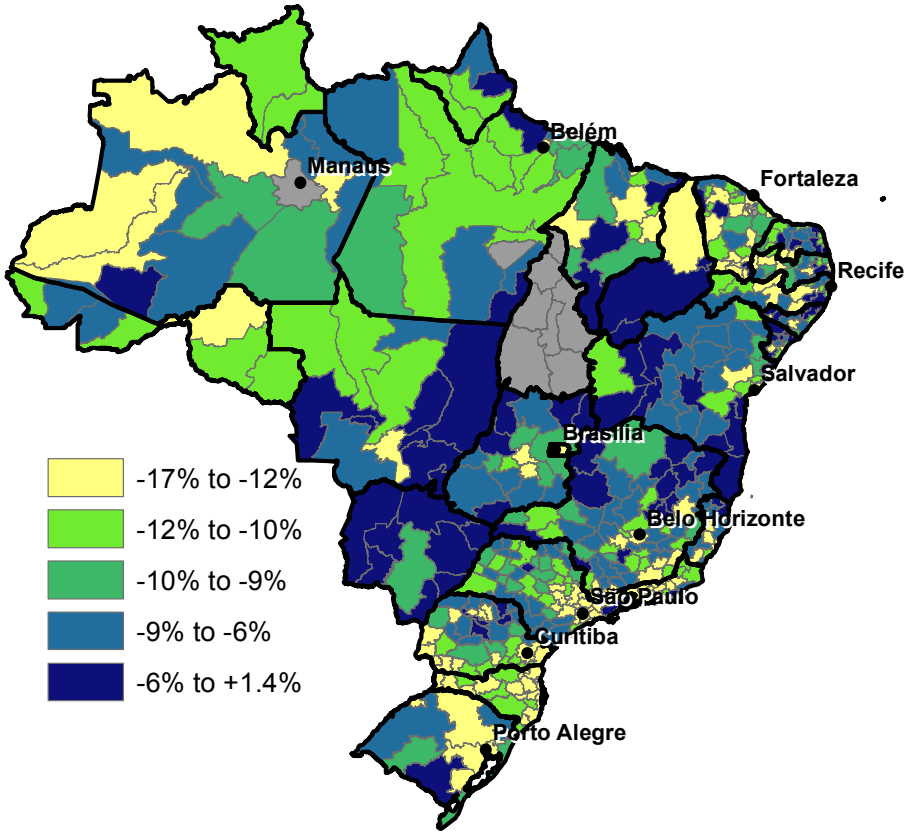
Industries sorted by 1991 employment (largest to smallest). Tariff data from Kume et al. (2003), aggregated to allow consistent industry definitions across data sources. See Kovak (2013) for details of the industry classification.

Figure 2: Variation Underlying Regional Tariff Change



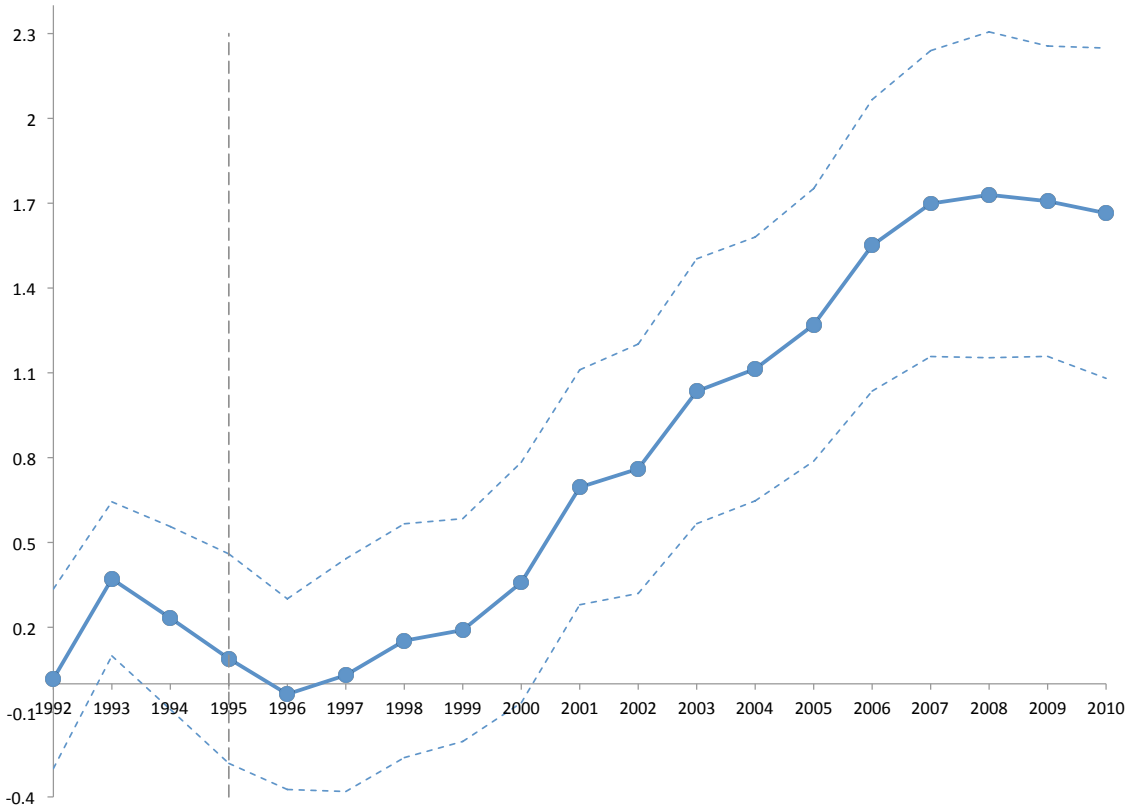
Industry distribution of 1991 employment in the two regions facing the most negative and most positive regional tariff changes. Industries sorted by the tariff change, shown in Figure 1. More weight on the left side of the figure leads to a more negative regional tariff change, and more weight on the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

Figure 3: Regional Tariff Changes



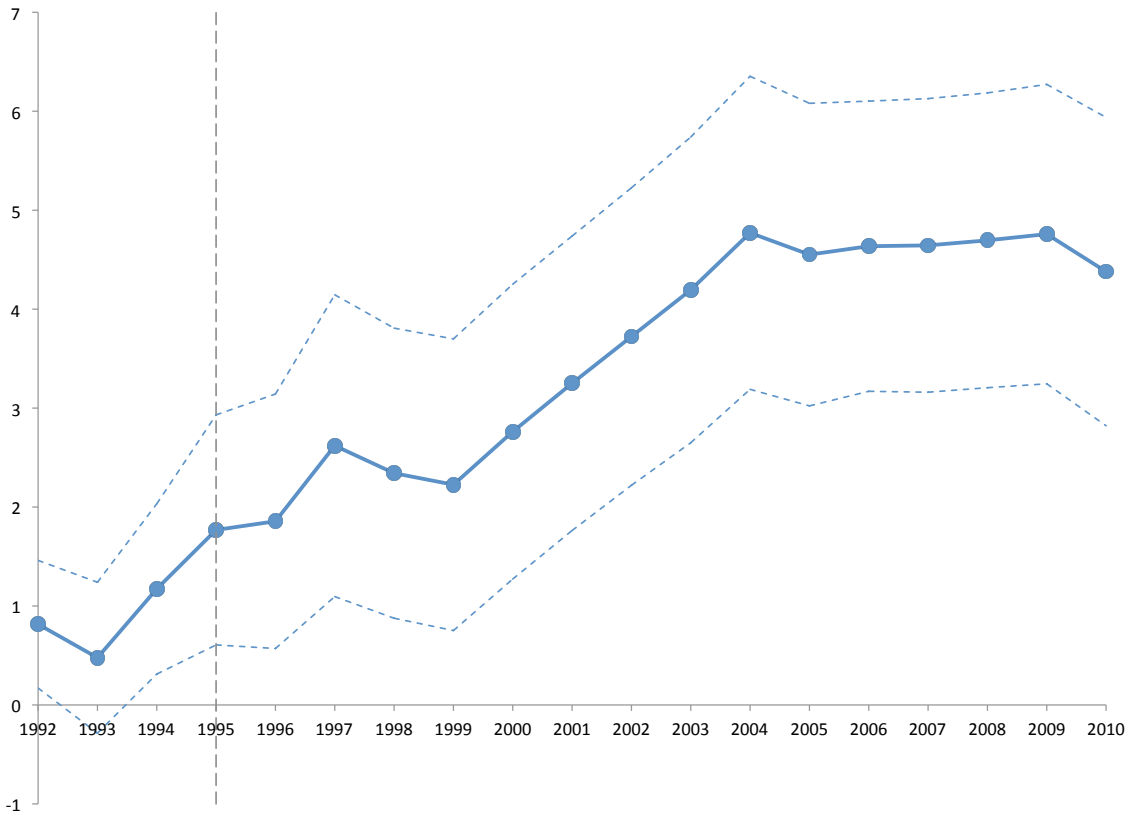
Local labor markets reflect microregions defined by IBGE, aggregated slightly to account for border changes between 1986 and 2010. Regions are colored based on the regional tariff change measure defined in (1). Dark lines represent state borders, gray lines represent consistent microregion borders, and gray regions show migroregions that 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



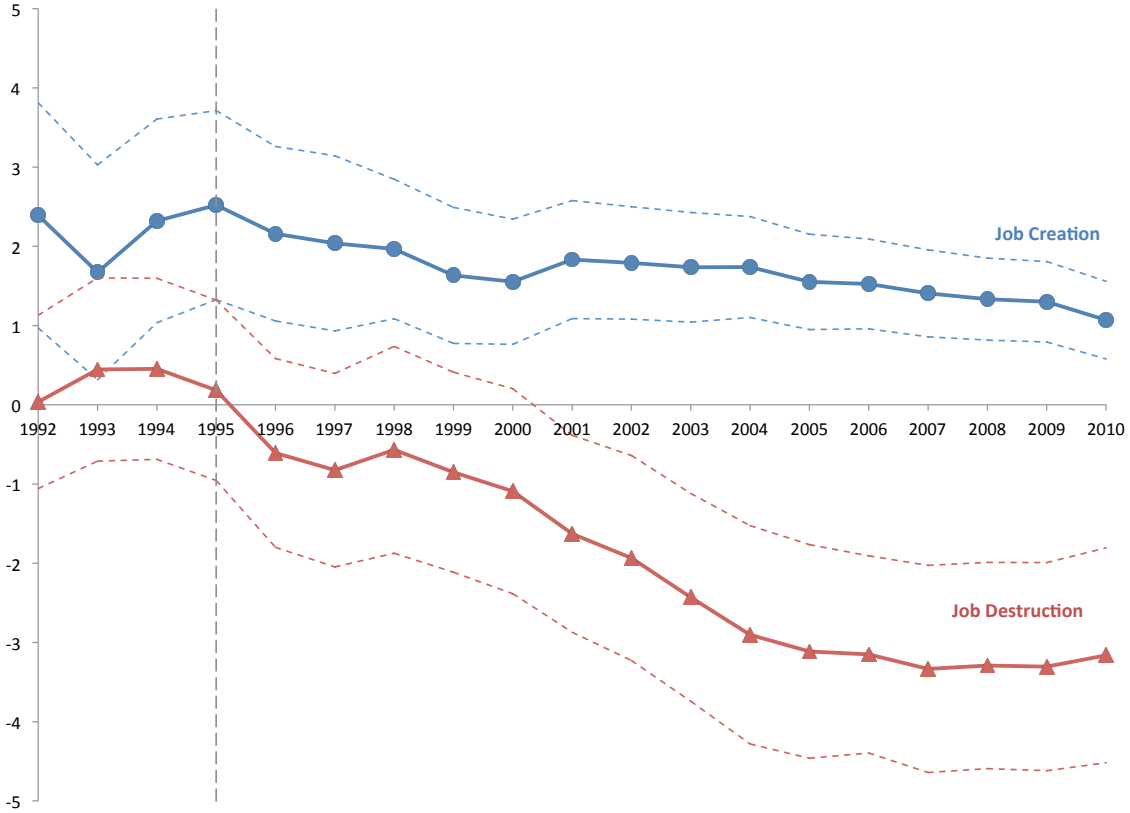
Each point reflects an individual regression coefficient where the dependent variable is the change in regional outcome from 1991 to the year listed on the x-axis. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with state fixed effects and pre-trend control. 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 Employment - 1992-2010



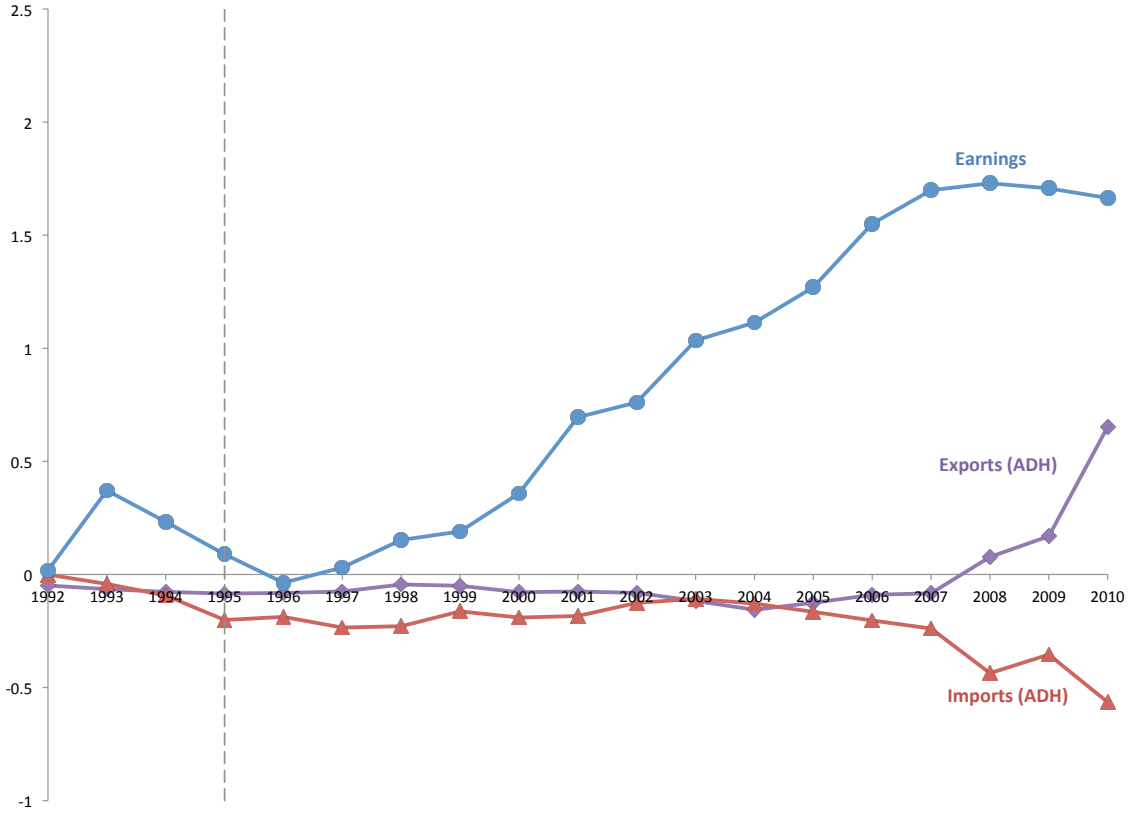
Each point reflects an individual regression coefficient where the dependent variable is the change in regional outcome from 1991 to the year listed on the x-axis. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with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

Figure 6: 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 destruction from 1991 to the year listed on the x-axis. 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. 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).

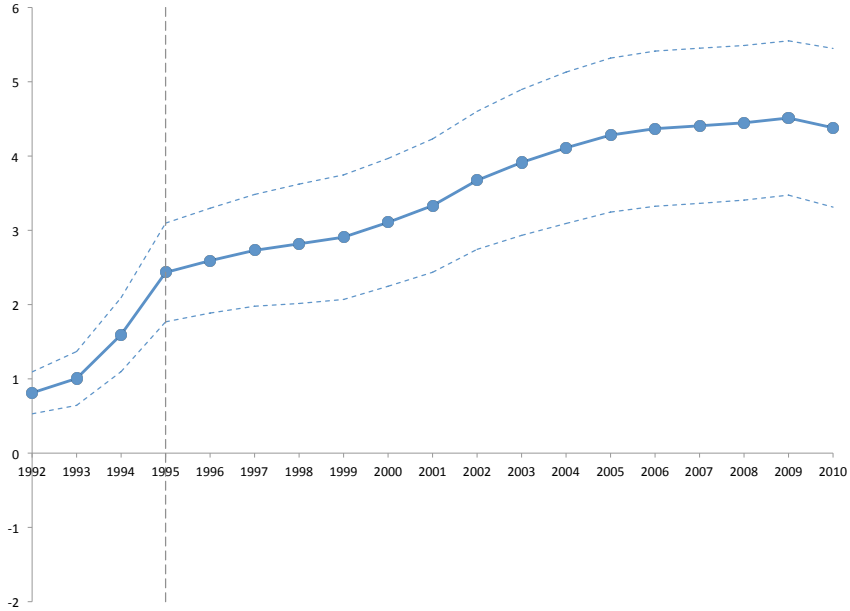
Figure 7: Regional log Employment, Imports per Worker, and Exports per Worker - 1992-2010



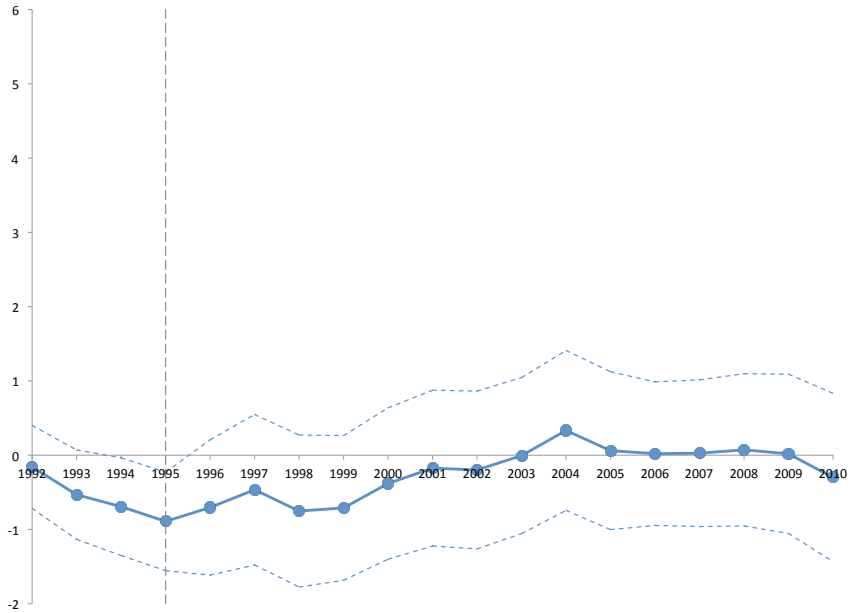
Each point reflects an individual regression coefficient where the dependent variable is the change in log employment, change in imports per worker, or change in exports per worker, calculated from 1991 to the year listed on the x-axis. For all years, the independent variable is the regional tariff change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

Figure 8: Capital Adjustment

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

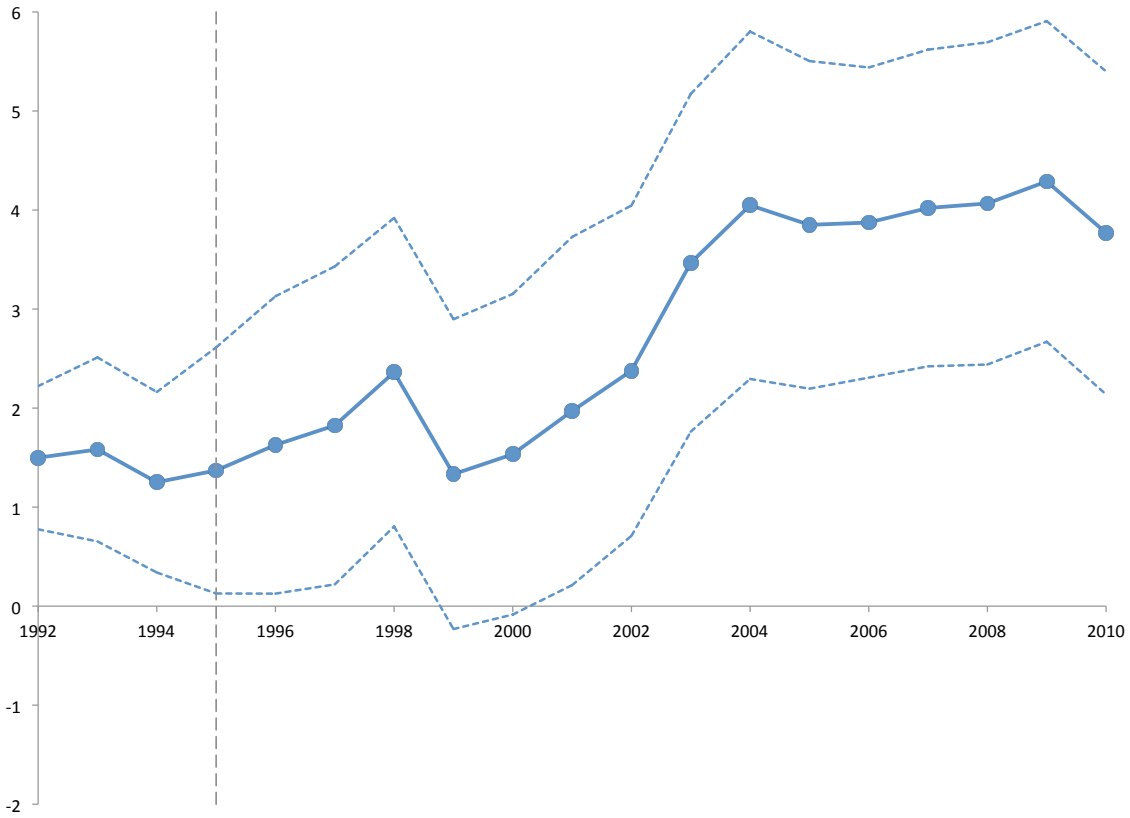


Panel B: Regional log Average Formal Establishment Size (Number of Workers) - 1992-2010



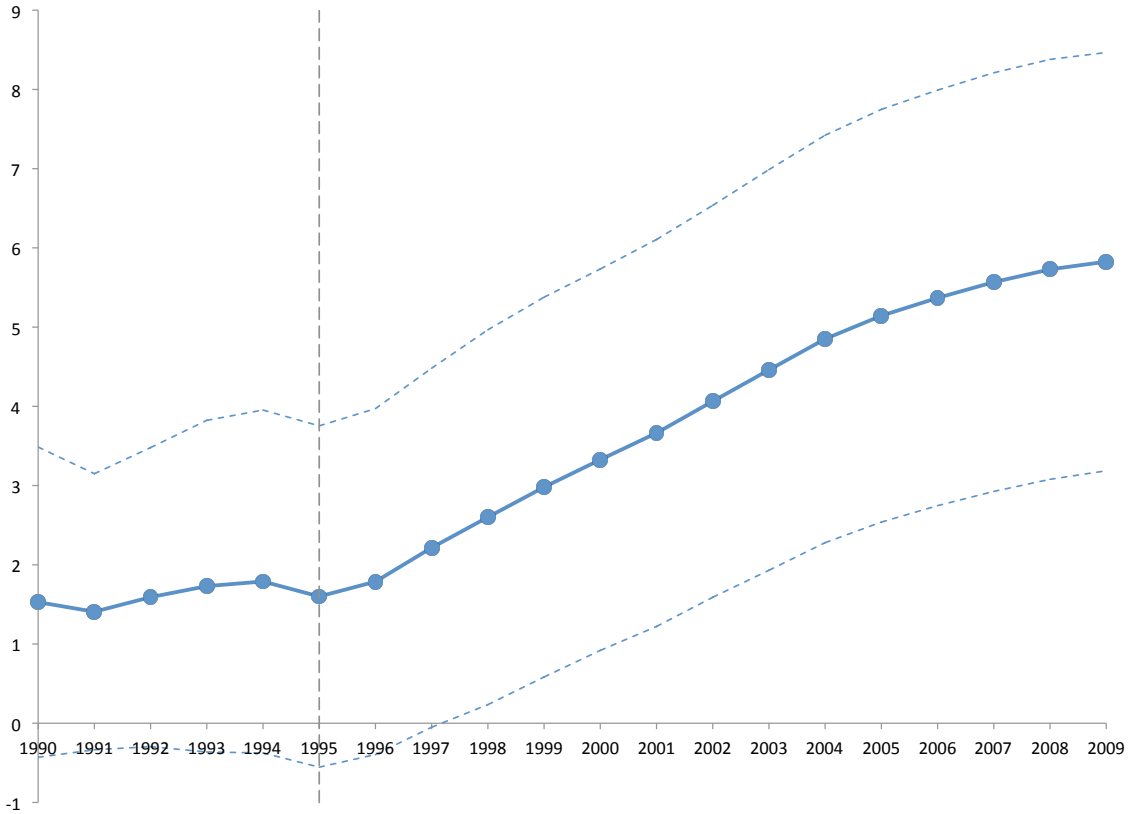
Each point reflects an individual regression coefficient where the dependent variable is the change in log number of establishments or log establishment size, calculated from 1991 to the year listed on the x-axis. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with state fixed effects and pre-trend control. 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 log Formal Wagebill - 1992-2010



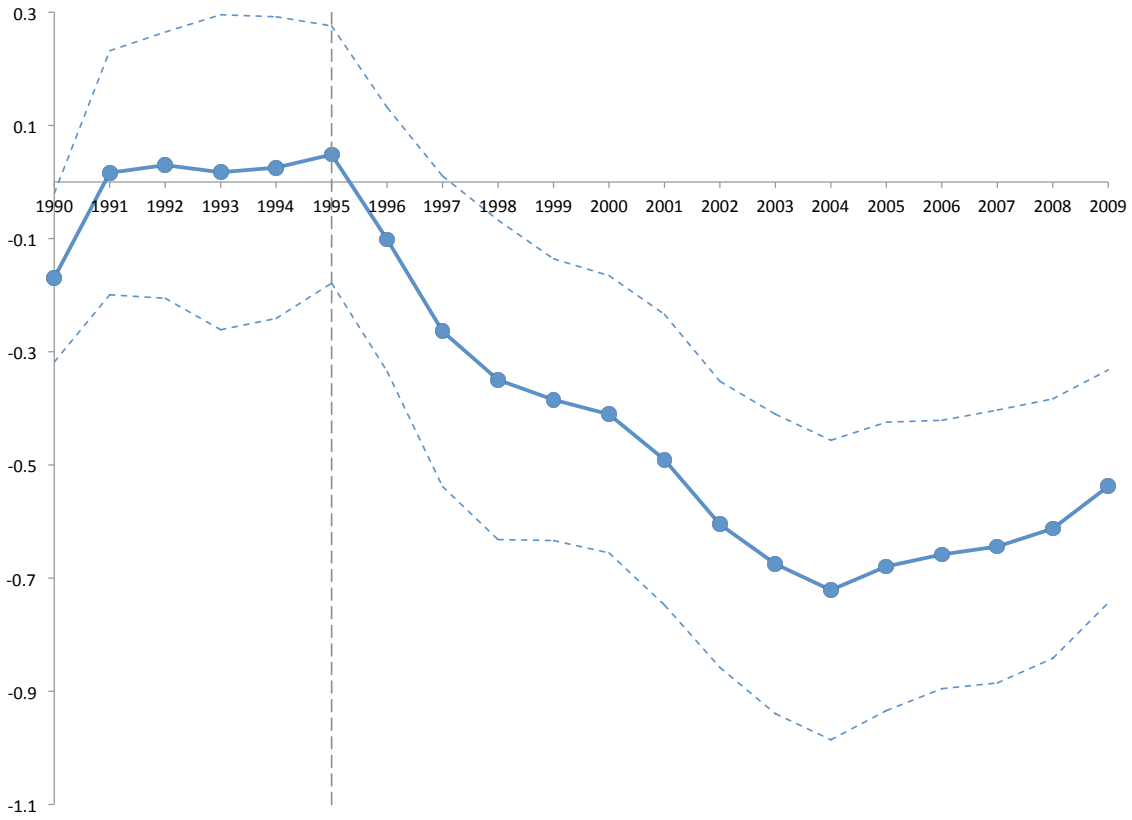
Each point reflects an individual regression coefficient where the dependent variable is the change in log formal wagebill from 1991 to the year listed on the x-axis. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text),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

Figure 10: Individual Average Months Formally Employed - Tradable Sector Worker Sample - 1990-2009



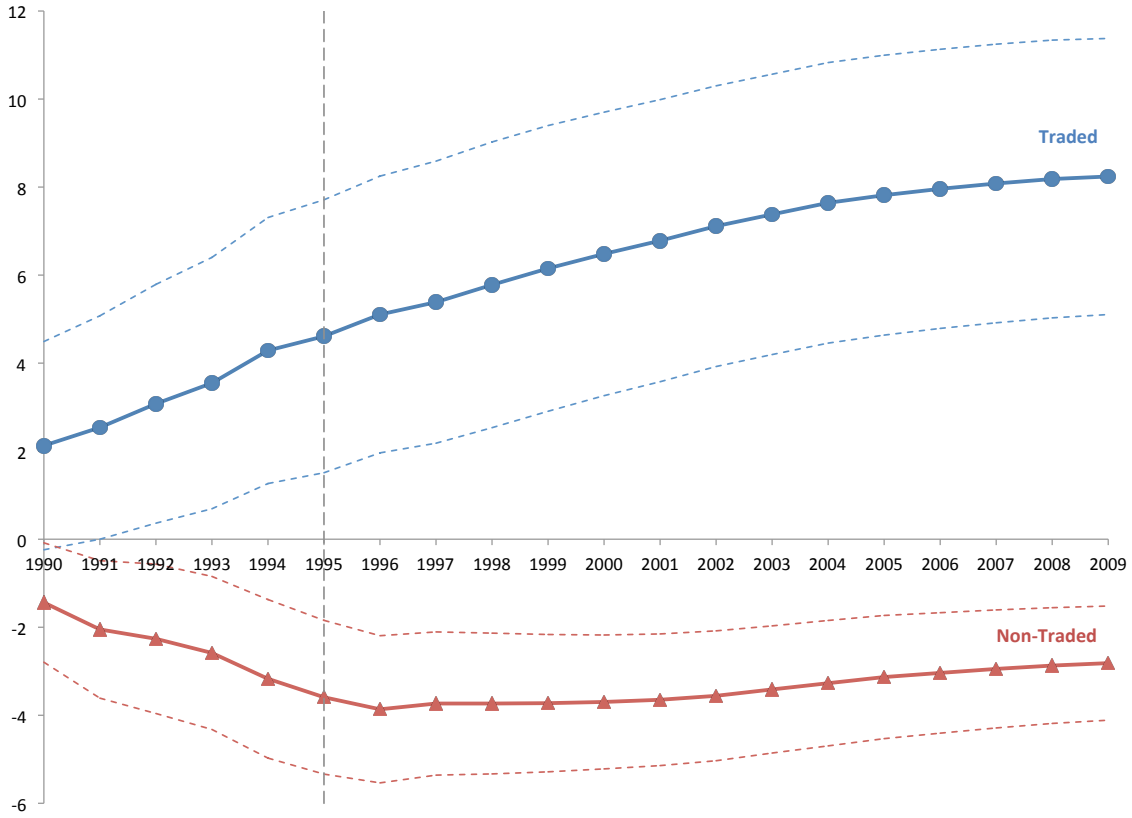
Each point reflects an individual regression coefficient where the dependent variable is the average number of months formally employed per year from 1990 to the year listed on the x-axis. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). 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: Fraction of the Year Outside Formal Employment - Tradable Sector Worker Sample - 1990-2009



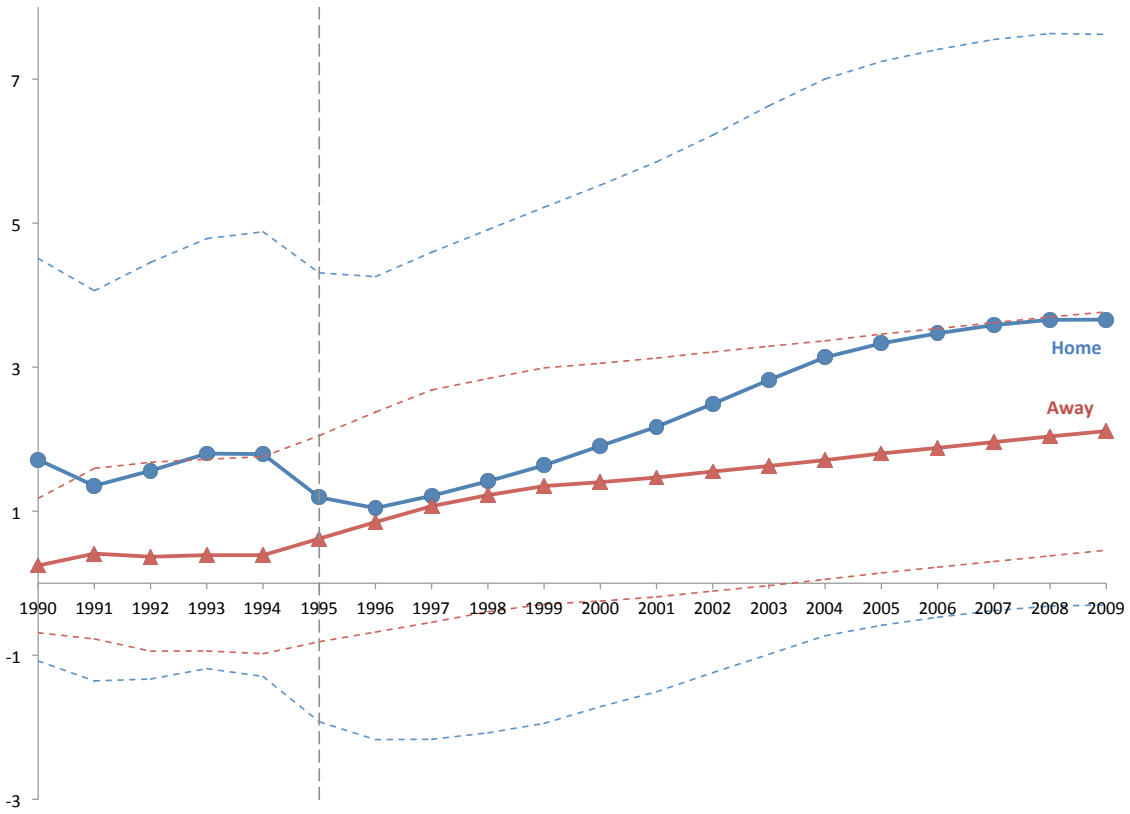
Each point reflects an individual regression coefficient where the dependent variable is the fraction of the year spent out of formal employment for the year listed on the x-axis. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described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

Figure 12: Individual Average Months Formally Employed, by Sector - Tradable Sector Worker Sample - 1990-2009



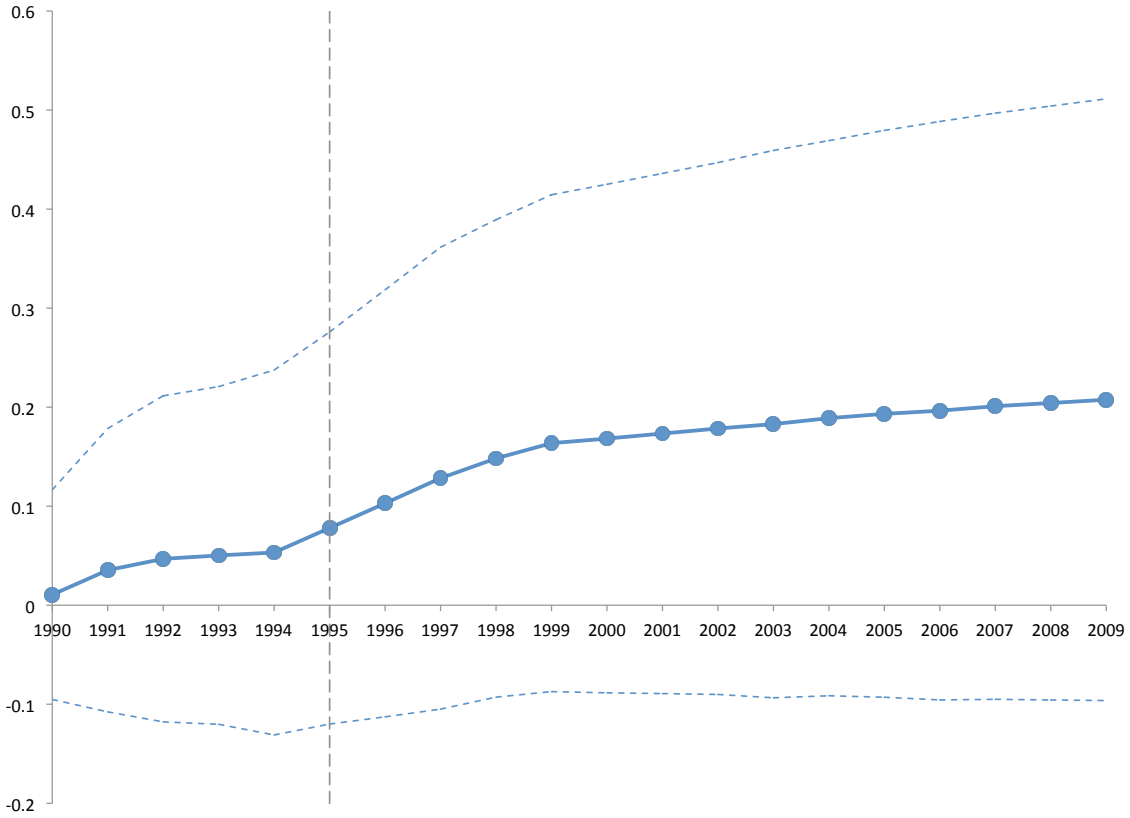
Each point reflects an individual regression coefficient where the dependent variable is the average number of months per year formally employed in the tradable or nontradable sector from 1990 to the year listed on the x-axis. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). 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, by Location - Tradable Sector Worker Sample - 1990-2009



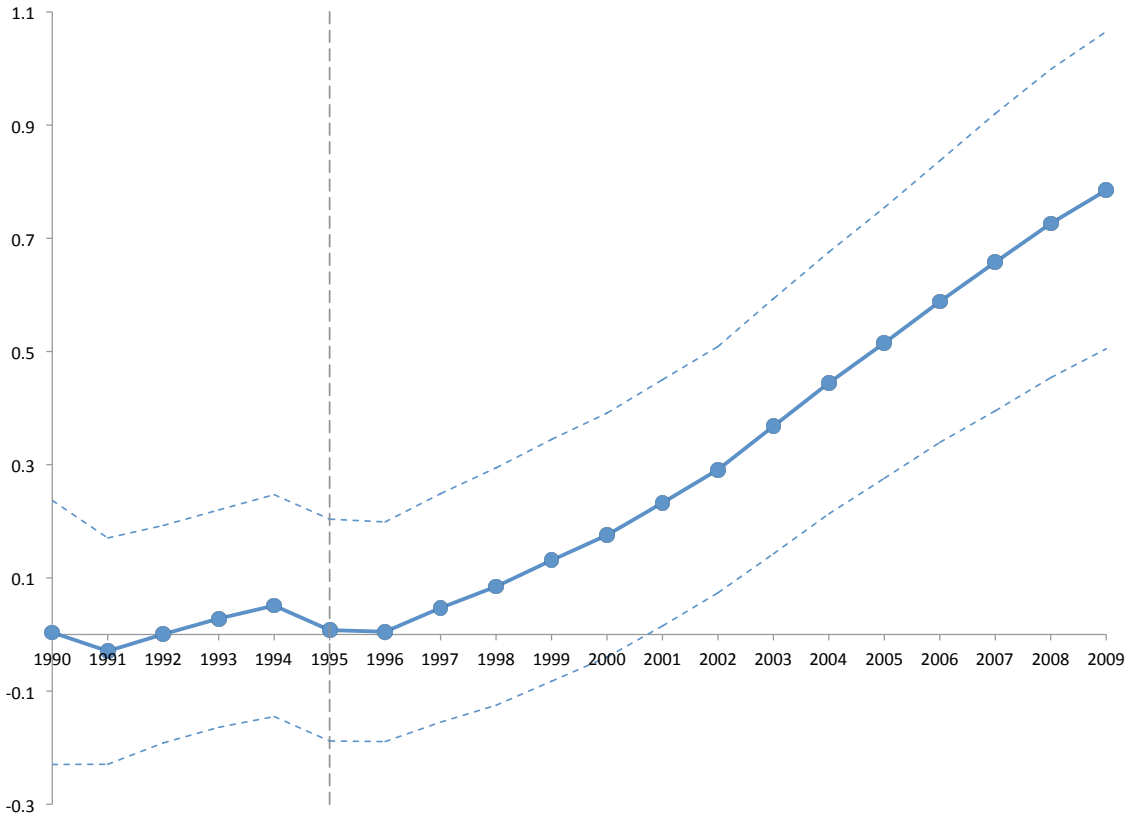
Each point reflects an individual regression coefficient where the dependent variable is the average number of months per year formally employed in the initial region or other region from 1990 to the year listed on the x-axis. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). 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: Individual Share of Formally Employed Months Away from Initial Region - Tradable Sector Worker Sample - 1990-2009



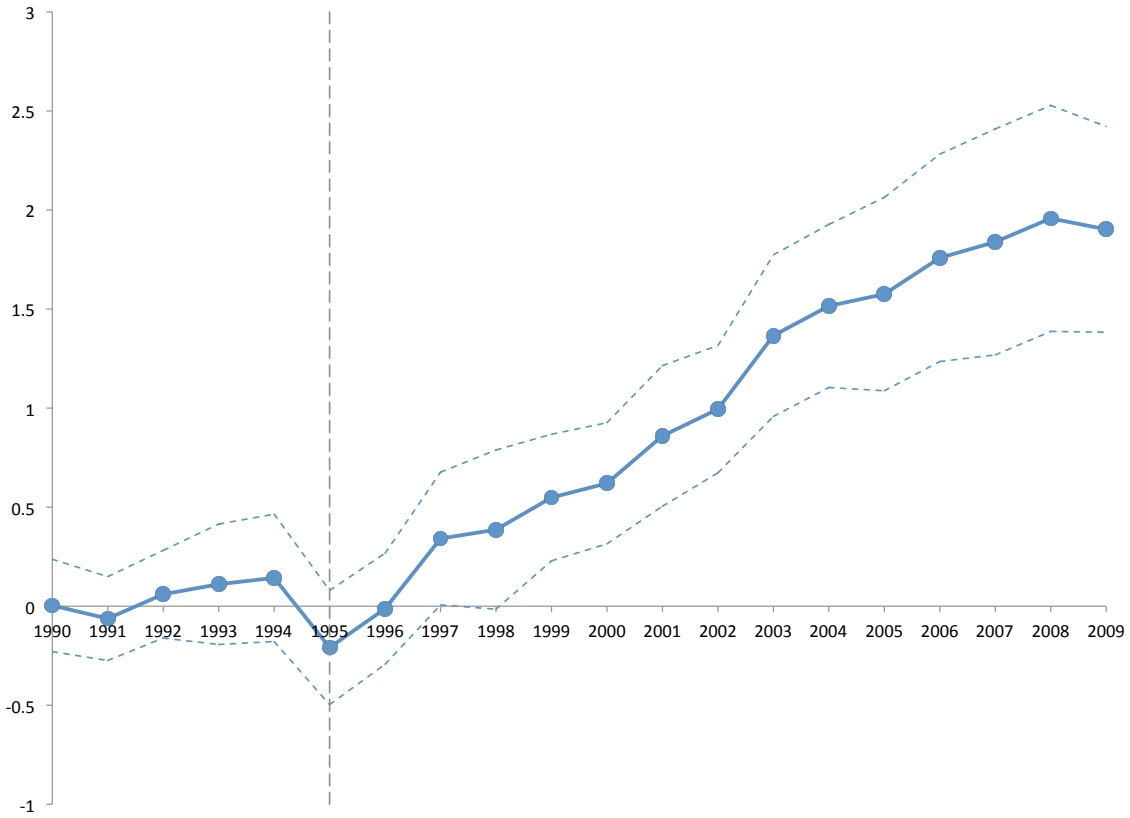
Each point reflects an individual regression coefficient where the dependent variable is the share of formally employed months in the year listed on the x-axis that the individual spent working in a region other than their initial region. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). 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 Earnings - Tradable Sector Worker Sample - 1990-2009



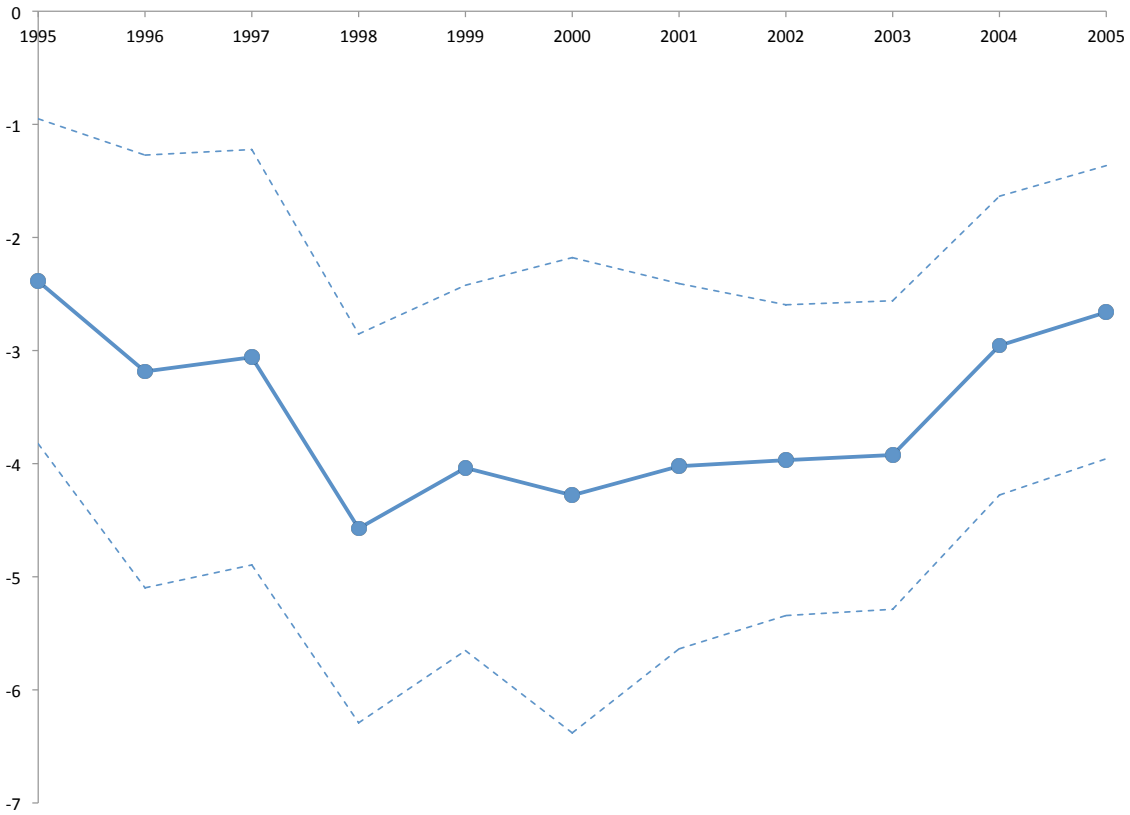
Each point reflects an individual regression coefficient where the dependent variable is the average earnings from 1990 to the year listed on the x-axis as a multiple of the worker's average pre-liberalization (1986-89) earnings. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). 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 Yearly Earnings - Tradable Sector Worker Sample - 1990-2009



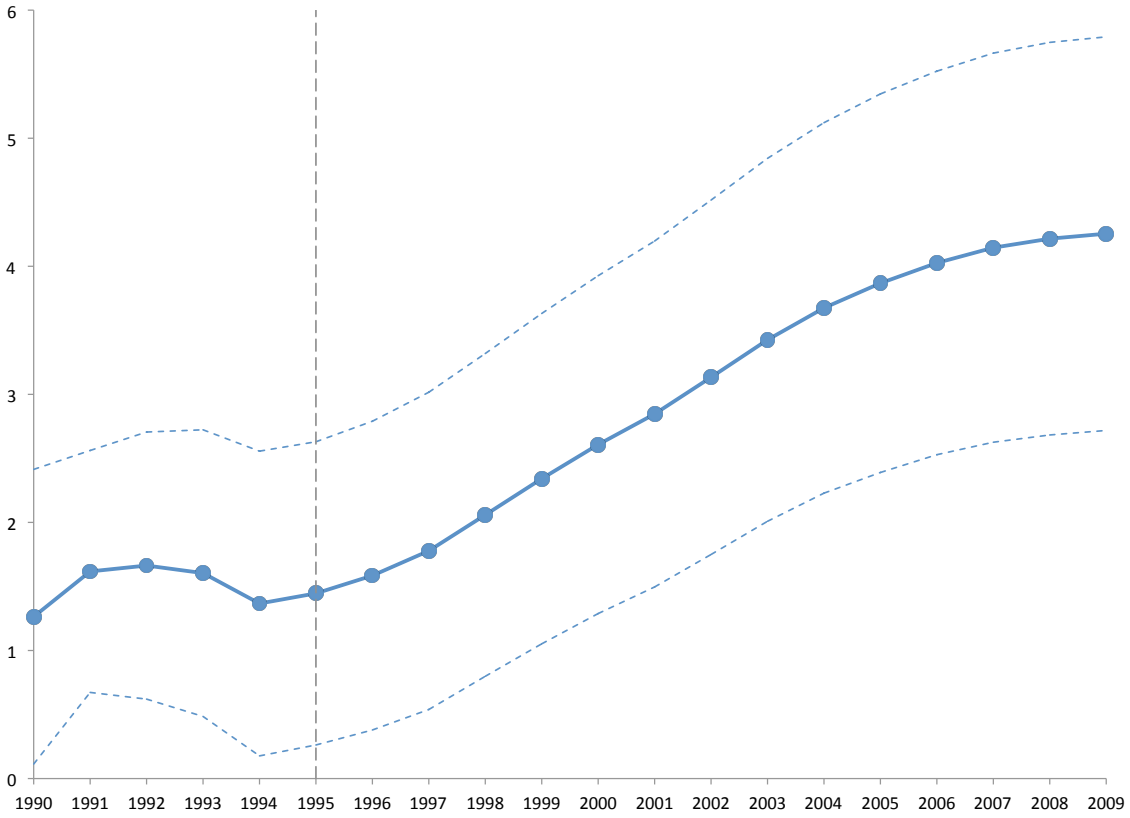
Each point reflects an individual regression coefficient where the dependent variable is the yearly earnings in the year listed on the x-axis as a multiple of the worker's average pre-liberalization (1986-89) earnings. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the text), with various controls (described in the text). 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: log Non-Formal Spell Length (Tobit) - Tradable Sector Worker Sample - 1995-2005



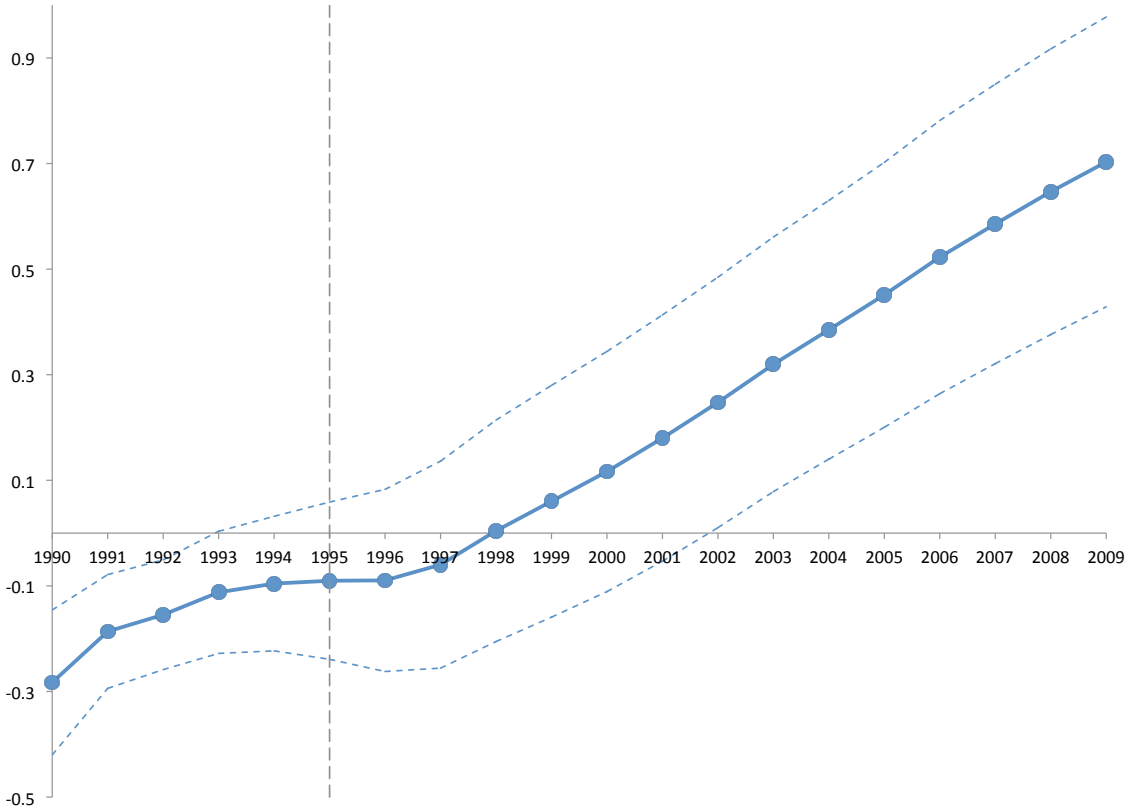
Each point reflects an individual Tobit regression coefficient where observations represent individual spells, and the dependent variable is the log spell length in years. For all years, the independent variable is the regional tariff change reflecting tariff changes from 1990-1995 (described in the 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). 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 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. 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). 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 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. 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). 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 Earnings						
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			-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	0.225	0.241	0.199	1.177	1.005***	0.958***
	(0.458)	(0.197)	(0.195)	(0.735)	(0.289)	(0.272)
Formal earnings pre-trend			-0.091*			-0.102
			(0.052)			(0.074)
State fixed effects (26)		✓	✓		✓	✓
R-squared	0.005	0.583	0.587	0.055	0.704	0.706
Panel C: Census Hourly Wages						
Regional tariff change	0.441	0.484**	0.427**	1.044	0.749***	0.691***
	(0.482)	(0.195)	(0.189)	(0.658)	(0.251)	(0.238)
Formal earnings pre-trend			-0.119			-0.128*
			(0.055)			(0.065)
State fixed effects (26)		✓	✓		✓	✓
R-squared	0.016	0.602	0.608	0.042	0.698	0.701

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)	(0.886)	(0.812)	(0.795)
Formal employment pre-trend			-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	2.318**	1.659***	1.423***	3.274**	2.321***	2.052**
	(0.936)	(0.561)	(0.536)	(1.412)	(0.886)	(0.862)
Formal employment pre-trend			0.223***			0.265***
			(0.071)			(0.100)
State fixed effects (26)		✓	✓		✓	✓
R-squared	0.113	0.542	0.563	0.096	0.546	0.559

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 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	1991-2000	1991-2005	1991-2010
	(1)	(2)	(3)	(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: Sample restriction - above median urban share (62%) - 238 obs				
Regional tariff change	0.072	0.272	1.297***	1.939***
	(0.205)	(0.232)	(0.277)	(0.343)
Panel E: Sample restriction - above median formal share (24%) - 238 obs				
Regional tariff change	0.182	0.412*	1.417***	1.907***
	(0.205)	(0.223)	(0.244)	(0.311)
Panel F: 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 G: 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	✓	✓	✓	✓
State fixed effects (26)	✓	✓	✓	✓

475 microregion observations, unless otherwise noted. 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. *** Significant at the 1 percent, ** 5 percent, * 10 percent level.

Table 4: Individual Analysis Summary Statsitics

	Tradable Sector Sample		Nontradable Sector Sample	
	mean	std. dev.	mean	std. dev.
Education				
Illiterate	0.02	0.15	0.02	0.13
4th grade incomplete	0.14	0.35	0.11	0.32
4th grade complete	0.26	0.44	0.20	0.40
8th grade incomplete	0.18	0.39	0.14	0.34
8th grade complete	0.14	0.34	0.14	0.35
High School incomplete	0.04	0.21	0.06	0.23
High School complete	0.12	0.32	0.19	0.39
College incomplete	0.02	0.15	0.03	0.18
College complete	0.07	0.25	0.12	0.32
Female	0.23	0.42	0.31	0.46
Age	34.9	7.5	35.1	7.6
Tenure at establishment (years)	4.4	5.1	4.4	5.3
December 1989 earnings (in 2010 R$)	1,951	2,573	1,890	2,828
1989 yearly yarnings (in 2010 R$)	19,721	25,435	19,349	27,890
Average yearly earnings 1986-1989 (in 2010 R$)	19,619	22,644	18,772	23,491
Firm employment	2,939	7,500	7,852	23,394
Months formally employed per year				
1990	10.2	3.4	9.9	3.8
1990-1995	8.1	3.8	8.2	3.9
1990-2000	6.9	3.7	7.1	3.9
1990-2005	6.2	3.7	6.3	3.8
1990-2010	5.7	3.7	5.9	3.8
Employed in different region in 1994 than in 1989	0.09	0.08	0.10	0.09
Observations	682,210		1,134,900	

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.

Table 5: Regional Employment Category Shares - 2000, 2010

	1991-2000			1991-2010		
Dependent variables	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Share Informally Employed						
Regional tariff change	-0.259	-0.210***	-0.248***	-0.515***	-0.453***	-0.523***
	(0.076)	(0.053)	(0.051)	(0.146)	(0.084)	(0.074)
Dependent variable pre-trend			0.073***			0.139***
			(0.025)			(0.046)
State fixed effects (26)		✓	✓		✓	✓
Panel B: Share Not Employed						
Regional tariff change	-0.176**	-0.195***	-0.198***	0.191**	0.114**	0.115**
	(0.072)	(0.053)	(0.052)	(0.078)	(0.057)	(0.054)
Dependent variable pre-trend			0.060			0.001
			(0.054)			(0.095)
State fixed effects (26)		✓	✓		✓	✓
Panel C: Share Informally Employed, Omitting Self-Employed						
Regional tariff change	-0.267***	-0.284***	-0.291***	-0.026	-0.120	-0.138
	(0.072)	(0.053)	(0.052)	(0.132)	(0.101)	(0.097)
Dependent variable pre-trend			0.139***			0.238***
			(0.033)			(0.059)
State fixed effects (26)		✓	✓		✓	✓
Panel D: Share Self-Employed						
Regional tariff change	0.029	0.083**	0.074**	-0.416***	-0.302***	-0.399***
	(0.039)	(0.036)	(0.037)	(0.114)	(0.084)	(0.111)
Dependent variable pre-trend			0.020			0.226**
			(0.027)			(0.117)
State fixed effects (26)		✓	✓		✓	✓
Panel E: Share Not Employed or Self-Employed						
Regional tariff change	-0.158*	-0.112*	-0.129**	-0.272**	-0.208**	-0.292***
	(0.083)	(0.054)	(0.057)	(0.136)	(0.091)	(0.097)
Dependent variable pre-trend			0.037			0.207**
			(0.052)			(0.089)
State fixed effects (26)		✓	✓		✓	✓

Decennial Census data. 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. All instruments' first stage F statistics ≥ 65. *** Significant at the 1 percent, ** 5 percent, * 10 percent level.

A Model

A.1 Setup and Equilibrium Restriction

This section introduces an extension of the model in Kovak (2013), which itself is based upon Jones (1975). 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

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

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

Perfect competition implies that the price equals factor payments,

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

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 θ_i be the cost share of the specific factor in industry i. Differentiating (A3),

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

which follows from the envelope theorem for cost minimization,

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

Now we differentiate the factor market clearing conditions, and deviate from Kovak (2013) by allowing for the possibility that $\hat{T}_i \neq 0$.

$$\hat{Y}_i = \hat{T}_i - \hat{a}_{Ti}. \tag{A6}$$

$$\sum_i \lambda_i(\hat{a}_{Li} + \hat{Y}_i) = \hat{L}, \tag{A7}$$

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

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

Combining (A6) - (A8) yields

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

Equations (A4) and (A9) can be expressed in matrix form as

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

For convenience of notation, rewrite this expression as

$$\left[\begin{array}{c|c} \boldsymbol{\Theta} & \boldsymbol{\theta}_{\mathbf{L}} \\ \hline \boldsymbol{\lambda}' & -\sum_i \lambda_i\sigma_i \end{array}\right] \left[\begin{array}{c} \hat{\mathbf{R}} \\ \hline \hat{w} \end{array}\right] = \left[\begin{array}{c} \hat{\mathbf{P}} \\ \hline \hat{L} - \sum_i \lambda_i \hat{T}_i \end{array}\right] \tag{A11}$$

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,

$$\hat{w} = \frac{\left(\hat{L} - \sum_i \lambda_i \hat{T}_i\right) - \boldsymbol{\lambda}'\boldsymbol{\Theta}^{-1}\hat{\mathbf{P}}}{-\sum_i \lambda_i\sigma_i - \boldsymbol{\lambda}'\boldsymbol{\Theta}^{-1}\boldsymbol{\theta}_{\mathbf{L}}} \tag{A12}$$

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

$$\hat{w}_r = \sum_i \beta_{ri}\hat{P}_i - \delta_r\left(\hat{L}_r - \sum_i \lambda_{ri}\hat{T}_{ri}\right), \tag{A13}$$

$$\text{where} \quad \beta_{ri} \equiv \frac{\lambda_{ri}\frac{1}{\theta_i}}{\sum_j \lambda_{rj}\frac{1}{\theta_j}} \qquad \text{and} \quad \delta_r \equiv \frac{1}{\sum_k \lambda_{rk}\frac{1}{\theta_k}}.$$

Plug this into (A4) to solve for the change in specific factor returns.

$$\hat{R}_{ri} = \frac{1-\theta_i}{\theta_i}\delta_r\left(\hat{L}_r - \sum_i \lambda_{ri}\hat{T}_{ri}\right) - \frac{1-\theta_i}{\theta_i}\sum_j \beta_{rj}\hat{P}_j + \frac{1}{\theta_i}\hat{P}_i. \tag{A14}$$

Equations (A13) and (A14) 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 (A13), 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 (A14) with factors held fixed yields

$$\begin{aligned}\sum_i \lambda_{ri}\hat{R}_{ri} &= \sum_i \lambda_{ri}\left(-\frac{1-\theta_i}{\theta_i}\sum_j \beta_{rj}\hat{P}_j + \frac{1}{\theta_i}\hat{P}_i\right) \\ &= \left(\sum_j \beta_{rj}\hat{P}_j\right)\sum_i \lambda_{ri}\left(-\frac{1-\theta_i}{\theta_i}\right) + \sum_i \lambda_{ri}\frac{1}{\theta_i}\hat{P}_i \\ &= \left(\sum_j \beta_{rj}\hat{P}_j\right)\sum_i \lambda_{ri}\left(-\frac{1-\theta_i}{\theta_i}\right) + RTC_r\sum_i \lambda_{ri}\frac{1}{\theta_i} \\ &= \left(\sum_j \beta_{rj}\hat{P}_j\right),\end{aligned} \tag{A15}$$

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.