Trade, Labor Market Concentration, and Wages*

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Abstract

I estimate how trade liberalization affects employers' wage-setting power using Brazil's tariff cuts, employer-employee data, and a model linking concentration to markdowns. Workers took home 53 cents per marginal dollar pre-shock. A 10% increase in import competition raised concentration by 22% but lowered the take-home share by under a cent, because workers substitute nearly as inelastically across firms as across markets. Incumbent firms' wages fell 1.2%, one third through higher markdowns; employment composition shifting toward higher-revenue-product firms cushioned average wages. Accounting for informality halves the markdown effect as self-employment curbs employers' power; effects are concentrated in the most formal regions.

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1 Introduction

A growing body of evidence suggests that import competition persistently reduces wages in more exposed labor markets relative to less exposed ones. These patterns have been documented in various contexts, including India (Topalova, 2010), Brazil (Kovak, 2013; Dix-Carneiro and Kovak, 2017, 2019), and the U.S. (Autor, Dorn and Hanson, 2013, 2021). Why? This paper revisits Brazil's liberalization episode in light of new data and theory to test one potential mechanism for these negative wage effects: trade-induced increases in firm labor market power.

A robust prediction of trade models with firm heterogeneity is that trade liberalization reallocates labor towards larger, more productive firms. In a *bilateral* liberalization, as in Melitz (2003), exporters expand in response to increased foreign demand, thickening the right tail of the firm size distribution, and eliminating the least productive firms on the left tail in equilibrium. In a *unilateral* liberalization, as in Melitz and Ottaviano (2008), increased import competition directly eliminates the least productive firms, cutting off the left tail of the firm size distribution, with the displaced labor being reallocated towards surviving firms in equilibrium. In either case, trade shifts the firm size distribution rightward, raising average productivity but also increasing labor market concentration. If labor markets are imperfectly competitive, with market power increasing in firm size—consistent with developing country evidence (Amodio et al., 2025)—increased concentration can strengthen surviving firms' wage-setting position, putting downward pressure on wages. How much of trade's negative wage effects can be accounted for by increased labor market power can only be assessed empirically.

This paper is an empirical study of the relationship between trade, labor market concentration, and wages in the context of Brazil's 1990s unilateral trade liberalization. I focus on the labor market power commanded by formal sector firms, and measure the degree of a firm's market power as the firm's wage markdown, the ratio of its marginal revenue product of labor to its wage. Formal firms are a natural focus. First, formal workers experienced the most negative effects of import competition on earnings, and these effects grew for more than a decade (Dix-Carneiro and Kovak, 2017, 2019). Second, formal firms are the largest employers in any one labor market, and thus those for which size-related labor market distortions are most relevant (Dix-Carneiro et al., 2026).

Using a sufficient statistics approach, employer-employee linked data, and import tariff reductions measured for 285 sectors, I test whether trade increased labor market concentration among formal employers in more exposed markets, and estimate the consequences for formal sector wages. I find that, within a decade, a 10% increase in import competition exposure increased concentration by 22%, driven by exporters gaining payroll share as fewer firms entered exposed markets. Yet the increase in concentration moved wage markdowns little: because Brazilian workers substitute across firms within a market nearly as inelastically as across markets, the wage take-home share—53 cents

on the dollar pre-shock—fell by less than half a cent. Wages fell as well, and higher markdowns account for one third of the decline at the firms that remained in exposed markets. Finally, incorporating informal wage work and self-employment leaves the level of markdowns unchanged but halves the effect of trade liberalization on them, because self-employment curbs formal firms' market power.

I develop a parsimonious model of imperfectly competitive labor markets in which firm size matters for wage setting. Workers have idiosyncratic preferences over jobs, yielding nested CES labor supply as in Berger, Herkenhoff and Mongey (2022), henceforth BHM, and firms compete for workers à la Cournot. Two elasticities govern how workers respond to wage changes: a *cross-market* elasticity of substitution, and a *within-market* cross-firm elasticity of substitution. The model delivers two results. First, a market's average wage markdown is determined by these two elasticities and by the market's payroll Herfindahl index, the sum of its firms' squared payroll shares, regardless of firms' production functions or the competition structure in product markets, on which I remain agnostic. Second, the effect of import competition on a market's average markdown can be quantified by two sufficient statistics: its effect on local labor market concentration, and the *gap* between workers' cross-market and within-market cross-firm inverse elasticities of substitution. The gap matters because, if workers substitute across markets as easily as across firms within a market, firms effectively operate in a single national market, where their relative size is inconsequential to market power.

I estimate these sufficient statistics using Brazil's trade liberalization. Between 1990 and 1994, Brazil reduced import tariffs from an average of 33% to an average of 13%, and because sectors differed in their pre-reform levels of protection, the tariff cuts varied substantially across sectors. I measure them for 285 sectors, an order of magnitude more than the about 20 industries of previous studies of this liberalization, and assign them to each firm in the employer-employee linked data. I define a local labor market as a microregion $\times$ occupational group cell, motivated by substantial and stable job-to-job transitions within these cells, and I measure its "import competition exposure" as a shift-share average of the tariff cuts faced by its tradable firms, in which each firm's share is its contribution to the market's baseline payroll Herfindahl.

I find that a 10 percent increase in import competition exposure increased local labor markets' payroll Herfindahl by 0.017 points between 1991 and 2000 relative to less exposed markets, with little evidence of pre-trends. This effect corresponds to a 22% increase relative to the 0.08 pre-reform mean across markets weighted by formal employment. I weight markets by their formal employment because doing so ensures that the effect is identified by the tariff cuts themselves: my estimate is unchanged when I recenter exposure on its expected value under random reassignment of tariff cuts across sectors. Concentration increased because the number of firms fell, entrants captured a smaller share of the market's payroll, and payroll concentrated among exporters, with no evidence that workers moved towards exporters. Concentration thus increased through selection

among firms, as in Melitz and Ottaviano (2008), rather than through the expansion of exporters into foreign markets, as in Melitz (2003).

Trade shocks that vary across firms within markets allow me to estimate both elasticities using IV, as opposed to BHM's method of indirect inference, adding transparency to the identifying source of variation, and dispensing with assumptions on production functions and product market structure. I estimate a within-market cross-firm inverse elasticity of substitution of 0.832, using within-market cross-firm variation in tariff cuts, and a cross-market inverse elasticity of substitution of 1.454, using cross-market variation in import competition exposure. Along with the pre-liberalization level of concentration, these elasticities imply that Brazilian workers took home only 53 cents of every marginal dollar they generated, less than the 65 to 74 cents estimated for the US, though consistent with estimates for other developing countries. This labor market power stems from workers' inelastic substitution across firms rather than from concentration: had employers been atomistic, workers would have taken home 55 cents. Brazil's within-market cross-firm elasticity of substitution is *nine* times as inelastic as BHM's estimate for the US.[1] Combined with the effect on concentration, the 0.622 gap between the two elasticities implies that a 10% increase in import competition exposure reduced the average wage take-home share by 0.37 cents on the dollar, or 0.7% of the average wage.

My theory and data also allow me to revisit existing estimates of the wage effects of Brazil's liberalization. Kovak (2013) and Dix-Carneiro and Kovak (2017) use an exposure measure derived from a specific-factors model with competitive labor markets and workers who are immobile across regions, and built from about 20 industries. My exposure measure is instead based on the equilibrium of oligopsonistic local labor markets, in which workers can substitute across firms, occupations, and regions: it weights the tariff cuts faced by a market's tradable firms, measured for 285 sectors, by their contributions to the market's payroll Herfindahl, which determines the market's average wage markdown under Cournot competition. These finer tariff shocks allow me to document, with the tools of the recent shift-share literature, what identifies each estimate. With my measure, the effects on concentration and on the number of firms are identified by differences in tariff cuts across sectors (the shifts) rather than by markets' sectoral composition (the shares). Their exposure measure varies across regions mainly with the share of agriculture, a sector with a small tariff cut. Despite these different sources of variation, the effects on employment and wages that I obtain as of 2000 are qualitatively similar to theirs, with a somewhat smaller effect on wages.

Specifically, at incumbent firms, those present in a market in both 1991 and 2000, a 10% increase in exposure reduced wage premia by 1.2%, or about 1% per standard deviation of exposure. For comparison, Dix-Carneiro and Kovak (2017) estimate a decline in formal earnings of about 2% per standard deviation of their exposure measure, which I reproduce in the RAIS data under their design.

[1] Brazil's inelastic cross-firm substitution may reflect high valuations of formal sector job amenities combined with search frictions: Brazilian workers are willing to pay 18–34% of wages for formal sector attributes (Felix et al., 2026), US workers below 10% for most job attributes (Maestas et al., 2023).

Beyond the part due to higher markdowns, this decline reflects a lower marginal revenue product of labor, which also accounts for the entire wage decline at the firms facing larger tariff cuts than their local competitors, a decline identified by those firms' own tariff cuts. Since firm productivity rose during the liberalization (Muendler, 2004), these reductions more likely reflect lower output prices than productivity losses. Across all of a market's firms, a shift in the composition of employment toward firms with a higher marginal revenue product of labor offset most of this decline within firms, so that the market's average marginal revenue product fell by less (Section 7).

My baseline results do not consider how informal wage work and self-employment, which together accounted for 58% of Brazil's employment in 1991, might increase or decrease formal sector firms' power over workers seeking formal employment. In Section 8, I extend the model to incorporate self-employment, as an alternative that workers choose, and informal wage work, as the consequence of involuntary separation from formal employment. The extension embeds both self-employment, as in Amodio, Medina and Morlacco (2025), and informal wage work, which may occur within or outside formal firms (Ulyssea, 2018), into a model of labor market oligopsony where finitely many differentiated formal firms compete for workers in dual local labor markets. Worker preferences, rather than comparative advantage as in Amodio, Medina and Morlacco (2025), drive sorting and are the source of formal firms' market power. Workers took home 53 cents of every marginal dollar in both models. I find that self-employment works as an elastic outside option that curbs formal firms' market power, while the risk of involuntary separation out of the formal sector slightly raises it. As a result, the effect of trade liberalization on markdowns is smaller in the extended model, 0.15 cents against 0.37, and concentrated in the most formal microregions.

The findings in this paper matter for our understanding of the effects of import competition on local labor markets, and of the extent of firms' labor market power. The effects of Brazil's trade liberalization on firm and worker outcomes have been widely studied,[2] with the overall takeaway being that tariff reductions raised firm productivity but brought persistent employment and earnings losses to the workers and regions most exposed to them. My focus is on mechanisms: I test whether trade-induced increases in labor market concentration reduce formal sector wages by increasing firm labor market power.[3] While some studies have documented that wages are lower in more concentrated local labor markets (e.g., Azar, Berry and Marinescu (2022); Azar, Marinescu and Steinbaum (2022)), and a few others have estimated that trade increases labor market concentration (e.g., Benmelech, Bergman and Kim (2022); Pham (2023)), to the best of my knowledge this is the first paper to provide a joint study of the relationship between trade, labor market concentration,

[2] See, for example, Muendler (2004), Gonzaga, Menezes Filho and Terra (2006), Krishna, Poole and Senses (2012), Dix-Carneiro and Kovak (2017), Dix-Carneiro, Soares and Ulyssea (2018), and Dix-Carneiro et al. (2026).

[3] Segerstrom and Sugita (2015) show that in a multi-industry extension of Melitz (2003), bilateral liberalization raises productivity more in nonliberalized industries, as cross-industry labor reallocation offsets within-industry selection. The critique does not apply to a unilateral liberalization, where the relevant framework is Melitz and Ottaviano (2008) and wages are pinned down by a freely traded outside good.

and wages. By showing when and how the key elasticities of substitution in models of oligopsony can be identified with IV rather than indirect inference, this paper also contributes to the literature on methods for estimating firm labor market power (e.g., Manning (2003); Dube et al. (2020); Lamadon, Mogstad and Setzler (2022); Yeh, Macaluso and Hershbein (2022)).

My findings show that Brazil's labor markets are much less competitive than markets in more advanced countries like the US,[4] and offer an explanation for why: Brazilian workers substitute across firms within a market far less elastically than US workers do. More attention should therefore be paid to what hinders substitution across firms within markets, such as search frictions or the value of formal job amenities, beyond the factors already accepted as important for mobility across markets. Finally, my findings also show that increased labor market power is not the leading mechanism behind trade-induced wage declines, at least in Brazil's context. Future research should focus on which components of the marginal revenue product of labor, such as output prices, productivity, or technology, account for them.

2 A model of labor market concentration and markdowns

In this section I introduce a model of Brazilian labor markets that provides the relationship between labor market concentration and wage markdowns. I write the model to mirror key features of the data and to deliver equations that I can bring to the data to estimate its key parameters. As in BHM, labor supply is nested CES, firms compete for workers à la Cournot, and there is a large number of labor markets.[5] Combined, these assumptions imply that the impact of trade liberalization on firm labor market power can be quantified by two key sufficient statistics only: the effect of trade liberalization on labor market concentration, and workers' cross-market vs. within-market cross-firm inverse elasticities of substitution. In the following sections I then estimate these sufficient statistics leveraging employer-employee linked data and Brazil's 1990s trade liberalization.

[4] Since the first release of this paper, Amodio et al. (2025) have estimated wage markdowns in 82 low- and middle-income countries, finding a median take-home share of 43 cents and markdowns that increase in firm size, as I do; Amodio, Medina and Morlacco (2025) find that self-employment curbs oligopsony power in Peru; and Sharma (2024) detects collusion among Indian textile firms.

[5] My model diverges from BHM's on two fronts: (i) I allow for arbitrary production functions and product market structure; and (ii) I allow wages to depend on firm-market-specific distaste shifters, which may reflect amenities. These divergences are possible because I estimate wage markdowns (set on the margin) rather than labor shares (which include infra-marginal revenues), and because my setting allows estimating elasticities via IV rather than indirect inference, which would require restrictions on (i) (see footnote 31). I also focus on relative effects of trade rather than general equilibrium counterfactuals. See Appendix B.2 for how this paper's results map into BHM's. The labor supply structure builds on the nested preference structures of international trade (Goldberg, 1995; Atkeson and Burstein, 2008) and on random utility models of worker preferences (Mas and Pallais, 2017; Azar, Berry and Marinescu, 2022).

2.1 Labor supply: Discrete choice

I follow a similar setup to BHM's micro-foundation of a nested CES labor supply system, which I extend to incorporate worker taste shifters for specific markets and for firm-market pairs. These taste shifters give structural interpretation to the regression residuals in the empirical specifications I use to estimate the model's key elasticities of substitution.

The economy consists of a continuum of homogeneous workers j, a large but finite number of local labor markets m, and a finite number of firms z within each local labor market. Each worker chooses to which firm-market pair zm they provide l^j_{zm} units of labor subject to making reservation earnings $y^j \sim F(y)$, solving the following discrete choice problem to minimize the disutility of work V_{zm}:

$$\min_{zm} V^j_{zm} = \ln l^j_{zm} + \ln \xi_m + \ln \xi_{zm} - \xi^j_{zm}$$
$$\text{s.t. } l^j_{zm} w_{zm} \geq y^j$$

where $\xi_{zm} > 0$ and $\xi_m > 0$ are firm-market- and market-specific taste shifters common to all workers, w_{zm} is the wage paid by firm z in local labor market m to identical workers, and ξ^j_{zm} is an idiosyncratic worker taste shifter with a General Extreme Value (GEV) distribution:[6]

$$G\left(\left\{\xi^j_{zm}\right\}\right) = \exp\left[-\sum_m \left(\sum_{z\in\Theta_m} e^{-(1+\eta)\xi^j_{zm}}\right)^{\frac{1+\theta}{1+\eta}}\right] \tag{1}$$

where Θ_m is the set of firms operating in market m.

The parameters $\theta > 0$ and $\eta > 0$ correspond to workers' cross-market and within-market cross-firm elasticities of substitution,[7] whose nesting structure is shown in Figure 1 from the point of view of worker j's decision. These are the two key elasticities of substitution whose estimates drive this paper's empirical findings.

Since ξ^j_{zm} follows a GEV distribution, the probability that worker j chooses firm z in market m can be written as a function of wages, taste-shifters, and the elasticities of substitution (see Appendix B for detailed derivations of all results in this section). Aggregating these probabilities to

[6] Equation 1 corresponds to the Gumbel distribution, a member of the GEV family. Per McFadden (1978), similar equations to those in this section can be derived for any member of the GEV family.

[7] BHM show that the nested discrete choice setup can be mapped into a representative worker problem where the representative worker has nested CES preferences over firms and markets, with θ wage elasticity of substitution across markets, and η wage elasticity of substitution within markets across firms.

the firm-market level gives the model's equation for residual labor supplied to firm z in market m:

$$l_{zm} = L \left(\frac{w_{zm}}{W_m} \right)^{\eta} \left(\frac{W_m}{W} \right)^{\theta} \left(\xi_{zm}^{1+\eta} \xi_m^{1+\theta} \right)^{-1} \tag{2}$$

where W_m, W, and L are CES wage and labor supply indices (i.e., "taste-adjusted" wages and employment indices), whose expressions can be found in Appendix B.1.

Equation 2 encapsulates the following intuition. The residual labor supplied to firm z in market m is increasing in how attractive its wage w_{zm} is relative to market m's wage level W_m, as well as in how attractive market m's wage level is relative to all other markets. It is also decreasing in the (dis)taste shifters ξ_{zm} and ξ_m, and larger if there is overall more (taste-adjusted) labor L supplied to all markets.

Finally, inverting equation 2 gives the model's equation for the wage w_{zm} firm z must pay in market m to obtain l_{zm} units of labor:

$$w_{zm} = W \left(\frac{l_{zm}}{L_m} \right)^{\frac{1}{\eta}} \left(\frac{L_m}{L} \right)^{\frac{1}{\theta}} \xi_{zm}^{1+\frac{1}{\eta}} \xi_m^{1+\frac{1}{\theta}} \tag{3}$$

where L_m is market m's taste-adjusted labor supply index, whose expression is in Appendix B.

Equation 3 encapsulates an intuition similar to that of equation 2, its counterpart. The wage w_{zm} needed to attract l_{zm} units of labor is increasing in the (dis)taste shifters ξ_{zm} and ξ_m—indicating workers must be compensated to move to a firm or market they dislike— as well as in the country-level wage index W. Sometimes referred to as the firm's wage equation, equation 3 is the firm's inverse residual labor supply, and it is the key equation underlying my empirical strategy to estimate $\frac{1}{\eta}$ and $\frac{1}{\theta}$, which I present in Section 5.

2.2 Labor demand: Cournot competition

Labor markets are imperfectly competitive. Firms compete à la Cournot, choosing their labor demand in each market to maximize their profits while taking as given the labor demand of other firms. Firm profits are:

$$\Pi_z = R_z \left(\{ l_{zm}, l_{-zm} \}, X \right) - \sum_m w_{zm} \left(\{ l_{zm}, l_{-zm} \} \right) l_{zm} \tag{4}$$

where R_z is the firm's revenue function—capturing both production function and goods market structure, on which I remain agnostic—and w_{zm} is the wage that firm z would need to pay to obtain l_{zm} units of labor in local labor market m. The expression $\{l_{zm}, l_{-zm}\}$ in curly braces denotes that, from firm z's perspective, both R_z and w_{zm} depend on the full profile of labor demanded by all firms

in all markets,[8] while X represents any exogenous shock to firm z's revenues.

To maximize profits, firm z looks at all local labor markets and considers, for each one, the effect that increasing employment in that market would have on its total revenues—holding labor demand at all other markets constant—and contrasts that marginal revenue gain to the marginal cost of this decision. This optimal tradeoff yields firm z's profit-maximizing wage setting formula in market m:

$$\underbrace{\partial R_z / \partial l_{zm}}_{\text{Marginal revenue}} = \underbrace{w_{zm} \times \overbrace{\left(1 + \varepsilon_{zm}^{-1}\right)}^{\text{Markdown}}}_{\text{Marginal cost}} \tag{5}$$

where $1 + \varepsilon_{zm}^{-1} \equiv \mu_{zm}$ is firm z's markdown in market m, which is a function of $\varepsilon_{zm}^{-1} \equiv \frac{\partial \ln w_{zm}}{\partial \ln l_{zm}}$, the inverse elasticity of residual labor supply the firm faces in that market.

The markdown μ_{zm} is a number, ranging from one to infinity, that equals the ratio of a firm's marginal revenue product to the wage. Therefore, the wage take-home share—the share of workers' marginal revenue product paid in wages—is simply the markdown inverse, $\mu_{zm}^{-1} = \left(1 + \varepsilon_{zm}^{-1}\right)^{-1}$, a number between zero and one. The question is: does the assumption of nested CES labor supply from Section 2.1 imply anything about ε_{zm}^{-1}?

It does. When worker preferences are nested CES as in Section 2.1, it is a standard result that differentiating equation 3 with respect to l_{zm} gives the following expression for ε_{zm}^{-1} that is solely a function of firm z's payroll share in market m and workers' key elasticities of substitution:

$$\varepsilon_{zm}^{-1} = \frac{1}{\theta} s_{zm} + \frac{1}{\eta}\left(1 - s_{zm}\right) \tag{6}$$

where

$$s_{zm} \equiv \frac{w_{zm} l_{zm}}{\sum_j w_{jm} l_{jm}} = \frac{\partial \ln L_m}{\partial \ln l_{zm}} \tag{7}$$

is firm z's payroll share in market m. This means that the markdown of firm z in market m can be written as

$$\mu_{zm} = 1 + \varepsilon_{zm}^{-1} = 1 + \frac{1}{\theta} s_{zm} + \frac{1}{\eta}\left(1 - s_{zm}\right) \tag{8}$$

Equation 7 is the key standard result that makes equation 6 hold. It states that a firm's marginal ef-

[8] l_{-zm} denotes labor employed by all other firms or in all other markets. The wage w_{zm} depends on all these components via L_m and L in equation 3. Similarly, R_z depends on them through firm z's production function (e.g., how it combines labor across markets) and through output equilibrium prices (e.g., competitors' production affects goods market structure).

fect on its market's taste-adjusted labor supply index L_m when hiring a marginal worker equals its payroll share.

As in BHM, a useful feature of equation 8 is that it encompasses perfect competition and monopsonistic competition as limiting cases. If $\frac{1}{\eta} = \frac{1}{\theta} = 0$, workers move instantaneously across firms anywhere in response to shocks. This is the perfect competition limiting case, and it implies that $\mu_{zm} = 1$: the full marginal revenue product of labor is paid in wages.[9] When $\frac{1}{\eta} = \frac{1}{\theta} > 0$, workers substitute across labor markets as strongly as they substitute across firms within markets, such that firms compete in a unified national labor market. This is the monopsonistic competition limiting case, where μ_{zm} is constant, and firm labor market power is therefore independent of firm size.

Finally, it is important to highlight that, in this model, the set of markets in which a firm operates is endogenous. This is a consequence of remaining agnostic about the firm's revenue function R_z, and specifically of not restricting it to be market-specific. Instead, the set of local labor markets in which a firm operates can be interpreted as part of its production function. This flexibility allows for wage markdowns based on this model to be consistent not only with the existence of multi-establishment firms, but also consistent with optimal firm behavior in granular labor markets.[10] For example, in the face of a negative shock, a firm might find it optimal to change its occupation mix or close a specific establishment. The effect of such restructuring would be captured both in the marginal revenue product of labor (via changes in productivity) and in wage markdowns (via relative sizes). At the local labor market level, Corollary 1 below isolates the effect on markdowns.

2.3 Labor market concentration and the average wage markdown

Aggregating the right-hand side of equation 8 across all firms in a local labor market, using payroll shares as weights, gives the key relationship between a market's average wage markdown and its concentration level:

Proposition 1. *When labor supply is nested CES, and firms compete for workers à la Cournot, as in the environment described in Sections 2.1–2.2, the average wage markdown at labor market m is given by:*

$$\mu_m \equiv \frac{\bar{r}_m}{\bar{w}_m} = 1 + \varepsilon_m^{-1} = 1 + \frac{1}{\theta} HHI_m + \frac{1}{\eta}(1 - HHI_m) \tag{9}$$

where $\bar{r}_m$ and $\bar{w}_m$ are market m's (employment-weighted) average marginal revenue product of

[9] Trade's negative effects on local wages could also be rationalized under perfect competition if workers cannot easily move across markets (Dix-Carneiro and Kovak, 2017). My paper instead considers the possibility, suggested by Manning (2003), that imperfect worker mobility is itself an outcome of firms exploiting workers' heterogeneous preferences to mark wages down.

[10] Around 3% of firms are multi-establishment (operating in different microregions) in this context. Appendix Figure A.1 shows nearly nonexistent within-firm cross-region movement in the data.

labor and average wage, respectively, ε_m^{-1} is the (payroll-weighted) average inverse elasticity of firm-specific residual labor supply across firms in market m, and $HHI_m = \sum_{z \in \Theta_m} s_{zm}^2$ is the market's payroll Herfindahl.

Proof. See Appendix B.2.3. □

In other words, a market's average wage markdown is linear in its level of concentration: it equals one plus a weighted average of workers' key inverse elasticities of substitution, whose weights are given by concentration. Because it is generally assumed (although not imposed later during estimation) that workers substitute more easily across firms within markets than across markets (i.e., $\frac{1}{\theta} \geq \frac{1}{\eta}$), equation 9 implies that the higher the level of concentration in a market, the larger is its average wage markdown. In addition, the larger are the *inverse* elasticities of substitution, the weaker is worker movement in response to wage shocks, and thus the larger is the wage markdown.[11]

This paper's key theoretical result, used later in Section 7 to quantify the causal effect of trade liberalization on local labor markets' average wage markdown, is a direct implication of Proposition 1:

Corollary 1. *In the labor market environment described in Proposition 1, the effect of an exogenous shock X on market m's average wage markdown μ_m at time t is given by:*

$$\gamma_t \equiv \frac{d\mu_{mt}}{dX} = \left(\frac{1}{\theta} - \frac{1}{\eta}\right)\beta_t \tag{10}$$

where $\beta_t \equiv \frac{dHHI_{mt}}{dX}$ is the effect of the exogenous shock on market m's payroll Herfindahl at time t, $\frac{1}{\theta}$ is workers' cross-market inverse elasticity of substitution, and $\frac{1}{\eta}$ is workers' within-market cross-firm inverse elasticity of substitution.

Proof. Differentiate equation 9 with respect to X. See Appendix B.2.5. □

To see the intuition behind Corollary 1, suppose that the exogenous shock X is trade liberalization, whose policy-induced shock variation I introduce later in Section 4. Then, two things must hold in order for trade liberalization to increase market m's average wage markdown, and thereby reduce wages in market m via firm labor market power.

First, trade must increase labor market concentration (i.e., $\beta_t > 0$). The reason is simple: labor market concentration is the only endogenous component of a market's average wage markdown.

[11] Proposition 1 refers to the average *wage markdown* at a *local labor market*, not the *country-level labor share* common in the labor and macro literatures. The wage markdown concerns wage-setting at the margin, whereas the labor share concerns payments to and revenues generated by *all* workers (including infra-marginal), requiring additional assumptions on production functions and goods market structure. Under such assumptions (e.g., Cobb-Douglas production and perfectly competitive goods and capital markets, as in BHM), a country's average markdown is closely related to its aggregate labor share. Appendix B.2.4 shows the country-level wage take-home share from equation 9 is mathematically equivalent to a sub-component of BHM's country-level labor share. Appendix B.2.6 shows that Proposition 1 still holds under these additional assumptions.

The other two components are simply labor supply parameters, which by assumption do not change. Intuitively, the source of market power in the labor market environment described in Section 2 is worker preference heterogeneity for markets and firms. Firms can "exploit" this preference heterogeneity to mark wages down. The bigger a firm is relative to its competitors, the more it can mark wages down without workers easily leaving because there are fewer employment options nearby, and workers tend to prefer switching locally across firms before switching markets completely. Thus, the degree of market power in a local labor market can only meaningfully change if the relative sizes of its firms meaningfully change, captured by changes in labor market concentration.

Second, there must be a gap between workers' key inverse elasticities of substitution (i.e., $\frac{1}{\theta} - \frac{1}{\eta} > 0$). If there is no gap, then workers move far away as easily as they move close by in response to shocks, such that to attract workers firms must compete in a unified country-level labor market, where their wage setting ability is independent of size. In this scenario, the effect of trade on labor market concentration would be irrelevant for changes in firm labor market power. Such is the case under my model's two limiting cases: monopsonistic competition (i.e., no gap to induce *effects* on market power, but because $\frac{1}{\theta} = \frac{1}{\eta} > 0$, there is still some *level* of market power); and perfect competition (i.e., no gap to induce effects, and because $\frac{1}{\theta} = \frac{1}{\eta} = 0$, no level of market power either).

3 Data and setting

This section describes the data and setting I leverage to estimate the sufficient statistics in equation 10. The analysis combines employer-employee linked administrative records (RAIS), product-level import tariffs, firm-level exporting activity, and Brazil's 1991 and 2000 population censuses. Appendix A describes all datasets and mapping procedures.

3.1 Data

First, rich labor market data come from Brazil's administrative employer-employee linked database Relações Anuais de Informações Sociais (RAIS), spanning years 1986–2000.[12] RAIS covers the universe of Brazilian formal sector workers. I focus on the sample of private sector workers aged 18 to 64. Second, data on tariffs come from UNCTAD TRAINS, downloaded from WITS, which I map to RAIS via the 5-digit economic activity code CNAE95, using product-to-sector concordances from IBGE. This gives tariff changes for 285 sectors, against about 20 industries in previous studies of Brazil's liberalization. The finer tariff shocks are what allow me to compare firms facing different tariff cuts within the same local labor market, and to establish that my main estimates are identified

[12] I end the analysis in 2000 for two reasons. Brazil's occupation classification changed in 2002, which alters the definition of local labor markets; and the set of firms present in both 1991 and the end year shrinks with each additional year, which weakens the identification of the within-market cross-firm elasticity from the 1990–1994 tariff cuts.

by differences in tariff cuts across sectors (Section 4.2). Third, exporting activity is mapped to RAIS using firms' unique identifier CNPJ. What I observe in terms of exporting activity is the list of exporting firms for years 1990–1994, which was provided upon request by the (extinct as of 2019) Ministry of Development, Industry, and Foreign Trade (MDIC), currently a part of the Ministry of the Economy. Finally, I use data from the 1991 and 2000 Brazilian census when discussing model extensions and heterogeneity by market characteristics in Sections 8.1 and 8, downloaded from the replication data for Dix-Carneiro and Kovak 2017.

3.2 Brazil's formal sector labor market structure

Three facts about Brazil's labor market guide how I model it. The first two motivate the baseline model of Section 2, in which formal firms compete for workers within local labor markets; the third explains why I can estimate the baseline on formal firms alone, and what the extension of Section 8 adds by incorporating informal wage work and self-employment.

First, formal labor markets have geographic and occupational boundaries. Conditional on switching jobs, most Brazilian workers remain within the same microregion, and switches within microregion × occupational group cells are the most stable share of firm-to-firm transitions throughout the entire period (Appendix Figures A.1 and A.2). In contrast, most firm switchers change sectors (Appendix Figure A.3). Based on these patterns, I define a formal sector local labor market as a microregion × occupational group cell,[13] and present effects with local labor markets defined by microregions only as robustness.

Second, these markets are concentrated, and wages are lower where they are more concentrated. Appendix Table A.1 presents summary statistics of the roughly 20,000 local labor markets in Brazil. In the baseline year of 1991, the unweighted average payroll Herfindahl across local labor markets was 0.29, and the median was 0.21. Many local labor markets are thus highly concentrated, but because most workers work in larger labor markets, the payroll-share-weighted average concentration is much smaller: 0.08 on a scale from zero (infinitely tiny firms) to one (one firm), equivalent to an average worker being in a market whose equilibrium is pinned down as if only $12.5 = 1/0.08$ equally sized firms operated it. Between 1991 and 2000, the unweighted average fell from 0.29 to 0.21 and the employment-weighted average from 0.08 to 0.06. Formal sector wages and employment decline with formal sector employment concentration (Appendix Figure A.4), consistent with labor market oligopsony.

Third, while the formal sector accounted for only 42% of total employment in 1991, most formal

[13]The literature on labor market power typically considers granular market boundaries, such as region × occupation (e.g., Azar, Marinescu and Steinbaum (2022); Azar et al. (2020); Schubert, Stansbury and Taska (2024)) or region × sector (e.g., BHM, Lamadon, Mogstad and Setzler (2022), and Alfaro-Ureña, Manelici and Vasquez (2021)), with few studies using region-only boundaries (e.g., Pham (2023)). Section 6 shows that finer boundaries yield similar estimates of the within-market cross-firm elasticity of substitution.

workers were nonetheless in labor markets that are primarily formal: the most formal quartile of microregions hosts 73% of formal employment (Figure 2 and Section 8.1). The rest of the workforce was split between informal wage work (34%), without a formal contract that guarantees formal sector benefits like unemployment insurance, and self-employment (25%); by 2000, these shares were 40% and 24%, and the formal share had fallen to 36% (Appendix Table A.2). Each year most formal workers remain at their prior year's formal employer, and most new formal sector hires come directly from other formal sector firms rather than from outside the formal sector (Appendix Figure A.5 and Appendix Table A.3). In the baseline model I therefore study formal firms' wage markdowns using data on formal firms alone. In Section 8, I add workers' options outside formal employment, which matter most where the formal sector is small.[14]

3.3 Setting: Brazil's 1990s trade liberalization

The key policy-induced variation I leverage throughout my analyses comes from Brazil's 1990s unilateral import tariff reductions. Dix-Carneiro and Kovak (2017) provide an in-depth discussion of Brazil's 1990s import tariff reform. Tariffs were reduced from a pre-liberalization average of 33% to a post-reform average of 13%,[15] with some sectors experiencing larger reductions than others because they were previously more protected, as shown in Figure 3.

These tariff reductions generated plausibly exogenous variation in labor demand shocks across firms and across markets, which I exploit to estimate the key sufficient statistics in equation 10. Kovak (2013) argues that the close correlation between pre-liberalization tariff levels and reform-induced tariff cuts, as documented in Figure 3, is the main support for exogeneity of the tariff cuts. The key argument is that, because the pre-liberalization levels of protection were set decades earlier (Kume, Piani and Souza, 2003), it is unlikely that the 1990s tariff cuts were correlated with counterfactual sector performance at the time. Instead, the reductions were motivated by the broader national goal to reduce all tariffs towards a much lower and much more equalized level of protection across all sectors. Agricultural tariffs changed little, and the tariff reform coincided with the removal of non-tariff barriers in 1990 and with the Real Plan stabilization of 1994; these features of the reform matter for comparisons between more and less tradable markets (Section 4.2).

The main identification concern posed by using Brazil's import tariff reductions as exogenous shocks is pre-trends. Despite the plausible exogeneity in tariff cuts, one might be concerned that the decades-long level of protection enjoyed by the sectors experiencing the largest tariff cuts might induce differential trends in sector outcomes. For example, if the most protected sectors were also

[14] Given measurement limitations with occupation codes in Census data, Section 8's analysis incorporating informal wage work and self-employment considers informal sector conditions more broadly at the microregion level, but allows them to vary by age, education, and gender.

[15] Simple 1990 averages of nominal tariffs at the CNAE95 level. See Appendix A.

the least productive ones, one might observe negative pre-trends in either payroll or employment, which could confound the negative estimates of the effect of trade liberalization on these outcomes. Reassuringly, Appendix Figure A.6 shows no correlation between sector-level import tariff cuts and sector-level changes in either employment or payroll in the years preceding the tariff cuts (1986–1990). This is different from the pattern observed during liberalization (1990–1994), when employment and payroll shrink more in the sectors with the largest tariff cuts. In the analyses that follow, I further check for pre-trends at the local labor market level by estimating year-specific regression coefficients for all outcomes of interest.

4 Effect of trade liberalization on concentration

My first step towards quantifying the effect of trade liberalization on firm labor market power is to estimate parameter β_t from equation 10. Specifically, I leverage the market-level exogenous labor demand shocks spurred by Brazil's trade liberalization to estimate β_t as the effect of trade liberalization on local labor markets' payroll Herfindahl indices.

4.1 Empirical strategy

My identification strategy for estimating the effect of trade liberalization on local labor market concentration follows the shift-share treatment intensity approach adopted by other papers on the regional incidence of trade (e.g., Kovak (2013); Dix-Carneiro and Kovak (2017)). The key idea is that the reduction in import tariffs spurred by Brazil's 1990s liberalization would have a differential effect across local labor markets depending on these markets' pre-liberalization sectoral composition. The precise functional form linking sector-level tariff reductions to market-level shocks is guided by the model I outlined in Section 2. Specifically, I define local labor market m's Import Competition Exposure (ICE) shock as

$$\Delta ICE_m \equiv -\sum_{z \in \Theta_m^T} \kappa_{zm} \ln\left(\frac{1+\tau_{i(z),1994}}{1+\tau_{i(z),1990}}\right) \tag{11}$$

$$\kappa_{zm} \equiv \frac{s_{zm,1991}^2}{\sum_{j \in \Theta_m^T} s_{jm,1991}^2}, \quad s_{zm,1991} \equiv \frac{w_{zm,1991} l_{zm,1991}}{\sum_j \left(w_{jm,1991} l_{jm,1991}\right)}$$

where Θ_m^T is the set of all tradable sector firms in market m in the baseline year of 1991,[16] $s_{zm,1991}$ is the 1991 payroll share of each of these firms as a fraction of all firms operating in the market, and $\tau_{i(z),t}$ is the import tariff faced by firm z's output sector in year t.

[16] Year-end wages and employment for 1990 might reflect the removal of non-tariff barriers in 1990; following Dix-Carneiro and Kovak (2017), I use 1991 as the base year for all analyses.

In other words, ΔICE_m is a weighted average of the firm-level shocks experienced by tradable sector firms,[17] where the weight κ_{zm} of each firm z is its contribution to the tradable sector's component of market m's pre-liberalization payroll Herfindahl, $HHI^T_{m,1991} \equiv \sum_{j \in \Theta^T_m} s^2_{jm,1991}$.[18] The functional form for κ is guided by equation 7, according to which the effect of a firm hiring a marginal worker on its market's labor supply index is precisely the firm's payroll share. This suggests that firm-level labor demand shocks should be aggregated to the market level in proportion to firms' baseline payroll shares. Finally, to further align a firm's weight with its contribution to the market's payroll Herfindahl, which determines the market's average markdown (Proposition 1), I construct κ_{zm} by placing firm z's squared baseline payroll share in the numerator, and dividing through by the tradable sector's component of market m's baseline Herfindahl. Appendix C derives the model's measure of a market's exposure, an average of the tariff reductions faced by its tradable firms with weights that depend only on the firms' baseline payroll shares, and shows the conditions under which ΔICE_m coincides with it; otherwise ΔICE_m approximates it. I report results with linear payroll shares in Appendix Table A.4, column (2). I also present robustness checks to alternative measures of tariff shocks.

Figure 4 displays the variation in ΔICE_m across geography for two example occupations, while Appendix Table A.1 provides the mean and key percentiles of the distribution of ΔICE_m across local labor markets. The mean change in import competition exposure was 13%, ranging from a 10th percentile of no exposure change (i.e., a local market made primarily of non-tradable sector firms) to a 90th percentile of 23% increase.

Having defined the import competition exposure shock, I proceed to estimate its effect on local labor market outcomes using a difference-in-differences strategy. Specifically, I estimate the cumulative effect (as of year k) of import competition on a local labor market's outcome Y_m as ζ_k from the following regression:

$$\Delta Y_{mt} = \sum_{k \neq 1991} \zeta_k \left(\Delta ICE_m \times 1_{t=k} \right) + \delta_m + \delta_t + \epsilon_{mt} \tag{12}$$

where ΔY_{mt} denotes the long difference in Y_m from year t back to the base year 1991,[19] and δ_m and δ_t are local labor market and year fixed effects. As the specification is in stacked differences,

[17] In 1991, 2,990 of the 19,757 microregion × occupation markets (15%) have no tradable sector firm with a tariff match. I set ΔICE_m to zero in those markets, which serve as controls in market-level regressions. When I define markets by microregion only, 6 of the 475 microregions have no such firm, and I likewise set ΔICE_m to zero there.

[18] By construction, the κ_{zm} weights sum to one in markets with tradable firms; in markets without them, $\Delta ICE_m = 0$, so that in the terminology of Borusyak, Hull and Jaravel (2022) the shares are incomplete, with a share sum equal to an indicator for having a tradable firm (see Section 4.2 and Appendix C.5).

[19] Following Dix-Carneiro and Kovak (2017), every long difference runs forward in time: for $t > 1991$, $\Delta Y_{mt} = Y_{mt} - Y_{m,1991}$, and for $t < 1991$, $\Delta Y_{mt} = Y_{m,1991} - Y_{mt}$. A positive pre-period coefficient thus means that the outcome grew faster in more exposed markets between year t and 1991.

note that the fixed effects absorb not only the constant, but also market-level secular trends over the entire period. I estimate this regression using years 1986 to 2000, clustering standard errors by local labor market.[20] In my main estimates I weight local labor markets by their 1991 formal employment (Section 4.2).

4.2 Estimates of effect of trade liberalization on concentration

Figure 5 and Table 1 present my main estimates of the effect of trade liberalization on local labor market concentration. Column (1) shows the main specification: a 10 percent increase in import competition exposure increased local labor markets' premium-based payroll Herfindahls, computed with each firm's payroll valued at its wage premium, by 0.017 points between 1991 and 2000. This effect corresponds to a 22% increase relative to the pre-liberalization average of 0.08 across markets weighted by their formal employment. Concentration starts rising in 1992, while tariffs were still being cut, and reaches three quarters of its 2000 effect by 1995 (0.129 vs. 0.172). The effect is also present at every level of local informality (Section 8.1).

The effect is robust to the use of wage levels (as opposed to wage premia) to compute payroll Herfindahls, and to measuring concentration using the employment (instead of payroll) Herfindahl (22% and 19% of the respective 1991 means; second and third rows of Table 1). It is also robust to constructing ΔICE_m from effective rates of protection, which net out tariff reductions on firms' inputs (21% of the 1991 mean per 10%, vs. 22%; Appendix E).[21] Finally, the effect is also present when I define labor markets more broadly by microregions only: weighting microregions by their 1991 formal employment, as in the extended model of Section 8, I estimate an effect of 0.005 points per 10% increase in exposure (Table 7, column (2)).[22]

Why do I weight markets by their 1991 formal employment? ΔICE_m is a shift-share measure: its shifts are the sectoral tariff cuts, and its shares are the payroll shares of the market's tradable firms in those sectors. In the 85% of markets with at least one tradable firm the shares sum to one, so that ΔICE_m varies across them only with the shifts and with how the market's payroll is distributed among its firms. In the remaining 15% of markets, which had no tradable firm in 1991, ΔICE_m is zero (footnote 17). Weighting by 1991 formal employment places most of the weight on large markets with many tradable firms, so that my estimates are identified by the shifts—the quasi-experimental variation in this setting. Without weights, the many small markets without tradable

[20] A positive ζ_k indicates that the import tariff reductions had a positive effect on the outcome.

[21] Effective rates of protection are output tariffs net of input tariffs, constructed using Brazil's 1985 input-output table at a broader sector level (Nível 50) than output tariffs (CNAE95, 285 sectors); see Appendix A. Appendix E reports every market-level outcome and the two elasticities with this shock.

[22] The effect is 0.421 when I construct ΔICE_m with linear payroll shares, vs. 0.172, and 0.174 when I drop the largest 1% of markets. The statistical significance of the effect on concentration is also robust to two-way clustering by occupation and region. See Appendix Tables A.4 and A.6.

firms count as much as the large ones, and the estimates are instead identified mostly by comparing markets with and without tradable firms. I report those unweighted estimates in Appendix D.[23]

I confirm that my main estimate of the effect on concentration is identified by the shifts in two ways. First, the estimate is 0.168 when I recenter ΔICE_m on its expected value under random reassignment of tariff reductions within IBGE subsectors, the check for non-random exposure proposed by Borusyak and Hull (2023) (Table 1, column (2)). Second, placebo coefficients from that reassignment exceed my actual estimate in only 1.6% of draws (Appendix Figure C.2, panel (a)).[24] Interpreting the effect as causal thus requires that tariff reductions were as good as randomly assigned across sectors (Section 3.3). I find little evidence of pre-trends (Figure 5), and the estimate is 0.152 when I control for state × year fixed effects and each market's 1986–1990 trend, as in Dix-Carneiro and Kovak (2017) (Table 1, column (3)).

I also estimate equation 12 for the number of firms, employment, and wage premia of local labor markets, presented in Panels B and C of Table 1. A 10% increase in import competition exposure reduced the number of firms in a market by 9.7% by 2000, and reduced its formal employment by 4.0%. Like the effect on concentration, the decline in the number of firms is identified by the shifts (−1.09 with the recentered measure vs. −0.97), and is robust to controlling for state × year fixed effects and 1986–1990 trends (−0.68, column (3)). The employment effect is less robust: it is smaller and imprecise both when I recenter ΔICE_m (−0.30, SE of 0.21) and when I control for state × year fixed effects and 1986–1990 trends (−0.19, SE of 0.15), and it is not negative when I drop the largest 1% of markets (0.19, SE of 0.10; Appendix Table A.4).[25]

My estimates of the effect on market wages are negative but imprecise. Through 2000, a 10% increase in exposure reduced the market's average firm wage premium—the employment-weighted mean of its firms' wage premia—by 1.2% (SE of 1.3), the market's wage premium conditional on

[23] Dix-Carneiro and Kovak (2017)'s exposure measure, RTR (regional tariff reductions), varies across regions mainly with the share of agriculture, a sector with a small tariff cut, whereas ΔICE_m is unrelated to that share (Appendix Figure C.1). Their employment and earnings effects thus rest on a contrast in regional composition, as my unweighted estimates do, which is why controlling for pre-trends is central in their paper and why my estimation of $1/\theta$ includes their state fixed effects and 1986–1990 trends (Section 5). Despite the different sources of variation, the effects on employment and wages that I obtain as of 2000 are qualitatively similar to theirs, the main difference being that my wage effect is somewhat smaller. Research for which it matters that effects be identified by differences in tariff cuts across sectors, rather than by baseline shares, should use ΔICE_m as the exposure measure and weight regressions by baseline employment, as I do in Table 1. See Appendix C.5 for a detailed discussion, and Appendix D for how the weights affect each estimate.

[24] The effect is 0.162 when I drop markets without tradable firms (Appendix Table C.1). Appendix Table D.1 shows that, within quintiles of 1991 market size, the weighted and unweighted estimates coincide and the effect is positive in every quintile.

[25] Appendix Table A.5 reports every market-level effect with state × year fixed effects, with these and 1986–1990 trends, and with microregion × year fixed effects. The effects on concentration and on the number of firms remain statistically significant in every specification; the effect on employment does not once I include state × year fixed effects. The reform coincided with the removal of import licenses and with the Real Plan, which fell on the same contrast between tradable and other sectors, so for the effects on employment and wages the parallel-trends assumption carries more weight than for the effect on concentration.

worker observables by 1.1% (SE of 0.9), and the average of its workers' log wages by 0.8% (SE of 1.1) (Table 1, Panel C).[26] Within markets, I find that wage premia fall at firms facing larger tariff cuts (Section 6), and that the wage premium index of the market's incumbent firms falls as well (−0.183, with state fixed effects and 1986–1990 trends; Section 6.2). I compare the wage decline at incumbent firms with the market average in Section 7.

Finally, I address an over-rejection concern uncovered by recent literature on shift-share instruments, which arises due to spatial correlation in sectoral composition across markets (Adao, Kolesár and Morales, 2019). I address this by increasing the number of industries (the sectors of firms' main activity, at which tariffs are measured) used to construct ΔICE_m by an order of magnitude, to 285, relative to those currently used in the literature—[27] adding further granularity in tariff shocks that mitigates the spatial correlation—and by reporting standard errors that account for this correlation, computed following the procedure described in Adao, Kolesár and Morales (2019). Column (3) of Appendix Table A.6 shows that, for the effect on concentration, these standard errors are similar to those clustered by local labor market (0.045 vs. 0.042), so that the effect remains statistically significant.

4.3 Source of increased concentration

What drives the increase in concentration shown in Figure 5? Concentration can increase because firms exit, because payroll shifts among the firms that remain, or because fewer firms enter. Table 1 shows that there were fewer firms. To separate these three margins, I decompose each market's change in the payroll Herfindahl since 1991 exactly, firm by firm, into three parts: the part due to firms present in 1991 but absent in year t (exit), the part due to the changing shares of firms present in both years, and the part due to firms present in t only (entry). I then estimate my main specification on each part (Appendix Table A.7).

I find that concentration increased because fewer firms entered and because payroll concentrated among the largest firms. Of the 0.018 point increase in the payroll Herfindahl per 10% increase in exposure,[28] 0.013 points are due to the shares of firms present in both years, which shifted towards the larger among them, and 0.012 points are due to entry: entrants captured 2.5 percentage points less of the market's 2000 payroll.[29] To see which firms gained, I classify firms into three

[26] With state × year fixed effects and 1986–1990 trends, these estimates are 1.9% (SE of 1.2), 0.8% (SE of 0.8), and 1.5% (SE of 1.0). Without the largest 1% of markets, the point estimates are positive (Appendix Table A.4, column (3)).

[27] Kume, Piani and Souza (2003) report tariffs at the sector levels of the IBGE input-output tables, 50 activities (Nível 50) and 80 products (Nível 80). Previous papers aggregate them to about 20 industries consistent across data sources (21 in Kovak (2013), including agriculture and non-traded goods). See Appendix A.

[28] The decomposition uses the Herfindahl computed with wage levels (Table 1, second row), whose total effect is 0.018 points, against 0.017 for the premium-based Herfindahl.

[29] Exit instead reduced the Herfindahl by 0.007 points, because the firms that left more exposed markets were larger

groups: exporters (those exporting at any point between 1991 and 1994), non-exporting tradable firms, and non-tradable firms. The increase is driven by exporters: their contribution to the market's Herfindahl—the sum of their squared payroll shares—increased by 0.014 points, two thirds of it among exporters present in both years, whereas the contributions of non-exporting tradable firms and of non-tradable firms changed little (−0.001 and 0.004). Exporters were the largest and highest-paying firms: in 1991 they employed 110 workers on average, compared to 5 at non-exporting tradable firms and 3 at non-tradable firms, and paid average December earnings of R$679, compared to R$204 and R$191 (Appendix Figure A.7).

Exporters did not gain payroll share by expanding. Figure 6 presents the effect of trade liberalization on each group's share of the market's employment. After the tariff cuts, the share of non-exporting tradable firms falls in more exposed markets and the share of exporters rises through the mid-1990s, but by 2000 neither is statistically distinguishable from its 1991 level (−0.3 and −0.4 percentage points per 10% increase in exposure, with SEs of 0.7 and 0.5). Worker flows lead to the same conclusion: hirings into exporters fell from every source, and the workers shed by non-exporting tradable firms were hired neither by exporters nor by their other competitors in the same market (Appendix Figure A.8).[30] Payroll thus concentrated among exporters because other firms contracted and fewer firms entered, not because workers moved to exporters.

Overall, fewer firms, less entry, and payroll concentrating among the largest firms are consistent with the pro-competitive effects of trade predicted by Melitz and Ottaviano (2008), as opposed to the expansion of exporters into foreign markets predicted by Melitz (2003).

5 Key labor supply parameters: Empirical strategy

Section 4.2 showed that local labor markets more exposed to import competition experienced an increase in labor market concentration following Brazil's trade liberalization. How did this affect wage markdowns? Per Corollary 1, the answer to this question depends on the gap between workers' within-market cross-firm elasticity of substitution $\frac{1}{\eta}$, and their cross-market elasticity of substitution $\frac{1}{\theta}$. This section describes my empirical strategy for estimating these key parameters.

My model provides the regression specifications, and my setting the exogenous variation. I use within-market cross-firm variation in import tariff reductions to estimate $\frac{1}{\eta}$, and cross-market variation in import competition exposure to estimate $\frac{1}{\theta}$. This section shows how—even in the presence of strategic firm interactions—this empirical strategy can be used to estimate $\frac{1}{\eta}$ and $\frac{1}{\theta}$, so long as within-market cross-firm shock variation is available. The elasticities of substitution can then

relative to their market.

[30]The lower panels of Appendix Figure A.8 include flows from and to other markets and into and out of the formal sector, which show the same pattern, so that differential movement across the formal-sector boundary does not account for the increase in concentration.

be combined with data on firm shares to compute firm-specific inverse elasticities of labor supply ε_{zm}^{-1}.

5.1 Within-market cross-firm inverse elasticity of substitution

5.1.1 Regression specification

To derive the regression equation for estimating $\frac{1}{\eta}$, I start by taking logs of a time-specific version of the model's equation for a firm's inverse residual labor supply function (i.e., equation 3), which gives:

$$\ln w_{zmt} = \frac{1}{\eta}\ln l_{zmt} + \underbrace{\left(\frac{1}{\theta}-\frac{1}{\eta}\right)\ln L_{mt} - \frac{1}{\theta}\ln L_t + \ln W_t + \ln \xi_{mt}^{1+\theta}}_{\text{Market x Year FE}} + \ln \xi_{zmt}^{1+\eta} \tag{13}$$

which simplifies to:

$$\ln w_{zmt} = \frac{1}{\eta}\ln l_{zmt} + \delta_{mt} + \epsilon_{zmt} \tag{14}$$

where δ_{mt} are market $\times$ year fixed effects (which absorb the constant), and $\epsilon_{zmt} = \ln \xi_{zmt}^{1+\eta}$ is the regression residual, which has a structural interpretation as workers' (scaled) taste shifter ξ_{zmt} for firm z in market m at time t. Anticipating that my empirical strategy for estimating $\frac{1}{\eta}$ will leverage Brazil's trade liberalization, whose key cross-firm exogenous variation is the 1990–1994 long-difference in tariffs, I take long-differences of equation 14, which becomes:

$$\text{[Second Stage]} \quad \Delta \ln w_{zm} = \frac{1}{\eta}\Delta \ln l_{zm} + \Delta\delta_m + \Delta\epsilon_{zm} \tag{15}$$

where $\Delta\delta_m$ is a market fixed effect in the already differenced regression, and its role is to absorb all market-level *changes* that feed into changes in firm z's wage in market m, shown explicitly in equation 13. Equation 15 is the regression specification I use to estimate $\frac{1}{\eta}$.

The key threat to identification of $\frac{1}{\eta}$ is that changes in labor supplied to firm z in market m (i.e., $\Delta \ln l_{zm}$) might be correlated with changes in workers' labor supply taste for firm z in market m (i.e., $\Delta\epsilon_{zm}$). I address this concern by instrumenting $\Delta \ln l_{zm}$ with a labor demand shock: $\Delta \ln\left(1+\tau_{i(z)}\right)$, the policy-induced change in import tariffs on firm z's output sector, using the following first stage regression:

$$\text{[First Stage]} \quad \Delta \ln l_{zm} = \lambda \Delta \ln\left(1+\tau_{i(z)}\right) + \Delta d_m + \Delta\nu_{zm} \tag{16}$$

where once again Δd_m is a market fixed effect. To see how $\frac{1}{\eta}$ is identified even in the presence of strategic interactions, consider what it means to instrument $\Delta \ln l_{zm}$ in equation 15 conditional

on a market fixed effect. By the Frisch–Waugh–Lovell Theorem, this is equivalent to shocking the partialled-out equation $\Delta \ln \tilde{w}_{zm} = \frac{1}{\eta} \Delta \ln \tilde{l}_{zm} + \Delta \epsilon_{zm}$, where $\tilde{x}$ indicates the residual from regressing x on the market fixed effect $\Delta \delta_m$. Shocking this equation gives $\partial \Delta \ln \tilde{w}_{zm} = \frac{1}{\eta} \partial \Delta \ln \tilde{l}_{zm} + \partial \Delta \epsilon_{zm}$, where ∂ indicates the effect of the shock. But since, by shock independence, $\partial \Delta \epsilon_{zm} = 0$, we have that $\frac{1}{\eta} = \partial \Delta \ln \tilde{w}_{zm} / \partial \Delta \ln \tilde{l}_{zm}$. This means that $\frac{1}{\eta}$ is identified by the effects of a firm-level shock on own-wage and own-employment *holding constant* the effects of the shock on other firms' decisions, which—under the assumption of nested CES—are entirely captured by changes in the market-level CES wage and labor supply indices, absorbed by the market fixed effect. In other words, $\frac{1}{\eta}$ is identified precisely by the *partial equilibrium* effects of a firm-level shock on own-wage and own-employment. This is different from attempting to directly estimate ε_{zm}^{-1}—a general equilibrium object—with regression, which cannot be done for Cournot competition, as shown by BHM.[31]

Identification of $\frac{1}{\eta}$ using IV relies on three assumptions: a) the shock is independent of firm potential outcomes, whose validity relies on the policy-driven nature of the shock; b) there is a first stage (i.e., $\lambda \neq 0$); and c) exclusion is satisfied, meaning that—conditional on market-level changes—import tariff shocks only affect workers' labor supply decision by changing wages, as opposed to changing workers' distaste ξ_{zm} for working at the particular firm-market pair. The main threat to exclusion is that worker tastes might be a function of non-wage amenities that a) change in response to trade; and b) are marginal to workers' labor supply decision.[32] Since I cannot test the exclusion restriction, I assume that amenities did not change in response to trade in a way that was marginal to workers' labor supply decision. This is similar to the approach in Lamadon, Mogstad and Setzler (2022), and is more flexible than most papers estimating elasticities of labor supply with labor demand shocks (e.g., BHM, Dube et al. (2020), etc.).[33]

5.1.2 Measurement

Estimating equations 15 and 16 requires measuring three model objects: the total units of labor l_{zmt} supplied to firm z in market m at year t, the wage w_{zmt} paid by that firm-market pair, and

[31] Under Cournot, ε_{zm}^{-1} cannot be directly estimated via regression, which holds competitors' reactions constant (partial equilibrium), whereas ε_{zm}^{-1} includes full equilibrium responses (BHM). However, its key parameters $\frac{1}{\eta}$ and $\frac{1}{\theta}$ can be identified with IV given within-market cross-firm shock variation. My "bottom-up" approach—estimate $\frac{1}{\eta}$ and $\frac{1}{\theta}$ with IV, then compute ε_{zm}^{-1}—contrasts with BHM's "top-down" approach: estimate a reduced-form ε_{zm}^{-1}, simulate firms' strategic behavior until model and data shares converge, and recover η and θ. The "top-down" approach works without within-market cross-firm shock variation; the "bottom-up" approach applies whenever firm-level shocks are available, as in Pham (2023) and Zavala (2022).

[32] For a non-wage amenity to be marginal to workers' labor supply decision, workers must be willing to pay for it (i.e., accept lower wages), not merely prefer it—e.g., schedule flexibility (Bustelo et al., 2023) or dignity (Dube, Naidu and Reich, 2022). See Kessler, Low and Sullivan (2019) on eliciting willingness to pay.

[33] Orthogonality between firm-specific labor demand shocks and amenities is implicit in most of the modern monopsony literature, where amenities do not enter the wage equation (e.g., Manning (2003), Ashenfelter, Farber and Ransom (2010)).

the tariff shock to the firm. I measure l_{zmt} (the total units of labor at firm z in market m in year t) as the total number of workers employed at firm z in market m during the entire month of December of year t.[34] This is equivalent to assuming that each worker provides one "effective monthly unit" of labor, whereas the model allows l^j_{zm} to be more generally pinned down by worker j's exogenous reservation earnings y^j.[35]

I measure w_{zmt} as firm z's wage *premium* in market m for the month of December of year t: the compensation received for the labor provided in December, conditional on worker characteristics.[36] A premium estimated on all of a firm's workers moves with the composition of its workforce, which a trade shock may change through selective hiring and separations. I therefore also estimate premia on the firm's stayers, workers employed in the same firm, microregion, and occupation in both 1991 and 2000. Because the same workers enter both years, their time-invariant characteristics, observed or unobserved, drop out of the change in a firm's premium, as they would with worker fixed effects (Abowd, Kramarz and Margolis, 1999). This approach does not rely on workers who move across firms, whose limited mobility can bias firm fixed effects (Bonhomme et al., 2023). It is important here to clarify that my use of the term "wage premia" follows papers in the literature on the regional incidence of trade, and is meant to indicate that the confounding effects of worker heterogeneity on wages have been netted out of cross-firm wage differences. Wage premia defined this way are the theory-consistent empirical measure for wages because my model assumes that all workers are equally productive. They still include, however, cross-firm differences in both components of the wage: the marginal revenue product of labor (productivity, production function, product market structure, etc.) and the wage markdown (the market power component).[37]

I measure the tariff shock to firm z as the policy-induced change in import tariffs on firm z's output sector:

$$\Delta \ln \left(1 + \tau_{i(z)}\right) \equiv - \ln \left(\frac{1 + \tau_{i(z),1994}}{1 + \tau_{i(z),1990}} \right) \tag{17}$$

[34] Workers employed as of December 31. This is the standard measure of firm-level employment in RAIS (e.g., Kovak (2013); Dix-Carneiro and Kovak (2017)). I estimate firm wage premia on the subset of these workers hired before December, whose December earnings cover the full month.

[35] Alternatively, one could measure l_{zmt} as total hours and w_{zmt} as the hourly wage premium. Hours data are unavailable for this period, but Dix-Carneiro and Kovak (2017) show that incorporating hours does not affect estimates of trade's effect on wages in later years.

[36] For each year, I estimate firm wage premia as firm × market fixed effects in a regression of workers' log December earnings on firm × market fixed effects plus controls for age, education, and gender (see Appendix A). Employment l_{zmt} counts all of the firm's workers.

[37] As robustness, I present estimates of the within-market cross-firm elasticity based on alternative wage measures, facilitating comparison with other papers that use average wages, as in BHM and Yeh, Macaluso and Hershbein (2022) for the US, or manufacturing plant data (e.g., Amodio and de Roux (2024); Pham (2023); Tortarolo and Zarate (2018)). To check whether unobservable worker characteristics that vary over time matter, I present results that also condition stayers' premia on worker fixed effects, as in Abowd, Kramarz and Margolis (1999).

where the minus sign is included to facilitate interpretation of regression coefficients (i.e., such that a positive coefficient means that the policy-induced import tariff reduction had a positive effect on the outcome variable). Appendix E reports results using effective rates of protection, which net out tariff reductions on firm inputs.

The identifying variation for equation 16 comes from firms of different output sectors operating in the same local labor market (i.e., hiring in the same microregion × occupation group pair), including firms in non-tradable sectors, for which the change in import tariffs is zero. I estimate equations 15 and 16 clustering standard errors at the firm level, and weighting the regression by the base-year employment of the firm's local labor market, so that markets count in proportion to their workforce. I then present robustness estimates to alternative clustering schemes, labor market boundaries, tariff shocks, and wage measurements.

5.2 Cross-market inverse elasticity of substitution

5.2.1 Regression specification

To derive the regression specification for estimating $\frac{1}{\theta}$, I start by returning to the long-differenced version of the model's logged inverse residual labor supply equation (i.e., equation 15), but this time I pay close attention to the market-level changes that are absorbed into the fixed effect $\Delta\delta_m$:

$$\Delta \ln w_{zm} = \frac{1}{\eta}\Delta \ln l_{zm} + \underbrace{\left(\frac{1}{\theta} - \frac{1}{\eta}\right)\Delta \ln L_m - \frac{1}{\theta}\Delta \ln L + \Delta \ln W + \Delta \ln \xi_m^{1+\theta}}_{\Delta\delta_m} + \Delta\epsilon_{zm} \tag{18}$$

It follows from equation 18 that, given estimates of $\Delta\delta_m$, $\frac{1}{\eta}$, and residuals $\Delta\epsilon_{zm}$—obtained by first estimating equation 15—the following regression can be used to estimate the gap $\left(\frac{1}{\theta} - \frac{1}{\eta}\right)$ between workers' key elasticities of substitution, and thus $\frac{1}{\theta}$:

$$\text{[Second Stage]} \quad \Delta\delta_m = \alpha + \left(\frac{1}{\theta} - \frac{1}{\eta}\right)\Delta \ln L_m + \Delta\epsilon_m \tag{19}$$

where the constant α absorbs country-level wage component changes common to all markets (i.e., $\alpha = \frac{1}{\theta}\Delta \ln\left(\frac{1}{L}\right) + \Delta \ln W$), $\Delta \ln L_m$ is the change in the CES market-level labor supply index, whose measurement I describe in Section 5.2.2, and $\Delta\epsilon_m = \Delta \ln \xi_m^{1+\theta}$ is the market-level regression residual, which also has a structural interpretation as the (scaled) change in workers' taste for market m.

The key threat to identification of $\left(\frac{1}{\theta} - \frac{1}{\eta}\right)$ in regression equation 19 is that changes in the taste-adjusted labor supplied to market m (i.e., $\Delta \ln L_m$) are correlated with changes in workers' taste for market m (i.e., $\Delta\epsilon_m = \Delta \ln \xi_m^{1+\theta}$). To address this concern, I instrument the market-level change in

labor supply with a market-level labor demand shock introduced earlier: ΔICE_m, the market-level policy-induced import competition exposure shock commonly felt by all firms in market m. My market-level first stage regression is thus:

$$\text{[First Stage]} \quad \Delta \ln L_m = \tilde{\alpha} + \lambda \Delta ICE_m + \Delta \nu_m \tag{20}$$

where $\tilde{\alpha}$ is a constant, and $\Delta \nu_m$ is a regression residual. In both stages I also include state fixed effects and the market's 1986–1990 changes in log employment and in the log wage premium, the controls that Dix-Carneiro and Kovak (2017) use in their regional regressions. These controls do not come from the model, where the taste shocks $\Delta \epsilon_m$ are the only source of variation in the wage index other than labor demand. However, workers' tastes for a market may have been changing before the reform—through regional policies, migration, and the state-level reforms that accompanied the 1988 Constitution—in ways correlated with the trade shock at the regional level, in which case $\Delta \epsilon_m$ would be correlated with ΔICE_m. State fixed effects absorb the part of this process that is common to a state's markets, and the market's own 1986–1990 changes absorb the part that is specific to the market; because these changes end in 1990, they share no base year with the 1991–2000 changes. I report estimates with and without each set of controls.

The two identifying assumptions are that there is a first stage (i.e., $\lambda \neq 0$), and that the instrument is excluded (i.e., ΔICE_m affects $\Delta \delta_m$, the market-level component of firm wages, only via market-level changes in employment, as opposed to changing workers' distaste ξ_m for market m). Once again, the first stage assumption is testable, and while the exclusion restriction is not testable, it might be amenable to exploration in future work by correlating estimates of ξ_m with market characteristics that might influence worker tastes.

Finally, I estimate $\frac{1}{\theta}$ by summing my estimate of $\left(\frac{1}{\theta} - \frac{1}{\eta}\right)$ from equation 19 with my estimate of $\frac{1}{\eta}$ from equation 15, taking into account the standard errors of each estimate in order to assess precision for $\frac{1}{\theta}$.

5.2.2 Measurement

To estimate equations 19 and 20, I need to measure three objects: $\Delta \delta_m$, the market-level component of the firm-level wage change; $\Delta \ln L_m$, the market-level change in the CES labor supply index; and ΔICE_m, whose measurement I have already introduced in Section 4. I measure $\Delta \delta_m$ as the market fixed effect from regression equation 15 in Section 5.1.1, and compute $\Delta \ln L_m$ given my

point-estimate for $\frac{1}{\eta}$ as follows:

$$\Delta \ln L_m = \Delta \ln \left\{ \left[\sum_{z \in \Theta_m} (\xi_{zm} l_{zm})^{\frac{1+\eta}{\eta}} \right]^{\frac{\eta}{1+\eta}} \right\}$$

where Θ_m is the set of all firms operating in market m, and the taste-shifters ξ_{zm} are calculated using equation 14 and my point-estimate for $\frac{1}{\eta}$.[38] I estimate equation 19 clustering standard errors at the local labor market level. I estimate this regression without weights, unlike my main market-level regressions of Section 4. Here, the choice of weights does not concern which markets the estimate describes, since $1/\theta$ is a structural elasticity common to all markets, but rather which variation identifies it: when I weight by 1991 employment, the first stage is driven by a few large markets and has no power (F statistic below 1). The variation that identifies $1/\theta$ differs across its two components: the response of market employment to ΔICE_m comes mostly from comparing markets with and without tradable firms, whereas the response of the market's wage premium index comes in part from differences in tariff reductions across exposed markets (Appendix C.5). The extended model of Section 8, which incorporates informal workers, including those in agriculture, and which I estimate on census cells, yields a cross-market inverse elasticity of 1.187 (SE of 0.481), compared to 1.454 (SE of 0.196) here.

6 Estimates of key elasticities of substitution

6.1 Within-market cross-firm inverse elasticity of substitution

Table 2 presents my estimate of $\frac{1}{\eta}$ based on equations 15 and 16, and Appendix Figure A.9 plots its identifying variation year by year. Column (1) presents the main specification. The first stage in Panel A shows that a 1 percent decrease in the import tariff on firms' output reduced employment by 0.736 percent. This is a strongly identified first stage, with an F-statistic of 94.9. Panel B shows that the proportional effect on firms' wage premia was somewhat smaller, at a 0.612 percent reduction. Combined, these effects imply a within-market cross-firm inverse elasticity of substitution of 0.832.

A within-market cross-firm inverse elasticity of substitution of 0.832 means that if a firm wished to poach from its local competitors 1 percent of its current employment, it would have to increase its wage premium by roughly 0.8 percent. This is a large estimate, about nine times larger than BHM's corresponding estimate of 0.09 for the US,[39] suggesting that Brazilian workers substitute much less

[38] Following equation 14, I compute the taste-shifters for each year as $\xi_{zmt} = \exp(\nu_{zmt}/(1+\eta))$, where ν_{zmt} are the residuals from a regression of $[\ln w_{zmt} - (1/\eta) \ln l_{zmt}]$ on a market fixed effect. This follows from the structural residual $\epsilon_{zmt} = \ln \xi_{zmt}^{1+\eta} = (1+\eta) \ln \xi_{zmt}$, so that $\ln \xi_{zmt} = \epsilon_{zmt}/(1+\eta) = \nu_{zmt}/(1+\eta)$.

[39] BHM report an η of 10.85, whose inverse is 0.09, based on local labor markets defined as commuting zone ×

readily across firms in response to wage changes than US workers do. This rather inelastic preference parameter places an upper bound of $1/(1+0.832) \approx 55\%$ on firms' wage take-home shares.

This point estimate is robust to alternative specifications, two of which are presented in the remaining columns of Table 2. A first concern is that such an inelastic within-market cross-firm response might be driven by local labor markets being defined too narrowly, such that within any one market there are too few firms for workers to substitute across. This does not appear to be the case: column (3) of Table 2 shows that defining local labor markets more broadly by microregion only yields a larger, not a smaller, within-market cross-firm inverse elasticity of substitution (0.870 against 0.832). Since the baseline estimate is only identified by variation within occupations, this suggests that barriers to occupational switching cannot account for Brazilian workers' rather inelastic within-market cross-firm elasticity of substitution.

I also find a similar elasticity when I use effective rates of protection instead of import tariffs as shocks (0.878 against 0.832; Appendix E), with first stage and reduced form effects about a third smaller, consistent with the greater noise in effective rates of protection (footnote 21). In column (2) of Table 2, I use stayers' premia conditional on worker fixed effects (Section 5.1.2). The inverse elasticity is 0.831, the same as in column (1). Across wage measures and samples, the implied upper bound on the take-home share ranges from 54 to 56 cents (Appendix Table A.8). At least in the Brazilian context, then, the labor market power exerted by local firms is not driven by worker sorting across firms.[40] One important limitation of the labor supply preference structure in Section 2 is that it assumes preference parameters are homogeneous across worker demographics and across regions, which might not be the case. If firms take advantage of demographic-based preference heterogeneity to set differential markdowns by demographics, the estimates thus far would not capture it.[41] In Section 8, I first examine, within the baseline model, how my estimated elasticities vary with local informality (Section 8.1). I then extend the model to incorporate informal wage work and self-employment, and I show how formal firms' markdowns and their response to trade vary across local labor markets.

Finally, $\frac{1}{\eta}$ is a preference parameter, and thus assumed stable over the period of study. My estimate nonetheless rises with the end year of the long difference (Appendix Figure A.10): it is 0.44 for 1991–1995, between 0.64 and 0.72 for end years 1996 to 1999, and 0.83 for 1991–2000.[42]

NAICS3 pairs.

[40] Appendix Table A.8 reports alternative wage measures and samples. Stayers' premia without worker fixed effects give 0.779, with the same first stage (−0.743 against −0.736) and a slightly smaller reduced form (−0.579 against −0.612), and December average wages instead of wage premia give 0.820. Including firms that exit by 2000, coding their employment and wages as zero at exit, gives 0.837. Appendix Table A.9 reports alternative clustering.

[41] Union bargaining could also compress wages across workers. Appendix Figure A.13 plots formal sector earnings against union density across microregions in 1991 and 2000.

[42] Appendix Figure A.9 also shows differential pre-trends in firm-level wage premia and employment, of the opposite sign to the post-liberalization effects. The pre-period coefficients do not come from the same set of firms as the 1991–2000 long difference: each year's coefficient uses the firm–market pairs observed in that year and in 1991, and the

6.2 Cross-market inverse elasticity of substitution

Table 3 presents my estimate of $\frac{1}{\theta}$ based on equation 19, with state fixed effects and 1986–1990 trends. The first stage in Panel A shows that a 1 percent increase in a market's import competition exposure reduced employment by 0.294 percent, whereas Panel B shows that the proportional effect on markets' wage premia indices was about 60 percent as large, at a 0.183 percent reduction. Combined, the first stage and reduced form produce an IV estimate of 0.622 (SE of 0.217) for the difference between $\frac{1}{\theta}$ and $\frac{1}{\eta}$, which implies a cross-market inverse elasticity of substitution of 1.454 given the estimate for $\frac{1}{\eta}$ from Section 6.1. Columns (2) and (3) repeat the estimation with the variants of Table 2: with wage premia conditional on worker fixed effects, $\frac{1}{\theta}$ is 1.135 (SE of 0.211), and with local labor markets defined as microregions, 1.034 (SE of 0.250); with the shock measured by effective rates of protection, it is 1.758 (SE of 0.421; Appendix E). In every case $\frac{1}{\theta}$ exceeds $\frac{1}{\eta}$, although the difference is not statistically significant in columns (2) and (3).

There are two important takeaways from Table 3. The first is that the standard error on the IV estimate allows us to reject that $\frac{1}{\theta}$ and $\frac{1}{\eta}$ are the same (p-value < 0.01), which means that we can reject the model's limiting case of monopsonistic competition. This means increases in labor market concentration do matter for firm labor market power. The gap grows as I add the controls of Section 5: it is 0.298 (SE of 0.127) with a constant only, 0.430 (SE of 0.151) with state fixed effects, and 0.622 with state fixed effects and the 1986–1990 trends; with stayers' premia it is smaller and less precise in every specification, 0.298 (SE of 0.182) with the full controls. With standard errors clustered two ways, by microregion and by occupational group, its standard error is 0.473 (Appendix Table A.9). The second is that, while I can reject the null that the within-market and cross-market elasticities are the same, the magnitude of their gap is moderate, suggesting that Brazilian workers find it somewhat harder to substitute globally (i.e., across markets) than locally (i.e., within markets, across firms). This matters for quantifying the effect of increased concentration on wage markdowns.

It is helpful to compare my estimate of 1.454 for Brazil's cross-market inverse elasticity to other contexts. A cross-market inverse elasticity of substitution of 1.454 means that a market's wage premium index (i.e., the taste-adjusted wage premium) would have to increase by 1.454 percent before one percent more workers were attracted from other markets. While relatively inelastic, this point estimate is about 60% of BHM's corresponding estimate of 2.4 for the US, suggesting Brazilian workers substitute more readily across local markets than US workers do.[43]

imbalance is concentrated in 1986–1988, the transition years between the military dictatorship and the 1988 Constitution. In my main specification I therefore restrict the estimation to firm–market pairs also observed in 1986, which were present before the 1988 Constitution, and I control for the pair's own 1986–1991 changes in log employment and in the log wage premium (Table 2). Without these controls, on the same pairs, I estimate 1.372; using the 1988–1991 changes as controls instead, I estimate 1.083. The controls share the 1991 level with the outcome and the endogenous regressor, so that transitory shocks to a pair's 1991 wage premium and employment enter the controls and the 1991–2000 changes with opposite signs, and the controlled estimate is the lower one.

[43] BHM report $\theta = 0.42$, whose inverse is 2.4, based on local labor markets defined as commuting zone × NAICS3

On net, this relatively inelastic cross-market elasticity of substitution places a lower bound of $1/(1+1.454) \approx 41\%$ on wage take-home shares.

6.3 Pre-liberalization average wage markdown

I now combine my estimates of $\frac{1}{\theta}$ and $\frac{1}{\eta}$ from Section 6 with data on local labor markets' payroll Herfindahl indices to estimate Brazil's pre-liberalization average markdown, along with its (more easily interpretable) inverse, the wage take-home share. Appendix B.2.4 shows that the country-level average markdown—that is, the country-level ratio of (employment-weighted) average MRPL to (employment-weighted) average wage—is a weighted average of the market-level markdowns in Proposition 1, where the weights are each market's payroll share of the country's total payroll.

In the baseline year of 1991, the payroll-weighted average concentration was 0.08 (Section 3.2). Combined with my estimates of $\frac{1}{\theta}$ and $\frac{1}{\eta}$, this level of concentration implies, per equation 9, that Brazil's formal sector pre-liberalization average wage markdown was approximately 1.9, whose inverse gives an average wage take-home share of approximately 53 percent.[44] That is, Brazilian formal sector workers took home 53 cents of every dollar of marginal revenue product of labor they generated. This places Brazil's formal sector as less competitive than US labor markets, where wage take-home shares are estimated at 65% for manufacturing (Yeh, Macaluso and Hershbein, 2022) and 74% with BHM's estimates,[45] but close to estimates for developing countries. Using a production function approach on a harmonized global panel of manufacturing firms across 82 low- and middle-income countries, Amodio et al. (2025) estimate a median take-home share of 43 cents on the dollar. For Peru, Amodio, Medina and Morlacco (2025) find that self-employment curbs formal firms' market power, while Pham (2023) report 47% for Chinese manufacturing.[46] Across these studies, markdowns are consistently found to increase with firm size.[47]

7 Wage effect decomposition

What does the trade-induced increase in labor market concentration documented in Section 4 imply for wages? In the model, a firm's wage is its marginal revenue product of labor times its take-home

pairs.

[44]The country's average wage take-home share does vary somewhat with alternative measures of labor market concentration due to the gap between the two inverse elasticities of substitution. For example, the country-level take-home share would have been about 50 percent if evaluated at the (unweighted) average payroll Herfindahl of 0.29.

[45]$74\% \approx 1/(1 + (1/0.42) * 0.11 + (1 - 0.11) * (1/10.85))$ using $\theta = 0.42$, $\eta = 10.85$, and payroll-share-weighted Herfindahl of 0.11 from BHM. See Appendix B.2.4 and B.2.6.

[46]Pham (2023)'s "pass-through" corresponds to the wage take-home share: $0.47 = 1/2.14$, where 2.14 is the average $MRPL_i/w_i$.

[47]In my model, markdowns increase with a firm's payroll share whenever $1/\theta > 1/\eta$; regressing $\ln(\mu_{zm})$ on log employment yields a small positive coefficient (+0.003) across specifications (Appendix Table A.10).

share, $w_{zm} = \mu_{zm}^{-1} r_{zm}$ (equation 5), so any change in a wage is a change in the take-home share plus a change in the marginal revenue product. I compute the take-home share from the estimated elasticities and observed payroll shares (equation 8), observe the wage, and obtain the marginal revenue product as the residual. I apply this identity first within markets, to the firms facing larger tariff cuts, and then across markets, to the incumbent firms present in both 1991 and 2000 and to the market average, which also reflects changes in the composition of employment across firms, including the firms that enter and exit. Across markets, I use one specification throughout: the 1991–2000 change on ΔICE_m, with each market counting once, state fixed effects, and the market's 1986–1990 changes in log employment and log wage premium, as in the estimation of $1/\theta$ (Table 3). Appendix B.3 states the identities and reports the estimates under other weights and controls.

At the market level, Figure 7 presents my estimates of γ_t from Corollary 1 for all sample years, and Table 4 the 2000 estimate. The 0.622 gap between the cross-market and within-market inverse elasticities of substitution implies that a 10% increase in import competition exposure increased the average wage markdown by 0.013, lowering the take-home share by 0.37 cents from its pre-liberalization level of 53 cents on the dollar, or 0.7% of the average wage.[48] The effect is small because the gap is moderate: with the 2.29 gap that BHM estimate for the US, the same increase in concentration would have lowered the take-home share by 2.8 cents, more than seven times as much (Appendix B.3).

Within markets, however, markdowns did not rise at the firms facing larger tariff cuts. A one log point larger tariff cut reduced a firm's wage premium by 0.61 log points (Table 2) without changing its take-home share, so a lower marginal revenue product of labor accounts for the whole decline (Appendix Table A.12). Since firm productivity rose during Brazil's liberalization (Muendler, 2004), and the within-market elasticity is nearly unchanged when I measure tariff shocks as effective rates of protection, which net out input tariff reductions (Appendix E), this decline more likely reflects lower output prices than productivity losses, consistent with the pro-competitive effects of import competition predicted by Melitz and Ottaviano (2008).

Higher markdowns instead lowered wages at incumbents, whose payroll shares rose as their markets became more concentrated. Table 5 applies the identity to these firms, weighted by their 1991 employment. A 10% increase in import competition exposure reduced incumbents' average wage premium by 1.2%: 0.38 percentage points, or 33%, through a lower take-home share, and 0.77 points, or 67%, through a lower marginal revenue product of labor.[49]

Across firms, the marginal revenue product of labor fell by less than at incumbents, because the composition of employment shifted toward firms with higher marginal revenue products. I split the

[48] Here I use the unweighted effect on concentration (Appendix Table A.4, column (4)) because I estimate $1/\theta$ with each market counting once (Section 6.2). With the weighted effect of Table 1, the take-home share falls by 0.30 cents.

[49] I compute standard errors using a microregion cluster bootstrap with 200 replications, each of which re-estimates $1/\eta$, $1/\theta$ and every object of the table.

change in a market's average marginal revenue product into a within-firm term, the change in each firm's marginal revenue product at its 1991 employment share, and a composition term, the change in employment shares across firms, including the firms that enter and exit (equation B.10). Figure 8 plots the effect of import competition exposure on the two terms year by year. By 2000, a 10% increase in exposure lowered the within-firm term by 2.4% and raised the composition term by 1.8%, so that the market's average marginal revenue product fell by 0.6% (SE of 0.4). The within-firm term is larger than the decline at incumbents in Table 5 because it adds up firms' changes in levels, in which firms with a high marginal revenue product weigh more, whereas Table 5 averages their log changes. Before 1991, both terms rose somewhat faster in more exposed markets, by at most 0.06 and 0.09 per unit of ΔICE_m, against −0.24 and 0.18 by 2000. Changes in employment shares among incumbents account for 2.1 percentage points of the composition term; the contributions of entrants (2.1) and exiters (−2.5) nearly offset each other. The market's average wage premium therefore fell by 1.0%, against 1.2% at incumbents.

Taken together, my estimates show that formal firms' labor market power is large but responds little to concentration. This power stems from workers' inelastic substitution across firms, not from concentration: workers would take home 55 cents, instead of 53, if employers were atomistic. Moreover, because the within-market and cross-market elasticities of substitution are close, the 22% increase in concentration lowered the take-home share by less than half a cent. This is a small change in the wage, but so is the wage decline through 2000, of which it accounts for one third at incumbents. Across firms, a shift in the composition of employment toward firms with higher marginal revenue products offset most of the decline in marginal revenue products within firms. In Section 8, I ask whether these conclusions hold once I incorporate informal wage work and self-employment.

8 Incorporating informal wage work and self-employment

How do the paper's conclusions change when the model incorporates informal wage work and self-employment? In this section, I explore this question in two steps. I first estimate, with the baseline model, how $1/\eta$ and the effect of trade liberalization on concentration vary with local formality (Section 8.1). I then extend the model of Section 2 to incorporate both margins and re-estimate each of the paper's main results (Sections 8.2 to 8.4; Appendix F presents the extended model in detail). In the extension, the object of interest is unchanged: the wage markdown commanded by formal sector firms. The labor demand side is also unchanged: as in the baseline, I model the wage-setting problem of formal firms, which compete for labor à la Cournot. What changes is the labor supply structure, which describes the preferences of workers seeking formal employment: it now includes the options that these workers have outside formal employment, and hence changes the labor supply that formal firms face.

I find that the results based on estimating equations from the baseline model hold, but adding informality lowers the effect of trade liberalization on wage markdowns by more than half and makes it vary across local labor markets. Trade raised concentration at every level of local formality, and $1/\eta$ in the most formal quartile of microregions, which hosts 73% of formal employment, is 0.824, against 0.832 in the pooled sample; elsewhere I cannot estimate it precisely. In the extended model, formal sector workers took home 53 cents of every marginal dollar, as in the baseline, while a 10% increase in the trade shock lowers the wage take-home share by 0.15 cents on the dollar, against 0.37 cents in the baseline (Table 7), with an effect that ranges from −0.16 cents in the most formal quartile to 0.00 cents in the least formal.

8.1 Heterogeneity by informality

In the baseline model, the within-market cross-firm inverse elasticity of substitution, $1/\eta$, is the key parameter for the level of wage markdowns, because most formal sector workers are in local labor markets with relatively low concentration: the employment-weighted average payroll Herfindahl is 0.08 (Table 1), compared with a median of 0.21 across local labor markets (Appendix Table A.1). Had employers been atomistic, workers would have taken home 55 cents instead of 53. Whether $1/\eta$ and the effect of trade liberalization on local labor market concentration vary with baseline formality thus gives a first indication of how not accounting for the informal sector might bias the baseline estimates. I classify microregions into quartiles of the formal share of total employment in the 1991 census, from most formal (quartile 1), with a median formal share of employment of 50%, to least formal (quartile 4), with a median of 10%, and estimate equations 12 and 15–16 separately for the local labor markets of each quartile. Table 6 presents the results.

Panel A shows that trade raised concentration in every quartile: a 10% increase in import competition exposure raised the payroll Herfindahl by 0.015 to 0.017 points in the three most formal quartiles and by 0.048 points in the least formal. Concentration itself rises with informality: the 1991 employment-weighted average Herfindahl is 0.06 in the most formal quartile and 0.29 in the least formal. Relative to these levels, a 10% increase in exposure raised concentration by 23% in the most formal quartile and by 9 to 17% in the others.

Panel B shows that $1/\eta$ is identified in the most formal quartile, which hosts 73% of national formal employment. The estimate there, 0.824, is nearly identical to the pooled estimate of 0.832. In quartiles 2 to 4, the firm-level tariff instrument has no power (first-stage F statistics below 3), and the estimates are too imprecise to be informative.[50]

These splits hold the model fixed and vary only the estimation sample, and they deliver two

[50] Appendix Table A.11 splits quartile 1 at the median formal share of its microregions and pools quartiles 2 to 4. The upper half of quartile 1, which alone hosts 57% of national formal employment, yields 0.821. The lower half and the pooled quartiles 2 to 4 yield imprecise estimates (1.650 and −1.252), with first-stage F statistics of 1.4 and 1.9.

messages. First, the baseline estimates come from the markets where formal employment is located (Figure 2): nearly three quarters of formal employment is in the most formal quartile, where the estimate of $1/\eta$ matches the pooled one and concentration is lowest. Since the country-level average markdown weights markets by their payroll (Section 6.3), this suggests that incorporating informal wage work and self-employment should have a limited effect on its level. Second, outside the most formal quartile, where informal wage work and self-employment account for most employment, concentration is higher, trade raised it further, and I cannot estimate the baseline elasticity. Assessing how these margins change the paper's conclusions thus requires a model of workers' options outside formal employment. I provide one in the rest of this section.

8.2 Model extension

I modify the labor supply structure of Section 2.1, which describes the preferences of people seeking formal employment, to account for the two forms of employment outside the formal sector: self-employment and informal wage work.

I motivate this structure with evidence on what workers at the margin of informality value in a job. Using discrete choice experiments in Brazil's largest slum complex, Felix et al. (2026) document two facts. First, workers who search for jobs search for formal jobs: among job seekers, 77% of formal employees, 84% of informal employees, 56% of the self-employed, and 63% of the unemployed look for formal jobs, and less than 1% of any group look for informal ones. Second, willingness to pay for the amenities mandated in formal jobs is high: 34% of wages for unemployment insurance, 24% for parental leave, and 18% for termination notice. Informal employees value unemployment insurance about as much as formal employees do (34% and 38% of wages), while the self-employed value it little (10%). Formal jobs thus carry a premium beyond wages, attributable to mandated amenities, that formal and informal employees value similarly and the self-employed do not. This points to sorting on preferences between wage work and self-employment, and to formal jobs being rationed among those who prefer wage work, consistent with the rationing of wage jobs documented by Breza, Kaur and Shamdasani (2021) in India. In a monopsonistic model of Brazil's labor market with a minimum wage, Parente, Brotherhood and Iachan (2026) show that low-productivity formal firms constrained by the minimum wage hire fewer workers than are willing to work for them, so that workers who prefer a formal job can be rationed out of it.

Combined, these facts lead me to model the two informal margins differently. I model self-employment as a sector that workers choose: each local labor market is split into wage work and self-employment, with an elasticity of substitution $\tilde{\rho}$ between the two, so that workers' choice set nests firms within wage work, wage work versus self-employment within markets, and markets within the economy (with within-market cross-firm elasticity $\tilde{\eta}$ and cross-market elasticity $\tilde{\theta}$). I model

informal wage work through the probability of involuntary separation from formal employment. I assume that a worker separated from a formal job finds informal wage work without frictions, in a perfectly competitive local labor market for informal labor. This assumption reflects how informal wage work operates in Brazil. Informal jobs are destroyed three to five times as often as formal jobs, and unemployed workers find them more frequently (Meghir, Narita and Robin, 2015). For younger workers in particular they resemble a queue for formal jobs, whereas entry into self-employment is largely voluntary (Bosch and Maloney, 2010). And they absorbed workers displaced from formal employment by Brazil's trade liberalization (Dix-Carneiro and Kovak, 2019).[51][52]

Expected wages. Formally, a job at formal firm z in the wage work sector $\overline{g}$ of market m carries a perceived probability $p_{z\overline{g}m}$ of involuntary separation, after which the worker earns the local informal wage w_m^o (in a variant, preceded by unemployment insurance benefits; Section 8.4). The wage relevant for attracting labor supply to the firm is therefore the expected wage

$$\bar{w}_{z\overline{g}m} = \left(1 - p_{z\overline{g}m}\right) w_{z\overline{g}m} + p_{z\overline{g}m}\, w_m^o = w_{z\overline{g}m} \left(1 + p_{z\overline{g}m}\, \sigma_{z\overline{g}m}\right), \tag{21}$$

where $w_{z\overline{g}m}$ is the firm's posted formal wage and $\sigma_{z\overline{g}m} \equiv (w_m^o - w_{z\overline{g}m})/w_{z\overline{g}m}$ is the wage gap between the local informal sector and the firm's formal wage, which may be negative or positive. I measure the separation probabilities $p_{z\overline{g}m}$ from firms' histories of separations in RAIS, which are themselves endogenous labor demand decisions. I assume that firms make labor demand decisions one period at a time and derive *no dynamic gains from influencing workers' priors*: a firm does not internalize that its hirings and separations at time t lead workers to update, at $t + 1$, their beliefs about separation risk at that firm. This assumption is strong, but it preserves the static wage-setting problem of Section 2; relaxing it would turn the separation probability into an additional, dynamic instrument of wage setting.

Labor supply to a formal firm. Workers' idiosyncratic tastes follow a nested GEV distribution with a sector level between the firm and market levels of the baseline model (Appendix F.1). The

[51] Appendix F.4 considers the alternative in which I treat informal wage work and self-employment symmetrically, as a single informal sector that workers choose as an alternative to formal employment within each local labor market. Appendix Table F.1 reports the results, which Section 8.4 compares with those of the extended model.

[52] An alternative extension would nest the choice between formal and informal wage work inside the wage work alternative, which requires estimating workers' elasticity of substitution between the two. The difficulty is that informal wage work in Brazil occurs both inside and outside formal firms (Ulyssea, 2018), and the two cannot be told apart in the data. A trade shock that lowers formal labor demand can lead formal firms to move workers off the books, in which case changes in informal employment and earnings trace firms' demand for informal labor, not workers' supply of it. Appendix Figures F.3 and F.4 show this pattern: across microregions, larger values of Dix-Carneiro and Kovak (2017)'s exposure measure, RTR, are associated with higher employment shares and lower earnings in informal wage work and in self-employment, and with a lower formal employment share but no change in the earnings of workers who report formal jobs. Estimates that treat informal employment as a sector workers choose therefore overstate the elasticity of substitution: under the symmetric treatment, the inverse elasticity of substitution between formal and informal employment is 0.023 (SE of 0.060).

inverse labor supply curve faced by formal firm z is the counterpart of equation 3,

$$\bar{w}_{z\bar{g}m} = W \left(\frac{l_{z\bar{g}m}}{L_{\bar{g}m}}\right)^{\frac{1}{\tilde{\eta}}} \left(\frac{L_{\bar{g}m}}{L_m}\right)^{\frac{1}{\tilde{\rho}}} \left(\frac{L_m}{L}\right)^{\frac{1}{\tilde{\theta}}} \xi_{z\bar{g}m}^{1+\frac{1}{\tilde{\eta}}} \xi_{\bar{g}m}^{1+\frac{1}{\tilde{\rho}}} \xi_m^{1+\frac{1}{\tilde{\theta}}}, \tag{22}$$

where $(W_{\bar{g}m}, L_{\bar{g}m})$, (W_m, L_m), and (W, L) are nested CES wage and labor supply indices over firms, sectors, and markets, and the ξ's are taste shifters. The wage needed to attract labor is increasing in the firm's employment relative to its sector, in the sector's employment relative to the market, and in the market's employment relative to the country, with the elasticities $\tilde{\eta}$, $\tilde{\rho}$, and $\tilde{\theta}$. The difference from the baseline model is that the relevant wage is the expected wage of equation 21.

Firm-specific wage markdown. Firms compete for labor à la Cournot, as in Section 2.2, holding constant competing firms' labor demand, workers' priors about separation probabilities, expected informal wages, and expected returns to self-employment. Differentiating the log of equation 22 with respect to the firm's own employment gives the inverse elasticity of residual labor supply that the firm faces, and hence its wage markdown $\mu_{z\bar{g}m} = 1 + \varepsilon_{z\bar{g}m}^{-1}$:

$$\varepsilon_{z\bar{g}m}^{-1} = \left[\frac{1}{\tilde{\rho}} s_{z\bar{g}m} + \frac{1}{\tilde{\eta}}(1 - s_{z\bar{g}m})\right] + \left(\frac{1}{\tilde{\theta}} - \frac{1}{\tilde{\rho}}\right) s_{zm}, \tag{23}$$

where $s_{z\bar{g}m}$ is firm z's share of the wage work sector's *expected* payroll and s_{zm} is its share of the market's total expected payroll, including self-employment. The elasticity of labor supply to a formal firm thus depends on how easily workers substitute between firms within markets ($\tilde{\eta}$), between wage work and self-employment within markets ($\tilde{\rho}$), and between markets ($\tilde{\theta}$); the relevant shares are based on expected wages, so that separation probabilities and local informal wages enter market structure.

Local average wage markdown. Aggregating equation 23 with firms' expected payroll shares as weights, exactly as in Section 2.3, yields the extended counterpart of equation 9. The average wage markdown in the wage work sector of market m is $\mu_{\bar{g}m} = 1 + \bar{\varepsilon}_{\bar{g}m}^{-1}$, with

$$\bar{\varepsilon}_{\bar{g}m}^{-1} = \left[\frac{1}{\tilde{\rho}} \widetilde{HHI}_{\bar{g}m} + \frac{1}{\tilde{\eta}}\left(1 - \widetilde{HHI}_{\bar{g}m}\right)\right] + \left(\frac{1}{\tilde{\theta}} - \frac{1}{\tilde{\rho}}\right) \widetilde{HHI}_{\bar{g}m} \times s_m, \tag{24}$$

where $\widetilde{HHI}_{\bar{g}m} \equiv \sum_z s_{z\bar{g}m}^2$ is the wage work sector's *expected payroll* Herfindahl—throughout, a tilde on a Herfindahl denotes one computed with expected wages—and s_m is the wage work share of market m's total expected payroll, with self-employment making up the remainder. Relative to the baseline model, equation 24 shows that formal firms' market power now also depends on how easily workers substitute into self-employment ($1/\tilde{\rho}$ replaces $1/\theta$ as the coefficient on concentration), and that concentration interacts with the local importance of wage work through $\widetilde{HHI}_{\bar{g}m} \times s_m$.

Import competition effects on wage markdowns. The effect of a trade shock X_m on markdowns follows from differentiating equation 24:

$$\frac{\partial \bar{\varepsilon}^{-1}_{\bar{g}m}}{\partial X_m} = \left(\frac{1}{\tilde{\rho}} - \frac{1}{\tilde{\eta}}\right)\beta_{HHI} + \left(\frac{1}{\tilde{\theta}} - \frac{1}{\tilde{\rho}}\right)\left[s_m\,\beta_{HHI} + \widetilde{HHI}_{\bar{g}m}\,\beta_{s_m}\right], \tag{25}$$

where β_{HHI} and β_{s_m} are the reduced-form effects of the shock on the expected payroll Herfindahl and on the wage work share, respectively. Equation 25 is the extended-model counterpart of Corollary 1: trade moves markdowns through concentration, as in the baseline model, but now also through the composition of market employment between wage work and self-employment.

The bias from ignoring involuntary separation. The within-market cross-firm inverse elasticity in the extended model decomposes into the baseline elasticity and a bias term. Under the extended model, $1/\tilde{\eta}$ is the second-stage coefficient from an IV regression of firms' log *expected* wages on log employment, instrumented by an exogenous labor demand shock $X_{z\bar{g}m}$. Expanding the expected wage into the formal wage and the expected wage gap per $\ln \bar{w}_{z\bar{g}m} = \ln w_{z\bar{g}m} + \ln\left(1 + p_{z\bar{g}m}\sigma_{z\bar{g}m}\right)$, and letting $(*)$ denote partialled-out variables to apply the Frisch-Waugh-Lovell theorem, gives

$$\frac{1}{\tilde{\eta}} = \underbrace{\frac{\mathrm{Cov}(\ln w^*_{z\bar{g}m}, X_{z\bar{g}m})}{\mathrm{Cov}(\ln l^*_{z\bar{g}m}, X_{z\bar{g}m})}}_{\frac{1}{\eta}\text{: Supply response to formal wage}} + \underbrace{\frac{\mathrm{Cov}[\ln(1 + p_{z\bar{g}m}\sigma_{z\bar{g}m})^*, X_{z\bar{g}m}]}{\mathrm{Cov}(\ln l^*_{z\bar{g}m}, X_{z\bar{g}m})}}_{\Omega\text{: Supply response to expected wage gap}} \tag{26}$$

where $\sigma_{z\bar{g}m} \equiv (w^o_m - w_{z\bar{g}m})/w_{z\bar{g}m}$ is the pay gap between the local informal sector and firm z's formal wage for an effective unit of labor.

The first term in equation 26 is the labor supply response to changes in formal wages, the baseline parameter $1/\eta$. The second term, Ω, is the additional labor supply response to changes in expected formal–informal wage gaps. It is the bias in the baseline model's elasticity if the extended model is true. Whether the option of frictionless informal work curbs formal firms' market power or raises it depends on the sign of Ω, which captures the covariance between expected formal–informal wage gaps and formal labor supply among people seeking formal employment. If $\Omega < 0$, informal wage work acts as an outside option that curbs formal firms' market power, as self-employment does. If $\Omega > 0$, the threat of involuntary separation out of the formal sector strengthens formal firms' wage setting, and the baseline model understates their market power. The asymmetric treatment of the two informal margins is thus a modeling choice, but I estimate their effects on markdowns.

Taking the extended model to the data requires variation that separates the self-employment margin from the cross-market margin. I obtain it by partitioning each local labor market into 18 demographic cells defined by gender (2), education (3), and age (3), each with its own wage work and self-employment sectors. The partition rests on four assumptions. First, the labor demand side

is unchanged: conditional on demographics, workers are perfect substitutes from firms' perspective, as in Section 2. Workers of different cells therefore compete for the same jobs, and firms substitute freely among them; the cells are not segmented labor markets. Second, the elasticities $\tilde{\eta}$, $\tilde{\rho}$, and $\tilde{\theta}$ are the same for all cells. Third, the cells of a market are not separate markets: they are subject to the same market-level effects of trade liberalization, so the cross-market margin is common to them. Fourth, the cells of a market differ in their response to the market's trade shock, because at baseline they differ in their formal and wage work shares, and these differences are unrelated to changes in their tastes for wage work or for the market. The cells are therefore a source of identifying variation for elasticities that are common to all workers, not a dimension of preference heterogeneity. Empirically, I measure wages net of demographics: I estimate firm wage premia and census earnings conditional on gender, education, and age, and I compute expected wage gaps within cell.

8.3 Estimation of key parameters

To estimate the extended model, I combine RAIS with Brazil's 1991 and 2000 population censuses, which measure employment and earnings in informal wage work and self-employment by microregion and demographic cell. Since the census's occupation codes are unreliable in this period, I define markets as microregions in the extension. This applies to every object of the extended model, including concentration and $1/\eta$, which I re-estimate with markets defined as microregions. I use one trade shock, defined next, and the same 1991–2000 long-difference design for all of the extended model's estimates, as in the baseline.

The trade shock must also cover the informal margins of each market. I define the extended model's import competition exposure as

$$\Delta ICE^{*}_{m} \equiv f_{m,1991}\, \Delta ICE_{m} + \left(1 - f_{m,1991}\right) RTR_{m}, \tag{27}$$

where ΔICE_m is the import competition exposure of Section 4 (equation 11), with markets defined as microregions; RTR_m is Dix-Carneiro and Kovak (2017)'s exposure measure, RTR (regional tariff reductions), computed with the 1991 industry composition of all of the market's workers; and $f_{m,1991}$ is the formal share of the market's 1991 wage bill, measured in the census as the earnings of formal employees over the earnings of all employed workers, including informal employees and the self-employed. I define the shock for each microregion, so it is common to the demographic cells of a microregion. The measure reflects that the formal wage bill is exposed through the tariff reductions faced by formal firms, and informal earnings and the returns to self-employment through the industry composition of the region. Appendix F.2 provides details.

I estimate the effects of trade liberalization on the two dimensions of market structure that

enter equation 24 from

$$\Delta Y_m = \beta_Y \, \Delta ICE^*_m + \delta_r + u_m, \qquad Y \in \left\{ \widetilde{HHI}_{\bar{g}m}, \, s_m \right\}, \tag{28}$$

where Δ denotes 1991–2000 long differences and δ_r is a fixed effect for each of Brazil's five major regions. I weight regressions by microregions' 1991 formal employment and cluster standard errors by microregion. I estimate the labor supply response to formal wages, $1/\eta$, as in Section 5, with firm $\times$ microregion observations and microregion fixed effects.

I estimate the bias term Ω in equation 26 from the second-stage regression

$$\Delta \ln \left(1 + p_{z\bar{g}m} \sigma^c_{z\bar{g}m} \right) = \Omega \, \Delta \ln l^c_{z\bar{g}m} + \delta_z + \delta_m + \delta_c + \epsilon^c_{z\bar{g}m}, \tag{29}$$

where an observation is a firm $\times$ microregion $\times$ demographic cell, $l^c_{z\bar{g}m}$ is firm z's employment of workers of cell c, $\sigma^c_{z\bar{g}m}$ is the gap between the cell's informal earnings in microregion m and its earnings at firm z, differences are taken over 1991–2000, and δ_z, δ_m, and δ_c are firm, microregion, and cell fixed effects. The firm fixed effect ensures that Ω captures the labor supply response to expected wage gaps *conditional on* firm-level wage changes—the variation that identifies $1/\eta$. The separation probability $p_{z\bar{g}m}$ is the share of the firm's workers in the microregion who were fired during the year, measured in RAIS. I instrument the change in employment with the interaction between the firm-level tariff reduction (the instrument for $1/\eta$) and the microregion's ΔICE^*_m, which shifts equilibrium informal wages. I interact the instrument with gender, education, and age group indicators, weight observations by the lagged log of the firm's employment of the cell, and cluster standard errors by microregion. With firm, microregion, and cell fixed effects, the instrument varies only across the microregions of firms that operate in several of them, through differences in ΔICE^*_m scaled by the firm's own tariff reduction, and across demographic cells within a firm through the interactions; the first stage is driven by the common component, with small differentials by education and age.

The self-employment and cross-market elasticities come from the market-level labor supply system. Aggregating labor supply to the wage work sector and taking long differences gives

$$\Delta \ln W_{\bar{g}m} = \alpha + \frac{1}{\tilde{\rho}} \Delta \ln \left(\frac{L_{\bar{g}m}}{L_m} \right) + \frac{1}{\tilde{\theta}} \Delta \ln \left(\frac{L_m}{L} \right) + \epsilon_{\bar{g}m}. \tag{30}$$

Equation 30 cannot separately identify $\tilde{\rho}$ from $\tilde{\theta}$, because each market has a single wage work sector and tariff reductions were primarily a shock to wage work labor demand. Partitioning markets into demographic cells provides the required within-market, cross-sector variation. Cells of the same market differ in their response to the trade shock, and thus in the change in their wage work share,

while sharing the market's cross-market margin:

$$\Delta \ln W^c_{\overline{g}m} = \frac{1}{\tilde{\rho}} \Delta \ln \left(\frac{L^c_{\overline{g}m}}{L^c_m} \right) + \frac{1}{\tilde{\theta}} \Delta \ln \left(\frac{L_m}{L} \right) + \delta_r + \epsilon^c_{\overline{g}m}, \tag{31}$$

where $W^c_{\overline{g}m}$ is the residual earnings of cell c's wage workers, formal and informal, in microregion m, $L^c_{\overline{g}m}/L^c_m$ is the wage work share of the cell's employment, L_m/L is the microregion's share of national employment, and δ_r is a region fixed effect absorbing regional general equilibrium effects. I instrument the two endogenous variables with ΔICE^*_m and its interactions with age group, gender, and education indicators and with the cell's 1991 formal employment share. The shock is the same for all cells of a microregion, and the interactions let the first-stage responses of the two regressors differ across cells. My sample consists of cells with positive 1991 formal employment; I weight the regression by the cell's 1991 formal employment and cluster standard errors by microregion.

Finally, I compute the effect of trade liberalization on markdowns from equation 25, with the elasticities above and the reduced-form effects β_{HHI} and β_{s_m} of equation 28, evaluated at the 1991 averages of $\widetilde{HHI}_{\overline{g}m}$ and s_m across microregions.

8.4 Results

Table 7 compares the extended model's estimates to the baseline, with one panel for each of the paper's main results. Panels A and B report the effect of trade liberalization on the two dimensions of market structure that enter the extended model's markdown. In the baseline model, a 10% increase in import competition exposure raises the payroll Herfindahl by 0.021 points (Appendix Table A.4, column (4)). In the extended model, a 10% increase in ΔICE^* raises the expected payroll Herfindahl by 0.028 points and lowers the wage work share of market payroll by 0.032.[53] Trade thus increased concentration among formal employers in both models, and reduced the weight of wage work in local labor markets.

Panels C and D report the within-market cross-firm margin. With markets defined as microregions, $1/\eta$ is 0.870, estimated with the same wage premia and controls as the baseline, against 0.832 in the baseline (Table 2). The bias from ignoring the risk of involuntary separation out of the formal sector is positive and small: $\Omega = 0.0091$, so that $1/\tilde{\eta} = 0.879$. The threat of involuntary separation out of the formal sector therefore increases formal firms' market power rather than curbing it, by

[53] The two shocks differ in scale, which matters when I compare effects per 10% across models: ΔICE averages 13% across microregion × occupation markets (Section 4.1), and ΔICE^* averages 8.4% across microregions. Both are measured over 1991–2000. I weight the microregion regressions by the microregion's 1991 formal employment: at the microregion level nearly every market has a tradable firm, so the contrast between markets with and without one, which drives the unweighted microregion × occupation estimates, is absent, and the weighted estimates are the ones identified by the cross-sector tariff cuts (Appendix D); the baseline model re-estimated at the microregion in column (2) uses the same weights.

1% of the within-market cross-firm inverse elasticity.[54]

Panel E reports the two remaining margins, which I estimate jointly from equation 31. The inverse elasticity of substitution between wage work and self-employment is $1/\tilde{\rho} = 0.684$, below the within-market cross-firm value: self-employment is the most elastic margin in the extended model, consistent with evidence for Peru in Amodio, Medina and Morlacco (2025).[55] The cross-market inverse elasticity is $1/\tilde{\theta} = 1.187$ (SE of 0.481), against the baseline's 1.454 (SE of 0.196; Table 3); the two are within one standard error of each other, and the census estimate is less precise because it rests on 475 microregions rather than 17,628 microregion × occupation markets.

Panel F combines these estimates. Evaluated at 1991 averages, equation 24 implies a wage take-home share of 53.1 cents on the dollar in the extended model, the same as in the baseline. Applying equation 25, a 10% increase in the trade shock raises the average inverse elasticity of residual labor supply by 0.0053 in the extended model, a reduction of 0.15 cents in the take-home share. In the baseline, it raises the average inverse elasticity by 0.013, a reduction of 0.37 cents (Section 7). Of the two channels of equation 25, higher concentration raises the average inverse elasticity by 0.0056, and the shift of market payroll out of wage work lowers it by 0.0004 (Appendix Table A.13). The effect is less than half of the baseline's because concentration matters less for markdowns once workers can substitute into self-employment, the most elastic margin: the coefficient on concentration in equation 24 is smaller than the baseline's gap between $1/\theta$ and $1/\eta$. In the extended model, the standard error in Panel F holds the elasticities fixed; in the baseline, it also reflects their uncertainty.

I include column (2) as an intermediate step to help interpret the differences between columns (1) and (3), which are at different levels of aggregation: it re-estimates the baseline model with markets defined as microregions. Because I still estimate column (2) per the baseline model, its effect on concentration uses the baseline's trade shock, ΔICE, and payroll Herfindahl, whereas column (3) uses the extended model's shock, ΔICE^*, and expected payroll Herfindahl; the effects on concentration in the two columns (0.005 and 0.028) therefore differ even though both are at the microregion level. Column (2)'s effect on concentration is small in levels because microregions are much less concentrated than microregion × occupation markets: relative to the 1991 employment-weighted mean payroll Herfindahl of 0.006, it is an increase of about 80%, against 22% at the

[54] Unemployment insurance paid formally separated workers roughly four months of salary-based benefits during the 1990s. With four of the twelve months after separation valued at the formal wage, the wage gap is two thirds of its unadjusted value, and Ω is 0.0061.

[55] At my estimated substitution parameters, labor supply to every formal firm slopes upward in every market (equation 23 is positive for any shares), so markdowns are well defined, even though the ordering $1/\tilde{\rho} < 1/\tilde{\eta}$ departs from the one under which a nested GEV model is consistent with random utility maximization *at all wages* (McFadden, 1978). A less elastic within-market cross-firm margin is consistent with frictions that dampen substitution among formal employers, such as the inertia and inattention documented in other discrete choice settings (Cantillo, Ortuzar and Williams, 2007; Heiss et al., 2021). The ordering also implies that concentration raises markdowns only where the wage work share of market payroll exceeds 0.39 (equation 24). More broadly, nested logit estimates that violate consistency with random utility maximization at all wages can remain consistent with it over the range of wages and shares observed in the data (Börsch-Supan, 1990; Herriges and Kling, 1996; Gil-Moltó and Hole, 2004).

microregion × occupation level (Table 1). The effect on the take-home share in column (2) is small and not statistically distinguishable from zero (0.02 cents), because at the microregion level both the effect of ΔICE on concentration (0.005, against 0.021) and the gap between $1/\theta$ and $1/\eta$ (0.16, against 0.62) are smaller, and $1/\theta$ is less precisely estimated (1.034, SE of 0.250). The other parameters and effects are similar across columns: $1/\eta$ is between 0.832 and 0.879, the 1991 take-home share is 53 cents, and column (3)'s effect on concentration and cross-market inverse elasticity are close to column (1)'s. Column (3) is thus a close comparison to column (1), except that column (1) does not incorporate informal wage work and self-employment.

The average effect in Panel F masks heterogeneity across local labor markets that the baseline model cannot capture. In equation 25, trade affects markdowns through two channels: higher concentration, which raises markdowns where wage work exceeds 39% of market payroll, and a lower wage work share of the market, which lowers them, because the weight on the cross-market margin falls as employment shifts into self-employment. Evaluating both channels at the 1991 averages of each quartile of microregions by formal employment share (the quartiles of Table 6) shows how the two channels trade off. The take-home share is about 53 cents in every quartile, and ranges from 49 to 56 cents across individual microregions (Appendix Table A.14). The effect of trade liberalization is not uniform: a 10% increase in the trade shock lowers the take-home share by 0.16 cents in the most formal quartile, where concentration is low and wage work dominates, by 0.11 and 0.06 cents in the second and third, and leaves it unchanged (0.00 cents) in the least formal quartile, where the self-employment channel offsets the concentration channel (Appendix Table A.13). Figure 9 plots the effect microregion by microregion, each evaluated at the microregion's own trade shock rather than at a common 10% increase: it ranges from −0.36 to +0.32 cents, with a formal-wage-bill-weighted average of −0.23 cents (Appendix Figure A.12 shows the same microregions at a common 10% increase).

The two informal margins affect formal firms' market power in opposite directions. Self-employment is an elastic outside option. Informal wage work is not: the risk of involuntary separation out of the formal sector slightly raises formal firms' wage-setting power. I estimate the sign of Ω; the model does not impose it. When I instead treat informal wage work and self-employment symmetrically (Appendix F.4), the take-home share is 53.3 cents on the dollar, and a 10% increase in the trade shock lowers it by 0.37 cents (Appendix Table F.1): under this alternative, the level of markdowns and the effect of trade liberalization are those of the baseline.

9 Conclusion

How much of the wage losses from trade liberalization reflects employers' increased labor market power? In Brazil's formal sector that power is large: before liberalization, workers took home

53 cents of every marginal dollar they generated. Within a decade, a 10% increase in import competition exposure raised local labor market concentration by 22%, as fewer firms entered and exporters gained payroll share, an effect identified by differences in tariff cuts across sectors. Yet the take-home share fell by less than half a cent, because workers substitute across firms within a market nearly as inelastically as across markets. Wage premia at incumbent firms fell by 1.2%, of which higher markdowns account for one third and a lower marginal revenue product of labor for the rest. Across firms, a shift in the composition of employment toward firms with a higher marginal revenue product of labor offset most of that decline within firms. Accounting for informal wage work and self-employment halves the markdown effect, because self-employment is an elastic outside option. These findings matter for three reasons. First, an increase in concentration need not translate into a comparable increase in labor market power: what matters is the gap between workers' cross-market and within-market elasticities of substitution, and with the gap estimated for the US the same increase in concentration would have lowered the take-home share more than seven times as much. Second, formal firms' labor market power in Brazil stems from workers' inelastic substitution across employers, about nine times as inelastic as in the US, rather than from concentration. Third, the wage decline through 2000 was modest and came mostly from a lower marginal revenue product of labor, which, given that firm productivity rose (Muendler, 2004), more likely reflects lower output prices than productivity losses. Three questions merit further research: why workers substitute so inelastically across formal employers; how output prices, productivity, and technology each contributed to the decline in the marginal revenue product of labor; and how concentration and markdowns evolved after 2000, as the effects on earnings grew (Dix-Carneiro and Kovak, 2017), which requires estimating the key elasticities with post-liberalization shocks.

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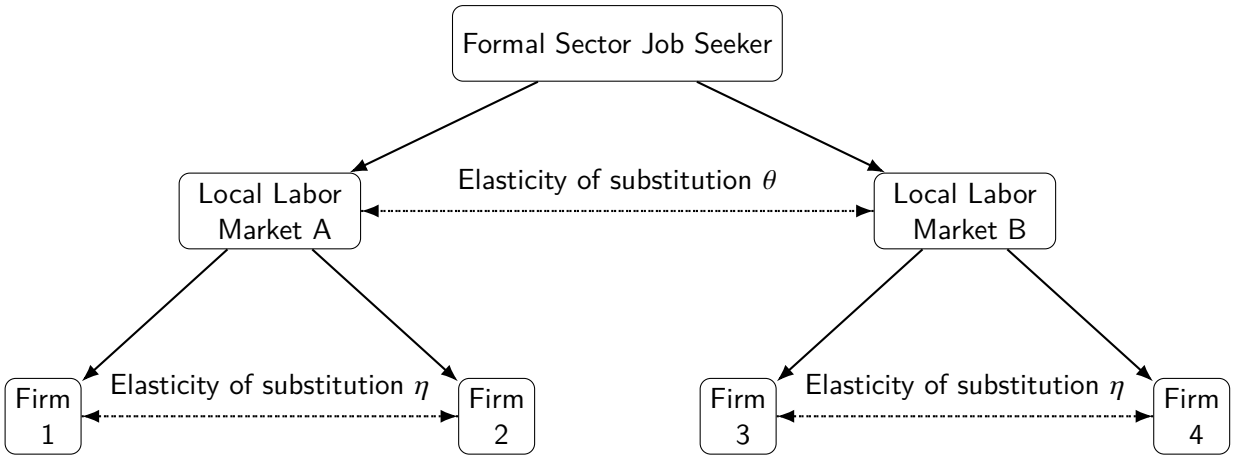
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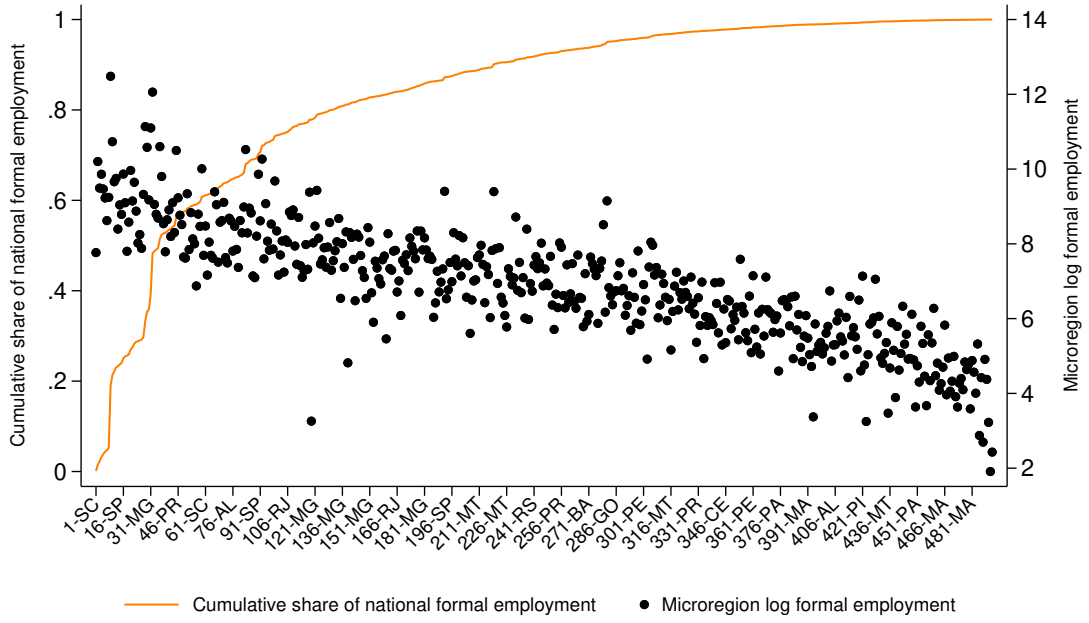
Figure 1: Discrete choice microfoundation of nested CES preferences over formal sector jobs



Notes: Workers' labor supply decision, microfounded in discrete choice (Section 2). In the empirical exercise, a local labor market is a microregion × 2-digit occupation pair (Appendix Figures A.1 through A.3). Section 8 and Appendix F extend this setup in two ways: 1) a within-market elasticity of substitution between wage work and self-employment; and 2) informal wage work as a potential outcome of involuntary separation from a formal sector job.

Figure 2: 1991 distribution of formal employment across microregions

(a) Cumulative distribution of formal employment, from most to least formal regions

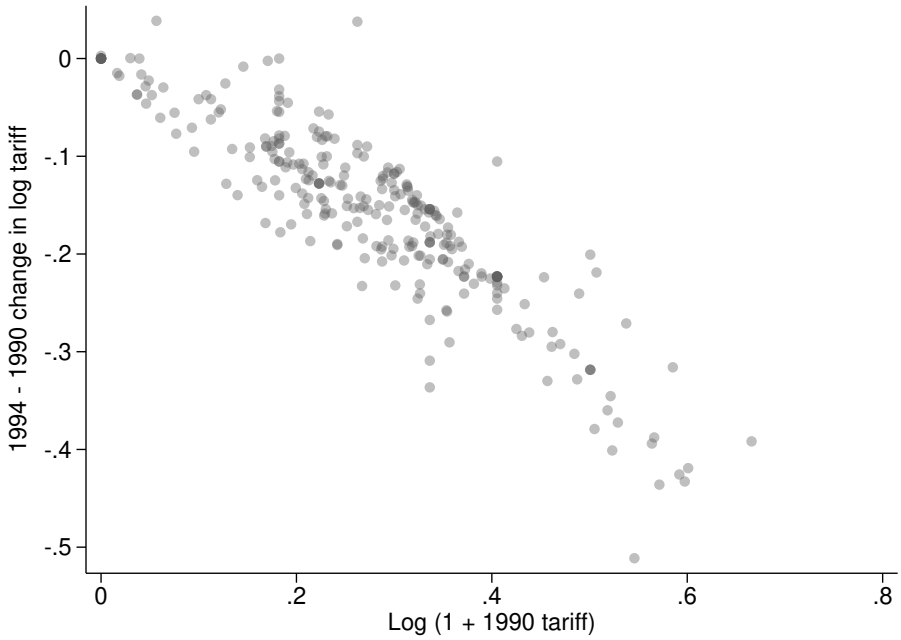


(b) Formal wage work, informal wage work, and self-employment



Notes: Distribution of formal employment across the 486 microregions of the 1991 Population Census. Panel (a): cumulative distribution of formal employment (red line, left axis) and log formal employment (black dots, right axis). Panel (b): shares of employment in formal wage work (black), informal wage work (gray), and self-employment (blue) on the left axis; log total employment (yellow dots) and log formal employment (black dots) on the right axis. Labels give the microregion's rank in the national sorting and its state (1-SC, the most formal microregion, is in Santa Catarina).

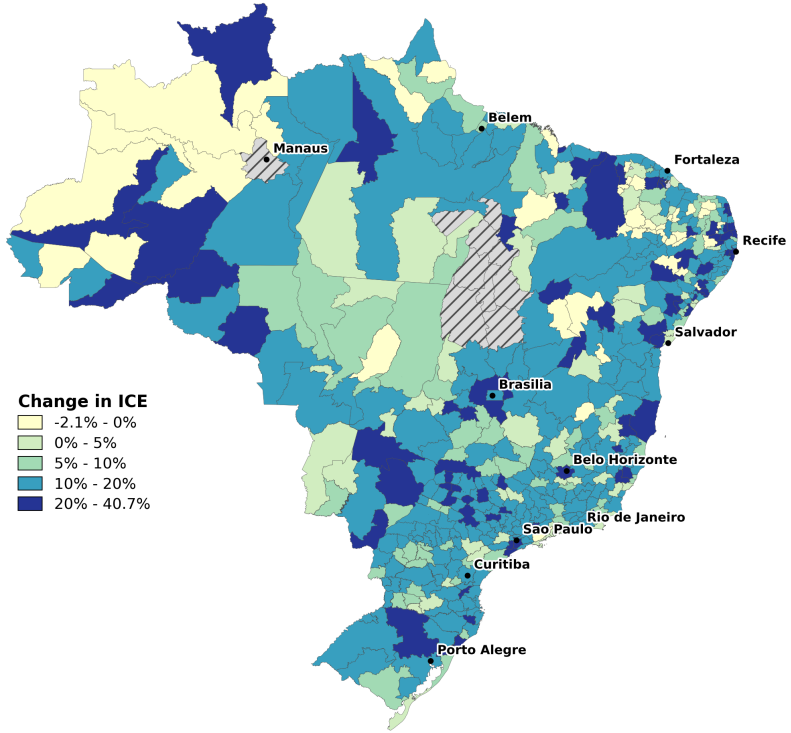
Figure 3: Brazil's 1990–1994 tariff reduction reform: Variation across 285 sectors



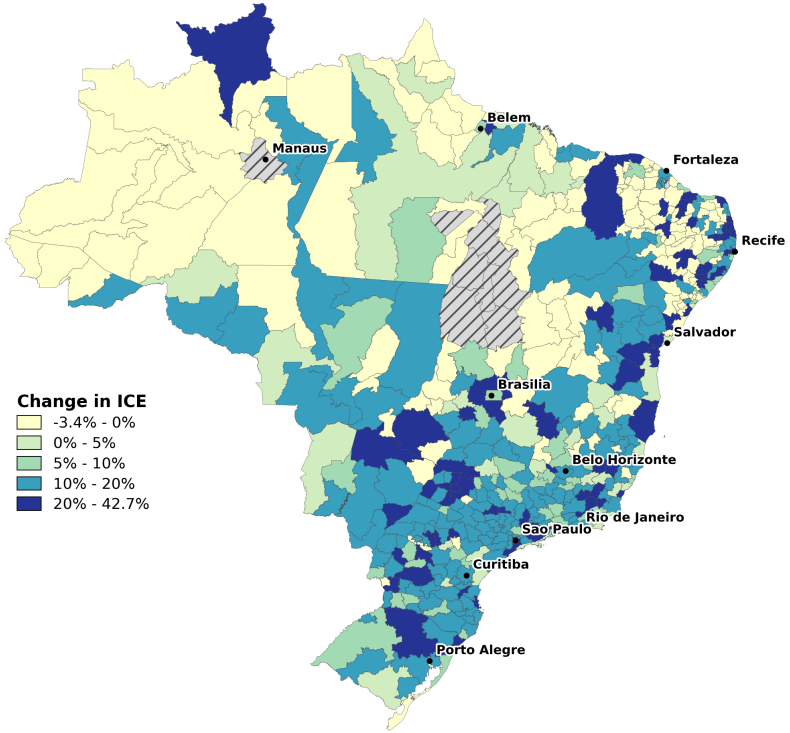
Notes: Import tariff reductions of Brazil's 1990–1994 reform across RAIS 5-digit sectors ("CNAE-95"): 285 tradable and 280 non-tradable sectors. Sector tariffs are simple averages of 6-digit product tariffs from UNCTAD TRAINS, mapped to CNAE-95 with IBGE's product-to-sector mappings (Section 3.3).

Figure 4: Variation in Change in Import Competition Exposure across local labor markets

(a) Office administration workers

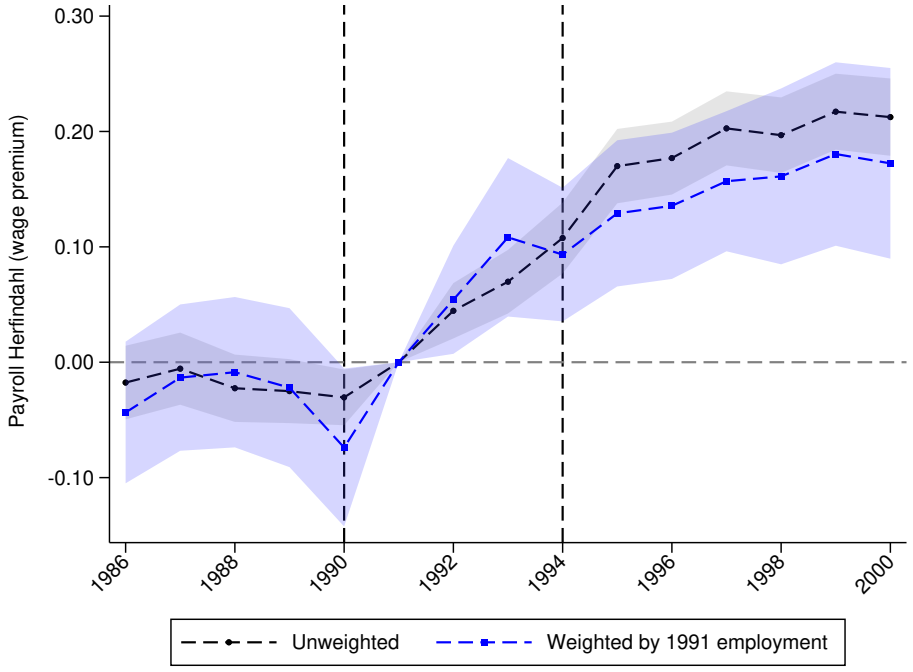


(b) Managers and supervisors of industrial workers



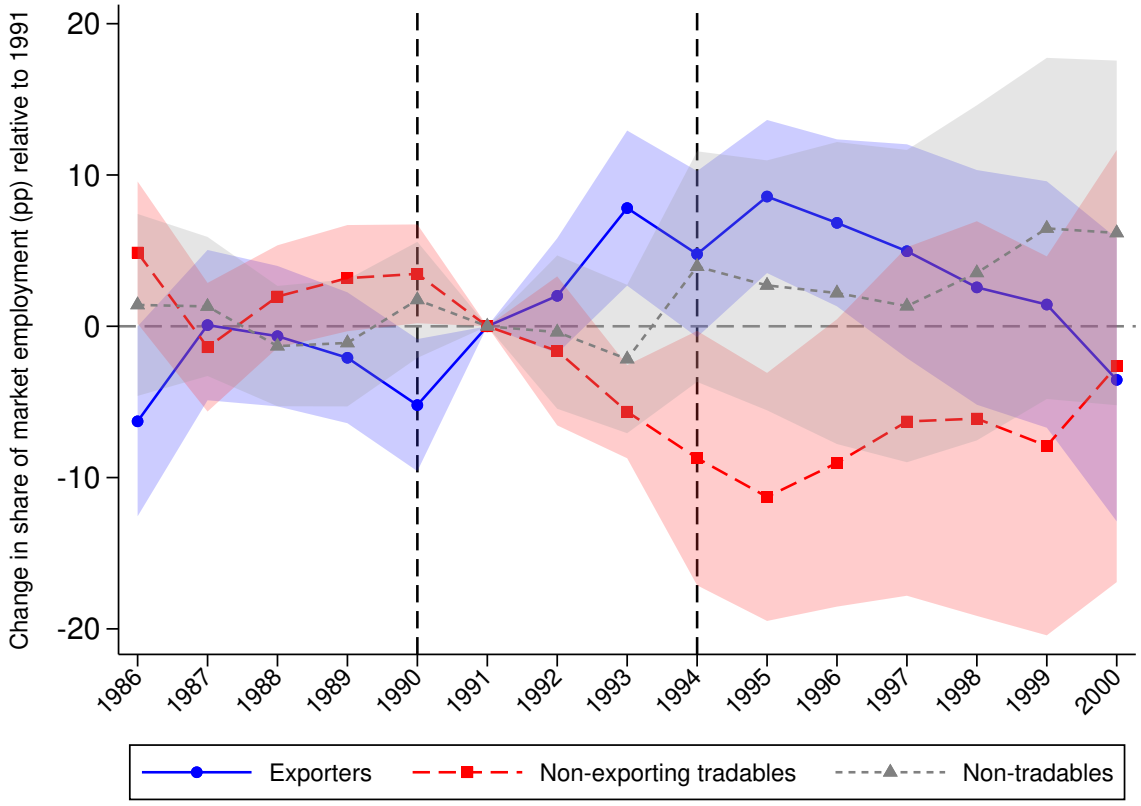
Notes: ΔICE_m, the change in import competition exposure, across local labor markets for two occupation groups. Microregion boundaries from Dix-Carneiro and Kovak (2017).

Figure 5: Effect of import competition on local labor market concentration



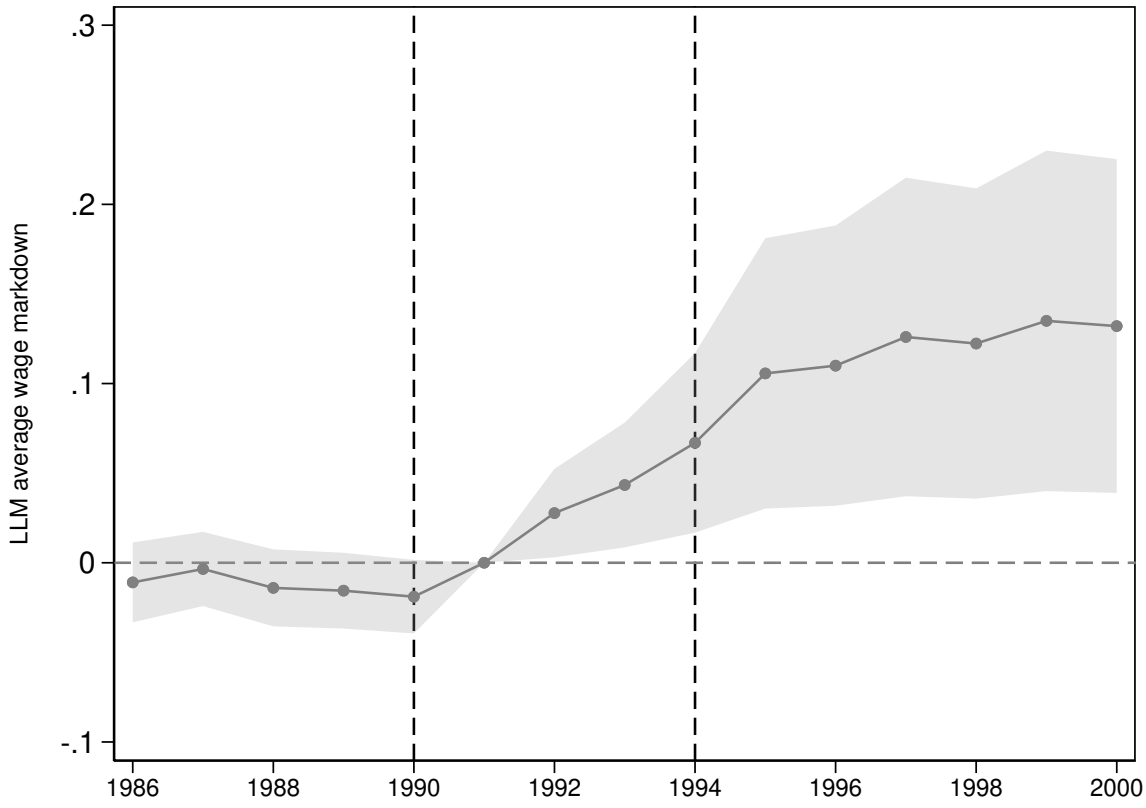
Notes: Coefficients ζ_k on ΔICE_m from equation 12; the outcome is the change in the premium-based payroll Herfindahl relative to 1991. A 10% increase in import competition exposure changes the outcome by $\zeta_k/10$ units. Shaded areas are 95% confidence intervals, with standard errors clustered by local labor market. The blue line is the main specification, weighted by local labor markets' 1991 formal employment (Table 1, column (1)); the black line is unweighted (Appendix Table A.4, column (4)).

Figure 6: Effect of import competition on the employment shares of non-exporting tradable firms, exporters, and non-tradable firms



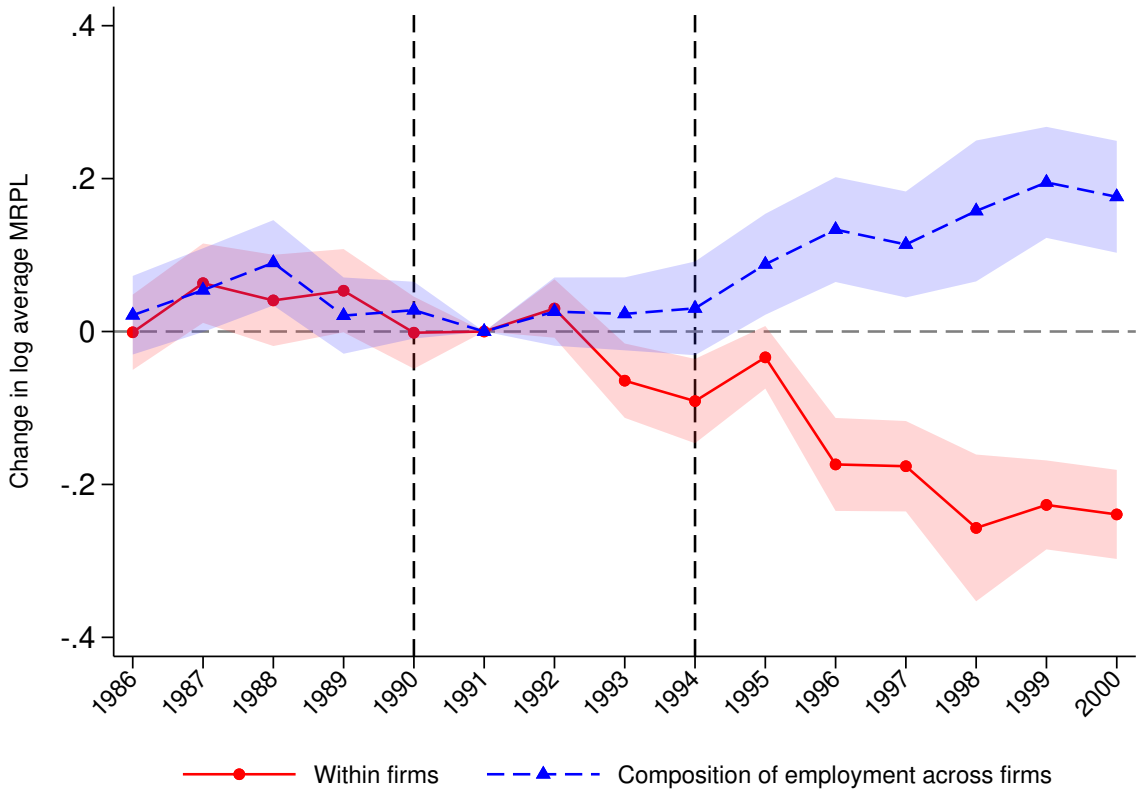
Note: This figure plots the ζ_k coefficients of equation 12 for three outcomes: the shares of the market's December employment in exporters, in non-exporting tradable firms, and in non-tradable firms, in percentage points. To keep the composition of the groups fixed, firms exporting at any time between 1991 and 1994 are tagged as exporters for the whole period. Dotted lines indicate the beginning and end of the tariff reductions. For pre-liberalization years the outcome is the 1991 share minus the year's share, and for post-reform years the year's share minus the 1991 share, so that all differences run from a past to a future year. Regressions include local labor market and year fixed effects, and weight local labor markets by their 1991 formal employment, as in Table 1, column (1); standard errors are clustered by local labor market. Shaded areas are 95% confidence intervals.

Figure 7: Effect of import competition on local labor market average wage markdowns



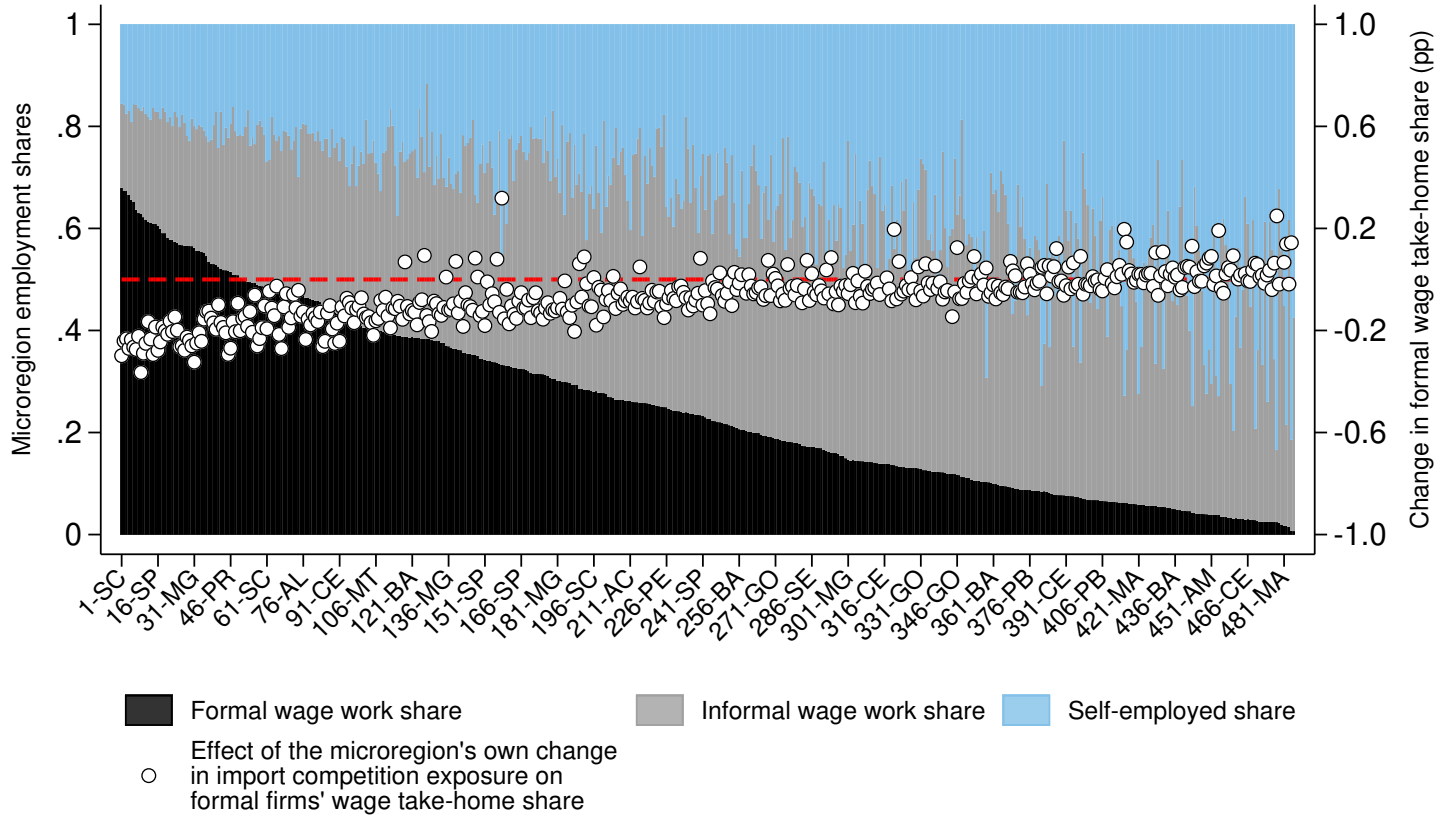
Notes: γ_t, the effect of import competition on local labor markets' average wage markdown in year t (equation 10): the product of $\left(\frac{1}{\theta} - \frac{1}{\eta}\right)$ (Table 3) and the unweighted β_t, the effect of import competition on concentration (Figure 5, black line). Standard errors assume that the two are independent (Appendix B).

Figure 8: Effect of import competition on the local labor market average marginal revenue product of labor: within firms and composition of employment across firms



Notes: Coefficients on ΔICE_m from regressions of the two terms of equation B.10 on ΔICE_m, one regression per year, with state fixed effects and the market's 1986–1990 changes in log employment and log wage premium, unweighted, as in Table 5. The within-firm term is the change in firms' marginal revenue product of labor at their 1991 employment shares; the composition term is the change in employment shares across firms, including entry and exit. Both are divided by the logarithmic mean of the market's average marginal revenue product in the year and in 1991, so that they add up to the change in its log. For years before 1991 the outcome is the 1991 value minus the year's value. Dashed lines mark the beginning and end of the tariff reductions. Shaded areas are 95% confidence intervals from heteroskedasticity-robust standard errors.

Figure 9: Effect of import competition on formal workers' wage take-home share (right y-axis), separately by microregion, based on the extended model, which incorporates self-employment and informal wage work into workers' labor supply decisions



Notes: Bars (left axis), as in Figure 2, Panel (b): the microregion's 1991 census shares of employment in formal wage work (black), informal wage work (gray), and self-employment (blue), for the 486 microregions of the census, sorted from the highest to the lowest formal wage work share; x-axis labels give the rank and the state. White circles (right axis), for the 478 microregions of the extended model: the change in the wage take-home share, in percentage points, implied by the microregion's own 1991–2000 trade shock ΔICE_m^* per equations 24 and 25, evaluated at the microregion's 1991 expected payroll Herfindahl and wage work share with the elasticities and reduced-form effects of Table 7, column (3); differences across microregions therefore reflect both their 1991 market structure and the size of their trade shock. The red dashed line marks zero.

Table 1: Effect of import competition on local labor market concentration, employment and wages

	Main specification	ΔICE recentered	State × year FE and 1986–1990 trends
	(1)	(2)	(3)
Panel A: Formal sector labor market concentration			
Δ Payroll Herfindahl (premium-based)	0.172	0.168	0.152
	(0.042)	(0.036)	(0.036)
1991 mean	0.079	0.079	0.079
Effect per 10%, percent of 1991 mean	22%	21%	19%
Δ Payroll Herfindahl	0.183	0.172	0.157
	(0.045)	(0.037)	(0.038)
1991 mean	0.082	0.082	0.082
Effect per 10%, percent of 1991 mean	22%	21%	19%
Δ Employment Herfindahl	0.109	0.110	0.110
	(0.024)	(0.027)	(0.023)
1991 mean	0.056	0.056	0.056
Effect per 10%, percent of 1991 mean	19%	20%	20%
Panel B: Number of firms and formal employment			
Δ Log number of firms	−0.971	−1.089	−0.681
	(0.188)	(0.219)	(0.129)
Δ Log formal employment	−0.398	−0.299	−0.194
	(0.186)	(0.214)	(0.145)
Panel C: Wages			
Δ Log average firm wage premium	−0.124	−0.126	−0.190
	(0.130)	(0.117)	(0.118)
Δ Log formal wage premium	−0.105	−0.119	−0.084
	(0.089)	(0.088)	(0.077)
Δ Log formal average wage	−0.075	−0.074	−0.154
	(0.110)	(0.110)	(0.098)
Weighted by 1991 formal employment	Yes	Yes	Yes
ΔICE recentered	No	Yes	No
State × year fixed effects and 1986–1990 trends	No	No	Yes
Observations	296,311	296,311	296,143
Local labor markets	19,757	19,757	19,745

Notes: Standard errors in parentheses, clustered by local labor market. Each cell reports the 2000 coefficient from a stacked differences-in-differences regression of long-differenced outcomes on import competition exposure (ICE) interacted with year dummies, with local labor market fixed effects. Column (1) is the main specification, with local labor markets (microregion × occupation) weighted by their 1991 formal employment. Column (2) subtracts from ΔICE_m its expected value when the 1990–1994 tariff changes are reassigned at random across CNAE sectors within IBGE subsectors (Borusyak and Hull, 2023), so that only differences in tariff reductions across detailed sectors identify the effect. Column (3) adds to the main specification state × year fixed effects and each market's 1986–1990 change in the outcome interacted with year, the controls of Dix-Carneiro and Kovak (2017). The premium-based Payroll Herfindahl is the sum of the squares of firms' shares of the market's premium-valued payroll, each firm's December employment times its December wage premium (Section 3); the Payroll Herfindahl values payroll at wages instead. In Panel A, the effect per 10% increase in ICE is $\zeta_{2000}/10$; the 1991 mean is that of the markets the column describes, and the effect per 10% is also expressed relative to it. In Panel C, the average firm wage premium is the employment-weighted mean of the market's firms' December wage premia; the wage premium is the market's December wage premium conditional on worker observables; the average wage is the mean of workers' log December earnings. Appendix Table A.4 reports the unweighted estimates and alternative shocks and samples; and Appendix Tables A.6 and A.5 alternative clustering and fixed effects.

Table 2: 2SLS estimates of workers' within-market cross-firm inverse elasticity of substitution $1/\eta$

	Main specification	Wage premia conditional on worker FE	Local labor market is microregion
	(1)	(2)	(3)
Panel A: First stage			
Δ Firm log employment in LLM	−0.736	−0.658	−0.803
	(0.076)	(0.126)	(0.056)
First stage F	94.915	27.124	207.063
Panel B: Reduced form			
Δ Firm wage premium in LLM	−0.612	−0.546	−0.699
	(0.036)	(0.071)	(0.021)
Panel C: 2SLS			
Labor supply within-market cross-firm	0.832	0.831	0.870
inverse elasticity of substitution	(0.093)	(0.173)	(0.062)
Implied upper bound on wage take-home share	55%	55%	53%
Local labor market (LLM) FE	Yes	Yes	Yes
Pre-trend controls (1986–91)	Yes	Yes	Yes
Observations	363,407	84,624	217,626
Firms	149,083	42,331	205,413
Local labor markets	11,632	6,628	466

Notes: Standard errors clustered by firm in parentheses. All regressions include local labor market (LLM) fixed effects and are weighted by baseline market employment. The instrument is the change in log import tariff faced by the firm between 1990 and 1994. Column (1): main specification with LLM defined as microregion × occupation, using the 1991–2000 change in the firm's December wage premium, estimated on all of its workers conditional on observables, on firm–market pairs present in 1986, with the pair's 1986–1991 changes in log employment and in the log wage premium as controls. Column (2): wage premium conditional on worker fixed effects and demographic-by-year controls, estimated on the firm's stayers, workers in the same firm, microregion, and occupation in both years. Column (3): defines LLM as microregion only (no occupation dimension). Appendix Table E.2 uses the effective rate of protection as the instrument instead of the import tariff.

Table 3: 2SLS estimate of workers' cross-market inverse elasticity of substitution $1/\theta$

	Main specification	Wage premia conditional on worker FE	Local labor market is microregion
	(1)	(2)	(3)
Panel A: First stage			
Δ LLM employment index	−0.294	−0.258	−1.157
	(0.044)	(0.048)	(0.290)
First stage F	45.6	28.6	15.9
Panel B: Reduced form			
Δ LLM wage premium index	−0.183	−0.078	−0.190
	(0.050)	(0.065)	(0.275)
Panel C: 2SLS			
$1/\theta - 1/\eta$	0.622	0.304	0.164
	(0.217)	(0.273)	(0.258)
Panel D: Cross-market inverse elasticity of substitution			
$1/\theta$	1.454	1.135	1.034
	(0.196)	(0.211)	(0.250)
Implied lower bound on wage take-home share	41%	47%	49%
State fixed effects and 1986–1990 trends	Yes	Yes	No
Local labor markets	17,628	11,455	475

Notes: Standard errors in parentheses, clustered by local labor market. First stage, reduced form, and two-stage least squares estimates of $\frac{1}{\theta} - \frac{1}{\eta}$, and the implied $\frac{1}{\theta}$, from equations 19 and 20. The implied lower bound on the wage take-home share is $\left(1 + \frac{1}{\theta}\right)^{-1}$, equation 9 with one firm per local labor market ($HHI_m = 1$). Column (2) uses wage premia conditional on worker fixed effects and the $\frac{1}{\eta}$ of Table 2, column (2); column (3) defines local labor markets as microregions, with the shock ΔICE^* of Section 8 and the $\frac{1}{\eta}$ of Table 2, column (3), without state fixed effects and 1986–1990 trends (with them, $\frac{1}{\theta}$ is 0.318, SE of 0.209, with a first-stage F statistic of 4.5).

Table 4: Effect of import competition on the average wage take-home share

	Effect per 10% increase in ICE
Effect of Δ Import Competition Exposure on market average wage markdown	0.0132 (0.0048)
β Effect of Δ Import Competition Exposure on payroll Herfindahl	0.021 (0.002)
$\frac{1}{\theta} - \frac{1}{\eta}$ Difference between key inverse elasticities of labor supply	0.622 (0.217)
Effect of Δ Import Competition Exposure on market average wage take-home share	-0.0037 (0.0013)
Local labor markets	19,757

Notes: Estimates of γ_{2000} per equation 10 and its two components: β_{2000} from Appendix Table A.4, column (4), with standard errors clustered by local labor market, and $(1/\theta - 1/\eta)$ from Table 3. Standard errors assume that the two are independent (Appendix B). The effect on the take-home share is $-\gamma_{2000}$ times the square of the 1991 country-level take-home share of 0.531 (Section 6.3).

Table 5: Effect of import competition on incumbent firms' wage premia

	Pre-reform level (R$ of Dec. 2000) (1)	Percent effect of a 10% increase in ICE (2)	Effect in R$ of Dec. 2000 at the 1991 level (3)	Percent of the total effect on incumbents' wage (4)
Average wage premium of incumbent firms	317.85	−1.15	−3.64	100%
		(0.44)	(1.40)	
Average wage take-home share	0.51	−0.38	−1.20	33%
		(0.15)	(0.49)	
Average marginal revenue product of labor	642.00	−0.77	−2.44	67%
		(0.39)	(1.24)	

Notes: Standard errors in parentheses, from a microregion cluster bootstrap with 200 replications, each of which re-estimates $1/\eta$, $1/\theta$, and every object of the table. Rows are means over the firms present in the market in both 1991 and 2000, weighted by the firm's 1991 employment: the firm's December wage premium, its take-home share, which follows from its payroll share and the estimated elasticities (equation 8), and its marginal revenue product of labor, the wage premium divided by the take-home share. Column (1) is the unweighted mean across markets of the 1991 level (wage premium and marginal revenue product in reais of December 2000). Column (2) is the coefficient on ΔICE_m in a regression of the 1991–2000 change in the row's object on ΔICE_m, with state fixed effects and the market's 1986–1990 changes in log employment and log wage premium, unweighted, as in Table 3, in percent per 10% increase in ICE. Column (3) expresses column (2) in reais of December 2000 at the 1991 level of column (1). Column (4) is each component's share of the effect on incumbents' average wage premium, which is exact because the identity of equation B.9 holds in logs.

Table 6: Heterogeneous effects of import competition on concentration and cross-firm substitution by microregion baseline informality

	All markets	Quartile 1 (most formal)	Quartile 2	Quartile 3	Quartile 4 (least formal)
	(1)	(2)	(3)	(4)	(5)
Share of national formal employment, 1991	0.995	0.726	0.171	0.073	0.026
Median formal share of employment, 1991	0.265	0.502	0.335	0.205	0.102
Panel A: Effect of 10% increase in import competition on concentration					
Δ Payroll Herfindahl (premium-based)	0.017	0.015	0.015	0.017	0.048
	(0.004)	(0.005)	(0.006)	(0.005)	(0.017)
Mean (unweighted), 1991	0.287	0.212	0.285	0.339	0.426
Mean (weighted), 1991	0.079	0.064	0.129	0.192	0.289
Observations	296,311	108,396	87,903	62,792	37,220
Local labor markets	19,757	7,227	5,861	4,187	2,482
Panel B: Within-market cross-firm inverse elasticity $1/\eta$ (2SLS)					
$1/\eta$	0.832	0.824	-1.064	0.315	0.014
	(0.093)	(0.091)	(0.733)	(0.339)	(0.286)
First-stage F	94.9	97.1	2.5	1.8	2.5
Implied upper bound on take-home share	0.546	0.548	–	–	–
Observations	363,407	287,985	54,414	17,342	3,666
Firms	149,083	115,676	24,988	8,067	1,830
Local labor markets	11,632	5,361	3,518	1,975	778

Notes: Standard errors in parentheses, clustered by local labor market in Panel A and by firm in Panel B. Column (1) pools all local labor markets (microregion × occupation); columns (2)–(5) split them by quartile of the microregion's formal share of total employment in the 1991 census. The first two rows report each group's share of national formal employment and the median formal share of employment across its microregions. Panel A: effect of a 10% increase in import competition exposure on the premium-based payroll Herfindahl as of 2000 (equation 12), with local labor markets weighted by their 1991 formal employment, as in Table 1; the 1991 means are across the group's markets, unweighted and weighted by 1991 market employment. Panel B: 2SLS estimates of $1/\eta$ (equations 15 and 16). The implied upper bound on the take-home share is $1/(1+1/\eta)$ and prints "–" when $1/\eta$ is not statistically significant at the 5% level.

Table 7: Effect of import competition on formal sector labor market concentration and wage markdowns: Baseline model vs. Extended model with informality

	Baseline model (RAIS Data)		Extended model (RAIS and Census Data)
	Microregion × occupation	Microregion	Microregion
	(1)	(2)	(3)
Panel A: Effect of import competition on formal labor market concentration			
Δ Concentration, per 10% increase in the trade shock	0.021	0.005	0.028
	(0.002)	(0.002)	(0.007)
Trade shock	ΔICE	ΔICE	ΔICE^*
Concentration measure	Payroll Herf.	Payroll Herf.	*Expected* Payroll Herf.
Level of analysis	Mmc × occ.	Mmc	Mmc
Panel B: Effect of import competition on the wage work share of market payroll			
Δ Sector share, per 10% increase in the trade shock	–	–	-0.032
			(0.005)
Share	–	–	Wage work
Trade shock	–	–	ΔICE^*
Level of analysis	–	–	Mmc
Panel C: Within-market cross-firm inverse elasticity of substitution			
$1/\eta$ (response to formal wages)	0.832	0.870	0.870
	(0.093)	(0.062)	(0.062)
Trade shock		Firm-level tariff reduction	
Level of analysis	Firm × mmc × occ.	Firm × mmc	Firm × mmc
Wage measure		Firm Dec. premium (all workers), pair pre-trend controls	
Panel D: Bias Ω in 2SLS estimates of $1/\eta$ due to risk of separation from formal job			
Ω (response to expected formal–informal wage gaps)	–	–	0.0091
			(0.0025)
$1/\tilde{\eta} = 1/\eta + \Omega$	–	–	0.879
			(0.062)
Ω, with unemployment insurance	–	–	0.0061
			(0.0017)
Trade shock	–	–	Firm tariff × ΔICE^*
Level of analysis	–	–	Firm × mmc × demog. cells
Fixed effects	–	–	Firm, mmc, demographic cell
Panel E: Cross-sector and cross-market inverse elasticities of substitution			
Within-market cross-sector inverse elasticity	–	–	0.684
			(0.234)
			$1/\tilde{\rho}$ (wage work vs. self-emp.)
Cross-market inverse elasticity	1.454	1.034	1.187
	(0.196)	(0.250)	(0.481)
	$1/\theta$	$1/\theta$	$1/\tilde{\theta}$
Trade shock	ΔICE	ΔICE^*	ΔICE^*
Fixed effects	State, 1986–90 trends	–	Region
Level of analysis	Mmc × occ.	Mmc	Mmc × demog. cells
Wage measure	Nested CES Index	Nested CES Index	Residual log earnings, wage workers
Panel F: Effect of import competition on the average wage take-home share			
Δ average inverse elasticity, per 10% increase in the trade shock	0.0132	0.0008	0.0053
	(0.0048)	(0.0013)	(0.0014)
Implied change in take-home share (pp)	-0.37	-0.02	-0.15
Baseline (1991) wage take-home share	0.531	0.535	0.531
Trade shock	ΔICE	ΔICE	ΔICE^*

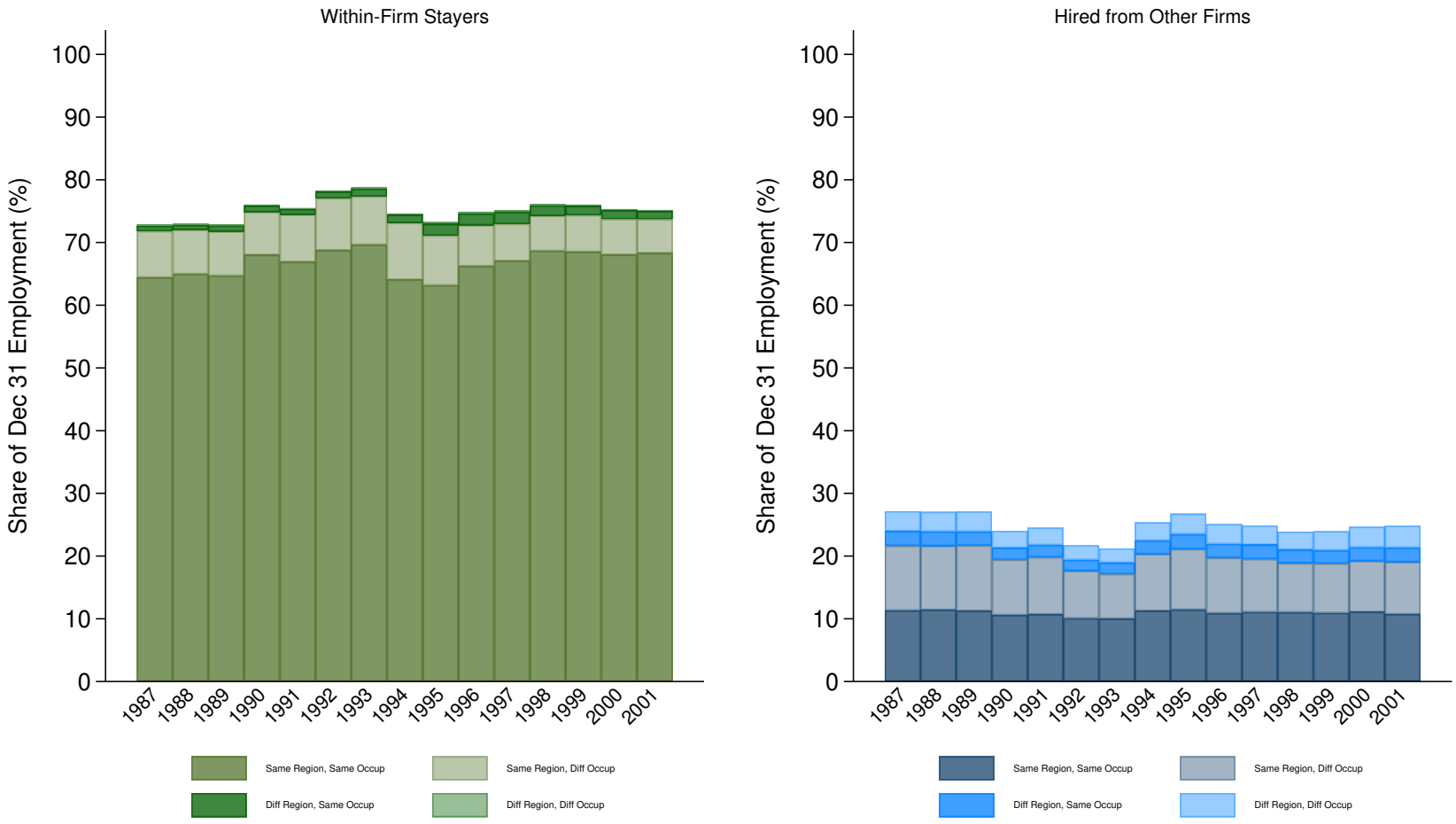
Notes: Standard errors in parentheses. Column (1) reproduces the paper's main estimates: Panel A from Appendix Table A.4, column (4), Panel C from Table 2, Panel E from Table 3, and Panel F from Table 4 and Section 7. Column (2) re-estimates the baseline model with local labor markets ("Mmc") defined as microregions, so that the extended model's geography is compared like for like; its $1/\theta$ uses ΔICE^* as the instrument, the only one with a first stage at the microregion level (Appendix F.3.6). Column (3) reports the extended model of Section 8, in which local labor markets are microregions and the trade shock is ΔICE^* (equation 27); Panels B and D exist only in the extended model. Panel D's unemployment insurance row values four of the twelve months after separation at the formal wage. Panel F applies Corollary 1 in columns (1)–(2) and equations 24 and 25 in column (3), at 1991 averages. The microregion regressions of Panels A–B and the 1991 means of Panel F in columns (2)–(3) are weighted by the microregion's 1991 formal employment (Section 8.4, footnote, and Appendix D). Cells marked "–" have no counterpart in that model. Appendix F.3.6 gives the specification, instruments, weights, clustering, and first-stage and overidentification statistics behind each panel.

Online Appendix for

"Trade, Labor Market Concentration, and Wages,"

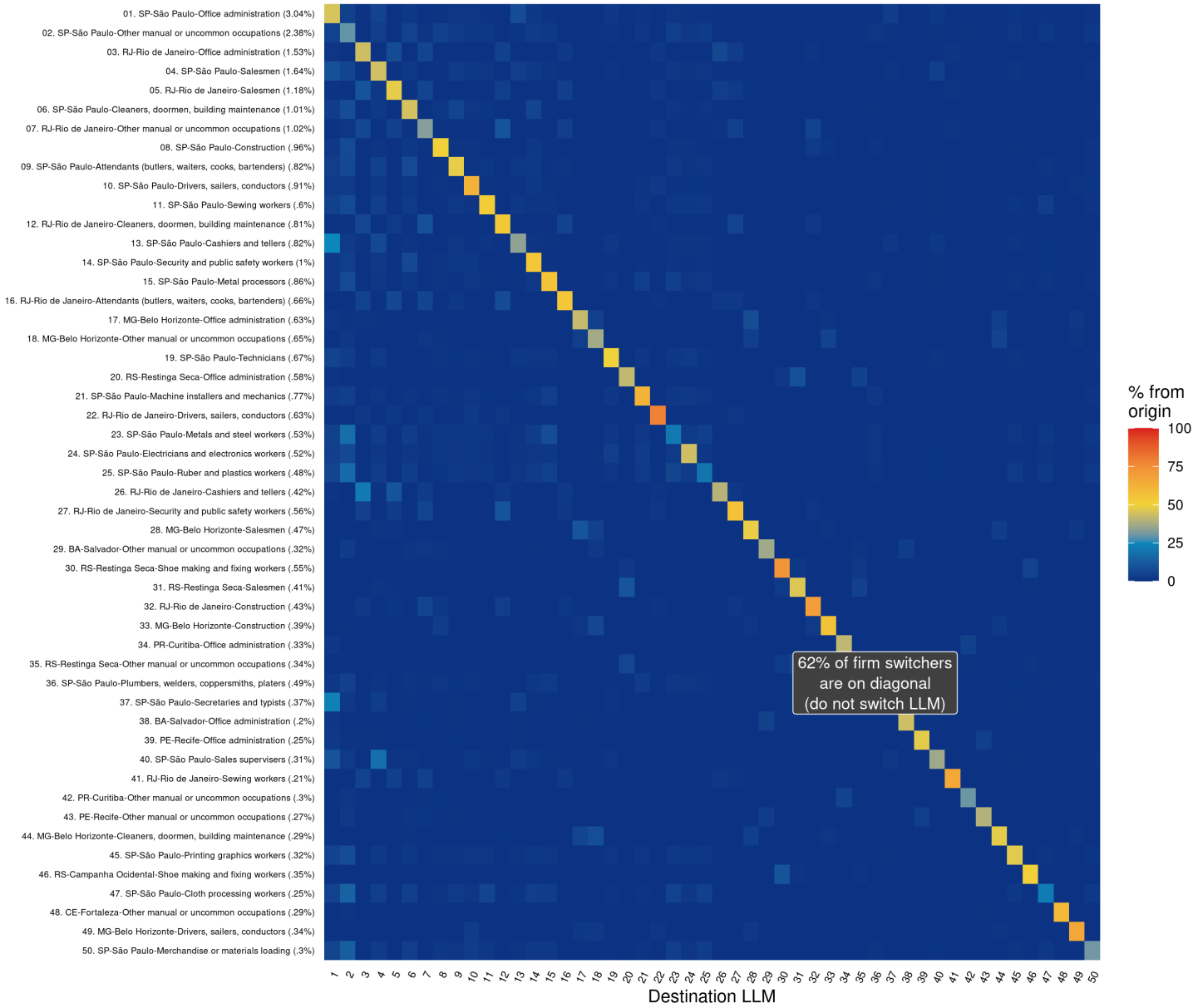
by Mayara Felix

Figure A.1: Region and occupation origins of firm stayers and firm switchers as of Dec 31 of each year



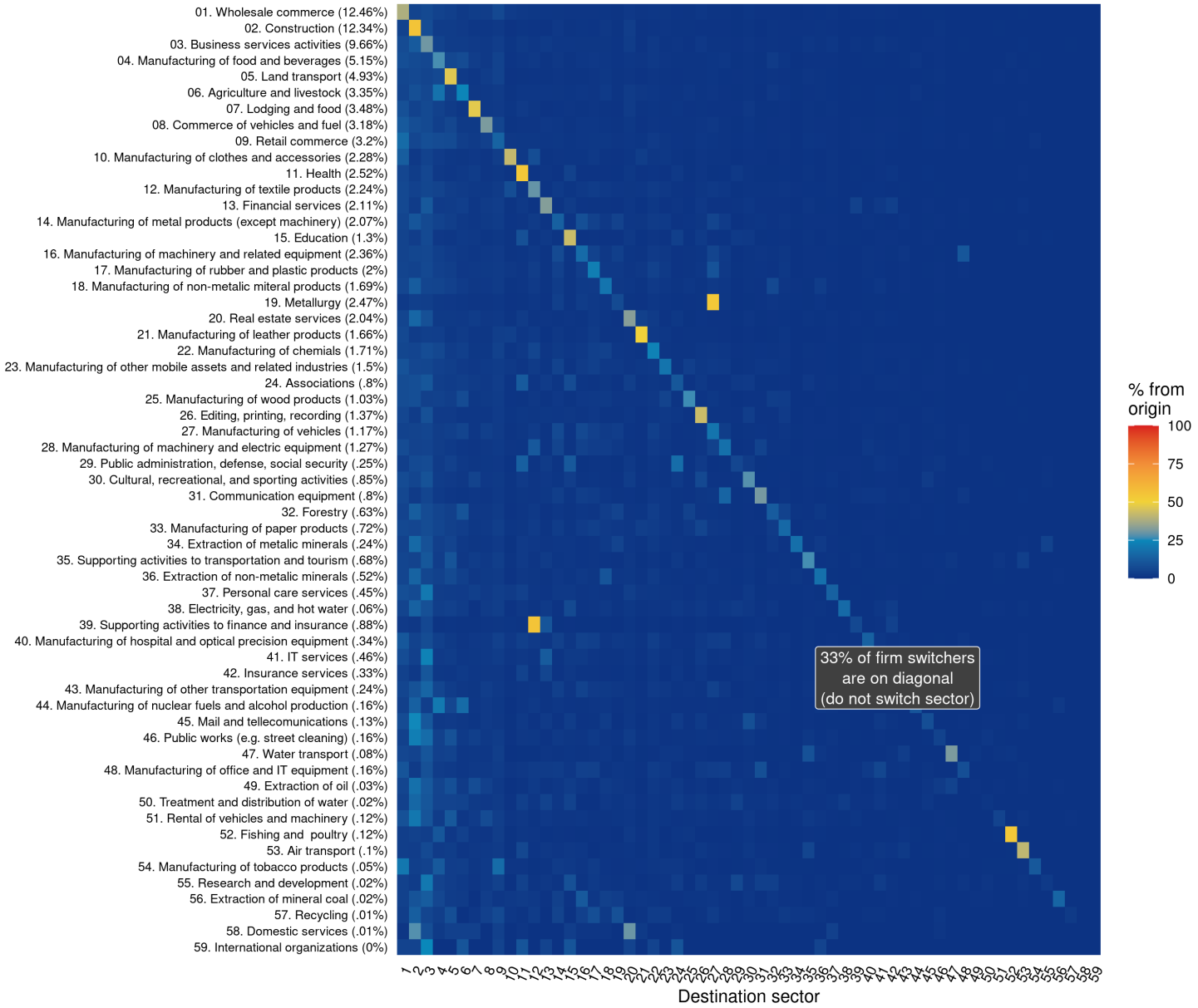
Notes: Composition of formal private sector employment in RAIS, 1986–2000. Each bar reports the December 31 employment shares of firm stayers (workers employed at the same firm on December 31 of the previous year) and of firm switchers, by region of origin in one panel and by occupation of origin in the other. Flows exclude retirements, deaths, and transitions to and from public administration.

Figure A.2: 1990-1991 local labor market transitions conditional on switching firms (Top 50)



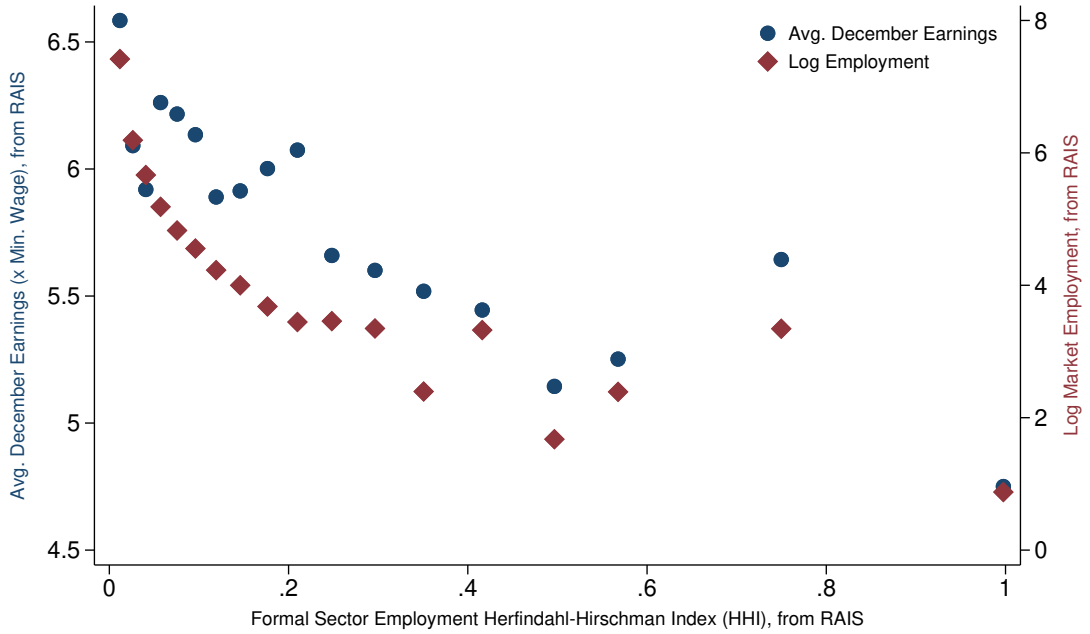
Note: This figure plots worker local labor market to local labor market transitions, among workers who switched employers between 1990 and 1991, for the top 50 local labor markets by number of workers at origin. A local labor market is a microregion × occupational group pair. Each row lists the origin microregion (with percent of total workers indicated in parentheses), while each column lists the destination microregion.

Figure A.3: 1990-1991 sector transitions conditional on switching firms (Top 50)



Note: This figure plots worker sector group to sector group transitions, among workers who switched employers between 1990 and 1991, for the top 50 sector (2-digit CNAE95) groups by number of workers at origin. Each row lists the origin sector group (with percent of total workers indicated in parentheses), while each column lists the destination sector group.

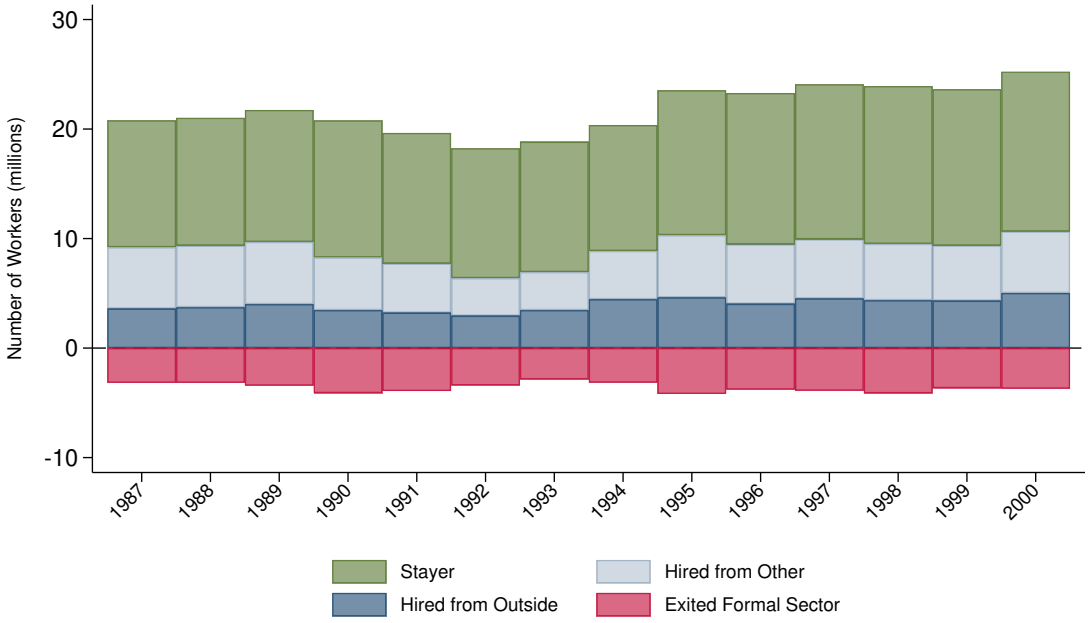
Figure A.4: 1991 wages, employment, and formal sector concentration



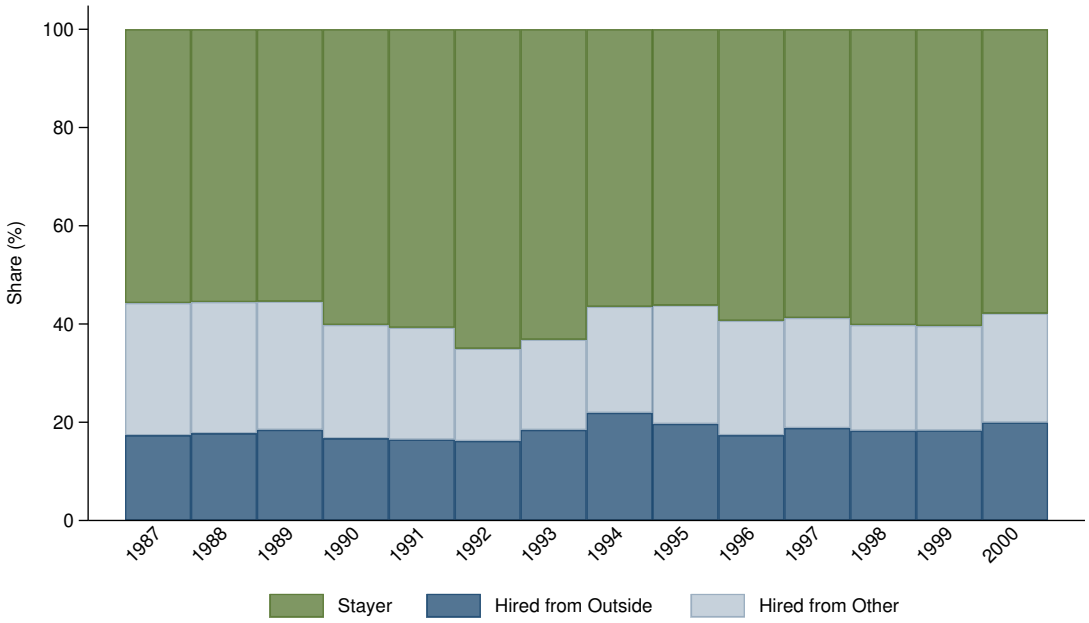
Notes: This figure plots wages and employment as a function of formal sector HHI in microregion x occupation pairs using RAIS data for 1991.

Figure A.5: Formal sector worker flows

(a) Total firm-to-firm and firm-to-outside flows



(b) Flow shares

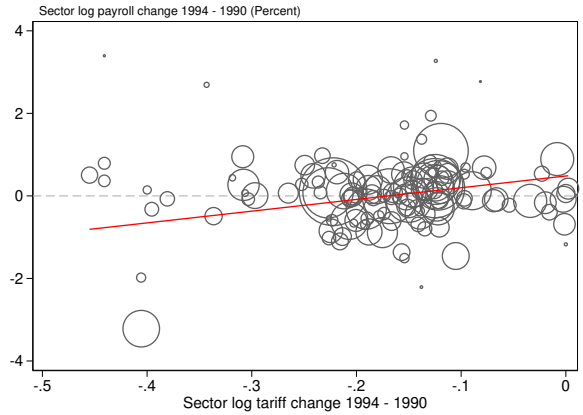


Notes: Worker flows as reported by formal sector firms in RAIS: hires directly from other formal firms ("Hired from Other"), separations followed within one year by a match to another formal firm ("Sep to Other"), hires from outside the formal sector ("Hired from Outside"), separations to outside the formal sector ("Sep to Outside"), and workers who do not switch employers within the year ("Stayer"). "Hired from Other" and "Sep to Other" are measured separately, the former from the records of the receiving firm and the latter from those of the origin firm. Flows include within-year transitions and exclude flows to and from public administration, retirements, and deaths.

Figure A.6: Changes in sector-level outcomes Before vs. After liberalization

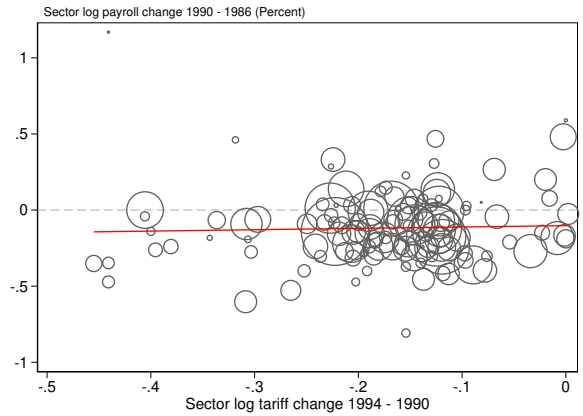
(a) After liberalization (1990–1994)





(b) Before liberalization (1986–1990)

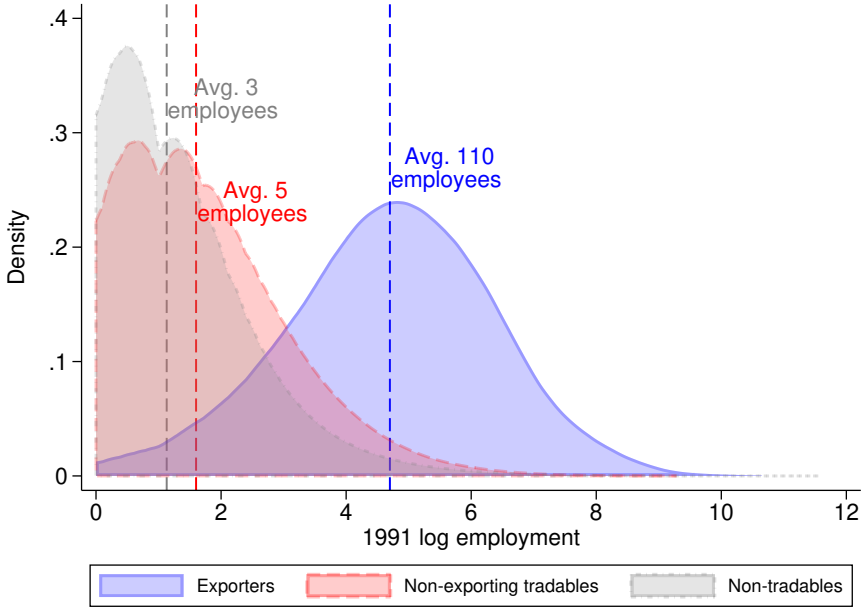




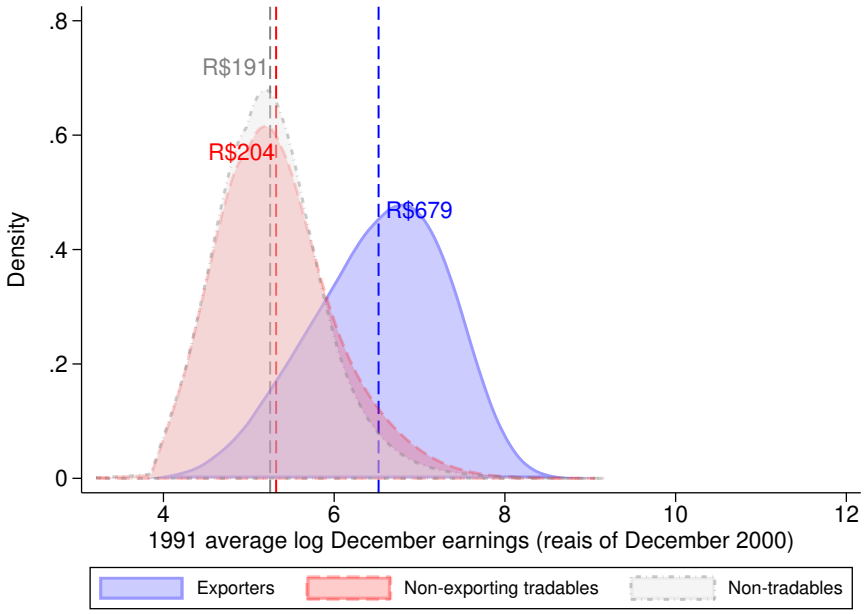
Notes: Changes in sector log employment and log payroll in the four years before and after liberalization against the 1990–1994 import tariff reduction on the sector's output. Sector totals are for the entire country, in 148 tradable sectors of RAIS's 4-digit "ibgesubatividade" variable, consistently reported throughout the period.

Figure A.7: Pre-liberalization distributions of firm size and wages by firm type

(a) Firm log employment



(b) Firm average log December earnings

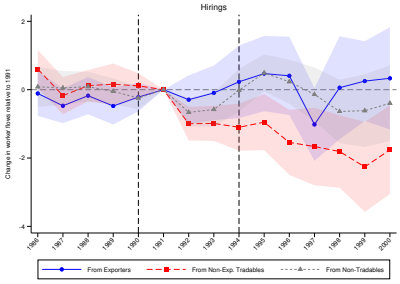


Notes: 1991 distributions of firm log employment and of firm average log December earnings (reais of December 2000) for exporters, non-exporting tradable firms, and non-tradable firms. Dashed lines mark each group's mean, annotated in levels. Exporters are firms that exported in 1991; tradable firms are in sectors with an import tariff.

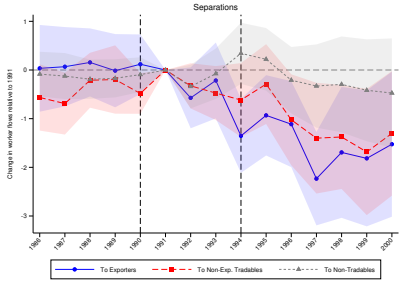
Figure A.8: Effect of import competition on worker flows into and out of non-exporting tradable firms and exporters

Within-market flows

Panel A: Hirings into non-exporting tradable firms

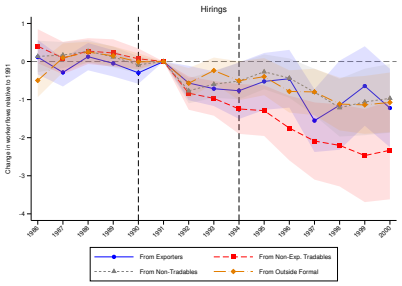


Panel B: Separations from non-exporting tradable firms

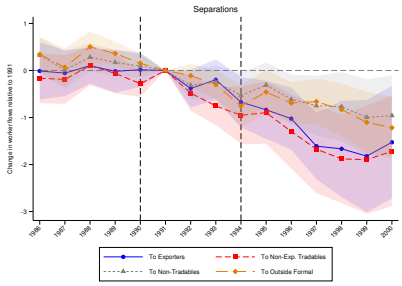


All flows

Panel A: Hirings into non-exporting tradable firms

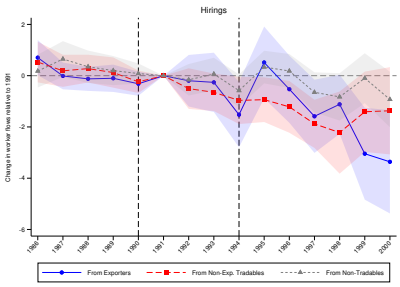


Panel B: Separations from non-exporting tradable firms

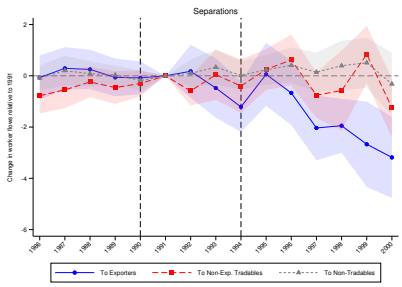


Within-market flows into and out of exporters

Panel C: Hirings into exporters



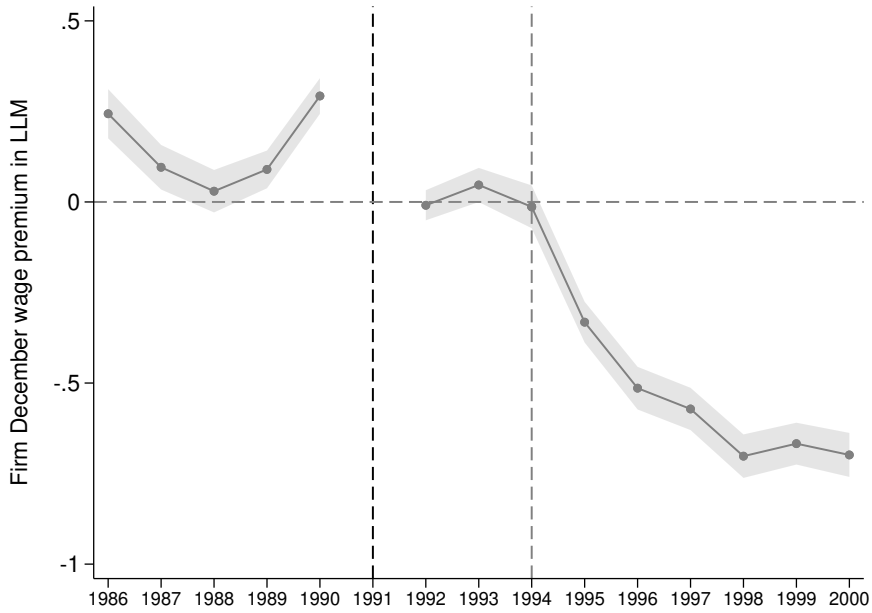
Panel D: Separations from exporters



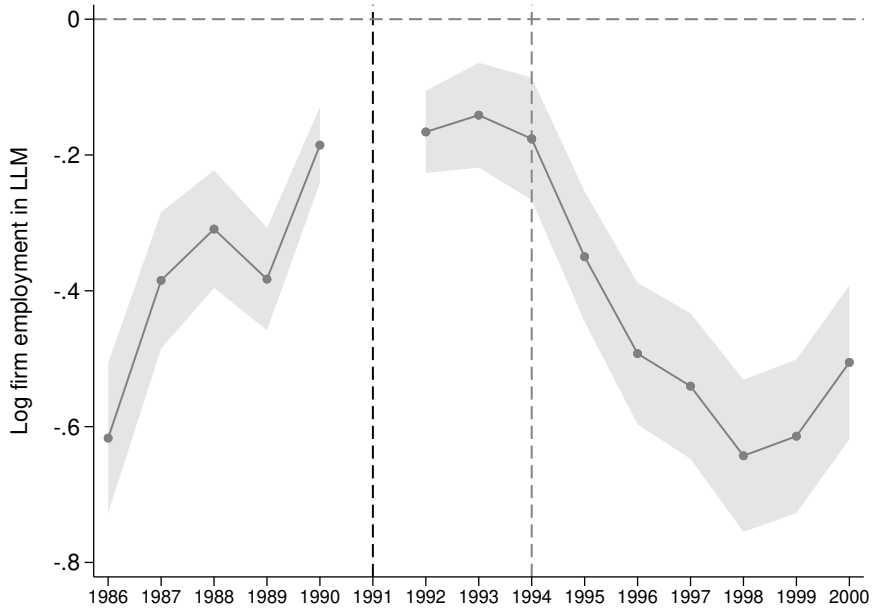
Note: This figure plots the ζ_k coefficients of equation 12 for worker flows, with local labor market (microregion × occupation) and year fixed effects and markets weighted by their 1991 formal employment, as in Table 1, column (1); standard errors are clustered by local labor market. Panel A shows hirings into non-exporting tradable firms from exporters, from other non-exporting tradable firms, and from non-tradable firms; Panel B shows separations from non-exporting tradable firms to each of these firm types. The top row restricts to flows within the same local labor market; the middle row includes all flows, including flows into and out of the formal sector. Panels C and D show within-market hirings into and separations from exporters, by the type of firm on the other side of the flow. Outcomes are inverse-hyperbolic-sine transformed counts, in long differences relative to 1991. Shaded areas are 95% confidence intervals.

Figure A.9: Year by year effects of firm-level import tariff reductions on changes in firms' local labor market wage premia and employment

(a) Wage premia

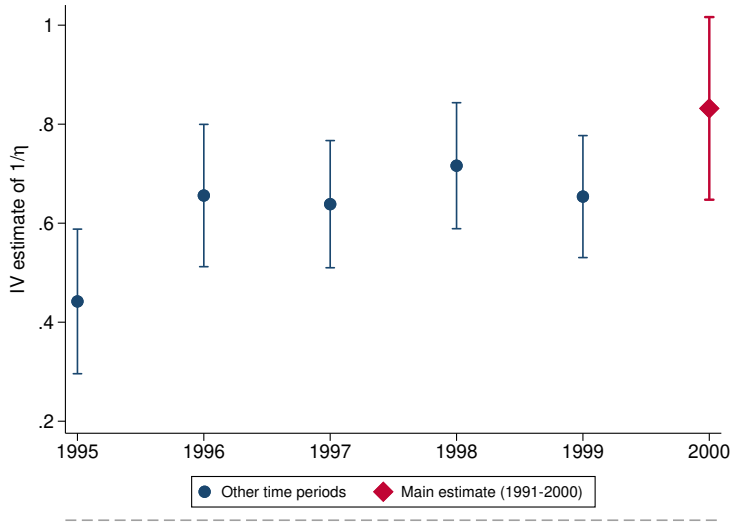


(b) Employment



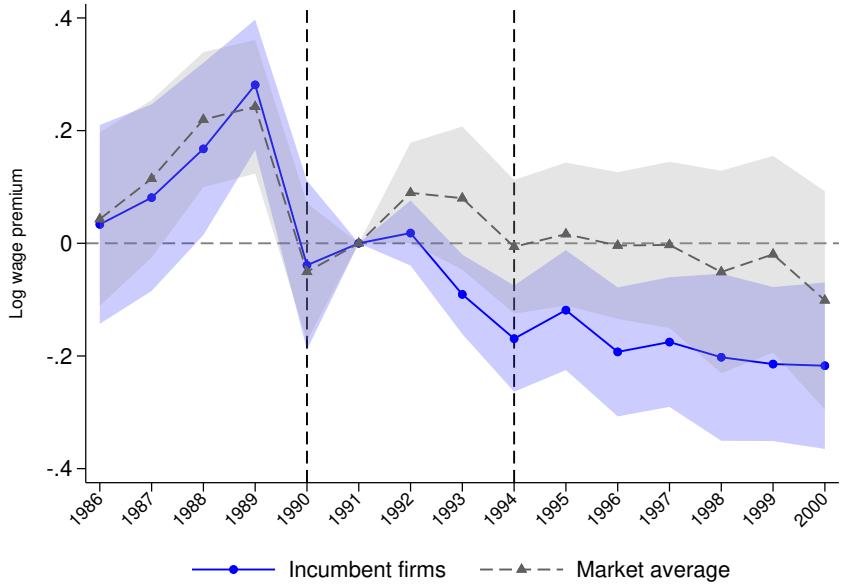
Notes: Coefficients from year-by-year regressions of the change in the firm's wage premium (Panel a) and in its log employment (Panel b) relative to 1991 on $-\ln\left(\frac{1+\tau_{1994}}{1+\tau_{1990}}\right)$, the firm's import tariff reduction. For years before 1991 the outcome is the 1991 value minus the year's value. Regressions include local labor market fixed effects and use the weights of Table 2; standard errors are clustered by firm. Dotted lines mark the beginning and end of the tariff reductions.

Figure A.10: Within-market cross-firm inverse elasticity of substitution $1/\eta$ by end year



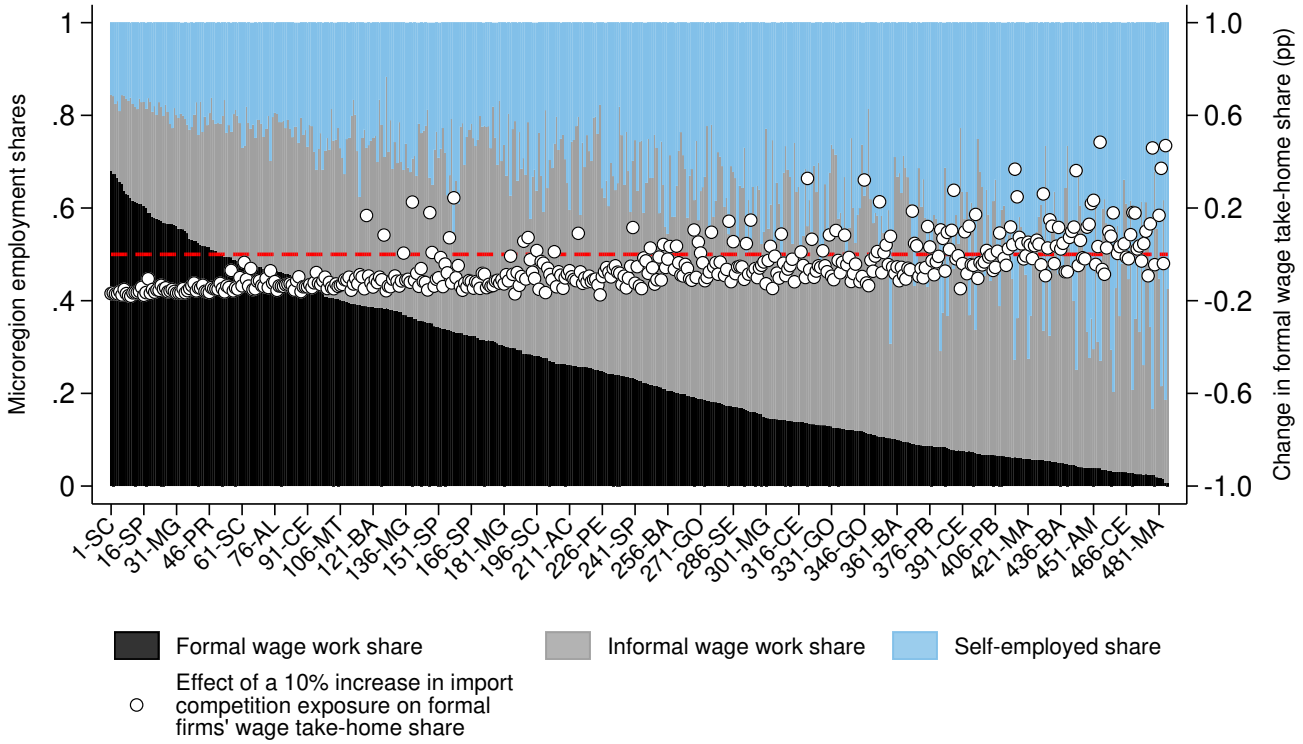
Notes: 2SLS estimates of $1/\eta$ and 95% confidence intervals from equations 15 and 16, with long differences from 1991 to each end year and the specification of Table 2, column (1). Standard errors clustered by firm.

Figure A.11: Effect of import competition on wage premia: incumbents and market average



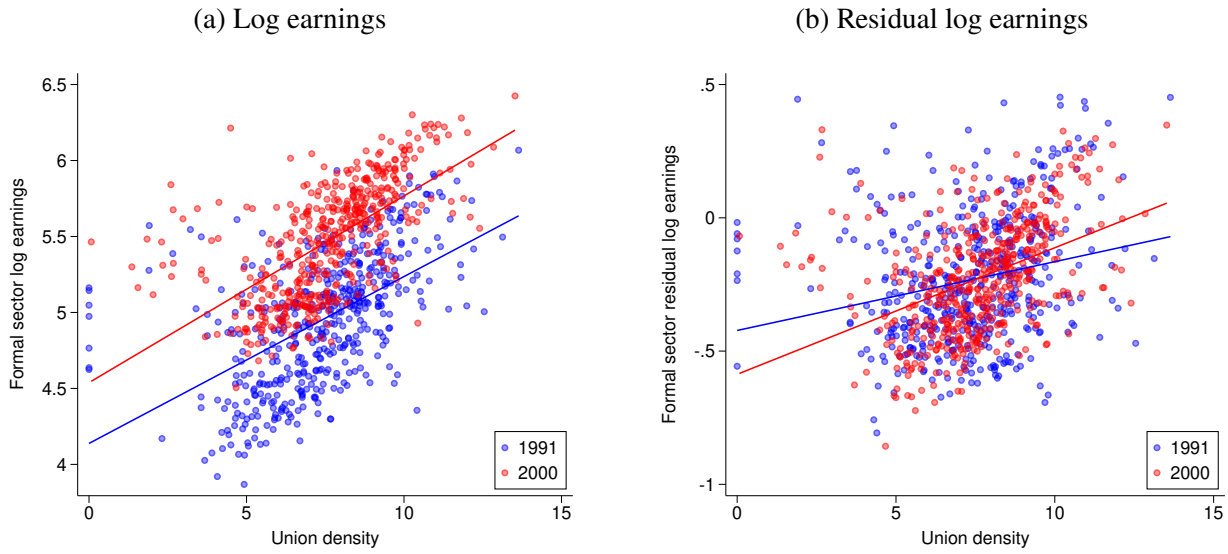
Notes: Coefficients on ΔICE_m and 95% confidence intervals from equation 12, weighted by the market's 1991 formal employment, with standard errors clustered by local labor market, in the markets of Table 5. "Incumbent firms" is the change in the wage premium of the firms present in the market in both 1991 and the year indicated, weighted by their 1991 employment. "Market average" is the change in the employment-weighted average wage premium of all of the market's firms. As in Figure 5, differences run forward in time: for the years before 1991, the outcome is its 1991 value minus its value in that year. Dashed lines indicate the beginning and end of the tariff reductions.

Figure A.12: Extended model: effect of a 10% increase in the trade shock on the formal sector wage take-home share, by microregion



Notes: As Figure 9, with every microregion receiving the same 10% increase in the trade shock ΔICE^* instead of its own 1991–2000 shock. Bars (left axis): the microregion's 1991 census shares of employment in formal wage work (black), informal wage work (gray), and self-employment (blue), for the 486 microregions of the census, sorted from the highest to the lowest formal wage work share. White circles (right axis), for the 478 microregions of the extended model: the change in the wage take-home share, in percentage points, implied by a 10% increase in ΔICE^* per equations 24 and 25, evaluated at the microregion's 1991 expected payroll Herfindahl and wage work share with the elasticities and reduced-form effects of Table 7, column (3); differences across microregions come only from their 1991 market structure. The red dashed line marks zero.

Figure A.13: Formal sector earnings and union density



Notes: Each dot is a microregion in 1991 (blue) or 2000 (red), with a linear fit for each year. Panel (a) plots the average log earnings of the microregion's formal sector workers against union density, and Panel (b) the same earnings net of observable worker characteristics. The horizontal axis is log(1 + union density). Union density data comes from https://academic.oup.com/jeea/article/23/1/236/7628306; it was collected for 1992 and 2001 and is matched to the 1991 and 2000 censuses.

Table A.1: 1991 descriptive statistics of formal local labor markets (microregion x occupation group pairs)

	Mean	Std. Dev.	p10	p25	Median	p75	p90	N
Number of formal workers	689	5,234	6	16	59	257	992	19,757
Number of formal tradable sector workers	251	1,917	0	2	15	83	346	19,757
Number of exporter workers	252	2,064	0	1	10	68	328	19,757
Number of formal non-tradable sector workers	438	3,789	2	6	27	129	533	19,757
Number of formal firms	114	790	3	6	16	54	180	19,757
Number of exporter firms	17	116	0	1	2	8	26	19,757
Payroll HHI (premium-based)	0.287	0.252	0.042	0.094	0.210	0.406	0.656	19,757
Payroll HHI (based on wage level)	0.292	0.252	0.045	0.098	0.216	0.412	0.663	19,757
Employment HHI	0.239	0.238	0.027	0.065	0.158	0.333	0.556	19,757
Average December wage (R$ of Dec. 2000)	483	495	137	193	317	571	1,019	19,757
Average December wage premium (R$ of Dec. 2000)	205	126	91	121	170	250	363	19,509
Import competition exposure (ΔICE)	0.126	0.094	0.000	0.048	0.132	0.179	0.234	19,757

Note: This table presents descriptive statistics across 19,757 Brazilian local labor markets defined as microregion × occupation group pairs, in the baseline year (1991). Means are unweighted. HHI denotes the Herfindahl-Hirschman Index of employment or payroll concentration; the premium-based Payroll HHI values each firm's payroll at its wage premium (Table 1). ΔICE is the change in import competition exposure.

Table A.2: Brazilian workforce during 1990s trade liberalization

	Population Census Year	
	1991	2000
N	51,332,184 (42.6%)	69,295,119 (57.4%)
Formal wage work	0.403 (0.491)	0.312 (0.463)
Informal wage work	0.324 (0.468)	0.344 (0.475)
Self-employed	0.239 (0.427)	0.205 (0.404)
Unemployed	0.044 (0.206)	0.138 (0.345)
Hours worked	43.534 (12.007)	43.950 (14.655)
Earnings (2000 reais)	331.088 (412.028)	553.742 (704.550)
Earnings (2000 reais) - Formal	383.827 (413.412)	596.306 (653.343)
Earnings (2000 reais) - Informal	279.714 (418.674)	497.428 (725.555)
Earnings (2000 reais) - Self-employed	309.382 (388.172)	568.641 (743.992)
Minimum wage (2000 reais)	126.514 (0.000)	151.000 (0.000)
Female	0.326 (0.469)	0.404 (0.491)
Age	34.262 (11.490)	34.717 (11.500)
Years of education	5.899 (4.522)	7.032 (4.392)

Notes: Descriptive statistics from the 1991 and 2000 population censuses, computed from the sample extracts in the supplementary materials of Dix-Carneiro and Kovak (2017) with individual sample weights. The sample is workers employed in the private sector. Formal wage work is having a signed *Carteira de Trabalho* (variables v350 and v0447 in the 1991 and 2000 censuses). Earnings and hours are for the main job; earnings statistics exclude the top and bottom 1% of the earnings distribution and are in 2000 reais (IPCA deflator). The minimum wage is the federal monthly minimum wage for July 1991 and for 2000, for a 44-hour work week.

Table A.3: Formal employment and worker flows by year

	1986	1987	1988	1989	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000
Panel A: December 31 Employment (millions)															
Total	10.64	20.80	21.04	21.74	20.79	19.65	18.25	18.84	20.37	23.54	23.29	24.09	23.93	23.64	25.25
Exporters	1.69	3.97	3.88	4.17	3.88	3.59	3.27	3.35	3.47	3.53	3.06	2.79	2.46	2.21	2.26
Non-Exp. Tradables	2.18	3.67	3.49	3.64	3.52	3.25	2.91	3.14	3.96	4.56	4.56	4.85	4.84	4.94	5.40
Non-Tradables	6.77	13.16	13.66	13.94	13.39	12.81	12.07	12.35	12.94	15.45	15.67	16.45	16.63	16.49	17.59
Panel B: Flows In/Out of Formal Sector (millions)															
Hired from Outside	7.84	3.60	3.73	4.01	3.47	3.23	2.95	3.47	4.46	4.63	4.04	4.53	4.37	4.33	5.03
Into Exporters	1.20	0.50	0.45	0.55	0.40	0.39	0.32	0.40	0.44	0.41	0.29	0.27	0.21	0.22	0.25
Into Non-Exp. Tradables	1.68	0.70	0.68	0.80	0.69	0.61	0.55	0.70	1.03	1.04	0.90	1.03	0.98	1.01	1.18
Into Non-Tradables	4.97	2.40	2.60	2.66	2.38	2.23	2.08	2.37	2.98	3.17	2.85	3.23	3.18	3.10	3.60
Separated to Outside	3.00	3.17	3.17	3.43	4.11	3.92	3.40	2.86	3.16	4.19	3.79	3.89	4.13	3.66	3.71
From Exporters	0.40	0.48	0.41	0.48	0.64	0.57	0.46	0.38	0.41	0.54	0.40	0.38	0.34	0.23	0.21
From Non-Exp. Tradables	0.59	0.66	0.59	0.64	0.78	0.74	0.62	0.53	0.67	0.92	0.82	0.84	0.90	0.81	0.85
From Non-Tradables	2.01	2.04	2.17	2.31	2.69	2.61	2.33	1.94	2.08	2.72	2.57	2.67	2.89	2.62	2.65
Panel C: Flows Between Firms (millions)															
From Exporters	0.88	0.96	0.79	0.82	0.83	0.63	0.51	0.52	0.65	0.74	0.53	0.49	0.41	0.36	0.35
From Non-Exp. Tradables	1.08	1.11	0.97	0.94	0.90	0.71	0.58	0.65	0.94	1.18	1.05	1.08	1.02	1.05	1.19
From Non-Tradables	3.83	3.73	3.75	3.55	3.27	2.80	2.36	2.59	3.19	3.92	3.57	3.58	3.46	3.55	3.92

Notes: This table summarizes formal private sector employment and worker flows in Brazil from 1986 to 2000 using matched employer-employee data from RAIS. Panel A reports December 31 employment (in millions) by firm trading status: exporters, non-exporting tradable firms, and non-tradable firms. Panel B reports annual flows into and out of the formal sector, separately by destination or origin firm type. "Hired from Outside" counts workers appearing in the formal sector on December 31 who were not employed in any formal firm on December 31 of the previous year. "Separated to Outside" counts workers employed in a formal firm on January 1 who are not found in any formal firm on December 31 of the same year. Panel C reports within-year job-to-job transitions between formal sector firms, by origin firm type. All values are in millions of workers.

Table A.4: Effect of import competition on market outcomes: Alternative specifications

	Main specification	ICE weights are firms' 1991 linear payroll shares	Dropping the largest 1% of markets	Unweighted
	(1)	(2)	(3)	(4)
Panel A: Formal sector labor market concentration				
Δ Payroll Herfindahl (premium-based)	0.172	0.421	0.174	0.212
	(0.042)	(0.074)	(0.031)	(0.017)
Δ Payroll Herfindahl	0.183	0.439	0.179	0.213
	(0.045)	(0.083)	(0.031)	(0.017)
Δ Employment Herfindahl	0.109	0.278	0.134	0.239
	(0.024)	(0.035)	(0.022)	(0.016)
Panel B: Number of firms and formal employment				
Δ Log number of firms	−0.971	−1.443	−0.424	−0.510
	(0.188)	(0.196)	(0.099)	(0.048)
Δ Log formal employment	−0.398	−0.048	0.194	−0.401
	(0.186)	(0.199)	(0.104)	(0.069)
Panel C: Wages				
Δ Log average firm wage premium	−0.124	0.200	0.112	−0.071
	(0.130)	(0.151)	(0.060)	(0.035)
Δ Log formal wage premium	−0.105	0.208	0.124	0.027
	(0.089)	(0.113)	(0.054)	(0.030)
Δ Log formal average wage	−0.075	0.255	0.119	0.014
	(0.110)	(0.131)	(0.062)	(0.035)
Weighted by 1991 formal employment	Yes	Yes	Yes	No
Observations	296,311	296,311	293,341	296,311
Local labor markets	19,757	19,757	19,559	19,757

Notes: Standard errors in parentheses, clustered at the local labor market level. Each cell reports the 2000 coefficient of the stacked differences-in-differences regression of Table 1, with local labor markets weighted by their 1991 formal employment except in column (4). Column (1) repeats Table 1, column (1). Column (2) constructs ΔICE_m with firms' 1991 linear payroll shares as weights, instead of squared payroll shares. Column (3) drops the largest 1% of local labor markets by 1991 formal employment. Column (4) is the main specification without weights, each local labor market counting once; its effect on concentration is identified mostly by comparing markets with and without tradable firms (Appendix D) and is the one that enters the effect on markdowns of Table 4. The sample is that of the concentration regressions.

Table A.5: Effects of import competition on market outcomes: Alternative fixed effects and controls

	Year FE (main) (1)	State × year FE (2)	State × year FE and 1986–1990 trends (3)	Microregion × year FE (4)
Panel A: Formal sector labor market concentration				
Δ Payroll Herfindahl (premium-based)	0.172	0.146	0.152	0.084
	(0.042)	(0.036)	(0.036)	(0.035)
Δ Payroll Herfindahl	0.183	0.152	0.157	0.092
	(0.045)	(0.038)	(0.038)	(0.038)
Δ Employment Herfindahl	0.109	0.107	0.110	0.044
	(0.024)	(0.023)	(0.023)	(0.021)
Panel B: Number of firms and formal employment				
Δ Log number of firms	−0.971	−0.795	−0.681	−0.440
	(0.188)	(0.138)	(0.129)	(0.122)
Δ Log formal employment	−0.398	−0.200	−0.194	−0.027
	(0.186)	(0.146)	(0.145)	(0.150)
Panel C: Wages				
Δ Log average firm wage premium	−0.124	−0.198	−0.190	−0.248
	(0.130)	(0.115)	(0.118)	(0.136)
Δ Log formal wage premium	−0.105	−0.096	−0.084	−0.085
	(0.089)	(0.076)	(0.077)	(0.084)
Δ Log formal average wage	−0.075	−0.154	−0.154	−0.196
	(0.110)	(0.096)	(0.098)	(0.108)

Notes: Standard errors in parentheses, clustered at the local labor market level. Each entry is the coefficient on ΔICE_m in the final year of the stacked differences-in-differences regression of equation (12). Column (1) is the main specification, with local labor market and year fixed effects. Column (2) replaces year fixed effects with state × year fixed effects. Column (3) adds the market's 1986–1990 change in the outcome, interacted with year. Column (4) uses microregion × year fixed effects. Markets are weighted by their 1991 formal employment, as in Table 1, column (1).

Table A.6: Effect of import competition on market outcomes: Alternative clustering

	Main specification	Two-way clustered by microregion and occupational group	AKM (2019) standard errors
	(1)	(2)	(3)
Panel A: Formal sector labor market concentration			
Δ Payroll Herfindahl (premium-based)	0.172	0.172	0.172
	(0.042)	(0.046)	(0.045)
Δ Payroll Herfindahl	0.183	0.183	0.183
	(0.045)	(0.048)	(0.048)
Δ Employment Herfindahl	0.109	0.109	0.109
	(0.024)	(0.035)	(0.033)
Panel B: Number of firms and formal employment			
Δ Log number of firms	−0.971	−0.971	−0.971
	(0.188)	(0.576)	(0.407)
Δ Log formal employment	−0.398	−0.398	−0.398
	(0.186)	(0.442)	(0.371)
Panel C: Wages			
Δ Log average firm wage premium	−0.124	−0.124	−0.124
	(0.130)	(0.086)	(0.122)
Δ Log formal wage premium	−0.105	−0.105	−0.105
	(0.089)	(0.099)	(0.120)
Δ Log formal average wage	−0.075	−0.075	−0.075
	(0.110)	(0.070)	(0.124)
Observations	296,311	296,311	296,311
Local labor markets	19,757	19,757	19,757

Notes: Standard errors in parentheses. Point estimates are identical across columns; only standard errors differ. Column 1 clusters at the local labor market level, column 2 two-way clusters by microregion and occupational group, and column 3 uses AKM (2019) shift-share standard errors. All columns report the 2000 coefficient from stacked differences-in-differences regressions, as in Table 1.

Table A.7: Effect of import competition on concentration: Sources

Panel A: Exit, firms present in both years, and entry	
Δ Payroll Herfindahl	0.183
	(0.045)
Via firms present in 1991 only (exit)	−0.066
	(0.035)
Via the shares of firms present in both years	0.127
	(0.053)
Via firms present in 2000 only (entry)	0.122
	(0.040)
1991 payroll share of firms absent in 2000	−0.020
	(0.070)
2000 payroll share of firms absent in 1991	−0.253
	(0.077)
Observations	296,370
Panel B: By firm type	
Δ Payroll Herfindahl	0.183
	(0.045)
Exporters	0.144
	(0.041)
Firms present in 1991 only (exit)	0.060
	(0.018)
Firms present in both years	0.095
	(0.038)
Firms present in 2000 only (entry)	−0.012
	(0.004)
Non-exporting tradable firms	−0.005
	(0.012)
Firms present in 1991 only (exit)	0.007
	(0.005)
Firms present in both years	0.012
	(0.003)
Firms present in 2000 only (entry)	−0.024
	(0.012)
Non-tradable firms	0.044
	(0.019)
Firms present in 1991 only (exit)	0.084
	(0.016)
Firms present in both years	0.020
	(0.011)
Firms present in 2000 only (entry)	−0.060
	(0.011)
Observations	296,370

Notes: Each cell is the 2000 coefficient on $\Delta ICE_m \times \mathbf{1}[t = 2000]$ from the main specification of Table 1, column (1) (equation 12, 1986–2000, market and year fixed effects, markets weighted by their 1991 formal employment, standard errors clustered by market in parentheses), with the outcome in each row. Panel A: the first row is the change in the market's December wage-bill Herfindahl since 1991, the Payroll Herfindahl of Table 1; the next three rows decompose it exactly, firm by firm, into the change in the Herfindahl of the 1991 shares when the firms absent in t are removed and the remaining shares renormalized (exit), the change between the renormalized 1991 shares and the renormalized t shares of the firms present in both years, and the change from adding the firms present in t only (entry). The last two rows are the 1991 payroll share of the firms absent in t and the payroll share in t of the firms absent in 1991. Panel B: the change in the Herfindahl is the sum over firms of the change in the firm's squared payroll share, with a share of zero in a year in which the firm is absent; each row sums these changes over the firms of one type (exporters, firms that exported at any point in 1991–1994; other tradable firms; non-tradable firms) and, within type, over the firms absent in 2000, present in both years, and absent in 1991. Shares are not renormalized in Panel B, so the rise in a surviving firm's share when other firms leave is in the row of the surviving firm. A 10% increase in exposure is 0.1 units of ΔICE_m.

Table A.8: Within-market inverse elasticity $1/\eta$: Robustness

	Main specification (all-worker premia)	Stayers' wage premia	December average wage	Including exiting firms, coding employment and wages at exit as zero
	(1)	(2)	(3)	(4)
Panel A: First stage				
Δ Firm log employment in LLM	-0.736	-0.743	-0.736	-0.737
	(0.076)	(0.091)	(0.076)	(0.075)
First stage F	94.915	66.474	94.915	96.505
Panel B: Reduced form				
Δ Firm wage premium in LLM	-0.612	-0.579	-0.604	-0.617
	(0.036)	(0.049)	(0.038)	(0.036)
Panel C: 2SLS				
Labor supply within-market cross-firm	0.832	0.779	0.820	0.837
inverse elasticity of substitution	(0.093)	(0.107)	(0.094)	(0.093)
Implied upper bound on wage take-home share	55%	56%	55%	54%
Local labor market (LLM) FE	Yes	Yes	Yes	Yes
Pre-trend controls (1986–91)	Yes	Yes	Yes	Yes
Observations	363,407	190,069	363,407	700,065
Firms	149,083	89,962	149,083	303,393
Local labor markets	11,632	9,502	11,632	14,175

Notes: Standard errors clustered by firm in parentheses. All regressions include local labor market (LLM) fixed effects and are weighted by baseline market employment. Column (1): main specification (Table 2, column (1)), using the December wage premium of all of the firm's workers, conditional on observables. Column (2): wage premium estimated on the firm's stayers, workers in the same firm, microregion, and occupation in 1991 and 2000; Table 2, column (2), reports the version conditional on worker fixed effects. Column (3): December average wage instead of the residualized wage premium. Column (4): includes firms that exited between 1991 and 2000, coding their exit-year employment and wages as zero (inverse hyperbolic sine transform). Table 2 reports microregions as local labor markets and Appendix Table E.2 the effective rate of protection as the shock.

Table A.9: Alternative standard errors for the within-market and cross-market inverse elasticities of substitution

	$1/\eta$ (Table 2, column (1))			$1/\theta$ (Table 3, column (1))	
	Clustered by firm (main)	Clustered by local labor market	Clustered by sector (CNAE)	Clustered by local labor market (main)	Two-way clustered by microregion and occupation
	(1)	(2)	(3)	(4)	(5)
Panel A: First stage					
Δ Firm log employment in LLM	-0.736	-0.736	-0.736		
	(0.076)	(0.088)	(0.148)		
Δ LLM employment index				-0.294	-0.294
				(0.044)	(0.100)
First stage F	94.9	70.0	24.8	45.6	8.7
Panel B: Reduced form					
Δ Firm wage premium in LLM	-0.612	-0.612	-0.612		
	(0.036)	(0.114)	(0.087)		
Δ LLM wage premium index				-0.183	-0.183
				(0.050)	(0.103)
Panel C: 2SLS					
$1/\eta$	0.832	0.832	0.832		
	(0.093)	(0.218)	(0.132)		
$1/\theta - 1/\eta$				0.622	0.622
				(0.217)	(0.473)
$1/\theta$				1.454	1.454
				(0.196)	(0.464)
Observations	363,407	363,407	363,407	17,628	17,628

Notes: Standard errors in parentheses. Point estimates are those of Table 2, column (1) (columns (1)–(3)) and of Table 3, column (1) (columns (4)–(5)); only the clustering of the standard errors changes. Column (1) clusters by firm, column (2) by local labor market (microregion × occupation), column (3) by sector (CNAE95); column (4) clusters by local labor market, column (5) two-way by microregion and occupational group.

Table A.10: Correlates of firm-level wage markdowns

	Dependent variable: $\ln(\mu_{zm})$			
	(1)	(2)	(3)	(4)
Panel A: Employment				
$\ln(\text{Employment}_{zm})$	0.003824	0.003599	0.003423	0.003882
	(0.000185)	(0.000171)	(0.000153)	(0.000173)
Observations	9,484,956	9,484,955	9,380,001	9,484,955
Panel B: Residual wage				
$\ln(\text{Residual wage}_{zm})$	0.000759	0.000666	0.000313	0.003355
	(0.000152)	(0.000121)	(0.000111)	(0.000156)
Observations	9,484,956	9,484,955	9,380,001	9,484,955
Panel C: Average wage				
$\ln(\text{Average wage}_{zm})$	0.000710	0.000553	0.000030	0.002825
	(0.000141)	(0.000113)	(0.000098)	(0.000142)
Observations	9,484,956	9,484,955	9,380,001	9,484,955
Year FE	Yes	Yes	Yes	Yes
Sector FE		Yes		
Firm FE			Yes	
Local labor market FE				Yes
Firms	580,576	580,575	475,621	580,576
Local labor markets	19,023	19,023	19,005	19,022

Notes: This table presents regressions of log firm-level wage markdowns on firm-level employment, residual wages, and average wages. All regressions include year fixed effects. Standard errors in parentheses, clustered by local labor market.

Table A.11: Heterogeneity by informality: quartile 1 split at its median and quartiles 2–4 pooled

	All markets	Quartile 1 (most formal)			Quartiles 2–4 (pooled)
		All	Upper half	Lower half	
	(1)	(2)	(3)	(4)	(5)
Share of national formal employment, 1991	0.995	0.726	0.573	0.153	0.269
Median formal share of employment, 1991	0.265	0.502	0.559	0.457	0.205
Panel A: Effect of 10% increase in import competition on concentration					
Δ Payroll Herfindahl (premium-based)	0.017	0.015	0.013	0.018	0.019
	(0.004)	(0.005)	(0.006)	(0.009)	(0.004)
Mean (unweighted), 1991	0.287	0.212	–	–	0.331
Mean (weighted), 1991	0.079	0.064	0.053	0.118	0.151
Observations	296,311	108,396	57,628	50,768	187,915
Local labor markets	19,757	7,227	3,842	3,385	12,530
Panel B: Within-market cross-firm inverse elasticity $1/\eta$ (2SLS)					
$1/\eta$	0.832	0.824	0.821	1.650	-1.252
	(0.093)	(0.091)	(0.091)	(1.401)	(0.977)
First-stage F	94.9	97.1	97.0	1.4	1.9
Implied upper bound on take-home share	0.546	0.548	0.549	–	–
Observations	363,407	287,985	237,348	50,637	75,422
Firms	149,083	115,676	95,720	20,808	34,513
Local labor markets	11,632	5,361	3,054	2,307	6,271

Notes: See notes to Table 6. Columns (3) and (4) split quartile 1 at the median formal share of total employment across its microregions. Column (5) pools quartiles 2 to 4.

Table A.12: Wage effects at firms facing larger tariff cuts vs the market average

	Δ Log wage premium	Δ Log employment	Δ Payroll share of the market	Δ Log wage take-home share	Δ Log marginal revenue product of labor
	(1)	(2)	(3)	(4)	(5)
Panel A. Firms facing larger tariff cuts, within local labor markets					
Δ in log import tariff faced by firm	-0.612	-0.736	-0.0004	0.0001	-0.613
	(0.036)	(0.076)	(0.0001)	(0.0000)	(0.036)
Share of the wage premium response	100%	–	–	0%	100%
Local labor market FE, pre-trend controls	Yes	Yes	Yes	Yes	Yes
Observations (firm × market)			363,407		
Panel B. Local labor market average, across local labor markets					
ΔICE_m	-0.124	-0.398	–	–	–
	(0.130)	(0.186)			
Local labor market and year FE	Yes	Yes			
Observations (market × year)			296,311		

Notes: Standard errors in parentheses, clustered by firm in Panel A and by local labor market in Panel B. All regressions are weighted by the market's 1991 formal employment. In Panel A, each column reports the coefficient on the firm's 1990–1994 change in log import tariff in a regression of the firm's 1991–2000 change in the column's outcome, with local labor market fixed effects and the pair's 1986–1991 changes in log employment and in the log wage premium, on the sample of Table 2, column (1). Column (1) is the reduced form of Table 2 for the firm's wage premium, and column (2) its first stage. Column (3) is the change in the firm's share of its market's payroll. Column (4) is the change in the firm's log wage take-home share implied by its payroll share and the estimated elasticities, $\ln \mu_{zm}^{-1} = -\ln(1 + (1/\eta)(1 - s_{zm}) + (1/\theta)s_{zm})$, with $1/\eta$ from Table 2, column (1) and $1/\theta$ from Table 3. Column (5) is column (1) minus column (4), the change in the log marginal revenue product of labor. The row "share of the wage premium response" divides columns (4) and (5) by column (1). Panel B reports the estimates of Table 1, column (1): the 2000 coefficient on ΔICE_m for the market's log average firm wage premium (column 1) and log employment (column 2), relative to 1991.

Table A.13: Extended model: Markdown effects by formality quartile

	All microregions	Quartile 1 (most formal)	Quartile 2	Quartile 3	Quartile 4 (least formal)
	(1)	(2)	(3)	(4)	(5)
Microregions	475	121	121	116	117
Share of national formal employment, 1991	0.997	0.726	0.171	0.073	0.027
Share of 1991 formal employment, RAIS (weight)	1.000	0.834	0.119	0.039	0.008
Panel A: Market structure, 1991 (weighted averages)					
Expected payroll Herfindahl	0.022	0.015	0.048	0.091	0.121
Wage work share of market payroll	0.780	0.800	0.707	0.638	0.537
Panel B: Wage take-home share at the 1991 averages					
Extended model	0.531	0.531	0.530	0.529	0.530
with all market payroll in wage work ($s_m = 1$)	0.530	0.531	0.528	0.524	0.522
with $\Omega = 0$ ($1/\eta$ in place of $1/\tilde{\eta}$)	0.533	0.534	0.532	0.531	0.532
Panel C: Effect of import competition on the average wage take-home share, at the 1991 averages					
Δ average inverse elasticity, per 10% increase in the trade shock	0.0053	0.0057	0.0038	0.0021	0.0002
	(0.0014)	(0.0015)	(0.0012)	(0.0009)	(0.0006)
through the change in concentration	0.0056	0.0059	0.0046	0.0036	0.0021
through the change in the wage work share	-0.0004	-0.0002	-0.0008	-0.0015	-0.0020
Implied change in take-home share (pp)	-0.15	-0.16	-0.11	-0.06	-0.00

Notes: Microregions are grouped into quartiles of the 1991 formal share of total employment, as in Table 6. The take-home share is equation 24 evaluated at the group's 1991 averages of the expected payroll Herfindahl and of the wage work share of market payroll, with the elasticities of Table 7; the two counterfactual rows set the wage work share to one and the separation-risk bias Ω to zero. The effect of a 10% increase in the trade shock is equation 25 evaluated at the same averages, with the effects on the expected Herfindahl and on the wage work share from the microregion regressions of Table 7, Panels A–B, weighted by 1991 formal employment and estimated separately by group; the two rows below it are its concentration and wage-work-share components. Standard errors by the delta method, holding the elasticities fixed.

Table A.14: Extended model: Distribution of the wage take-home share and of the effect of trade across microregions

	Min	p5	p10	p25	p50	p75	p90	p95	Max
Panel A: 1991									
Wage take-home share	0.492	0.521	0.525	0.527	0.530	0.531	0.532	0.534	0.562
Expected payroll Herfindahl	0.002	0.012	0.017	0.031	0.070	0.145	0.292	0.410	0.930
Wage work share of market payroll	0.070	0.324	0.437	0.587	0.694	0.771	0.806	0.823	0.874
N (microregions)					475				
Panel B: 2000									
Wage take-home share	0.494	0.519	0.525	0.528	0.531	0.532	0.532	0.532	0.532
Expected payroll Herfindahl	0.001	0.006	0.008	0.014	0.032	0.081	0.173	0.277	1.000
Wage work share of market payroll	0.475	0.600	0.637	0.686	0.731	0.766	0.787	0.801	0.894
N (microregions)					475				
Panel C: Effect of a 10% increase in the trade shock (1991 market structure)									
Change in take-home share (pp)	-0.18	-0.16	-0.15	-0.13	-0.09	-0.01	0.09	0.17	0.48
N (microregions)					475				

Notes: Unweighted distribution across the 475 microregions of the extended model. The take-home share is equation 24 evaluated at each microregion's expected payroll Herfindahl and wage work share of market payroll in the year, with the elasticities of Table 7, which are common to all microregions. Panel C is equation 25 evaluated at each microregion's 1991 market structure, with the effects on the expected Herfindahl and on the wage work share of Table 7, Panels A and B.

A Data and Methods Appendix

Data on workers and firms: RAIS

Overview. I use Brazil's Relação Anual de Informações Sociais (RAIS) for years 1986 to 2000 as my source of information on workers and firms. RAIS is an administrative employer-employee linked dataset collected by the federal government for the purposes of administering workers' social security. Thus, RAIS covers all workers with signed worker cards (Carteira do Trabalho), namely the entirety of formal sector employment. Firms report to RAIS once a year, listing all workers who ever worked for the firm in the prior calendar year. Firms are required to report a rich set of information about each employment contract (e.g., occupation, admission date, separation date, etc.), as well as worker demographics (i.e., education, date of birth, and gender), separately for each establishment. The municipality of each establishment as well as the economic sector of the firm are also reported.

Wages. RAIS includes two wage variables for years 1986–2000: average monthly earnings and December monthly earnings. Both variables are reported as multiples of the national minimum wage.

Occupation codes. RAIS' occupation codes are 5-digit variables "CBO" (prior to 1994) and "CBO94" (1994 onwards). I focus on the first 2 digits to group workers into occupation groups. Both variables share the same data dictionary, with the only difference between them being phased-out and phased-in occupation codes. I have compiled a complete list of all raw occupation codes, along with the total number of workers in each of them, labels, and flags for which codes were either "phased-out" or phased-in, which I identified based on whether the number of workers changed by a factor of more than 100 between any two years. I then re-classified the first two digits of all phased-out and phased-in codes as "99 - Other occupations," a reclassification that affects roughly 10% of all workers.

Sector codes. RAIS' finest sector codes for 1986–2000 are 4-digit "IBGESUBATIVIDADE" (prior to 1995) and 5-digit "CNAE95" (1995 onwards). I focus on the 5-digit CNAE95 codes to map tariff shocks to firms in RAIS. For firms that exit the data prior to reporting any CNAE95 codes, I assign a CNAE95 code using a correspondence table I constructed using the pre-1995 and post-1995 codes of firms in business in both periods. To each IBGESUBATIVIDADE code I assign the most commonly reported CNAE95 code. Finally, throughout all years I use the first CNAE95 code ever reported by a firm as its official CNAE95 code.

Sample restrictions. I focus on workers employed as of December 31 of each year, aged 18–64, and with positive December earnings. I exclude all workers in the public sector or with unknown sector. To make sure all public sector workers are excluded, I further exclude workers whose employer's economic activity was not marked as government, but who hold public sector occupations (i.e., Diplomats, Civil servants, and Post office). Finally, following Dix-Carneiro and

Kovak (2017) I exclude from all analyses the microregion of Manaus.[56]

Data on tariff shocks: TRAINS

I use tariff data from UNCTAD's Trade Analysis Information System (TRAINS), which I download from the World Integrated Trade Solution (WITS) website (https://wits.worldbank.org/). The raw tariff data are available for Brazil at the 8-digit HS product level for years 1988 (the first year the data are available) through 2000. As outlined in Section 4, I compute a firm's tariff reduction shock as the change in log one plus the firm's CNAE95 sector code's nominal tariff between years 1990 and 1994.

To map product-level tariffs to CNAE95, I use Brazil's NCM (Nomenclatura Comum do Mercosul) classification, whose first six digits correspond to the HS. I map NCM codes to CNAE95 using official IBGE correspondence tables.[57] Over 90% of products map to a single activity code; I randomly assign the remainder across the handful of activity codes to which they correspond per the concordance tables. I compute CNAE95-level tariffs as averages of product-level nominal tariffs, weighted by the number of tariff lines. For robustness, I compute effective rates of protection (ERP) using the Corden formula: $ERP_j = (t_j - \sum_i a_{ij} t_i)/(1 - \sum_i a_{ij})$, where t_j is sector j's nominal tariff, t_i are input tariffs, and a_{ij} are Leontief coefficients from Brazil's 1985 intersectoral coefficients matrix ("Tabela 20") at Nível 50.[58] I map Nível 50 sectors to CNAE95 using CONCLA correspondence tables.

Other data

List of exporters. I classify firms as exporters during the reform period (1990–1994) by matching the list of exporters during that period to RAIS using firms' unique identifiers (CNPJ). The list of exporters was provided by the (extinct as of 2019) Ministry of Development, Industry, and Foreign Trade, currently a part of the Ministry of the Economy, in October 2018.

Census. I use data from the 1991 and 2000 census, available in the supplementary materials from Dix-Carneiro and Kovak (2017).

[56] Law 8.387 of December 1991 changed the rules of the Manaus Free Trade Zone during the liberalization, introducing a minimum set of production operations as a condition for its tax benefits and an 88 percent reduction of the import duty on approved projects' foreign inputs, so changes in Manaus over the period combine the liberalization with a change in the zone's own regime. Using synthetic controls, Possebom (2017) finds that the zone's effects on the region's GDP per capita were largest between 1975 and 1990 and weakened after 1991, as national tariff cuts reduced its relative advantage.

[57] https://concla.ibge.gov.br/classificacoes/correspondencias/atividades-economicas.

[58] https://www.ibge.gov.br/estatisticas-novoportal/economicas/contas-nacionais/9085-matriz-de-insumo-produto.html?&t=downloads.

Methods: wage premia regressions

For each year, I estimate each firm's wage premium in its local labor market as firm × market fixed effects in a regression of worker log December earnings on the firm × market fixed effects and the same worker observable controls as Dix-Carneiro and Kovak (2017), namely: a dummy for female; 4 age group dummies (25–29; 30–39; 40–49; 50–64); 8 education group dummies (primary school, incomplete primary school, middle school, incomplete middle school, high school, incomplete high school, college, incomplete college). The omitted category is therefore males aged 18–24 with no formal education. Similarly, for each year, I estimate each market's wage premia from a regression of worker log December earnings on the market fixed effects and the previously mentioned worker observable controls.

For my main estimates of the within-market cross-firm elasticity (column (1) of Table 2), I use as firm wage premia the firm × market fixed effects of these regressions, estimated on all of the firm's workers, and my outcome is their 1991–2000 change. As robustness, I instead estimate the premia on the firm's stayers—workers employed in the same firm, microregion, and occupation in both 1991 and 2000—by regressing stayers' log December earnings on firm × market fixed effects and the worker observable controls above for each of the two years (Appendix Table A.8); in column (2) of Table 2, I additionally condition on worker fixed effects. In the extended model of Section 8, I use the analogous premia for firm × microregion pairs.

For the robustness exercise of the within-market cross-firm elasticity using wage premia that condition on worker fixed effects (Table 2, column (2)), I estimate each firm's wage premia in 1991 and 2000 as firm × market × year fixed effects in a regression—containing years 1991 and 2000, on stayers—of worker log December earnings on worker fixed effects, the firm × market × year fixed effects, and worker observable-characteristics-by-year controls.

B Model Appendix

This Appendix provides detailed derivations for various expressions in Section 2.

B.1 Derivation of labor supply equation

Consider an economy consisting of a continuum of homogeneous workers j, a large but finite number of local labor markets m, and a finite number of firms z within each local labor market. Each worker chooses to which firm-market pair zm they provide h^j_{zm} units of labor by minimizing their indirect disutility of work V_{zm} subject to making reservation earnings $y^j \sim F(y)$:

$$\begin{aligned}\min_{zm} V^j_{zm} &= \ln l^j_{zm} + \ln \xi_m + \ln \xi_{zm} - \xi^j_{zm} \\ &\text{s.t. } l^j_{zm} w_{zm} \geq y_j\end{aligned}$$

where ξ^j_{zm} is an idiosyncratic taste for working at firm z in market m, and ξ_m and ξ_{zm} are taste shifters common to all workers. This is equivalent to

$$\max_{zm} \ln w_{zm} - \ln y_j - \ln \xi_m - \ln \xi_{zm} + \xi^j_{zm}$$

Now suppose ξ^j_{zm} has the following Gumbel distribution, a member of the General Extreme Value (GEV) family:

$$G\left(\left\{\xi^j_{zm}\right\}\right) = \exp\left[-\sum_m \left(\sum_{z\in B_m} e^{-\left(1+\frac{\sigma}{1-\sigma}\right)\xi^j_{zm}}\right)^{\frac{1+\frac{\varphi}{1-\varphi}}{1+\frac{\sigma}{1-\sigma}}}\right]$$

where $0 \leq \sigma < 1$ is the index of similarity across firms within a market, $0 \leq \varphi < 1$ is the index of similarity across markets, and B_m is the set of firms in market m.

The probability that worker j chooses firm z in market m is

$$P\left(\xi^j_{zm} > \ln w_{zm} - \ln y_j - \ln \xi_m - \ln \xi_{zm}\right),$$

which can be decomposed as:

$$P^j_{zm} = P(z|B_m)\, P(B_m) \quad \forall j$$

where $P(z|B_m)$ is the probability of choosing firm z conditional on choosing market m with set B_m of firms, and $P(B_m)$ is the probability of choosing market m. By the results in McFadden

(1978), P_{zm} can be computed as:

$$\begin{aligned}
P\left(z|B_m\right) =& \frac{\exp\left[\left(\ln w_{zm} - \ln y^j - \ln \xi_m - \ln \xi_{zm}\right)/(1-\sigma)\right]}{\sum_{k\in B_n}\exp\left[\left(\ln w_{kn} - \ln y^j - \ln \xi_m - \ln \xi_{km}\right)/(1-\sigma)\right]} \\
=& \frac{\exp\left[\left(\ln w_{zm}^{\frac{1}{1-\sigma}} - \ln y^{j\,\frac{1}{1-\sigma}} - \ln \xi_m^{\frac{1}{1-\sigma}} - \ln \xi_{zm}^{\frac{1}{1-\sigma}}\right)\right]}{\sum_{k\in B_m}\exp\left[\left(\ln w_{km}^{\frac{1}{1-\sigma}} - \ln y^{j\,\frac{1}{1-\sigma}} - \ln \xi_m^{\frac{1}{1-\sigma}} - \ln \xi_{km}^{\frac{1}{1-\sigma}}\right)\right]} \\
=& \frac{\left(\frac{w_{zm}}{y^j \xi_m \xi_{zm}}\right)^{\frac{1}{1-\sigma}}}{\sum_{k\in B_n}\left(\frac{w_{km}}{y^j \xi_m \xi_{km}}\right)^{\frac{1}{1-\sigma}}} \\
=& \frac{\left(\frac{w_{zm}}{\xi_{zm}}\right)^{\frac{1}{1-\sigma}}}{\sum_{k\in B_n}\left(\frac{w_{km}}{\xi_{km}}\right)^{\frac{1}{1-\sigma}}}
\end{aligned}$$

and

$$\begin{aligned}
P\left(B_m\right) =& \frac{\left\{\sum_{z\in B_m}\exp\left[\left(\ln w_{zm} - \ln y^j - \ln \xi_m - \ln \xi_{zm}\right)/(1-\sigma)\right]\right\}^{\frac{1-\sigma}{1-\varphi}}}{\sum_l\left\{\sum_{k\in B_l}\exp\left[\left(\ln w_{kl} - \ln y^j - \ln \xi_l - \ln \xi_{kl}\right)/(1-\sigma)\right]\right\}^{\frac{1-\sigma}{1-\varphi}}} \\
=& \frac{\left[\sum_{z\in B_m}\left(\frac{w_{zm}}{y^i \xi_m \xi_{zm}}\right)^{\frac{1}{1-\sigma}}\right]^{\frac{1-\sigma}{1-\varphi}}}{\sum_l\left[\sum_{k\in B_l}\left(\frac{w_{km}}{y^i \xi_m \xi_{km}}\right)^{\frac{1}{1-\sigma}}\right]^{\frac{1-\sigma}{1-\varphi}}} \\
=& \frac{\left[\left(\frac{1}{\xi_m}\right)^{\frac{1}{1-\sigma}}\sum_{z\in B_m}\left(\frac{w_{zm}}{\xi_{zm}}\right)^{\frac{1}{1-\sigma}}\right]^{\frac{1-\sigma}{1-\varphi}}}{\sum_l\left[\left(\frac{1}{\xi_l}\right)^{\frac{1}{1-\sigma}}\sum_{k\in B_l}\left(\frac{w_{kl}}{\xi_{kl}}\right)^{\frac{1}{1-\sigma}}\right]^{\frac{1-\sigma}{1-\varphi}}}
\end{aligned}$$

Putting them together

$$P_{zm}^j = \frac{\left(\frac{w_{zm}}{\xi_{zm}}\right)^{\frac{1}{1-\sigma}}}{\sum_{k\in B_n}\left(\frac{w_{km}}{\xi_{km}}\right)^{\frac{1}{1-\sigma}}} \times \frac{\left[\left(\frac{1}{\xi_m}\right)^{\frac{1}{1-\sigma}}\sum_{z\in B_m}\left(\frac{w_{zm}}{\xi_{zm}}\right)^{\frac{1}{1-\sigma}}\right]^{\frac{1-\sigma}{1-\varphi}}}{\sum_l\left[\left(\frac{1}{\xi_l}\right)^{\frac{1}{1-\sigma}}\sum_{k\in B_l}\left(\frac{w_{kl}}{\xi_{kl}}\right)^{\frac{1}{1-\sigma}}\right]^{\frac{1-\sigma}{1-\varphi}}} \quad \forall j$$

Let $\eta \equiv \frac{\sigma}{1-\sigma} > 0$, $\theta \equiv \frac{\varphi}{1-\varphi} > 0$, and denote $P^j_{zm} = P_{zm}$ for simplicity. Then:

$$P_{zm} = \frac{\left(\frac{w_{zm}}{\xi_{zm}}\right)^{1+\eta}}{\sum_{k\in B_n}\left(\frac{w_{km}}{\xi_{km}}\right)^{1+\eta}} \times \frac{\left[\left(\frac{1}{\xi_m}\right)^{1+\eta}\sum_{z\in B_m}\left(\frac{w_{zm}}{\xi_{zm}}\right)^{1+\eta}\right]^{\frac{1+\theta}{1+\eta}}}{\sum_l\left[\left(\frac{1}{\xi_l}\right)^{1+\eta}\sum_{k\in B_l}\left(\frac{w_{kl}}{\xi_{kl}}\right)^{1+\eta}\right]^{\frac{1+\theta}{1+\eta}}}$$

Finally, define the following wage indices:

$$W_m \equiv \left[\sum_z\left(\frac{w_{zm}}{\xi_{zm}}\right)^{1+\eta}\right]^{\frac{1}{1+\eta}}, \quad W \equiv \left[\sum_m\left(\frac{W_m}{\xi_m}\right)^{1+\theta}\right]^{\frac{1}{1+\theta}}$$

Then

$$P^j_{zm} = \frac{\left(\frac{w_{zm}}{\xi_{zm}}\right)^{1+\eta}}{W_m^{1+\eta}} \times \frac{\left(\frac{W_m}{\xi_m}\right)^{1+\theta}}{W^{1+\theta}} = \left(\frac{w_{zm}/\xi_{zm}}{W_m}\right)^{1+\eta} \times \left(\frac{W_m/\xi_m}{W}\right)^{1+\theta} \tag{B.1}$$

With equation B.1 at hand, total labor supplied to firm z in market m can be found by integrating probabilities P^j_{zm} (times $h^j_{zm} = y^j/w_{zm}$ supplied by each worker) over the continuum of workers:

$$l_{zm} = \int_0^1 P^j_{zm}\left(\frac{y^j}{w_{zm}}\right) dF(y) = w_{zm}^{-1} P_{zm} Y \tag{B.2}$$

where $Y \equiv \int_0^1 y^j dF(y)$ is the country-level labor income. To obtain an expression for l_{zm} that is a function of w_{zm}, parameters, and market-level aggregates, I define the following employment indices:

$$L_m \equiv \left[\sum_z (\xi_{zm} l_{km})^{\frac{1+\eta}{\eta}}\right]^{\frac{\eta}{1+\eta}}, \quad L \equiv \left[\sum_m (\xi_m L_m)^{\frac{1+\theta}{\theta}}\right]^{\frac{\theta}{1+\theta}}$$

which, together with equation B.2 and the previously defined wage indices, implies $Y = \sum_{zm} w_{zm} l_{zm} =$

WL and

$$
\begin{aligned}
l_{zm} =& w_{zm}^{-1} P_{zm} Y \\
=& w_{zm}^{-1} \left[\left(\frac{w_{zm}}{\xi_{zm} W_m} \right)^{1+\eta} \times \left(\frac{W_m}{\xi_m W} \right)^{1+\theta} \right] Y \\
=& \cancel{w_{zm}^{-1}} \left[\left(\frac{\cancel{w_{zm}}}{\xi_{zm} \cancel{W_m}} \right) \left(\frac{w_{zm}}{\xi_{zm} W_m} \right)^{\eta} \times \left(\frac{W_m}{\xi_m W} \right)^{\theta} \left(\frac{\cancel{W_m}}{\xi_m \cancel{W}} \right) \right] \cancel{W} L \\
=& (\xi_{zm} \xi_m)^{-1} \left(\frac{w_{zm}/\xi_{zm}}{W_m} \right)^{\eta} \left(\frac{W_m/\xi_m}{W} \right)^{\theta} L
\end{aligned}
$$

Rearranging:

$$
l_{zm} = L \left(\frac{w_{zm}}{W_m} \right)^{\eta} \left(\frac{W_m}{W} \right)^{\theta} \left(\xi_{zm}^{1+\eta} \xi_m^{1+\theta} \right)^{-1} \tag{B.3}
$$

B.2 Other proofs and derivations

B.2.1 Equation 3: $w_{zm} = W \left(\frac{l_{zm}}{L_m} \right)^{\frac{1}{\eta}} \left(\frac{L_m}{L} \right)^{\frac{1}{\theta}} \xi_{zm}^{1+\frac{1}{\eta}} \xi_m^{1+\frac{1}{\theta}}$

The inverse function of the residual labor supply equation 2 (same as Appendix equation B.3) is the wage w_{zm} at which l_{zm} units of labor are supplied to firm z at market m. To check that equation 3 satisfies this criterion, plug it into equation B.2 to obtain the identity $l_{zm} = l_{zm}$. I show this in two steps.

First, plug in the expression for w_{zm} into equation B.2 to get:

$$
\begin{aligned}
l_{zm} =& w_{zm}^{-1} P_{zm} Y \\
=& w_{zm}^{-1} \left(\frac{w_{zm}/\xi_{zm}}{W_m} \right)^{1+\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{1+\theta} Y \\
=& w_{zm}^{\eta} \left(\frac{1}{W_m \xi_{zm}} \right)^{1+\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{1+\theta} Y \\
=& \left[W \left(\frac{l_{zm}\xi_{zm}}{L_m} \right)^{\frac{1}{\eta}} \left(\frac{L_m \xi_m}{L} \right)^{\frac{1}{\theta}} \xi_m \xi_{zm} \right]^{\eta} \left(\frac{1/\xi_{zm}}{W_m} \right)^{1+\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{1+\theta} Y \\
=& \left[W \left(\frac{l_{zm}\xi_{zm}}{L_m} \right)^{\frac{1}{\eta}} \left(\frac{L_m \xi_m}{L} \right)^{\frac{1}{\theta}} \xi_m \xi_{zm} \right]^{\eta} \left(\frac{1/\xi_{zm}}{W_m} \right)^{\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{\theta} \cancel{W} L \left(\frac{1/\xi_{zm}}{\cancel{W_m}} \right) \times \left(\frac{\cancel{W_m}/\xi_m}{\cancel{W}} \right) \\
=& l_{zm} \xi_m^{\eta} \cancel{\xi_{zm}^{\eta}} \left(\frac{\cancel{\xi_{zm}}}{L_m} \right) \left[W \left(\frac{L_m \xi_m}{L} \right)^{\frac{1}{\theta}} \right]^{\eta} \left(\frac{1/\cancel{\xi_{zm}}}{W_m} \right)^{\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{\theta} \left(\frac{L}{\cancel{\xi_{zm}}\xi_m} \right) \\
=& l_{zm} \left(\frac{\xi_m^{\eta}}{L_m} \right) \left[W \left(\frac{L_m \xi_m}{L} \right)^{\frac{1}{\theta}} \right]^{\eta} \left(\frac{1}{W_m} \right)^{\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{\theta} \left(\frac{L}{\xi_m} \right) \\
=& l_{zm} \left(\frac{1}{L_m} \right) \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta}{\theta}} \left(\frac{W \xi_m}{W_m} \right)^{\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{\theta} \left(\frac{L}{\xi_m} \right) \\
=& l_{zm} \left(\frac{W \xi_m}{W_m} \right)^{\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{\theta} \left(\frac{L}{L_m \xi_m} \right) \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta}{\theta}} \\
=& l_{zm} \left(\frac{W}{W_m/\xi_m} \right)^{\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{\theta} \left(\frac{L}{L_m \xi_m} \right) \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta}{\theta}} \\
=& l_{zm} \left(\frac{W_m/\xi_m}{W} \right)^{\theta-\eta} \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta}{\theta}-1} \\
=& l_{zm} \left(\frac{W}{W_m/\xi_m} \right)^{\eta-\theta} \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta-\theta}{\theta}} \\
=& l_{zm} \left(\frac{W}{W_m/\xi_m} \right)^{\eta-\theta} \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta-\theta}{\theta}}
\end{aligned}
$$

Second, I show that $\left(\frac{W}{W_m/\xi_m} \right)^{\eta-\theta} \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta-\theta}{\theta}} = 1$ by expressing the CES wage index W_m as a function of labor and taste shifters, which can be done by first plugging in the expression for

w_{zm} into the definition of W_m:

$$
\begin{aligned}
W_m &= \left[\sum_{k\in B_n}\left(\frac{w_{zm}}{\xi_{km}}\right)^{1+\eta}\right]^{\frac{1}{1+\eta}} \\
&= \left[\sum_{k\in B_n}\left(\frac{\left[W\left(\frac{l_{km}\xi_{km}}{L_m}\right)^{\frac{1}{\eta}}\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\xi_m\xi_{km}\right]}{\xi_{km}}\right)^{1+\eta}\right]^{\frac{1}{1+\eta}} \\
&= \left[W^{1+\eta}\left(\frac{1}{L_m}\right)^{\frac{1+\eta}{\eta}}\left[\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\xi_m\right]^{1+\eta}\sum_{k\in B_n}\left(\frac{\left[(l_{km}\xi_{km})^{\frac{1}{\eta}}\cancel{\xi_{km}}\right]}{\cancel{\xi_{km}}}\right)^{1+\eta}\right]^{\frac{1}{1+\eta}} \\
&= \frac{W}{L_m^{\frac{1}{\eta}}}\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\xi_m\left[\sum_{k\in B_n}(l_{km}\xi_{km})^{\frac{1+\eta}{\eta}}\right]^{\frac{1}{1+\eta}} \\
&= \frac{W}{\cancel{L_m^{\frac{1}{\eta}}}}\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\xi_m\cancel{L_m^{\frac{1}{\eta}}} \\
&= W\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\xi_m
\end{aligned}
$$

Thus, $W_m = W\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\xi_m$. Recall from the first step that completing the proof requires showing that $\left(\frac{W}{W_m/\xi_m}\right)^{\eta-\theta}\left(\frac{L_m\xi_m}{L}\right)^{\frac{\eta-\theta}{\theta}} = 1$. Plugging in the expression for W_m into this equation gives:

$$
\begin{aligned}
\left(\frac{W}{W_m/\xi_m}\right)^{\eta-\theta}\left(\frac{L_m\xi_m}{L}\right)^{\frac{\eta-\theta}{\theta}} &= \left(\frac{\cancel{W}}{\left[\cancel{W}\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\cancel{\xi_m}\right]/\cancel{\xi_m}}\right)^{\eta-\theta}\left(\frac{L_m\xi_m}{L}\right)^{\frac{\eta-\theta}{\theta}} \\
&= \left(\frac{L_m\xi_m}{L}\right)^{-\frac{(\eta-\theta)}{\theta}}\left(\frac{L_m\xi_m}{L}\right)^{\frac{\eta-\theta}{\theta}} \\
&= 1
\end{aligned}
$$

which completes the proof that w_{zm} is the inverse function of l_{zm}.

B.2.2 Equation 7: $s_{zm} \equiv \frac{w_{zm}l_{zm}}{\sum_k (w_{km}l_{km})} = \frac{\partial \ln L_m}{\partial \ln l_{zm}}$.

To see why this holds, start from the definition of the labor market index L_m in Section 2 to derive $\partial \ln L_m / \partial \ln l_{zm}$ as

$$\frac{\partial \ln L_m}{\partial \ln l_{zm}} = \frac{(\xi_{km} l_{km})^{\frac{1+\eta}{\eta}}}{\sum_{j=1}^{N_m} (\xi_{jm} l_{jm})^{\frac{1+\eta}{\eta}}}$$

Now set this aside. Plug in equation 3 to the definition $s_{zm} \equiv w_{zm}l_{zm} / \sum_k (w_{km}l_{km})$ to obtain

$$s_{zm} = \frac{(\xi_{km} l_{km})^{\frac{1+\eta}{\eta}}}{\sum_{j=1}^{N_m} (\xi_{jm} l_{jm})^{\frac{1+\eta}{\eta}}}$$

Therefore, $s_{zm} = \partial \ln L_m / \partial \ln l_{zm}$.

B.2.3 Proposition 1: $\mu_m \equiv \frac{\bar{r}_m}{\bar{w}_m} = 1 + \varepsilon_m^{-1} = 1 + \frac{1}{\theta} HHI_m + \frac{1}{\eta}(1 - HHI_m)$,

In this expression, $\bar{w}_m$ and $\bar{r}_m$ are the (employment-weighted) average wage and average marginal revenue product of labor in market m, respectively.

First, I show that $1 + \varepsilon_m^{-1} = 1 + \frac{1}{\theta} HHI_m + \frac{1}{\eta}(1 - HHI_m)$. To see why this holds, let Θ_m denote the set of firms operating in labor market m, and take the (payroll-share-weighted) average of equation 8:

$$\begin{aligned}
\underbrace{\sum_{z \in \Theta_m} s_{zm} \left(1 + \varepsilon_{zm}^{-1}\right)}_{\equiv 1 + \varepsilon_m^{-1}} &= 1 + \sum_{z \in \Theta_m} s_{zm} \left[\frac{1}{\eta}(1 - s_{zm}) + \frac{1}{\theta} s_{zm}\right] \\
&= 1 + \sum_z \left[\frac{1}{\theta} s_{zm}^2 + \frac{1}{\eta}\left(s_{zm} - s_{zm}^2\right)\right] \\
&= 1 + \frac{1}{\theta} HHI_m + \frac{1}{\eta}(1 - HHI_m)
\end{aligned}$$

Second, I show that $\sum_{z \in \Theta_m} s_{zm} \left(1 + \varepsilon_{zm}^{-1}\right) = \frac{\bar{r}_m}{\bar{w}_m}$. To see that this equality holds, aggregate the

firm-level markdown equation $\frac{r_{zm}}{w_{zm}} = 1 + \varepsilon_{zm}^{-1}$ using payroll shares as weights to get:

$$\begin{aligned}
\underbrace{\sum_{z\in\Theta_m} s_{zm}\left(1+\varepsilon_{zm}^{-1}\right)}_{\equiv 1+\varepsilon_m^{-1}} &= \sum_{z\in\Theta_m} s_{zm}\left(\frac{r_{zm}}{w_{zm}}\right) \\
&= \sum_{z\in\Theta_m} \frac{\cancel{w_{zm}} l_{zm}}{\sum_j w_{jm} l_{jm}}\left(\frac{r_{zm}}{\cancel{w_{zm}}}\right) \\
&= \frac{\sum_{z\in\Theta_m} r_{zm} l_{zm}}{\sum_{j\in\Theta_m} w_{jm} l_{jm}} \\
&= \frac{\left(\sum_{z\in\Theta_m} r_{zm} l_{zm}\right) / \left(\sum_{z\in\Theta_m} l_{zm}\right)}{\left(\sum_{j\in\Theta_m} w_{jm} l_{jm}\right) / \left(\sum_{z\in\Theta_m} l_{zm}\right)} \\
&= \frac{\bar{r}_m}{\bar{w}_m} \equiv \mu_m
\end{aligned}$$

B.2.4 Country-level average wage markdown

I show that a particular country-level average of the market-level average wage markdown (i.e., equation 9) equals the country-level average (employment-weighted) wage markdown. The reader can then directly verify that the resulting expression is the inverse of Berger, Herkenhoff and Mongey (2022)'s expression for the "labor market power adjustment" component of the country-level labor share (see authors' equation 10).

Corollary. *Consider the market-level average wage markdown expression from Proposition 1. Then the country-level (employment-weighted) average wage markdown is given by:*

$$\mu \equiv \frac{\bar{r}}{\bar{w}} = \sum_m s_m \mu_m = 1 + \frac{1}{\theta} H\tilde{H}I + \frac{1}{\eta}\left(1 - H\tilde{H}I\right)$$

where $s_m = \frac{\bar{w}_m l_m}{\sum_m \bar{w}_m l_m}$ is market m's payroll share, $H\tilde{H}I = \sum_m s_m HHI_m$ is the country-level payroll-share-weighted average payroll Herfindahl, and $\bar{w}$ and $\bar{r}$ are the (employment-weighted) average wage and average marginal revenue product of labor at the country level, respectively.

Proof. Having provided a more detailed proof for Proposition 1, I use the same steps to show the

country-level aggregation result more directly. In particular:

$$\begin{aligned}
\mu \equiv \frac{\bar{r}}{\bar{w}} =& \frac{\left(\sum_m \bar{r}_m l_m\right) / \left(\sum_m l_m\right)}{\left(\sum_m \bar{w}_m l_m\right) / \left(\sum_m l_m\right)} \\
=& \frac{\sum_m \bar{r}_m l_m}{\sum_m \bar{w}_m l_m} \\
=& \frac{\sum_m \left(\frac{\bar{r}_m}{\bar{w}_m}\right) \bar{w}_m l_m}{\sum_m \bar{w}_m l_m} \\
=& \sum_m \left(\frac{\bar{r}_m}{\bar{w}_m}\right) \frac{\bar{w}_m l_m}{\sum_m \bar{w}_m l_m} \\
=& \sum_m \mu_m s_m \\
=& \sum_m s_m \left[1 + \frac{1}{\theta} HHI_m + \frac{1}{\eta}\left(1 - HHI_m\right)\right] \\
=& 1 + \frac{1}{\theta} \tilde{HHI} + \frac{1}{\eta}\left(1 - \tilde{HHI}\right)
\end{aligned}$$

□

B.2.5 Corollary 1: $\gamma_t \equiv \frac{d\mu_{mt}}{dX} = \left(\frac{1}{\theta} - \frac{1}{\eta}\right)\beta_t$

In this equation, β_t is the effect of an exogenous shock on the payroll Herfindahl. To derive the expression, plug in $\mu_{mt} \equiv 1 + \varepsilon_{mt}^{-1}$ and differentiate:

$$\begin{aligned}
\gamma_t \equiv \frac{d\mu_{mt}}{dX} =& \frac{d\left(1 + \varepsilon_{mt}^{-1}\right)}{dX} \\
=& \left[\frac{d\left(1 + \varepsilon_{mt}^{-1}\right)}{dHHI_{mt}} \cdot \frac{dHHI_{mt}}{dX}\right] \\
=& \left[\frac{d\left(1 + \varepsilon_{mt}^{-1}\right)}{dHHI_{mt}} \cdot \beta_t\right] \\
=& \left(\frac{1}{\theta} - \frac{1}{\eta}\right)\beta_t
\end{aligned}$$

I then compute standard errors for γ_t under the assumption that the effect on concentration and the labor supply parameters are independent. It follows that:

$$\begin{aligned}
\text{Var}(\gamma_t) =& \text{Var}\left[\left(\frac{1}{\theta}-\frac{1}{\eta}\right)\cdot\beta_t\right] \\
=& E\left[\left(\frac{1}{\theta}-\frac{1}{\eta}\right)^2\right]E\left[\beta_t^2\right]-\left[E\left(\frac{1}{\theta}-\frac{1}{\eta}\right)\right]^2\left[E\left(\beta_t\right)\right]^2 \\
=& \left[\text{Var}\left(\frac{1}{\theta}-\frac{1}{\eta}\right)+\left[E\left(\frac{1}{\theta}-\frac{1}{\eta}\right)\right]^2\right]\left[\text{Var}\left(\beta_t\right)+\left[E\left(\beta_t\right)\right]^2\right]-\left[E\left(\frac{1}{\theta}-\frac{1}{\eta}\right)\right]^2\left[E\left(\beta_t\right)\right]^2
\end{aligned}$$

whose components can all be plugged in using sample estimates.

B.2.6 Equation 9 under the setup in BHM

I show that equation 9 holds under the additional assumptions on production function and goods market structure in Berger, Herkenhoff and Mongey (2022), henceforth BHM. In that environment, μ_m should be interpreted as the ratio of the average marginal revenue (net of expenditures in non-labor inputs) to the average wage. I show this in two steps.

To start, consider the environment in BHM. Goods markets are perfectly competitive, with $p_{zm} = 1$ for all firms and markets. Firms compete for labor à la Cournot, solving:

$$\max_{k_{zm},l_{zm}} \pi_{zm} = \underbrace{A_{zm}\left(k_{zm}^{1-\gamma}l_{zm}^{\gamma}\right)^{\alpha}}_{\equiv y_{zm}} - Rk_{zm} - w_{zm}\left(\{l_{zm}, l_{-zm}\}\right)l_{zm} \tag{B.4}$$

where y_{zm} is firm revenues, k_{zm} is capital, A_{zm} is a general firm-market-specific productivity term, R is the rental rate of capital (in perfectly competitive capital markets), and w_{zm} is the wage firm z would have to pay to obtain l_{zm} units of labor, given nested CES labor supply preferences that yield the same expression for ε_{zm}^{-1}, the firm-specific inverse elasticity of residual supply, as derived in Section 2.1.

First, I show that equation 9 holds when the firm optimizes labor *holding capital constant*, denoting this corresponding average wage markdown by $\mu_m^{k-fixed}$. To avoid confusion due to differences in notation, let $mrpl_{zm}^{k-fixed}$ denote BHM's expression for the marginal revenue product of labor of firm z in market m *holding capital constant*, and continue to use the Greek letter μ to denote the wage markdown.[59] Computing the definition of market m's average wage markdown

[59] In BHM, the Greek letter μ refers to the wage take-home share (i.e., the inverse of the wage markdown) holding capital constant.

holding capital constant gives:

$$\mu_m^{k-fixed} \equiv \frac{mrpl_m^{\bar{k}-fixed}}{\bar{w}_m} = \frac{\left(\sum_z mrpl_{zm}^{k-fixed} l_{zm}\right) / \sum_z l_{zm}}{\left(\sum_z w_{zm} l_{zm}\right) / \sum_z l_{zm}} \tag{B.5}$$

$$= \frac{\sum_z \alpha\gamma \left(y_{zm}/l_{zm}\right) l_{zm}}{\sum_z w_{zm} l_{zm}} \tag{B.6}$$

$$= \alpha\gamma \frac{\sum_z y_{zm}}{\sum_z w_{zm} l_{zm}} \tag{B.7}$$

where $mrpl_{zm}^{k-fixed} = \partial y_{zm}/\partial l_{zm}|_k = \alpha\gamma \left(y_{zm}/l_{zm}\right)$.

Simplification of equation B.7 can now be done using the equalities in BHM's Proposition 1.1 (with special care given to note the difference in notation across the two papers). Equation B.7 becomes:

$$\begin{aligned}
\mu_m^{k-fixed} = \alpha\gamma \frac{\sum_z y_{zm}}{\sum_z w_{zm} l_{zm}} &= \alpha\gamma \left[\frac{1}{\alpha\gamma} \sum_z s_{zm} \mu_{zm}^{k-fixed}\right] \\
&= \sum_z s_{zm} \mu_{zm}^{k-fixed} \\
&= \sum_z s_{zm} \left[1 + \frac{1}{\theta} s_{zm} + \frac{1}{\eta}\left(1 - s_{zm}\right)\right] \\
&= 1 + \frac{1}{\theta} HHI_m + \frac{1}{\eta}\left(1 - HHI_m\right)
\end{aligned}$$

where HHI_m is similarly defined as the payroll Herfindahl of labor market m.

Second, I show that, in equilibrium, $\mu_{zm}^{k-fixed} = \mu_{zm}^{k-adjust}$ for all firms z and markets m. In other words, equation 9 holds whether or not optimization of capital is taken into account, *so long as expenditures on capital are netted out of firm revenues*. In this case, $\mu_{zm} = \mu_{zm}^{k-adjust}$ should be interpreted as the ratio of the marginal revenue (net of expenditures on capital) product of labor to the wage.

To show this, note first that since $\mu_{zm} = mrpl_{zm}/w_{zm}$, it suffices to show that in equilibrium $mrpl_{zm}^{k-fixed} = mrpl_{zm}^{k-adjust}$. Letting y_{zm}^{net} denote total firm revenues net of capital expenditures,

it follows that:

$$\begin{aligned} mrpl_{zm}^{k-adjust} \equiv \frac{dy_{zm}^{net}}{dl_{zm}} =& \frac{d\left[f\left(k_{zm}, l_{zm}\right) - Rk_{zm}\right]}{dl_{zm}} \\ =& \frac{\partial f\left(k_{zm}, l_{zm}\right)}{\partial k_{zm}} \frac{dk_{zm}}{dl_{zm}} + \frac{\partial f\left(k_{zm}, l_{zm}\right)}{\partial l_{zm}} - R\frac{dk_{zm}}{dl_{zm}} \\ =& \frac{\partial f\left(k_{zm}, l_{zm}\right)}{\partial l_{zm}} + \frac{dk_{zm}}{dl_{zm}} \underbrace{\left(\frac{\partial f\left(k_{zm}, l_{zm}\right)}{\partial k_{zm}} - R\right)}_{\text{=0 by firm's FOC for } k_{zm}} \\ =& mrpl_{zm}^{k-fixed} \end{aligned}$$

This result, $mrpl_{zm}^{k-fixed} = mrpl_{zm}^{k-adjust}$, also follows directly from the envelope theorem, as the firm is optimizing its non-labor inputs.

B.3 Wage effect decomposition: identities and additional estimates

This appendix states the identities behind the decomposition of Section 7 and reports the estimates that the main text summarizes.

The identity holds for a market's average wage, $\bar{w}_{mt} = \mu_{mt}^{-1}\bar{r}_{mt}$ (equation 9), where μ_{mt}^{-1} is the market average wage take-home share, which depends on the market's payroll Herfindahl, and $\bar{r}_{mt}$ is the market average marginal revenue product of labor. I can therefore decompose the effect of trade liberalization on average wages as:

$$\frac{d\bar{w}_{mt}}{dICE_m} = \frac{d\mu_{mt}^{-1}}{dICE_m}\bar{r}_{mt} + \frac{d\bar{r}_{mt}}{dICE_m}\mu_{mt}^{-1} = -\underbrace{\frac{d\mu_{mt}}{dICE_m}}_{\gamma_t}\frac{1}{\mu_{mt}^2}\bar{r}_{mt} + \frac{d\bar{r}_{mt}}{dICE_m}\mu_{mt}^{-1} \tag{B.8}$$

where $\gamma_t = \left(\frac{1}{\theta} - \frac{1}{\eta}\right)\beta_t$ from Corollary 1 is the effect of changes in import competition exposure on the average wage markdown in local labor market m. The first term is the markdown channel, which Table 4 reports; the second nets changes in the marginal revenue product within firms against changes in the composition of employment across firms, including the firms that enter and exit.

The wage effect that I identify in Section 6 is defined not on the market average, which changes with the firms that enter and exit, but rather on firms present in both years: the market wage premium index ψ_m, the outcome of the reduced form that identifies $1/\theta$, is the 1991-employment-share-weighted mean over incumbent firms of the change in the firm's log wage premium net of its move along its own labor supply curve, $\psi_m = \sum_{z \in S_m} s_{zm0}^e \left[\Delta \ln w_{zm} - \frac{1}{\eta}\Delta \ln l_{zm}\right]$, with S_m the set of incumbents and the shares normalized to sum to one over S_m. For these firms, the wage-setting

equation gives the decomposition firm by firm, and averaging it at baseline shares gives, in logs:

$$\sum_{z\in S_m} s^e_{zm0}\,\Delta\ln w_{zm} = \underbrace{\sum_{z\in S_m} s^e_{zm0}\,\Delta\ln \mu^{-1}_{zm}}_{\text{Via markdowns}} + \underbrace{\sum_{z\in S_m} s^e_{zm0}\,\Delta\ln r_{zm}}_{\text{Via MRPL}} = \psi_m + \frac{1}{\eta}\sum_{z\in S_m} s^e_{zm0}\,\Delta\ln l_{zm}, \qquad \text{(B.9)}$$

where each firm's take-home share μ^{-1}_{zm} follows from its payroll share and the estimated elasticities (equation 8), and its marginal revenue product is its wage premium divided by its take-home share. The identity is exact, so the effect of trade liberalization on incumbents' average wage premium is the sum of its effects on the two terms, and the index differs from that average by incumbents' moves along their labor supply curves.

Across all of a market's firms, the change in the average marginal revenue product, $\bar{r}_{mt} = \sum_z s^e_{zmt} r_{zmt}$, splits into a within-firm and a composition term:

$$\bar{r}_{mt} - \bar{r}_{m0} = \underbrace{\sum_{z\in S_m} s^e_{zm0}\,(r_{zmt} - r_{zm0})}_{\text{Within firms}} + \underbrace{\sum_{z} r_{zmt}\left(s^e_{zmt} - s^e_{zm0}\right)}_{\text{Composition of employment, including entry and exit}}, \qquad \text{(B.10)}$$

where entrants have $s^e_{zm0} = 0$ and exiters $s^e_{zmt} = 0$, and the composition term splits into the change in incumbents' employment shares and the contributions of entrants and exiters. I divide both terms by the logarithmic mean of $\bar{r}_{mt}$ and $\bar{r}_{m0}$, so that they add up to $\Delta\ln\bar{r}_{mt}$ in every market.

The markdown channel and the US benchmark. The market's average take-home share depends on the payroll Herfindahl of all of its firms, including those that exited, and thus falls by more (0.7%) than the average take-home share of incumbent firms (0.38%; Table 5). With the elasticities that BHM estimate for the US, the take-home share would have fallen by 2.8 cents, calculated as $2.8\% \approx \left[\mu_{0,US} + (1/0.42 - 1/10.85) * 0.021\right]^{-1} - \mu^{-1}_{0,US}$, where $\mu_{0,US}$ is the implied pre-liberalization average wage markdown evaluated at Brazil's payroll-weighted average labor market concentration of 0.08 but US elasticities of substitution from BHM ($\theta = 0.42, \eta = 10.85$): $\mu_{0,US} = 1 + (1/0.42) * 0.08 + (1 - 0.08) * (1/10.85)$.

The wage decline at firms facing larger tariff cuts. Within markets, firms whose tariffs fell by more reduced their wage premia relative to their competitors: a one log point larger tariff cut reduced a firm's premium by 0.61 log points (Table 2) and its employment by 0.74. In Appendix Table A.12, I decompose this wage response using the firm-level identity $\Delta\ln w_{zm} = \Delta\ln\mu^{-1}_{zm} + \Delta\ln r_{zm}$, where the firm's take-home share follows from its payroll share and the estimated elasticities (equation 8). Under the weights of Table 2, the market's 1991 employment, the largest markets dominate and a firm's payroll share is close to zero, so these firms' take-home share did not change, and their marginal revenue product of labor fell by the whole 0.61 log points (Panel A). Panel B

reports, for comparison, the effect of the market's exposure on its average firm wage premium and employment, from Table 1. The wage decline at firms facing larger tariff cuts is thus driven by a lower marginal revenue product of labor rather than by greater labor market power. Since firm productivity rose during Brazil's liberalization (Muendler, 2004), and the within-market elasticity is nearly unchanged when I measure tariff shocks as effective rates of protection, which net out input tariff reductions (Appendix E), these MRPL reductions more likely reflect lower output prices than productivity losses, consistent with the pro-competitive effects of import competition predicted by Melitz and Ottaviano (2008) and with the productivity gains documented in India's comparable episode by Topalova and Khandelwal (2011).

Appendix Figure A.11 plots the effect on incumbents' wage premia year by year, with markets weighted by their 1991 formal employment as in Table 1, rather than counted once as in Table 5. After 1991, incumbents' wage premia decline through 1994 and remain lower through 2000 (−0.22 per unit of ΔICE_m, SE of 0.08). Before liberalization, wages in more exposed markets are not stable: they grew faster between 1988 or 1989 and 1991, but not between 1986 or 1990 and 1991. The decline keeps its sign with state-by-year fixed effects and the market's 1986–1990 wage change (−0.16, SE of 0.07) and with microregion-by-year fixed effects (−0.13, SE of 0.08), but these controls do not remove the pre-liberalization movements. The two components of Table 5 are thus not equally well established. The markdown component, 0.4 percentage points, follows from the increase in concentration, which is stable before liberalization (Figure 5), and from the two elasticities. The total wage decline is less certain: across these specifications it ranges from 1.1% to 2.2% per 10% increase in exposure, so that higher markdowns account for between one sixth and one third of it.

C Import Competition Exposure: Theory and Identification

ΔICE_m weights the tariff reductions faced by a market's tradable firms by the firms' contributions to the market's baseline payroll Herfindahl, which determines the market's average markdown (Proposition 1). These are intuitive weights for studying the effects of trade liberalization on concentration: the tariff reductions that count the most are those faced by the firms that account for most of the market's concentration. A reasonable question is how close these weights are to those of the measure of exposure that the model itself implies. In this Appendix, I derive the model's measure and relate ΔICE_m to it.

The derivation requires assumptions that the rest of the paper does not make. The model of Section 2 specifies labor supply and how firms set wages given their labor demand, but it leaves firms' revenue functions unrestricted. It is therefore silent on how much a tariff reduction shifts a firm's labor demand, and on how much the firm's employment responds to that shift. A fully theory-driven, equilibrium-based measure of a market's exposure to trade liberalization, however, must aggregate the tariff reductions of the market's firms according to their effects on the market's equilibrium, which requires taking a stand on both. The assumptions below do so.

Note, however, that my estimates of the paper's key parameters do not depend on these additional assumptions. Per the model, any shock to labor demand that is unrelated to changes in workers' tastes identifies $1/\eta$ and $1/\theta$: the shock need not be a sufficient statistic for the effect of trade liberalization on a local labor market. Proposition 1 and Corollary 1 hold without them as well. The derivation is instead useful for interpreting the effects of ΔICE_m on local labor markets, because it shows what a model-based measure of exposure looks like and how ΔICE_m differs from it.

In the derivation, markdowns respond to the tariff reductions, as in the model of Section 2. The model's measure, ΔICE_m^{ω}, is an average of the tariff reductions faced by the tradable firms of a market, with weights that depend only on the firms' 1991 payroll shares. ΔICE_m coincides with it under conditions that I state, and approximates it otherwise. I also discuss what the derivation implies for concentration.

I then discuss how ΔICE_m and the model's measure relate to the most commonly used measure of exposure in analyses of Brazil's trade liberalization, the RTR (regional tariff reductions) of Kovak (2013) and Dix-Carneiro and Kovak (2017). Finally, I document which variation in ΔICE_m identifies my market-level estimates, and compare it with the variation in the RTR. Proofs are at the end of this Appendix.

C.1 The model's measure of exposure

This section derives ΔICE_m^{ω}, the model's measure of market m's exposure to the tariff reductions: the average of the tariff reductions faced by the market's tradable firms to which the changes in the

market's labor supply and wage indices are proportional. I proceed in two steps. Proposition 2 shows how a market's indices respond to any labor demand shocks at its firms. It uses Assumptions 1 and 2, which are consistent with the model of Section 2 and narrow it to a case in which this response has a closed form. Corollary 2 then obtains ΔICE_m^ω with Assumptions 3 to 5, which are new: they specify how tariff reductions shift firms' labor demand, which Section 2 leaves unspecified. The next section compares ΔICE_m^ω with ΔICE_m.

I use the model of Section 2 as is: labor supply (equation 3), wage setting (equation 5), and markdowns μ_{zm} (equation 8). Let $r_{zm} \equiv \partial R_z / \partial l_{zm}$ denote firm z's marginal revenue product of labor in market m. I compare the equilibria before and after the tariff reductions, to first order. Wages, employment, payroll shares, and markdowns all change between the two equilibria; the payroll shares s_{zm} that appear in the weights below are those of 1991.

Assumption 1 (No other changes). The taste shifters ξ_{zm} and ξ_m, the country-level indices W and L, and the set of firms operating in each market do not change with tariffs.

Assumption 2 (Labor demand). The marginal revenue product of labor of firm z in market m satisfies

$$d \ln r_{zm} = x_{zm} - \frac{1}{e_{zm}} d \ln l_{zm}, \tag{C.1}$$

where x_{zm} is the firm's labor demand shock, the change in its marginal revenue product at a given level of employment, and $e_{zm} > 0$ is the elasticity of its labor demand.

The first part of Assumption 1 is the exclusion restriction of Section 5. Its second part is a normalization when the elasticity of labor demand is the same for all firms: the country-level terms are then common to all markets and absorbed by the constants of the market-level regressions. Its third part excludes the entry and exit of firms, and with them the changes in the set of markets in which a firm operates that Section 2.2 allows. Assumption 2 restricts the revenue function R_z of equation 4, on which the rest of the paper remains agnostic: the marginal revenue product of labor in market m depends neither on the employment of the other firms in that market, as when the firms that share a local labor market sell their output in national or world markets, nor on firm z's employment in its other markets.

Proposition 2. *Under Assumptions 1 and 2,*

$$d \ln L_m = \frac{\sum_{z \in \Theta_m} \omega_{zm} \, x_{zm}}{\sum_{z \in \Theta_m} \omega_{zm} / e_{zm} + \frac{1}{\theta}}, \qquad d \ln W_m = \frac{1}{\theta} \, d \ln L_m, \tag{C.2}$$

where

$$\omega_{zm} \equiv \frac{s_{zm}/D_{zm}}{\sum_{k \in \Theta_m} s_{km}/D_{km}}, \qquad D_{zm} \equiv \frac{1}{e_{zm}} + \frac{1}{\eta} + \left(\frac{1}{\theta} - \frac{1}{\eta}\right)\left(1 + \frac{1}{\eta}\right) \frac{s_{zm}}{\mu_{zm}}, \tag{C.3}$$

provided that $D_{zm} > 0$ for all firms, which holds when $1/\theta \geq 1/\eta$.

Proof. See Appendix C.6. □

The labor demand shocks of a market's firms affect the market's labor supply and wage indices only through their weighted average. A firm's weight is its payroll share, its marginal effect on the market's labor supply index (equation 7), divided by D_{zm}, whose inverse is the response of the firm's employment to its own shock for a given L_m: smaller when the firm's labor demand is less elastic, when labor supply to the firm is less elastic, and when the firm's markdown rises with its payroll share. Since the markdown μ_{zm} is a function of the payroll share, the weights depend only on the firms' 1991 payroll shares, given the elasticities.

For the second step, let $\Delta\tau_z \equiv \Delta \ln(1 + \tau_{i(z)})$ denote the tariff reduction faced by firm z (equation 17), positive for a tariff cut, and Θ_m^T the tradable firms of market m, as in Section 4.1.

Assumption 3 (Pass-through). For tradable firms, $x_{zm} = -\psi \, \Delta\tau_z$, where $\psi > 0$ is the pass-through of tariff reductions to labor demand, common to all firms.

Assumption 4 (Common elasticity). The elasticity of labor demand is the same for all firms, $e_{zm} = e$.

Assumption 5 (Non-tradable firms). For firms outside Θ_m^T, $x_{zm} = \bar{x}_m^T \equiv \sum_{k \in \Theta_m^T} (\omega_{km}/\omega_m^T) \, x_{km}$, the average labor demand shock of the market's tradable firms with the weights of equation C.3, where $\omega_m^T \equiv \sum_{k \in \Theta_m^T} \omega_{km}$, with $x_{zm} = 0$ in markets without tradable firms.

Assumption 5 specifies the shocks of non-tradable firms, whose output demand is local, without modeling local goods markets.

Corollary 2. *Under Assumptions 1 to 5, $d \ln W_m = \frac{1}{\theta} \, d \ln L_m$ and*

$$d \ln L_m = -\frac{\psi}{\frac{1}{e} + \frac{1}{\theta}} \, \Delta ICE_m^{\omega}, \tag{C.4}$$

where

$$\Delta ICE_m^{\omega} \equiv \sum_{z \in \Theta_m^T} \frac{\omega_{zm}}{\omega_m^T} \, \Delta\tau_z \tag{C.5}$$

in markets with tradable firms, and $\Delta ICE_m^{\omega} \equiv 0$ in markets without them.

Proof. See Appendix C.6. □

By Corollary 2, ΔICE_m^{ω} is the model's measure of the market's exposure: the changes in the market's wage and labor supply indices are proportional to it, with the same factor in all markets. Equation C.4 is the model's counterpart of the first stage in equation 20. With a common elasticity of labor demand, D_{zm} varies across the firms of a market only with s_{zm}/μ_{zm}, so that when $1/\theta > 1/\eta$ a firm's weight increases less than proportionally with its payroll share.

C.2 Relation to ΔICE_m

Let $s_m^T \equiv \sum_{z \in \Theta_m^T} s_{zm}$ denote the share of tradable firms in the market's payroll. The payroll-share-weighted average of the tariff reductions of the market's tradable firms,

$$\Delta ICE_m^{\text{lin}} \equiv \sum_{z \in \Theta_m^T} \frac{s_{zm}}{s_m^T} \Delta\tau_z, \tag{C.6}$$

replaces D_{zm} with a value common to the firms of a market. It equals ΔICE_m^{ω} when $1/\theta = 1/\eta$, when the tradable firms of a market have equal payroll shares, or when they face the same tariff reduction, and it approximates it otherwise: at the estimates of Section 6, the term of D_{zm} that varies with the payroll share is 0.06 for a firm with a payroll share of 10% and 0.27 for a share of 50%, against $1/e + 1/\eta \geq 0.83$. The measure of Section 4, ΔICE_m of equation 11, averages the same tariff reductions over the same firms with weights κ_{zm} proportional to squared payroll shares, and

$$\Delta ICE_m = \Delta ICE_m^{\text{lin}} + \frac{s_m^T}{HHI_m^T} \text{Cov}_m^T(s_{zm}, \Delta\tau_z), \tag{C.7}$$

where $HHI_m^T \equiv \sum_{z \in \Theta_m^T} s_{zm}^2$ and Cov_m^T is the payroll-share-weighted covariance across the tradable firms of market m.

Assumption 6 (Tariff reductions and firm size). Across the tradable firms of each market, tariff reductions are uncorrelated with payroll shares: $\text{Cov}_m^T(s_{zm}, \Delta\tau_z) = 0$.

Under Assumption 6, $\Delta ICE_m = \Delta ICE_m^{\text{lin}}$; in markets whose tradable firms face the same tariff reduction, the three measures coincide. Where Assumption 6 does not hold, ΔICE_m gives the tariff reductions of the market's larger tradable firms more weight than their payroll shares, whereas the model's measure gives them somewhat less. ΔICE_m is thus an approximation to the model's measure of exposure, not the measure itself. Column (2) of Appendix Table A.4 reports the estimates of Section 4 with $\Delta ICE_m^{\text{lin}}$ in place of ΔICE_m.

C.3 Concentration

Proposition 3. *Under Assumptions 1 and 2,*

$$dHHI_m = 2\left(1+\frac{1}{\eta}\right)\sum_{z\in\Theta_m}\frac{s_{zm}^2}{D_{zm}}\left[x_{zm}-\left(\frac{1}{e_{zm}}+\frac{1}{\theta}\right)d\ln L_m\right], \tag{C.8}$$

provided that $D_{zm} > 0$ *for all firms.*

Proof. See Appendix C.6. □

Concentration rises when employment falls by less at firms with larger payroll shares. With a common pass-through and elasticity of labor demand (Assumptions 3 to 5), equation C.8 becomes

$$dHHI_m = -2\psi\left(1+\frac{1}{\eta}\right)\sum_{z\in\Theta_m^T}\frac{s_{zm}^2}{D_{zm}}\left(\Delta\tau_z-\Delta ICE_m^{\omega}\right):$$

a tariff reduction common to the tradable firms of a market leaves the Herfindahl unchanged, and concentration rises when the tradable firms with larger payroll shares face smaller tariff reductions than the market's average. The increase in concentration estimated in Section 4 can also arise from departures from these assumptions. One is firm heterogeneity in pass-through or in labor demand elasticities, which Propositions 2 and 3 accommodate: a lower pass-through at larger firms, for example because they export part of their output, or less elastic labor demand at larger firms, as with the linear demand of Melitz and Ottaviano (2008). The other is the entry and exit of firms, which Assumption 1 excludes and which Section 4.3 shows to be important.

C.4 Relation to Kovak (2013)

I now relate the model's measure to the exposure measure of Kovak (2013), also used by Dix-Carneiro and Kovak (2017): RTR (regional tariff reductions). In Kovak (2013), a region produces goods i with labor and industry-specific factors in fixed supply; labor is mobile across the region's industries, immobile across regions, and supplied inelastically; technology has constant returns to scale, all markets are perfectly competitive, and the region takes the prices P_i of tradable goods as given. Clearing of the regional labor market implies

$$d\ln w = \sum_i \beta_i\, d\ln P_i, \qquad \beta_i \equiv \frac{\lambda_i\, e_i}{\sum_{i'}\lambda_{i'}\, e_{i'}}, \tag{C.9}$$

where w is the regional wage, λ_i is industry i's share of regional employment, and e_i is the elasticity of the industry's labor demand. Kovak (2013) measures price changes of tradable goods

with tariff changes, $d \ln P_i = d \ln(1 + \tau_i)$, drops the non-tradable sector, whose price moves with the region's tradable prices under identical Cobb-Douglas preferences, and rescales the weights of tradable industries to sum to one. The result, RTR, is his model's measure of the region's exposure to liberalization.

ΔICE_m^ω has the same form, with firms in place of industries. In both models, exposure is a weighted average of tariff reductions, and the weight of each unit is the product of two terms: its size in the labor market, and how much its employment responds to a shock to its own labor demand. In equation C.9, these are the industry's share of regional employment, λ_i, and the elasticity of its labor demand, e_i: the tariff reduction of an industry counts more when the industry employs more of the region's workers, and when its employment falls by more for a given decline in its price. In equation C.3, they are the firm's payroll share, s_{zm}, which is the firm's marginal effect on the market's labor supply index (equation 7), and $1/D_{zm}$, the response of the firm's employment to its own labor demand shock.

The two models differ in what determines this response. In Kovak (2013), an industry hires any number of workers at the regional wage, so its employment response depends only on the elasticity of its labor demand. In my model, a firm faces an upward-sloping labor supply curve and sets its wage, which adds two terms to D_{zm}. The first, $1/\eta$, reflects that a firm that reduces employment moves down its labor supply curve: its wage falls, which dampens the decline in its employment. The second, $\left(\frac{1}{\theta} - \frac{1}{\eta}\right)\left(1 + \frac{1}{\eta}\right)\frac{s_{zm}}{\mu_{zm}}$, reflects that the firm's markdown increases with its payroll share when $1/\theta > 1/\eta$ (equation 8): a firm that loses payroll share marks wages down by less, which further dampens the decline in its employment, the more so the larger the firm. This second term is why the weights of ΔICE_m^ω rise less than proportionally with payroll shares.

The two models also differ in how a market absorbs the average shock to its firms. In Kovak (2013), labor does not move across regions, so the regional wage absorbs the shock in full. In my model, workers substitute across local labor markets with a finite elasticity, so the shock moves both the market's labor supply index and its wage index, in the proportion $1/\theta$ (equation C.2).

Equation C.9 is the limit of Proposition 2 in which these differences vanish. When firms take wages as given, the markdown term drops out and $D_{zm} = 1/e_{zm} + 1/\eta$. If, in addition, workers view the firms of a market as identical ($1/\eta \to 0$, $\xi_{zm} = 1$), then $D_{zm} = 1/e_{zm}$ and payroll shares equal employment shares; and if labor does not move across markets ($1/\theta \to \infty$), $d \ln W_m = \sum_z s_{zm} e_{zm} x_{zm} \Big/ \sum_k s_{km} e_{km}$, which is equation C.9 with firms in place of industries, where x_{zm} is the change in the price of the industry's good and e_{zm} is e_i. The assumptions that take me from Proposition 2 to ΔICE_m^ω also have counterparts in Kovak (2013): Assumption 3 corresponds to his measurement of price changes with tariff changes, and Assumption 5, when markets are microregions, to his result for the price of the non-tradable good. Assumption 4 is more restrictive than his treatment of labor demand: he lets the elasticity of labor demand differ across industries and

calibrates it with each industry's factor shares, whereas I assume an elasticity common to all firms.

The two measures therefore differ in what they average: RTR averages tariff reductions over the industries of a region, with weights given by the industries' shares of the region's employment and their labor demand elasticities, and ΔICE_m over the tradable formal firms of a local labor market, with weights given by their payroll shares. Appendix C.5 documents how these differences translate into different sources of identifying variation.

C.5 Sources of identifying variation

ΔICE_m is an alternative to the exposure measure of Kovak (2013), also used by Dix-Carneiro and Kovak (2017): RTR (regional tariff reductions). In this section, I use tools for shift-share designs (Goldsmith-Pinkham, Sorkin and Swift, 2020; Borusyak, Hull and Jaravel, 2022; Borusyak and Hull, 2023) to identify which variation in ΔICE_m drives each of my market-level estimates, and thus the assumption under which each can be interpreted as a causal effect. I also report the effects of the liberalization under the RTR, using the specification of Dix-Carneiro and Kovak (2017), in my RAIS data. All estimates are relative effects: equation (12) compares more with less exposed markets, and the year fixed effects absorb effects common to all markets. Unless otherwise noted, I weight local labor markets by their 1991 formal employment, as in my main specification; Appendix D discusses the unweighted estimates. In short, I find that the effects of ΔICE_m on concentration and firm exit are identified by differences in tariff cuts across sectors, and that its effect on employment rests more on the contrast between markets with and without tradable firms in 1991. The RTR varies across regions mainly with the share of agriculture, and the estimates under the RTR do not stand out from those that I obtain when I reassign tariff cuts at random across industries.

What ΔICE_m compares

ΔICE_m averages the tariff cuts faced by a market's tradable firms, weighted by their squared payroll shares, so it varies with which sectors those firms are in and with how the market's payroll is distributed among them. It is unrelated to the microregion's 1991 share of agriculture and mining in employment. It is zero in the 15 percent of microregion $\times$ occupation markets without a tradable firm in 1991. I call these markets never-exposed and the others exposed.

To separate the sources of variation in ΔICE_m, I use three versions of it: (a) ΔICE_m; (b) ΔICE_m recentered on its expected value when I reassign tariff cuts at random across CNAE sectors within IBGE subsectors (Borusyak and Hull, 2023), which keeps only differences in tariff cuts across detailed sectors; and (c) an indicator for exposed markets, which keeps only the contrast with never-exposed markets. An effect identified by the tariff cuts survives in (b). An effect identified by composition is close to zero in (b) and appears in (c). In Tables 1 and C.2, I implement (b) with the

recentered measure as the regressor. Controlling instead for the expected value of ΔICE_m interacted with year, an alternative implementation of the same correction (Borusyak and Hull, 2023), gives 0.151 against 0.168 for the effect on concentration (Table D.3).

Concentration and firm exit under ΔICE_m

My estimates of the effects on concentration and firm exit are identified by the tariff cuts. Recentering ΔICE_m on its expected value leaves my estimates nearly unchanged: by 2000 the Herfindahl rises by 0.168 against 0.172, and the number of firms falls by 1.09 against 0.97 log points (Table 1, column (2)). When I reassign the 1990–1994 tariff changes at random across CNAE sectors, holding each market's firms and shares fixed, the placebo estimates center near zero, and the actual estimate exceeds all but 1.6 percent of them (Figure C.2, panel (a)). The effect changes little with the RTR or the microregion's agriculture share, interacted with year, as controls (0.130 and 0.126 against 0.172; Table C.1). Interpreting it as causal requires that tariff cuts be as good as randomly assigned across sectors (Borusyak, Hull and Jaravel, 2022). The Herfindahl rises by 0.054 in 1992 and by 0.129 by 1995, three quarters of its 2000 level, whereas formal employment falls only after 1995: concentration increases while tariffs are being cut, and the decline in employment follows.

Table C.1: Controls for agriculture intensity and tradable sector intensity

	Δ Payroll Herfindahl (premium-based) (1)	Δ Employment Herfindahl (2)	Δ Log formal employment (3)
Main specification (Table 1)	0.172	0.109	-0.398
	(0.042)	(0.024)	(0.186)
Control: 1991 census agricultural share × year	0.130	0.071	-0.201
	(0.040)	(0.021)	(0.163)
ΔICE without agricultural firms	0.162	0.096	-0.299
	(0.042)	(0.024)	(0.184)
Dropping top-quartile agricultural microregions	0.163	0.097	-0.411
	(0.043)	(0.024)	(0.190)
Control: Dix-Carneiro and Kovak (2017)'s RTR × year	0.126	0.068	-0.207
	(0.040)	(0.020)	(0.167)
Control: 1991 payroll share of tariff firms × year	0.180	0.116	-0.313
	(0.042)	(0.024)	(0.175)
Control: 1{tariff firms in 1991} × year	0.162	0.093	-0.390
	(0.044)	(0.025)	(0.192)
Markets with tariff firms only	0.162	0.093	-0.390
	(0.044)	(0.025)	(0.192)
Microregion × year fixed effects	0.084	0.044	-0.027
	(0.035)	(0.021)	(0.150)
Occupation × year and size decile × year FE	0.128	0.067	-0.195
	(0.040)	(0.021)	(0.142)

Notes: Standard errors in parentheses, clustered at the local labor market level. All entries report the 2000 coefficient on ΔICE_m from the stacked differences-in-differences regressions of Table 1, weighted by the local labor market's 1991 formal employment. The agricultural employment share is the microregion's 1991 share of tradable employment in agriculture, from the census. The RTR is Dix-Carneiro and Kovak (2017)'s RTR exposure measure, computed from the 1991 census. The tariff firms' payroll share is the 1991 payroll share of the market's firms in sectors with a tariff. ΔICE_m without agricultural firms excludes agricultural firms from the market's tariff firms. The top quartile of agricultural share is defined across microregions.

Employment and wages under ΔICE_m

Employment and wages. Formal employment falls in more exposed markets by 2000, mostly after 1994; the formal wage premium changes little (−0.105, SE of 0.089, per unit of ΔICE_m; Table 1), and its 1986–1989 coefficients show no trend (p-value of 0.31; Table D.3).

Source of variation and identifying assumption. The employment effect rests more on 1991 composition than the effects on concentration and on the number of firms. With the recentered measure it is −0.30 (SE of 0.21), against −0.40 (SE of 0.19) for ΔICE_m (Table 1, column (2)). To the extent that it rests on that contrast, reading it as causal requires that exposed and never-exposed markets would have followed the same paths without the liberalization (Goldsmith-Pinkham, Sorkin and Swift, 2020). Its pre-period coefficients differ from zero (Table D.3), and the removal of import licenses in 1990, the 1990–1992 recession, the 1994 Real Plan and the 1999 devaluation all fell on

the same contrast between tradable and other sectors; this is why the estimation of $1/\theta$ includes state fixed effects and 1986–1990 trends, as in Dix-Carneiro and Kovak (2017). State × year fixed effects and the market's 1986–1990 trend halve the employment effect, to −0.194 (SE of 0.145), while the effects on concentration and on the number of firms remain statistically significant (Appendix Table A.5). The effect on the wage premium changes little when I recenter ΔICE_m (−0.119 against −0.105; Table 1) and is imprecise in both versions.

The wage premium index. Table C.2 applies the same three versions of ΔICE_m to the two objects of the $1/\theta$ estimation, on its own design (1991–2000 long differences, state fixed effects and 1986–1990 trends): the market wage premium index ψ_m, the reduced form of Table 3, and the market employment index, its first stage. Unweighted, the index falls by 0.183 (SE of 0.050) per unit of ΔICE_m, by 0.106 (SE of 0.063) with the recentered measure, and by 0.103 (SE of 0.057) among exposed markets alone; weighted, by 0.214 (SE of 0.065), 0.122 (SE of 0.077) and 0.217 (SE of 0.067). The wage decline at incumbent firms is thus identified in part by the differences in tariff cuts across sectors. The employment index is not: its first stage is −0.294 (SE of 0.044) unweighted and −0.008 (SE of 0.060) with the recentered measure, so it comes from the contrast between markets with and without tradable firms, and $1/\theta$, the ratio of the two, rests on that contrast under the parallel-trends assumption that the pre-period controls address.

Table C.2: The wage premium index and the employment index by source of variation in ΔICE_m

		ΔICE_m	Recentered ΔICE_m	Exposed indicator	ΔICE_m, exposed indicator as control	ΔICE_m, exposed markets only
		(1)	(2)	(3)	(4)	(5)
Δ Market wage premium index, ψ_m (reduced form)	Unweighted	-0.183	-0.106	-0.054	-0.094	-0.103
		(0.050)	(0.063)	(0.014)	(0.057)	(0.057)
	Weighted	-0.214	-0.122	-0.023	-0.217	-0.217
		(0.065)	(0.077)	(0.035)	(0.067)	(0.068)
Δ Market employment index (first stage)	Unweighted	-0.294	-0.008	-0.154	0.076	0.078
		(0.044)	(0.060)	(0.012)	(0.052)	(0.052)
	Weighted	0.013	0.066	-0.130	0.051	0.050
		(0.104)	(0.124)	(0.030)	(0.109)	(0.109)
State FE and 1986–1990 trends		Yes	Yes	Yes	Yes	Yes
Local labor markets		17,628	15,058	17,628	17,628	15,058

Notes: Standard errors in parentheses, clustered by local labor market. Each cell reports the coefficient on the column's regressor in a regression of the 1991–2000 change in the row's object on that regressor, with state fixed effects and the market's 1986–1990 changes in log employment and wage premium, as in Table 3; unweighted, or weighted by the market's 1991 formal employment. The market wage premium index ψ_m is the outcome of the reduced form of Table 3, and the market employment index the outcome of its first stage (Section 6). Column (2) subtracts from ΔICE_m its expected value when the 1990–1994 tariff changes are reassigned at random across CNAE sectors within IBGE subsectors (Borusyak and Hull, 2023), on the markets with a tariff firm in 1991. Column (3) replaces ΔICE_m with an indicator for these markets; column (4) adds the indicator as a control; column (5) restricts the sample to them.

ΔICE_m and the RTR

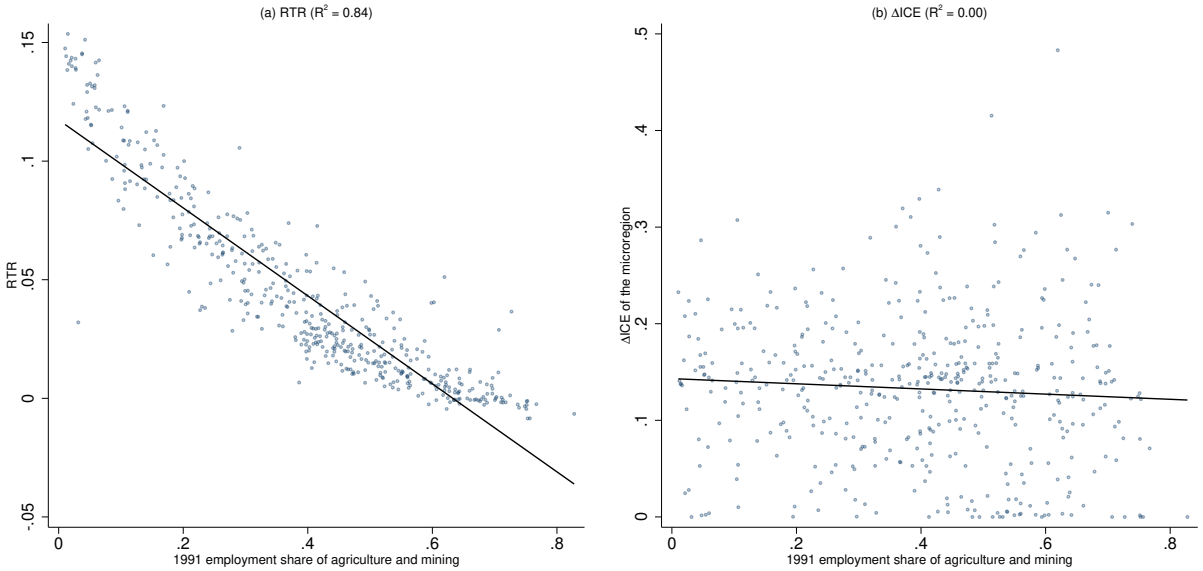
With the specification of Dix-Carneiro and Kovak (2017) at the microregion level, the RTR as the measure of exposure, state fixed effects and 1986–1990 trends, regional formal earnings premia

estimated in RAIS as in their paper give −0.442 (SE of 0.140) by 2000, against their −0.529 (SE of 0.141) (Table C.3; their Table 2). The data of this paper thus reproduce their findings under their design; the two measures differ in what they compare across regions.

The random reassignment of tariff cuts that I apply to ΔICE_m (Figure C.2, panel (a)) gives a different result for the RTR. When I reassign the 1990–1995 tariff changes at random across the tradable industries of Dix-Carneiro and Kovak (2017), holding each microregion's 1991 industry composition fixed, and re-estimate their specification for formal earnings premia, the placebo estimates are widely dispersed around zero, and 88 percent of them are larger in absolute value than the actual estimate of −0.442 (Figure C.2, panel (b)).

What the two designs identify. Dix-Carneiro and Kovak (2017)'s RTR exposure measure and ΔICE_m both rest on a composition contrast for the wage and employment effects as of 2000: for the RTR, the microregion's industry composition, in which the share of agriculture, an industry with a small tariff cut, dominates the cross-regional variation (a regression of the RTR on the census agricultural share of tradable employment has an R^2 of 0.93, and of 0.84 on the employment share of agriculture and mining, Figure C.1; agriculture is 55% of census tradable employment and 6% of RAIS employment); for ΔICE_m, whether the market had tradable firms in 1991. Both designs assess the contrast with 1986–1990 trends. ΔICE_m is unrelated to the agricultural share (an R^2 of 0.0002 in the same regression), so the two measures rank regions differently and their wage effects are not estimates of the same comparison, and under ΔICE_m part of the decline in the wage premium index survives recentering (Table C.2). The two designs answer different questions. The RTR asks how a region whose workers were in industries with larger tariff cuts fared relative to one whose workers were in industries with smaller cuts, agriculture above all; ΔICE_m asks how a formal labor market whose employers faced tariff cuts fared relative to one whose employers faced none. Under their design I reproduce their earnings decline in RAIS; under mine, incumbent firms' wage premia fall, by an amount of the same sign, and I estimate the decline in the market average imprecisely under my main market-level specification (Table 1). The effects on concentration and firm exit, which are the paper's subject, are identified by the tariff cuts themselves and do not depend on this contrast: with tariffs measured for 285 sectors and firms' payroll shares as the shares, ΔICE_m varies across markets with tradable firms only with the tariff cuts their firms faced.

Figure C.1: Dix-Carneiro and Kovak (2017)'s RTR exposure measure and ΔICE_m against the microregion's 1991 agricultural employment share

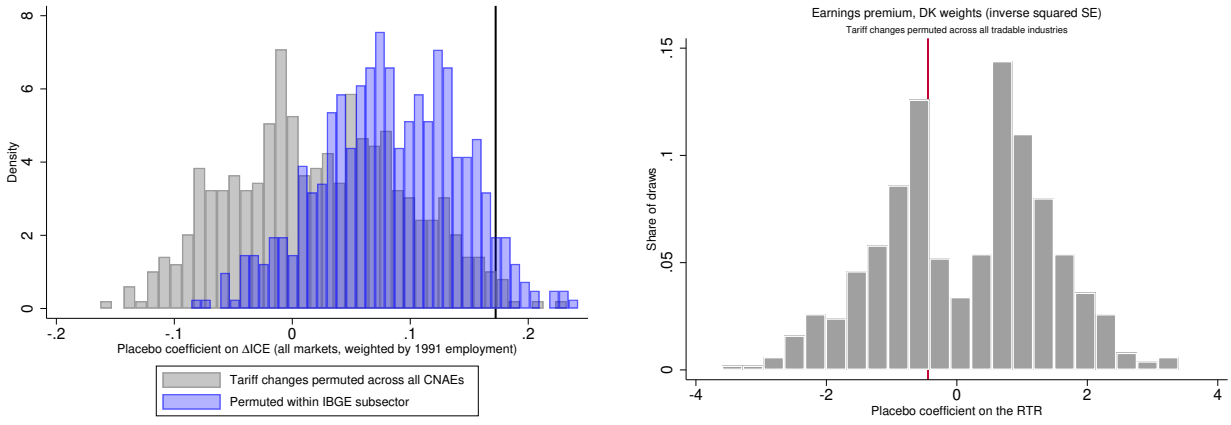


Notes: Each dot is a microregion. The horizontal axis is the 1991 share of agriculture and mining in the microregion's employment, from the census. Panel (a) plots Dix-Carneiro and Kovak (2017)'s RTR exposure measure, computed from the 1991 census (R^2 of the fitted line, 0.84); panel (b) plots ΔICE_m with the microregion as the local labor market (R^2 of 0.00). Lines are OLS fits.

Figure C.2: Placebo estimates under random reassignment of tariff cuts: ΔICE_m and Dix-Carneiro and Kovak (2017)'s RTR exposure measure

(a) ΔICE_m: payroll Herfindahl

(b) RTR: formal earnings premia



Notes: Panel (a) plots the 2000 coefficients of equation (12) for the premium-based payroll Herfindahl, with markets weighted by their 1991 formal employment, when ΔICE_m is recomputed after reassigning the 1990–1994 tariff changes at random across all CNAE sectors (gray) or across CNAE sectors within IBGE subsectors (blue), holding each market's firms and payroll shares fixed. Panel (b) plots the coefficients on Dix-Carneiro and Kovak (2017)'s RTR exposure measure in their specification for the 1991–2000 change in microregions' formal earnings premia (Table C.3, column (3)) when the RTR is recomputed after reassigning the 1990–1995 tariff changes at random across tradable industries, holding each microregion's 1991 industry composition fixed (500 draws). The vertical lines are the estimates with the actual tariff changes.

Table C.3: Replication of Dix-Carneiro and Kovak (2017) in RAIS: effects of their RTR exposure measure on microregions, 1991–2000

Change in outcome, 1991–2000:	(1)	(2)	(3)
Panel A. Log formal earnings premia			
Dix-Carneiro and Kovak (2017)'s RTR exposure measure	-0.353	-0.566	-0.442
	(0.159)	(0.162)	(0.140)
Panel B. Log formal employment			
Dix-Carneiro and Kovak (2017)'s RTR exposure measure	-3.752	-3.570	-3.555
	(0.512)	(0.554)	(0.576)
State fixed effects (26)		Yes	Yes
Pre-trend, 1986–1990			Yes
Microregions	475	474	474

Notes: Microregion-level regressions of the 1991–2000 change in the outcome on Dix-Carneiro and Kovak (2017)'s RTR exposure measure, computed from the 1991 census following their replication code. Panel A uses the microregion's log formal earnings premium, estimated in RAIS for each year as in Dix-Carneiro and Kovak (2017), from a regression of workers' log December earnings on age, sex, and education, industry fixed effects, and microregion fixed effects; regressions in Panel A are weighted by the inverse of the squared standard error of the change in the premium. Panel B uses the microregion's log formal employment, unweighted. Column (2) adds state fixed effects; column (3) also adds the microregion's 1986–1990 change in the outcome, the specification of Dix-Carneiro and Kovak (2017). Standard errors clustered by mesoregion in parentheses.

C.6 Proofs

Propositions 2 and 3 (equation C.8). Log-differentiating the first-order condition $r_{zm} = w_{zm}\mu_{zm}$ gives $d\ln r_{zm} = d\ln w_{zm} + d\ln \mu_{zm}$. By Assumption 1 and equations 3, 7, and 8,

$$d\ln w_{zm} = \frac{1}{\eta}\, d\ln l_{zm} + \left(\frac{1}{\theta} - \frac{1}{\eta}\right) d\ln L_m, \qquad d\ln L_m = \sum_{k\in\Theta_m} s_{km}\, d\ln l_{km},$$

$$d\ln \mu_{zm} = \left(\frac{1}{\theta} - \frac{1}{\eta}\right)\frac{ds_{zm}}{\mu_{zm}}, \qquad ds_{zm} = \left(1 + \frac{1}{\eta}\right) s_{zm}\left(d\ln l_{zm} - d\ln L_m\right),$$

where the last equality uses $s_{zm} = (\xi_{zm} l_{zm})^{\frac{1+\eta}{\eta}} \,/\, \sum_{k\in\Theta_m}(\xi_{km} l_{km})^{\frac{1+\eta}{\eta}}$ (Appendix B.2). Substituting these expressions and Assumption 2 into the differentiated first-order condition and rearranging gives

$$d\ln l_{zm} - d\ln L_m = \frac{1}{D_{zm}}\left[x_{zm} - \left(\frac{1}{e_{zm}} + \frac{1}{\theta}\right) d\ln L_m\right]. \tag{C.10}$$

Multiplying equation C.10 by s_{zm} and summing over the firms of the market gives zero on the left-hand side, and solving for $d\ln L_m$ gives the first part of equation C.2; the second part follows from $W_m = W\,(L_m\xi_m/L)^{\frac{1}{\theta}}\,\xi_m$ (Appendix B.2). For equation C.8, $dHHI_m = 2\sum_z s_{zm}\, ds_{zm} = 2\left(1+\frac{1}{\eta}\right)\sum_z s_{zm}^2\,(d\ln l_{zm} - d\ln L_m)$, and the result follows from equation C.10.

Corollary 2, the Herfindahl under Assumptions 3 to 5, and equation C.7. Under Assumption 4, $\sum_z \omega_{zm}/e_{zm} = 1/e$, so that equation C.2 becomes $(1/e + 1/\theta)\, d\ln L_m = \bar{x}_m \equiv \sum_{z\in\Theta_m}\omega_{zm}x_{zm}$.

By Assumption 5, $\bar{x}_m = \omega_m^T \bar{x}_m^T + (1 - \omega_m^T)\, \bar{x}_m^T = \bar{x}_m^T$, and by Assumption 3, $\bar{x}_m^T = -\psi\, \Delta ICE_m^\omega$, which gives equation C.4. Substituting $(1/e + 1/\theta)\, d \ln L_m = \bar{x}_m^T$ into equation C.8, the term in brackets is $x_{zm} - \bar{x}_m^T$, which is zero for non-tradable firms and $-\psi(\Delta\tau_z - \Delta ICE_m^\omega)$ for tradable firms. For equation C.7, the weights of equation 11 are $\kappa_{zm} = s_{zm}^2 / HHI_m^T$, so that $\frac{s_m^T}{HHI_m^T} \mathrm{Cov}_m^T(s_{zm}, \Delta\tau_z) = \frac{s_m^T}{HHI_m^T} \left[\sum_{z \in \Theta_m^T} \frac{s_{zm}^2}{s_m^T} \Delta\tau_z - \frac{HHI_m^T}{s_m^T} \Delta ICE_m^{\mathrm{lin}} \right] = \Delta ICE_m - \Delta ICE_m^{\mathrm{lin}}$.

D Weighting: Summary of Estimates by Weighting Scheme

In my main market-level estimates I weight local labor markets by their 1991 formal employment, so that they describe the market of the average formal worker and are identified by the tariff cuts. In this appendix, I summarize what changes when each market instead counts once, so that the estimates describe the average market, and which weights I use to estimate each of the model's parameters.

Table D.2 summarizes. Weighting changes which comparison identifies my market-level effects, but not their sign: weighted, the effects on concentration and on the number of firms come from differences in tariff reductions across sectors and survive the recentering and permutation checks, while the effect on employment rests more on the contrast between markets with and without tradable firms; unweighted, the effects on concentration and employment come from the contrast between the 85% of markets with a tradable firm in 1991 and the 15% without one, and rest on parallel trends between the two groups, which hold for concentration and not for employment (unweighted, the effect on employment is 0.01, SE of 0.09, with the recentered measure, and −1.27, SE of 0.12, with the indicator for markets with a tradable firm; Table D.3). The two agree within quintiles of market size (Table D.1). Relative to the baseline level of concentration, the unweighted effect is the smaller (7% against 22%), because the markets where formal employment is located are much less concentrated than the average market. At the microregion level nearly every market has a tradable firm (6 of 475 do not), so the contrast is absent and I weight by 1991 formal employment, as in the extended model. The model's elasticities are not averages over markets, so I choose the weights in their estimation for identification and precision: I estimate $1/\eta$ with market fixed effects and markets weighted by 1991 employment (with each firm-market pair counting once, I estimate 1.226, SE of 0.160); I estimate $1/\theta$ unweighted, because with employment weights its first stage has no power; and I estimate the extended model's $1/\tilde{\rho}$, $1/\tilde{\theta}$ and Ω with the weights under which their first stages are strong (Appendix F.3). Because I estimate $1/\theta$ with each market counting once, I use the unweighted effect on concentration to compute the effect on markdowns in Table 4; with the weighted effect it is 0.011 instead of 0.013, a reduction in the take-home share of 0.30 instead of 0.37 cents.

What the unweighted effect on concentration captures. Unweighted, a 10% increase in exposure raises the payroll Herfindahl by 0.021 points, 7% of the 0.29 mean across markets (Appendix Table A.4, column (4)). This effect is identified by the shares: it is close to zero when I control for the expected value of ΔICE_m under random reassignment of tariff cuts across sectors (−0.005, SE of 0.021, per unit of ΔICE_m, against 0.212; Table D.3), and it falls to 0.002 points per 10% increase in exposure when I drop the markets without a tradable firm (Table D.4). Interpreting it as causal requires that markets with and without tradable firms would have followed the same trends absent the reform, which the flat 1986–1989 coefficients support (p-value of 0.35 on their trend; Table D.3). The comparison captures a smaller decline in concentration: between 1991 and

2000 the payroll Herfindahl fell by 0.15 points on average in the markets with no tradable firm and by 0.06 points in the markets with one. The gap is in non-tradable firms, whose contribution to the Herfindahl fell by 0.17 points in the markets with no tradable firm, where they are all of the firms, and by 0.02 points in the others.

Table D.1: Effect of import competition on concentration by quintile of 1991 market size

	Δ Payroll Herfindahl (premium-based)		Δ Employment Herfindahl	
	Unweighted	Weighted	Unweighted	Weighted
	(1)	(2)	(3)	(4)
Quintile 1 (smallest)	0.274	0.203	0.326	0.245
	(0.045)	(0.044)	(0.044)	(0.042)
Quintile 2	0.128	0.130	0.178	0.177
	(0.033)	(0.034)	(0.030)	(0.032)
Quintile 3	0.086	0.085	0.120	0.114
	(0.030)	(0.031)	(0.027)	(0.028)
Quintile 4	0.026	0.044	0.042	0.049
	(0.033)	(0.034)	(0.029)	(0.030)
Quintile 5 (largest)	0.148	0.173	0.120	0.097
	(0.033)	(0.049)	(0.029)	(0.027)

Notes: Standard errors in parentheses, clustered at the local labor market level. Each cell reports the 2000 coefficient of the stacked differences-in-differences regression of Table 1, estimated separately within each quintile of local labor markets' 1991 formal employment, unweighted (columns (1) and (3)) and weighted by 1991 formal employment (columns (2) and (4)).

Table D.2: Market-level estimates and parameters by weighting scheme

	Weighted by 1991 employment (average worker)	Unweighted (average market)	Identified by
Panel A: Market-level effects, per unit of ΔICE_m, 1991–2000			
Δ Payroll Herfindahl	0.172	0.212	Weighted: tariff cuts; unweighted: exposed vs. never exposed
	(0.042)	(0.017)	
Δ Average wage markdown, γ, per 10%	0.011	0.013	via β
Implied change in take-home share, cents per 10%	−0.30	−0.37	
Microregion level (extended model, ΔICE^*), per 10%			
Δ Expected payroll Herfindahl	0.028	–	
	(0.007)		
Δ Wage work share of market payroll	−0.032	–	
	(0.005)		
Δ Average inverse elasticity (take-home, cents)	0.0053 (−0.15)	–	tariff cuts (Table 7)
	(0.0014)		
Panel B: Parameters and the weights used in their estimation			
$1/\eta$ (firm × market; weights: market 1991 employment)	0.832 (0.093)	1.226 (0.160)	tariff cuts across firms within markets
$1/\theta$ (market; unweighted)	no first stage (F below 1)	1.454 (0.196)	exposed vs. never exposed (employment); tariff cuts (wage index)
$1/\tilde{\rho}$, $1/\tilde{\theta}$ (census cells; weights: 1991 formal employment)	0.684 (0.234), 1.187 (0.481)	–	ΔICE^* across cells
Ω (firm × cell; weights: lagged cell employment)	0.0091 (0.0025)	–	firm tariff × ΔICE^*

Notes: Panel A reports the 2000 coefficient of equation (12) on ΔICE_m, weighted by the market's 1991 formal employment and unweighted; standard errors clustered by local labor market. The microregion rows are the extended model's reduced forms on ΔICE^*, with the take-home share change in cents in parentheses. Panel B lists the sample, the weights and the variation that identifies each parameter; the extended model's parameters are in Table 7. Sources: Tables 1, 2, 3, 4, and 7, and Appendix Tables A.4, column (4), and D.3.

Table D.3: Effects by source of variation and pre-period tests

		ΔICE_m (1)	Recentered ΔICE_m (2)	Exposed indicator (3)	Pre-period trend (p-value) (4)
Δ Payroll Herfindahl (premium-based)	Weighted	0.172	0.151	0.464	0.016
		(0.042)	(0.036)	(0.062)	
	Unweighted	0.212	−0.005	0.590	0.346
		(0.017)	(0.021)	(0.033)	
Δ Log number of firms	Weighted	−0.971	−1.021	−0.754	0.000
		(0.188)	(0.214)	(0.296)	
	Unweighted	−0.510	−0.300	−0.873	0.469
		(0.048)	(0.063)	(0.085)	
Δ Log formal employment	Weighted	−0.398	−0.242	−0.625	0.000
		(0.186)	(0.206)	(0.213)	
	Unweighted	−0.401	0.010	−1.266	0.027
		(0.069)	(0.090)	(0.123)	
Δ Log formal wage premium	Weighted	−0.105	−0.112	−0.186	0.308
		(0.089)	(0.082)	(0.109)	
	Unweighted	0.027	0.170	−0.343	0.004
		(0.030)	(0.038)	(0.060)	

Notes: Coefficient on ΔICE_m for the change between 1991 and 2000 in equation (12). Column (2) adds as a control, interacted with year, the expected value of ΔICE_m when tariff changes are reassigned at random across CNAE sectors within IBGE subsectors (Borusyak and Hull, 2023). Column (3) replaces ΔICE_m with an indicator for exposed markets (markets with firms facing tariff cuts in 1991). Column (4): p-value of the slope of a linear trend fitted to the 1986–1989 coefficients of column (1). Standard errors clustered by local labor market in parentheses.

Table D.4: Controls for agriculture intensity and tradable sector intensity, unweighted

	Δ Payroll Herfindahl (premium-based) (1)	Δ Employment Herfindahl (2)	Δ Log formal employment (3)
Main specification, unweighted	0.212	0.239	-0.401
	(0.017)	(0.016)	(0.069)
Control: 1991 census agricultural share × year	0.174	0.201	-0.254
	(0.017)	(0.016)	(0.069)
ΔICE without agricultural firms	0.216	0.244	-0.342
	(0.017)	(0.016)	(0.068)
Dropping top-quartile agricultural microregions	0.176	0.195	-0.306
	(0.018)	(0.018)	(0.074)
Control: Dix-Carneiro and Kovak (2017)'s RTR × year	0.165	0.191	-0.219
	(0.017)	(0.016)	(0.069)
Control: 1991 payroll share of tariff firms × year	0.230	0.255	-0.338
	(0.017)	(0.016)	(0.069)
Control: 1{tariff firms in 1991} × year	0.020	0.016	0.039
	(0.018)	(0.017)	(0.080)
Markets with tariff firms only	0.020	0.016	0.039
	(0.018)	(0.017)	(0.080)
Microregion × year fixed effects	0.163	0.190	-0.041
	(0.019)	(0.018)	(0.075)
Occupation × year and size decile × year FE	0.129	0.151	0.270
	(0.018)	(0.017)	(0.065)

Notes: Standard errors in parentheses, clustered at the local labor market level. All entries report the 2000 coefficient on ΔICE_m from the stacked differences-in-differences regressions of Table 1, with each local labor market counting once. The agricultural employment share is the microregion's 1991 share of tradable employment in agriculture, from the census. The RTR is Dix-Carneiro and Kovak (2017)'s RTR exposure measure, computed from the 1991 census. The tariff firms' payroll share is the 1991 payroll share of the market's firms in sectors with a tariff. ΔICE_m without agricultural firms excludes agricultural firms from the market's tariff firms. The top quartile of agricultural share is defined across microregions.

E Effective Rates of Protection

The tariff shocks of the paper are changes in output tariffs. Brazil's liberalization also cut tariffs on firms' inputs, which lowers their costs and works against the import competition they face. In this appendix, I repeat the market-level estimates of Table 1 and the two elasticities of Tables 2 and 3 with the shock measured as the change in effective rates of protection (ERP), output tariffs net of input tariffs using Brazil's 1985 input-output table (Appendix A). Input tariffs are available for about 50 sectors, against 285 for output tariffs, so the ERP shock is noisier and its effects attenuated; the ERP version of ΔICE_m also weights firms by their linear payroll shares rather than by their contributions to the payroll Herfindahl. In Table E.1 I therefore report, for each weighting of markets, the output-tariff shock of Table 1, the output-tariff shock with linear payroll shares, and the ERP shock with the same shares, so that I can tell the two differences apart.

Weighted by 1991 formal employment (columns (1)–(3)), the effect on concentration is the same with the ERP shock as in the main specification (21% against 22% of the 1991 mean), the number of firms falls by less, formal employment does not fall, and the three wage measures rise by 0.16 to 0.18 log points per unit of the shock (SE of 0.06 to 0.08). Column (2) shows that part of the difference comes from the share weights rather than from input tariffs: with output tariffs and linear payroll shares, the wage effects are also positive, though imprecise (0.20 to 0.26, SE of 0.11 to 0.15). Unweighted (columns (4)–(6)), the ERP shock gives effects on concentration, the number of firms and employment of the same sign as the output-tariff shock and about half as large relative to the 1991 mean (4% against 7% for the payroll Herfindahl), and a positive effect on the wage premium (0.065, SE of 0.020), where the output-tariff shock gives zero. The within-market elasticity is nearly unchanged with the ERP shock (0.878 against 0.832; Table E.2), with first stage and reduced form both about a third smaller, and the cross-market inverse elasticity is larger and less precise (1.758, SE of 0.421, against 1.454), so that the gap between them and the implied effect on markdowns are larger. Because the positive wage effects of the ERP shock partly reflect the weighting of firms within the exposure measure, and because input tariffs are measured at a much coarser level than output tariffs, I keep output tariffs as the shock in the main text and leave the separate roles of input and output tariff reductions for future work.

Table E.1: Effect of import competition on market outcomes: effective rates of protection

	Weighted by 1991 formal employment			Unweighted		
	Main specification	Linear payroll shares	Effective rates of protection	Main specification	Linear payroll shares	Effective rates of protection
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Formal sector labor market concentration						
Δ Payroll Herfindahl (premium-based)	0.172	0.421	0.164	0.212	0.262	0.121
	(0.042)	(0.074)	(0.039)	(0.017)	(0.020)	(0.011)
1991 mean	0.079	0.079	0.079	0.287	0.287	0.287
Effect per 10%, percent of 1991 mean	22%	54%	21%	7%	9%	4%
Δ Payroll Herfindahl	0.183	0.439	0.173	0.213	0.262	0.122
	(0.045)	(0.083)	(0.043)	(0.017)	(0.020)	(0.011)
1991 mean	0.082	0.082	0.082	0.292	0.292	0.292
Effect per 10%, percent of 1991 mean	22%	53%	21%	7%	9%	4%
Δ Employment Herfindahl	0.109	0.278	0.101	0.239	0.294	0.134
	(0.024)	(0.035)	(0.019)	(0.016)	(0.019)	(0.011)
1991 mean	0.056	0.056	0.056	0.239	0.239	0.239
Effect per 10%, percent of 1991 mean	19%	50%	18%	10%	12%	6%
Panel B: Number of firms and formal employment						
Δ Log number of firms	−0.971	−1.443	−0.307	−0.510	−0.625	−0.277
	(0.188)	(0.196)	(0.111)	(0.048)	(0.054)	(0.032)
Δ Log formal employment	−0.398	−0.048	0.128	−0.401	−0.479	−0.194
	(0.186)	(0.199)	(0.100)	(0.069)	(0.078)	(0.046)
Panel C: Wages						
Δ Log average firm wage premium	−0.124	0.200	0.157	−0.071	−0.085	0.022
	(0.130)	(0.151)	(0.075)	(0.035)	(0.040)	(0.024)
Δ Log formal wage premium	−0.105	0.208	0.180	0.027	0.039	0.065
	(0.089)	(0.113)	(0.056)	(0.030)	(0.035)	(0.020)
Δ Log formal average wage	−0.075	0.255	0.183	0.014	0.020	0.067
	(0.110)	(0.131)	(0.065)	(0.035)	(0.040)	(0.023)
Observations	296,311	296,311	296,311	296,311	296,311	296,311
Local labor markets	19,757	19,757	19,757	19,757	19,757	19,757

Notes: Standard errors in parentheses, clustered at the local labor market level. Each cell reports the 2000 coefficient from a stacked differences-in-differences regression of long-differenced outcomes on the trade shock interacted with year dummies, with local labor market fixed effects, as in Table 1. Columns (1) and (4) repeat Table 1, with ΔICE_m built from output tariffs with weights given by firms' contributions to the market's 1991 payroll Herfindahl (Table 1, column (1), and Appendix Table A.4, column (4)); columns (2) and (5) build ΔICE_m from output tariffs with firms' linear 1991 payroll shares as weights (Appendix Table A.4, column (2)); columns (3) and (6) build it from effective rates of protection with linear payroll shares. Columns (1)–(3) weight local labor markets by their 1991 formal employment; columns (4)–(6) count each local labor market once. Effective rates of protection are the output tariff net of the input-output-weighted average of input tariffs, over one minus the input share, computed with Brazil's 1985 input-output table at the Nível 50 sector level (Appendix A).

Table E.2: Labor supply elasticities with effective rates of protection

	$1/\eta$ (Table 2)		$1/\theta$ (Table 3)	
	Import tariffs (1)	Effective rates of protection (2)	Import tariffs (3)	Effective rates of protection (4)
Panel A: First stage				
Δ Firm log employment in LLM	−0.736	−0.447		
	(0.076)	(0.059)		
Δ LLM employment index			−0.294	−0.114
			(0.044)	(0.029)
First stage F	94.9	57.4	45.6	15.7
Panel B: Reduced form				
Δ Firm wage premium in LLM	−0.612	−0.392		
	(0.036)	(0.029)		
Δ LLM wage premium index			−0.183	−0.100
			(0.050)	(0.035)
Panel C: 2SLS				
$1/\eta$	0.832	0.878		
	(0.093)	(0.124)		
$1/\theta - 1/\eta$			0.622	0.880
			(0.217)	(0.438)
$1/\theta$			1.454	1.758
			(0.196)	(0.421)
Implied bound on wage take-home share	55%	53%	41%	36%
Observations	363,407	362,392	17,628	17,628
Firms	149,083	148,857		
Local labor markets	11,632	11,603	17,628	17,628

Notes: Columns (1) and (3) reproduce column (1) of Tables 2 and 3. Columns (2) and (4) replace the 1990–1994 change in the firm's import tariff by the change in its effective rate of protection as the instrument, in the firm-level regressions of Section 6.1 (standard errors clustered by firm; local labor market fixed effects; weighted by baseline market employment; 1986–1991 pre-trend controls) and in the market-level regressions of Section 6.2 (standard errors clustered by local labor market; state fixed effects and 1986–1990 trends), with ΔICE_m in column (4) built from effective rates of protection as in Appendix Table E.1, column (6). Column (4) uses the $1/\eta$ of column (2). The implied bound on the take-home share is $1/(1+1/\eta)$ in columns (1)–(2), an upper bound, and $1/(1+1/\theta)$ in columns (3)–(4), a lower bound.

F Extension: Oligopsony in Dual Labor Markets

F.1 Model

I model the labor supply preferences of *people seeking formal employment*. Firms exploit these preferences to set wages below workers' marginal revenue product when maximizing profits while holding local competitors' labor demand decisions constant. The starting point is the labor supply preference structure of Section 2.1, which I extend in two directions. First, I split each local labor market into two sectors, wage work and self-employment. Second, I allow workers to take into account the probability of involuntary separation out of the formal sector—which varies by firm—when making formal sector labor supply decisions.

F.1.1 Discrete choice formal labor supply in dual labor markets

There is a continuum of homogeneous workers j, each choosing where to supply l^j effective units of labor.[60] Each worker chooses a single market m from a large but finite number of local labor markets. Within each market, workers choose a sector g, which is either wage work $\overline{g}$ or self-employment $\underline{g}$. When considering wage work, workers compare expected earnings at a finite number of formal sector firms z, which reflect their perceived probability of involuntary separation out of the formal sector o in market m. When considering self-employment, they consider expected earnings from starting a small business in market m. Workers choose the employment option that minimizes their disutility of labor given their preference parameters, expected earnings, and a minimum earnings requirement, as in Section 2.1.

Two features distinguish the extension. First, workers value formal jobs at *expected* wages. A job at formal firm z carries a perceived probability $p_{z\overline{g}m}$ of involuntary separation from formal employment. I assume that a separated worker finds informal wage work without frictions, in a perfectly competitive local labor market for informal labor that pays w_m^o. The wage relevant for attracting labor to the firm is therefore

$$\bar{w}_{z\overline{g}m} = \left(1 - p_{z\overline{g}m}\right) w_{z\overline{g}m} + p_{z\overline{g}m}\, w_m^o = w_{z\overline{g}m}\left(1 + p_{z\overline{g}m}\, \sigma_{z\overline{g}m}\right), \tag{F.1}$$

where $w_{z\overline{g}m}$ is the firm's posted formal wage and $\sigma_{z\overline{g}m} \equiv (w_m^o - w_{z\overline{g}m})/w_{z\overline{g}m}$ is the wage gap between the local informal sector and the firm's formal wage. This gap may be negative (the informal sector pays less) or positive (the informal sector pays more). Workers likewise value self-employment at its expected returns $W_{\underline{g}m}$, since the returns to starting a small business are uncertain ex ante. Introducing earnings uncertainty is a natural way to incorporate both self-employment and informal

[60] As in the baseline model, the assumption of homogeneous productivity is mapped to the data by conditioning wages on worker demographics.

wage work into the labor supply decisions of people seeking formal sector employment.

No dynamic gains from influencing workers' priors. I assume that firms do not internalize that their hiring and firing decisions at time t influence workers' priors about the probability of separation from the firm at $t+1$ or after. The assumption shuts off $p_{z\overline{g}m}$ as an additional, dynamic lever of wage setting, and thereby preserves the baseline model's static wage-setting problem while letting priors enter labor supply. Workers may form priors from firms' past involuntary separations, which the administrative data record, but neither workers nor firms weigh the future costs or benefits of decisions made at t. Dynamic gains to labor market power from firms' ability to shape workers' priors about separation, and about other within-firm dynamics such as promotion and tenure, merit further research.

Second, idiosyncratic tastes follow a nested GEV distribution with a sector level between the firm and market levels of the baseline model:

$$G(\xi) = \exp\left(-\sum_m \left[\sum_{g\in\{\overline{g},\underline{g}\}} \left(\sum_z e^{-(1+\tilde{\eta})\,\xi^j_{zgm}}\right)^{\frac{1+\tilde{\rho}}{1+\tilde{\eta}}}\right]^{\frac{1+\tilde{\theta}}{1+\tilde{\rho}}}\right). \tag{F.2}$$

The parameters $\tilde{\eta}$, $\tilde{\rho}$, and $\tilde{\theta}$ are workers' elasticities of substitution across firms within the wage work sector, between wage work and self-employment within markets, and across markets, respectively. $\tilde{\eta}$ and $\tilde{\theta}$ are the extended counterparts of the baseline model's η and θ, and all three elasticities are common to all workers.

F.1.2 Labor supply curves faced by individual formal firms

As in the baseline model, the discrete choice formulation of workers' labor supply decisions with GEV idiosyncratic tastes yields analytic expressions for the probability that worker j chooses formal sector firm z in the wage work sector $\overline{g}$ of market m. Integrating these probabilities over the continuum of workers gives total labor supply to formal sector firm z:

$$l_{z\overline{g}m} = L\left(\frac{\bar{w}_{z\overline{g}m}}{W_{\overline{g}m}}\right)^{\tilde{\eta}}\left(\frac{W_{\overline{g}m}}{W_m}\right)^{\tilde{\rho}}\left(\frac{W_m}{W}\right)^{\tilde{\theta}} \times \left[\xi^{1+\tilde{\eta}}_{zgm}\xi^{1+\tilde{\rho}}_{gm}\xi^{1+\tilde{\theta}}_{m}\right]^{-1}. \tag{F.3}$$

whose inverse is:

$$\bar{w}_{z\overline{g}m} = W\left(\frac{l_{z\overline{g}m}}{L_{\overline{g}m}}\right)^{\frac{1}{\tilde{\eta}}}\left(\frac{L_{\overline{g}m}}{L_m}\right)^{\frac{1}{\tilde{\rho}}}\left(\frac{L_m}{L}\right)^{\frac{1}{\tilde{\theta}}} \xi^{1+\frac{1}{\tilde{\eta}}}_{z\overline{g}m}\xi^{1+\frac{1}{\tilde{\rho}}}_{\overline{g}m}\xi^{1+\frac{1}{\tilde{\theta}}}_{m}. \tag{F.4}$$

Here $(W_{\overline{g}m}, L_{\overline{g}m})$, (W_m, L_m), and (W, L) are nested CES wage and labor supply indices over firms, sectors, and markets, defined at the end of this subsection. Equation F.4 has the same form as its

baseline counterpart: the wage needed to attract labor is increasing in the firm's employment relative to its sector, in the sector's employment relative to the market, and in the market's employment relative to the country. The difference is that the relevant wage is the expected wage of equation F.1, which depends on the probability of separation and on local informal sector conditions.

In addition, because I can only estimate the elasticity of substitution between self-employment and wage work with census data—which do not include firm boundaries but do include regions—I introduce the market-level aggregate of equation F.4, the inverse labor supply curve to wage work in each market:

$$W_{\bar{g}m} = W \left(\frac{L_{\bar{g}m}}{L_m} \right)^{\frac{1}{\bar{\rho}}} \left(\frac{L_m}{L} \right)^{\frac{1}{\bar{\theta}}} \xi_{\bar{g}m}^{1+\frac{1}{\bar{\rho}}} \xi_m^{1+\frac{1}{\bar{\theta}}}. \tag{F.5}$$

Finally, the wage and labor supply indices in equations F.3–F.5 are nested CES:

$$W_{\bar{g}m} = \left[\sum_z \left(\frac{\bar{w}_{z\bar{g}m}}{\xi_{z\bar{g}m}} \right)^{1+\tilde{\eta}} \right]^{\frac{1}{1+\tilde{\eta}}}, \qquad L_{\bar{g}m} = \left[\sum_z \left(\xi_{z\bar{g}m} l_{z\bar{g}m} \right)^{\frac{1+\tilde{\eta}}{\tilde{\eta}}} \right]^{\frac{\tilde{\eta}}{1+\tilde{\eta}}} \tag{F.6}$$

$$W_m = \left[\left(\frac{W_{\underline{g}m}}{\xi_{\underline{g}m}} \right)^{1+\tilde{\rho}} + \left(\frac{W_{\bar{g}m}}{\xi_{\bar{g}m}} \right)^{1+\tilde{\rho}} \right]^{\frac{1}{1+\tilde{\rho}}}, \qquad L_m = \left[\left(\xi_{\underline{g}m} l_{\underline{g}m} \right)^{\frac{1+\tilde{\rho}}{\tilde{\rho}}} + \left(\xi_{\bar{g}m} L_{\bar{g}m} \right)^{\frac{1+\tilde{\rho}}{\tilde{\rho}}} \right]^{\frac{\tilde{\rho}}{1+\tilde{\rho}}} \tag{F.7}$$

$$W = \left[\sum_m \left(\frac{W_m}{\xi_m} \right)^{1+\tilde{\theta}} \right]^{\frac{1}{1+\tilde{\theta}}}, \qquad L = \left[\sum_m \left(\xi_m L_m \right)^{\frac{1+\tilde{\theta}}{\tilde{\theta}}} \right]^{\frac{\tilde{\theta}}{1+\tilde{\theta}}} \tag{F.8}$$

where $W_{\underline{g}m}$ are expected earnings from self-employment in market m.[61]

F.1.3 Firm-specific wage markdowns

Firms compete for labor à la Cournot, as in Section 2.2. Each firm chooses employment while holding constant competing firms' labor demand, workers' priors about separation probabilities, expected informal sector wages, and expected returns to self-employment. Taking logs of equation F.4 and differentiating with respect to the firm's own employment gives the inverse elasticity of

[61] In a nested CES preference structure, $W_{\underline{g}m}$ equals expected log earnings in the limit where the elasticity of substitution between self-employment options is zero. At the micro level, this means that no one worker can provide self-employment for another worker. At the macro level, this is the CES limiting case of Cobb-Douglas preferences over self-employment options. Formally, let $\eta_{\underline{g}}$ denote the elasticity of substitution between self-employment options and let $\mathbb{E}[\pi_{\underline{g}m}^j]$ denote worker j's expected earnings from self-employment in market m. Then $W_{\underline{g}m} = \left(\int_0^1 \mathbb{E}[\pi_{\underline{g}m}^j]^{1+\eta_{\underline{g}}} \, dj \right)^{1/(1+\eta_{\underline{g}})}$ is the CES aggregator over self-employment options in a given market. When $\eta_{\underline{g}} \to 0$, we have $W_{\underline{g}m} \to \exp\left(\int_0^1 \log \mathbb{E}[\pi_{\underline{g}m}^j] \, dj \right)$, the Cobb-Douglas form. Note that $W_{\underline{g}m} > 0$ because workers must satisfy their minimum earnings requirement in equilibrium.

residual labor supply that the firm faces:

$$\varepsilon_{z\bar{g}m}^{-1} = \left[\frac{1}{\tilde{\rho}} s_{z\bar{g}m} + \frac{1}{\tilde{\eta}}(1 - s_{z\bar{g}m})\right] + \left(\frac{1}{\tilde{\theta}} - \frac{1}{\tilde{\rho}}\right) s_{zm}. \tag{F.9}$$

Here $s_{z\bar{g}m}$ is firm z's share of the wage work sector's expected payroll and s_{zm} is its share of the market's total expected payroll, including self-employment. The elasticity of labor supply to a formal firm thus depends on how easily workers substitute between firms within markets ($\tilde{\eta}$), between markets ($\tilde{\theta}$), and between wage work and self-employment within markets ($\tilde{\rho}$), and the relevant shares are based on expected wages.

Local average wage markdown. As in the baseline model, the average wage markdown is $\mu_{\bar{g}m} \equiv 1 + \varepsilon_{\bar{g}m}^{-1}$ and the wage take-home share is $\mu_{\bar{g}m}^{-1}$. Taking a weighted average of the firm-specific inverse elasticities in equation F.9, with $s_{z\bar{g}m}$ as weights, gives the market's average inverse elasticity of residual labor supply:

$$\bar{\varepsilon}_{\bar{g}m}^{-1} = \frac{1}{\tilde{\rho}} \widetilde{HHI}_{\bar{g}m} + \frac{1}{\tilde{\eta}}(1 - \widetilde{HHI}_{\bar{g}m}) + \left(\frac{1}{\tilde{\theta}} - \frac{1}{\tilde{\rho}}\right) \widetilde{HHI}_{\bar{g}m} \cdot s_m. \tag{F.10}$$

Here $\widetilde{HHI}_{\bar{g}m} \equiv \sum_z s_{z\bar{g}m}^2$ is the wage work sector's expected payroll Herfindahl and s_m is the wage work sector's share of market m's expected payroll.

F.1.4 Policy effects on wage markdowns

Let X_m denote an exogenous shock to labor demand in local labor market m. Then this shock's effect on local markdowns is given by its effect on average local inverse elasticities of labor supply, which follows:

$$\frac{\partial \bar{\varepsilon}_{\bar{g}m}^{-1}}{\partial X_m} = \left(\frac{1}{\tilde{\rho}} - \frac{1}{\tilde{\eta}}\right) \times \beta_{\text{HHI}} + \left(\frac{1}{\tilde{\theta}} - \frac{1}{\tilde{\rho}}\right) \left[s_m \times \beta_{\text{HHI}} + \widetilde{HHI} \times \beta_{s_m}\right] \tag{F.11}$$

where $\widetilde{HHI}$ is short for $\widetilde{HHI}_{\bar{g}m}$, and β_{HHI} and β_{s_m} are the reduced-form effects of the shock on the expected payroll Herfindahl and on the wage work share. Without self-employment ($s_m = 1$ and $\beta_{s_m} = 0$), the expression reduces to Corollary 1. In the first term, concentration shifts weight between the cross-firm margin $\tilde{\eta}$ and the self-employment margin $\tilde{\rho}$. The second term arises because the weight of the cross-market margin in equation F.10 is $\widetilde{HHI}_{\bar{g}m} \times s_m$, which rises with concentration and falls when the wage work share falls.

F.2 Trade shock

The baseline measure of import competition exposure, ΔICE_m (equation 11), averages the tariff reductions faced by the formal firms of a market. In the extended model, the market also includes informal wage work and self-employment. For the extended model, I use a measure with two components, weighted by the formal share of the market's wage bill.

The first component is ΔICE_m of equation 11, computed with microregions as markets: the average tariff reduction faced by the microregion's tradable sector firms, weighted by their 1991 squared payroll shares as in equation 11. It is zero in the 6 of the 475 microregions of the sample without a tradable sector firm with a tariff match, which remain in all estimations.[62]

The second component is RTR_m, Dix-Carneiro and Kovak (2017)'s exposure measure, RTR, with industry weights computed from the 1991 census employment of all workers in tradable industries, as in Kovak (2013), where the return to labor in a regional market moves with the industry composition of all of the region's workers.

The extended model's shock weights the two components by the formal share of the market's 1991 wage bill:

$$\Delta ICE^*_m \equiv f_{m,1991}\, \Delta ICE_m + \left(1 - f_{m,1991}\right) RTR_m, \tag{F.12}$$

where I measure $f_{m,1991}$ in the 1991 census as the earnings of formal employees over the earnings of all employed workers: formal employees, informal employees, and the self-employed. The shock is defined for each microregion and assigned to all demographic cells of the microregion. Across microregions, ΔICE^*_m has a mean of 0.084 and a standard deviation of 0.037.

F.3 Estimation strategy

I define local labor markets as microregions for all of the extension's estimates.[63] Unless noted, Δ denotes 1991–2000 long differences and standard errors are clustered by microregion.

F.3.1 Substitution within markets, across firms: $1/\eta$

The labor supply response to firms' posted wages is the baseline model's within-market cross-firm inverse elasticity. I estimate it as in Section 5, with markets defined as microregions:

$$\Delta \ln w_{zm} = \frac{1}{\eta} \Delta \ln l_{zm} + \delta_m + \epsilon_{zm}, \tag{F.13}$$

[62] Appendix C.2 relates the squared-share weights to payroll-share weights and to the baseline model's measure of exposure.

[63] While the 1991 and 2000 censuses contain occupation codes, they are missing for a large share of observations, and the two years also follow different definitions from each other (and from RAIS).

where Δ denotes 1991–2000 long differences, w_{zm} is firm z's December wage premium in microregion m, estimated on all of the firm's workers in the microregion as in the baseline, l_{zm} is its employment, and δ_m is a microregion fixed effect. I instrument the change in employment with the firm-level tariff reduction. Firms in non-tradable sectors remain in the sample with a zero tariff change. I weight the regression by markets' 1991 employment and cluster standard errors by firm.

F.3.2 The bias from involuntary separation out of the formal sector: Ω and $1/\tilde{\eta}$

If the extended model is true, workers respond to expected rather than posted wages, and the IV estimate of equation F.13 recovers only part of the relevant elasticity. Taking logs of equation F.1 gives

$$\ln \bar{w}_{z\bar{g}m} = \underbrace{\ln w_{z\bar{g}m}}_{\text{Formal wage}} + \underbrace{\ln\left(1 + p_{z\bar{g}m}\sigma_{z\bar{g}m}\right)}_{\text{Expected wage gap}} \tag{F.14}$$

and taking logs of equation F.4, with δ_m denoting a market fixed effect, gives

$$\ln \bar{w}_{z\bar{g}m} = \frac{1}{\tilde{\eta}} \ln l_{z\bar{g}m} + \delta_m + \ln \xi_{z\bar{g}m}^{1+\tilde{\eta}} \tag{F.15}$$

Bias formula. Given equation F.15, $1/\tilde{\eta}$ is the second-stage coefficient from an IV regression in which $\ln l_{z\bar{g}m}$ is instrumented by an exogenous labor demand shock $X_{z\bar{g}m}$. Expanding expected wages per equation F.14, and letting $(*)$ denote partialled-out variables to apply the Frisch-Waugh-Lovell theorem, gives

$$\frac{1}{\tilde{\eta}} = \frac{\text{Cov}(\ln \bar{w}^*_{z\bar{g}m}, X_{z\bar{g}m})}{\text{Cov}(\ln l^*_{z\bar{g}m}, X_{z\bar{g}m})} = \underbrace{\frac{\text{Cov}(\ln w^*_{z\bar{g}m}, X_{z\bar{g}m})}{\text{Cov}(\ln l^*_{z\bar{g}m}, X_{z\bar{g}m})}}_{\frac{1}{\eta}\text{: Supply response to formal wage}} + \underbrace{\frac{\text{Cov}[\ln(1 + p_{z\bar{g}m}\sigma_{z\bar{g}m})^*, X_{z\bar{g}m}]}{\text{Cov}(\ln l^*_{z\bar{g}m}, X_{z\bar{g}m})}}_{\Omega\text{: Supply response to expected wage gap}} \tag{F.16}$$

The first term in equation F.16 is the response to formal wages, $1/\eta$, which equation F.13 estimates. The second term, Ω, is the labor supply response to changes in expected formal–informal wage gaps, conditional on changes in formal wages. It is the bias in the baseline model's elasticity if the extended model is true. If workers do not take the possibility of separation into account, the baseline model holds and $1/\eta$ is the relevant elasticity. The sign of Ω depends on local informal wages and on the joint distribution, across the formal firms of a market, of separation probabilities and formal wages.

Interpretation. If Ω is *negative*, *less* labor is supplied to firms with higher expected wage gaps relative to the same market's informal sector wage. That is, workers dislike firms that in expectation pay less than the local informal sector.[64] Conversely, if Ω is *positive*, *more* labor is

[64] The possibility that informal sector options might pay more than formal sector options should not be discarded.

supplied to firms with higher expected wage gaps relative to the same market's informal sector wage. In this case, the inverse elasticity of substitution $1/\eta$ estimated based on the baseline model is *too small* relative to the true inverse elasticity $1/\tilde{\eta}$, meaning that *firms have more labor market power* than originally estimated.

Regression. I estimate Ω from the second-stage regression

$$\Delta \ln \left(1 + p_{z\bar{g}m}\, \sigma^{c}_{z\bar{g}m}\right) = \Omega\, \Delta \ln l^{c}_{z\bar{g}m} + \delta_z + \delta_m + \delta_c + \epsilon^{c}_{z\bar{g}m}, \tag{F.17}$$

where an observation is a firm × microregion × demographic cell, the left-hand side is the 1991–2000 change in the expected wage gap of cell c at firm z, the regressor is the change in the firm's log employment of that cell, and the δ terms are firm, microregion, and cell fixed effects. Including a firm fixed effect ensures that Ω captures the supply response to changes in expected wage gaps relative to the informal sector conditional on firm-level changes in wage levels, the identifying variation for η in both the baseline and extended models. I weight observations by the lagged log of the firm's employment of the cell and cluster standard errors by microregion.

Measurement. RAIS records the reason for every separation, distinguishing firings from quits. I proxy workers' prior about the probability of involuntary separation from firm z in microregion m with the firm's fired-worker share,

$$p_{z\bar{g}m} = \frac{\text{fired}_{zm}}{\text{fired}_{zm} + \text{stayers}_{zm}}, \tag{F.18}$$

where fired_{zm} is the number of workers fired from firm z in microregion m during the calendar year and stayers_{zm} is the firm's employment in the microregion on December 31. The wage gap $\sigma^{c}_{z\bar{g}m}$ compares the informal monthly earnings of cell c in microregion m, measured in the census, with the cell's December earnings at firm z, measured in RAIS and scaled by the ratio of mean census to mean RAIS formal earnings in each year. Earnings are in constant 2000 reais. I drop firm × cell earnings above the 99th percentile and wage gaps outside the 1st–99th percentile range. I merge RAIS and census data by microregion, demographic cell, and year.

Panel A of Figure F.1 plots a binned scatter of firm-level probabilities of involuntary separation against informal–formal wage gaps. The probability is on average 3%. It is highest in local labor markets where the informal sector pays less than the formal sector (for the same demographic cell). This means that workers who are at higher risk of separation would also face lower-paying informal sector jobs. Panel B of Figure F.1 shows that a substantial number of jobs are in local labor markets where the informal sector pays less, but there is a sizable number of jobs for which informal wage work pays higher wages than the formal sector. This is consistent with the residual

Appendix Figure F.1 shows that a non-trivial share of regions exhibit larger monthly earnings for workers of the same demographic composition than monthly earnings at formal sector firms.

wage distributions plotted in Figure F.2.

Instruments. Identification requires labor demand shocks that vary across firms within markets, as for η, and that shift expected wage gaps conditional on firm-level changes in formal wages. I instrument $\Delta \ln l^c_{z\bar{g}m}$ with the interaction between the firm-level tariff reduction and the microregion's trade shock,

$$X^c_{z\bar{g}m} = \left(1 + \Delta \ln \left(1 + \tau_{i(z)}\right)\right) \times \Delta ICE^*_m, \tag{F.19}$$

where $\Delta \ln(1 + \tau_{i(z)})$ is the firm-level tariff reduction of equation 17. I interact the instrument with gender, education, and age group indicators. The firm-level tariff reduction changes the firm's labor demand and workers' perceived probability of separation from it. The microregion's shock changes equilibrium informal wages. Because the regression includes firm and microregion fixed effects, the instrument's variation comes from firms that operate in several microregions, through differences in ΔICE^*_m across a firm's microregions scaled by its own tariff reduction, and from the firm's differential exposure across demographic cells through the interactions. In the first stage, the common component has coefficients of -13.5 for men and -13.4 for women, and the education and age differentials are small. The exclusion restriction is that, conditional on firm, microregion, and cell fixed effects, tariff reductions affect labor supply only through separation probabilities and equilibrium wages, and not through workers' tastes for individual firms. The standard error of $1/\tilde{\eta} = 1/\eta + \Omega$ combines those of its two components as $\sqrt{\mathrm{SE}(1/\eta)^2 + \mathrm{SE}(\Omega)^2}$, treating the two estimates as independent. Since both use the same firms and tariff reductions, they need not be independent; under any correlation between them, the standard error is at most $\mathrm{SE}(1/\eta) + \mathrm{SE}(\Omega) = 0.100$, against 0.097 under independence.

Unemployment insurance. Finally, I consider an alternative measure of the expected wage in which a separated worker receives unemployment insurance before moving to informal wage work. I value earnings after separation at the formal wage for four of the twelve months, the typical benefit at the time, and at the local informal wage for the remaining eight, so that the wage gap is two thirds of its unadjusted value.[65]

F.3.3 Substitution to self-employment and across markets

To arrive at the regression equation to estimate the elasticity of substitution between wage work and self-employment, take logs of equation F.5, express it in long-differences, add a constant to absorb

[65] Brazil's unemployment insurance was instituted by the 1990 law in Brasil (1990). Its rules in the 1990s were involved, with earnings brackets limiting benefit amounts and eligibility contingent on proof of time of service, but the typical benefit was four months of salary, paid monthly. Benefit brackets and their changes over time are documented at https://www.debit.com.br/tabelas/seguro-desemprego.

country-level changes, and group taste shifters into an error term $\epsilon_{\overline{g}m}$ to get:

$$\Delta \ln W_{\overline{g}m} = \alpha + \frac{1}{\tilde{\rho}} \Delta \ln \left(\frac{L_{\overline{g}m}}{L_m} \right) + \frac{1}{\tilde{\theta}} \Delta \ln \left(\frac{L_m}{L} \right) + \epsilon_{\overline{g}m} \tag{F.20}$$

A challenge in estimating equation F.20 is that each labor market has a single wage work sector, and import tariff reductions were primarily a labor demand shock to wage work. As a result, I cannot separately identify $\tilde{\rho}$ from $\tilde{\theta}$ unless I partition markets into groups, each with its own wage work and self-employment sectors, that have differential shocks to wage work labor demand (the within-market, cross-sector variation that identifies $\tilde{\rho}$) but are subject to the same market-level shock (the cross-market variation that identifies $\tilde{\theta}$).[66]

Regression. I partition labor markets into demographic cells, each with its own wage work and self-employment sectors. Let c denote the 18 cells defined by gender (men; women), education (primary; secondary; tertiary), and age (18–29; 30–49; 50–64). I estimate

$$\Delta \ln W^c_{\overline{g}m} = \frac{1}{\tilde{\rho}} \Delta \ln \left(\frac{L^c_{\overline{g}m}}{L^c_m} \right) + \frac{1}{\tilde{\theta}} \Delta \ln \left(\frac{L_m}{L} \right) + \delta_r + \epsilon^c_{\overline{g}m}, \tag{F.21}$$

where Δ denotes long differences within microregion $\times$ cell pairs and δ_r is a fixed effect for each of Brazil's five major regions, which absorbs regional general equilibrium effects of trade liberalization. The cells of a market share the market's employment share L_m/L, so the cross-market margin is common to them. The partition provides the within-market, cross-sector variation that identifies $1/\tilde{\rho}$.

Measurement. I estimate equation F.21 using Brazil's 1991 and 2000 censuses. The indices $W^c_{\overline{g}m}$, $L^c_{\overline{g}m}$, L^c_m, and L_m reflect lower-nest elasticities and taste shifters. To keep the analysis executable with census data, I use their directly observable counterparts. I measure $L^c_{\overline{g}m}/L^c_m$ as the wage work share of cell c's employment in microregion m, L_m/L as the microregion's share of national employment, and $\ln W^c_{\overline{g}m}$ as the mean log residual real monthly earnings of the cell's wage workers. Wage workers are formal and informal employees. I residualize earnings at the individual level on a fully saturated vector of gender, education, age, and year indicators, so that neither demographic composition nor demographic-specific wage trends enter the identifying variation. My sample consists of microregion $\times$ cell pairs with positive 1991 formal employment, and I weight the regression by the cell's 1991 formal employment.

Instruments. I instrument the two endogenous variables in equation F.21 with the microregion's trade shock ΔICE^*_m and its interactions with (a) age group, gender, and education indicators and (b) the cell's 1991 formal employment share. The shock is common to the cells of a microregion. The interactions let the first-stage responses of the cell's wage work share and of the market's

[66] Note that identification of lower-nest elasticities also requires that each partition of the wage work sector has within-sector cross-firm variation in tariff reductions, which precludes partitioning the wage work sector by industry.

employment share differ with the cell's characteristics. The exclusion restriction is that the shock and its interactions with baseline characteristics are orthogonal to changes in workers' taste shifters for wage work and for the market.

F.3.4 Reduced-form effects of trade liberalization on market structure

The two derivatives in the policy effect formula come from long-difference regressions at the microregion level,

$$\Delta\widetilde{HHI}_{\bar{g}m} = \beta_{HHI}\,\Delta ICE^*_m + \delta_r + u_m, \qquad \Delta s_m = \beta_{s_m}\,\Delta ICE^*_m + \delta_r + v_m, \tag{F.22}$$

with region fixed effects and 1991 formal employment weights. I compute the expected payroll Herfindahl in RAIS, with each firm's wage bill valued at expected wages, and the wage work share s_m in the census, as the wage bill of formal and informal wage workers over the market's total wage bill, which includes self-employment earnings. I report effects per 10% increase in the shock.

F.3.5 Wage markdowns

I combine the estimated elasticities with the measured market shares to compute the average inverse elasticity of residual labor supply from equation F.10, the wage take-home share $1/(1+\bar{\varepsilon}^{-1}_{\bar{g}m})$, and, with the reduced forms in equation F.22, the effect of trade liberalization on the take-home share. I evaluate all three at the 1991 averages of $\widetilde{HHI}_{\bar{g}m}$ and s_m across microregions, weighted by 1991 formal employment (0.025 and 0.782), and I compute standard errors by the delta method over the two reduced-form coefficients, holding the elasticities fixed.

F.3.6 Specifications behind Table 7

Column (1) of Table 7 reproduces the paper's main estimates, with local labor markets defined as microregion × occupation pairs: Panel A from Appendix Table A.4, column (4), Panel C from Table 2, Panel E from Table 3, and Panel F from Table 4 and Section 7, where the standard error also reflects the uncertainty in the elasticities. Column (2) re-estimates the baseline model with local labor markets defined as microregions, weighted by the microregion's 1991 formal employment, with ΔICE^*_m as the instrument for $1/\theta$. Column (3) reports the extended model with local labor markets defined as microregions, estimated as follows.

Panels A and B. 1991–2000 long differences of the expected payroll Herfindahl and of the wage work share of market payroll on ΔICE^*_m (equation F.22), per 10% increase in the shock, weighted by the microregion's 1991 formal employment, with region fixed effects and standard errors clustered by microregion.

Panel C. 2SLS estimates of $1/\eta$ from 1991–2000 long differences of firm × microregion wage premia on employment (equation F.13), instrumented with the firm-level tariff reduction, with microregion fixed effects and standard errors clustered by firm; first-stage F statistic of 207.1. "Firm Dec. premium" is the firm × market fixed effect on workers' December monthly earnings, estimated conditional on demographics (Section 5).

Panel D. 2SLS estimate of Ω from equation F.17, over 1991–2000, with the instrument of equation F.19 interacted with gender, education, and age group indicators (six instruments remain after the collinear interactions are dropped); first-stage F statistic of 13.5, Hansen J p-value of 0.034. The standard error of $1/\tilde{\eta} = 1/\eta + \Omega$ combines those of its two components as $\sqrt{\mathrm{SE}(1/\eta)^2 + \mathrm{SE}(\Omega)^2}$, treating the two estimates as independent. The unemployment insurance row values four of the twelve months after separation at the formal wage, so that the wage gap is two thirds of its unadjusted value.

Panel E. Joint 2SLS estimates of $1/\tilde{\rho}$ and $1/\tilde{\theta}$ from equation F.21, over microregion × demographic cells with positive 1991 formal employment (7,425 cells), instrumented with ΔICE_m^* and its interactions with age group, gender, education, and the cell's 1991 formal employment share, with region fixed effects, weighted by the cell's 1991 formal employment, with standard errors clustered by microregion; first-stage F statistics of 228.5 for the sector share and 20.3 for the market share, Kleibergen-Paap Wald F of 10.9, Hansen J p-value of 0.0002. "Residual log earnings" are census log monthly earnings residualized on demographics × year at the individual level and averaged to the cell.

Panel F. Equations F.10 and 25 evaluated at the 1991 averages of the expected payroll Herfindahl and the wage work share across microregions, weighted by 1991 formal employment; the standard error is computed by the delta method over the two reduced-form coefficients of Panels A and B, holding the elasticities fixed.

F.4 Symmetric treatment of informal wage work and self-employment

The extended model treats the two informal margins asymmetrically. This subsection considers the alternative in which workers make no distinction between them: each market has two sectors that workers choose between, formal employment, populated by formal firms, and informal employment, which pools informal wage work and self-employment. Workers value formal jobs at posted wages, so the within-market cross-firm elasticity is the baseline η and there is no bias term Ω, and the average inverse elasticity of residual labor supply takes the form of equation F.10, with $\hat{\rho}$ the elasticity of substitution between formal and informal employment within markets, $\hat{\theta}$ the cross-market elasticity, HHI the payroll Herfindahl of formal firms, and the formal share of the market's wage bill s_m^f in

place of s_m. Estimation follows Appendix F.3 with three changes. First, the trade shock is

$$\Delta ICE_m^{\dagger} \equiv s_{m,1991}^{f} \, \Delta ICE_m^{f} + \left(1 - s_{m,1991}^{f}\right) RTR_m^{*}, \tag{F.23}$$

where $\Delta ICE_m^{f} = -\sum_z s_{zm,1991} \Delta \ln(1 + \tau_{i(z)})$ sums over all of the market's formal firms, weighted by their (linear) 1991 payroll shares, with a zero tariff change for non-tradable firms and firms without a tariff match, and RTR_m^{*} is Dix-Carneiro and Kovak (2017)'s exposure measure, RTR, with industry weights computed from the 1991 census employment of informal wage workers and the self-employed only; $\Delta ICE^{\dagger}$ averages 5.6% across microregions. Second, equation F.21 uses the residual earnings of the cell's formal workers and the formal share of the cell's employment in place of the wage work objects, is estimated over the cells for which the cell-level $\Delta ICE^{\dagger}$ is defined (cells with 1991 formal employment in RAIS and tradable-industry informal or self-employment in the 1991 census), weighted by the cell's 1991 population, and instrumented with the cell-level counterpart of $\Delta ICE^{\dagger}$ and its interactions with age group, gender, education, and the cell's 1991 formal employment share. Third, the outcomes of equation F.22 are the payroll Herfindahl and the formal share of the market's wage bill.

Table F.1 reports the results next to those of the extended model. A 10% increase in the trade shock raises the payroll Herfindahl by 0.016 points (SE of 0.004) and lowers the formal share of the market's wage bill by 0.120 (SE of 0.016); the cross-market inverse elasticity is 3.526 (SE of 1.790); the implied take-home share is 53.3 cents on the dollar, and a 10% increase in the trade shock lowers it by 0.37 cents. The inverse elasticity of substitution between formal and informal employment is 0.023 (SE of 0.060): taken at face value, formal and informal employment are close to perfect substitutes. This estimate is biased toward zero if the trade shock moved workers into informal employment within firms (Ulyssea, 2018; Derenoncourt et al., 2025), since the resulting changes in formal and informal employment trace firms' labor demand rather than workers' labor supply. Figures F.3 and F.4 show the pattern: across microregions, larger values of RTR are associated with lower employment shares and lower earnings in wage work, the same-sign pattern of a labor demand shock that traces out labor supply, but with higher employment shares and lower earnings in self-employment and in informal wage work. Absent data that separate the intensive and extensive margins of informality in this period, I cannot estimate the elasticity of substitution between formal and informal wage work from these moments.

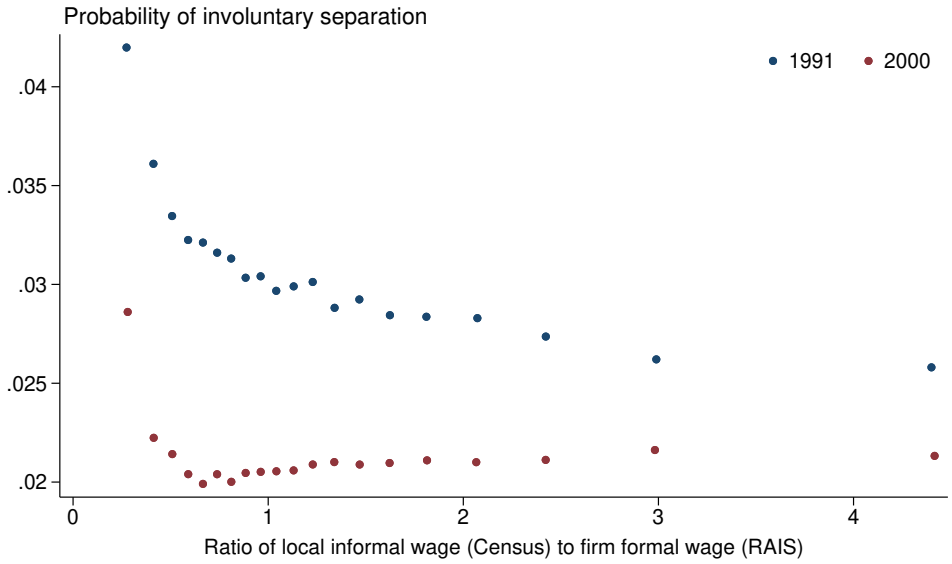
Table F.1: Extended model and symmetric treatment of informal wage work and self-employment

	Extended model (main)	Assuming no preference distinction between informal wage work and self-employment
	(1)	(2)
Panel A: Effect of import competition on formal labor market concentration		
Δ Concentration, per 10% increase in the trade shock	0.028	0.016
	(0.007)	(0.004)
Trade shock	ΔICE^{*}	$\Delta ICE^{\dagger}$
Concentration measure	*Expected* Payroll Herf.	Payroll Herf.
Level of analysis	Mmc	Mmc
Panel B: Effect of import competition on the formal / wage work share of market payroll		
Δ Sector share, per 10% increase in the trade shock	-0.032	-0.120
	(0.005)	(0.016)
Share	Wage work	Formal employment
Trade shock	ΔICE^{*}	$\Delta ICE^{\dagger}$
Level of analysis	Mmc	Mmc
Panel C: Within-market cross-firm inverse elasticity of substitution		
$1/\eta$ (response to formal wages)	0.870	0.870
	(0.062)	(0.062)
Trade shock	Firm-level tariff reduction	
Level of analysis	Firm × mmc	Firm × mmc
Wage measure	Firm Dec. premium (all workers), pair pre-trend controls	
Panel D: Bias Ω in 2SLS estimates of $1/\eta$ due to risk of separation from formal job		
Ω (response to expected formal–informal wage gaps)	0.0091	–
	(0.0025)	
$1/\tilde{\eta} = 1/\eta + \Omega$	0.879	–
	(0.062)	
Trade shock	Firm tariff × ΔICE^{*}	–
Level of analysis	Firm × mmc × demog. cells	–
Fixed effects	Firm, mmc, demographic cell	–
Panel E: Cross-sector and cross-market inverse elasticities of substitution		
Within-market cross-sector inverse elasticity	0.684	0.023
	(0.234)	(0.060)
	$1/\tilde{\rho}$ (wage work vs. self-emp.)	$1/\hat{\rho}$ (formal vs. informal)
Cross-market inverse elasticity	1.187	3.526
	(0.481)	(1.790)
	$1/\tilde{\theta}$	$1/\hat{\theta}$
Trade shock	ΔICE^{*}	Cell $\Delta ICE^{\dagger}$
Fixed effects	Region	Region
Level of analysis	Mmc × demog. cells	Mmc × demog. cells
Wage measure	Residual log earnings, wage workers	Residual log earnings, formal workers
Panel F: Effect of import competition on the average wage take-home share		
Δ average inverse elasticity, per 10% increase in the trade shock	0.0053	0.0132
	(0.0014)	(0.0042)
Implied change in take-home share (pp)	-0.15	-0.37
Baseline (1991) wage take-home share	0.531	0.533
Trade shock	ΔICE^{*}	$\Delta ICE^{\dagger}$

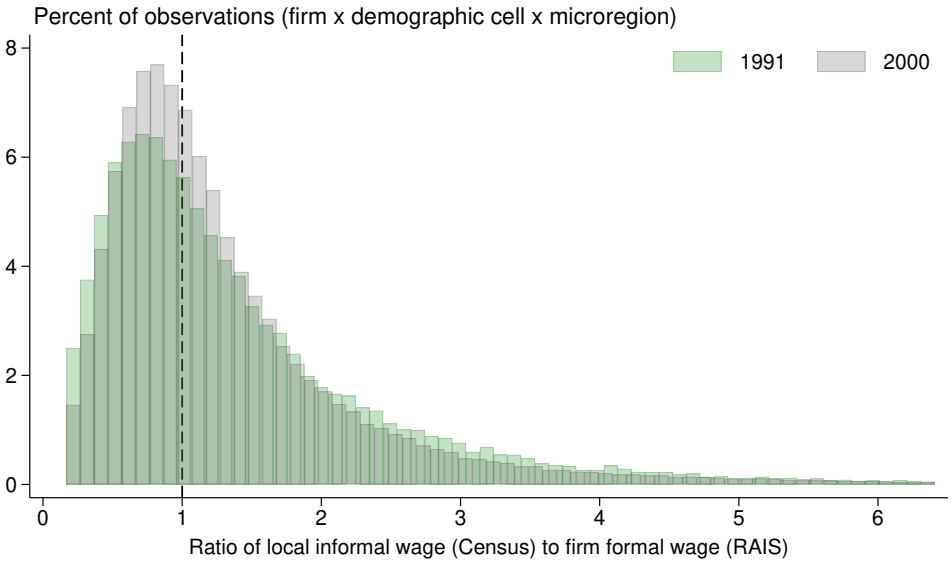
Notes: Standard errors in parentheses. Column (1) repeats column (3) of Table 7. Column (2) reports the model of Appendix F.4, in which informal wage work and self-employment form a single informal sector that workers choose as an alternative to formal employment at posted wages. Specifications are as in Table 7, with $\Delta ICE^{\dagger}$ as the trade shock, the payroll Herfindahl as the concentration measure, the formal share of the market's wage bill as the sector share, and the residual log earnings of formal workers as the wage measure in Panel E. Panel E, column (2), differs in three further respects (Appendix F.4): it covers the cells for which the cell-level $\Delta ICE^{\dagger}$ is defined (6,291 cells in 469 microregions), is weighted by the cell's 1991 population, and is instrumented with the cell-level $\Delta ICE^{\dagger}$ and its interactions. Panels A and B, column (2), include 469 microregions: those of column (1) except the microregions dropped by Dix-Carneiro and Kovak (2017). First-stage F statistics in Panel E, column (2), are 60.7 and 11.2.

Figure F.1: Probability of involuntary separation from formal firm and informal-formal wage gaps

Panel A: Probability of involuntary separation and wage gap



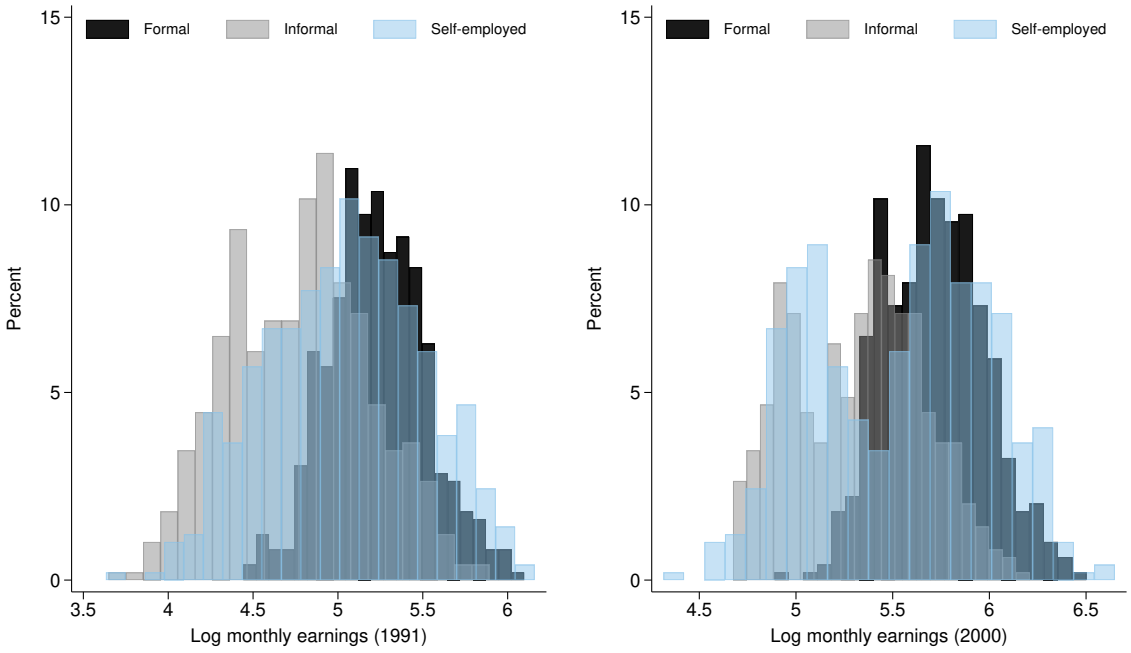
Panel B: Wage gap distribution



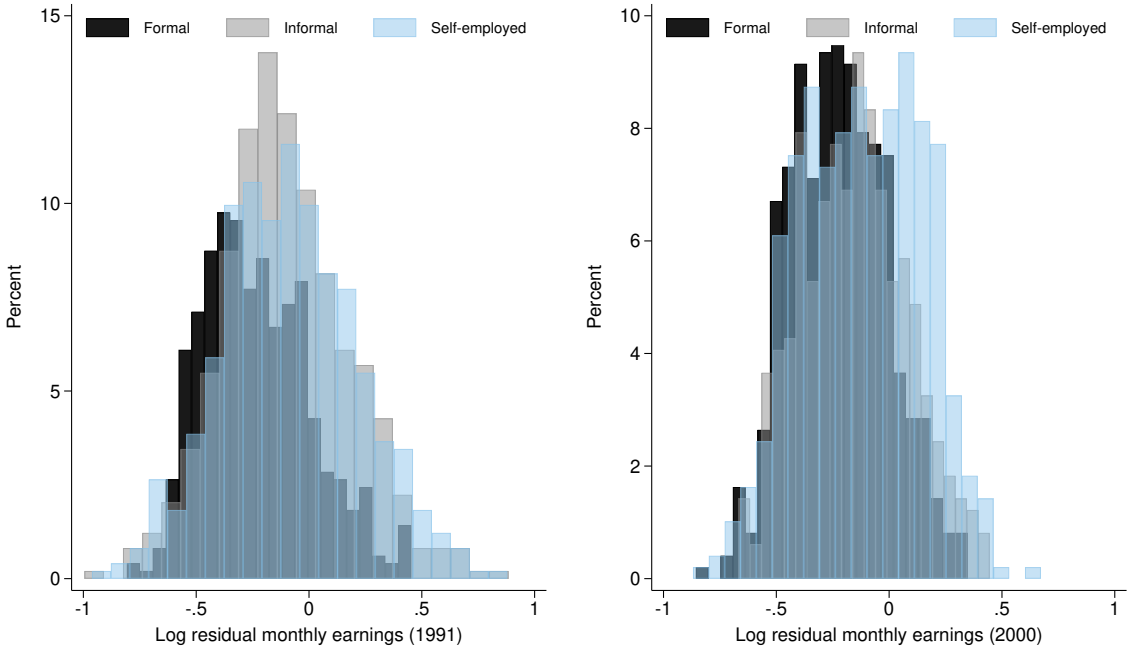
Notes: This figure shows the relationship between probabilities of involuntary separation from a formal sector firm in RAIS and the gap between local informal sector monthly earnings and the firm's formal monthly earnings, separately computed for each demographic cell. The sample includes all formal sector firms in RAIS in 1991 and 2000, merged at the microregion X demographic cell X year level to the corresponding census data. Each observation is a firm X microregion X demographic cell (gender X education group X age group dummies). Monthly earnings gaps are calculated by first converting all reported earnings to 2000 constant Reais. The probability of involuntary separation is calculated using separations between January 1 and December 31 of each year. Firm total formal employment is measured as of December 31 of each year.

Figure F.2: Formal, informal, and self-employment earnings

Panel A: Raw earnings



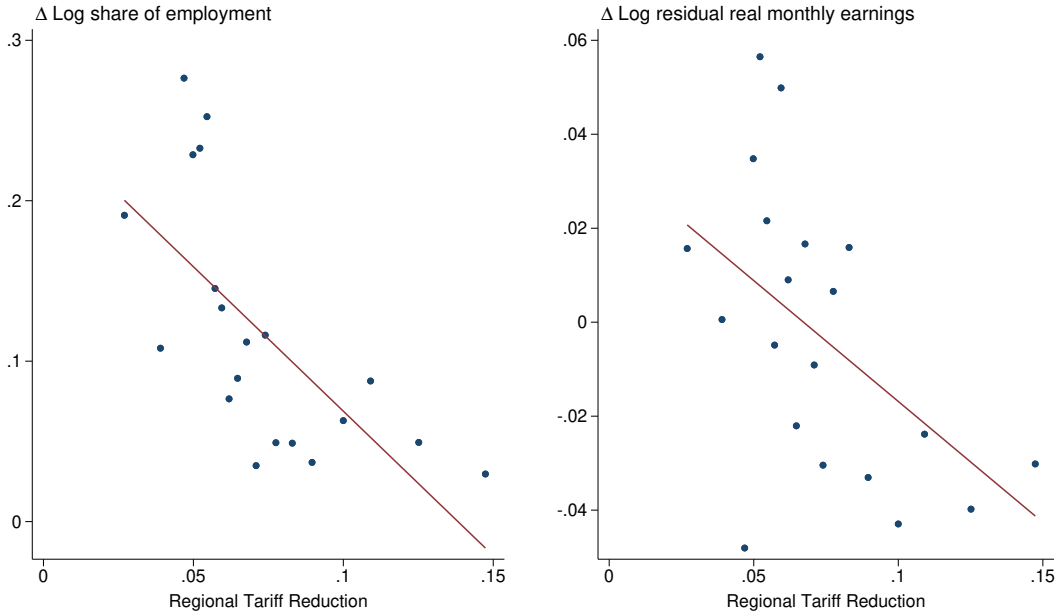
Panel B: Residual earnings



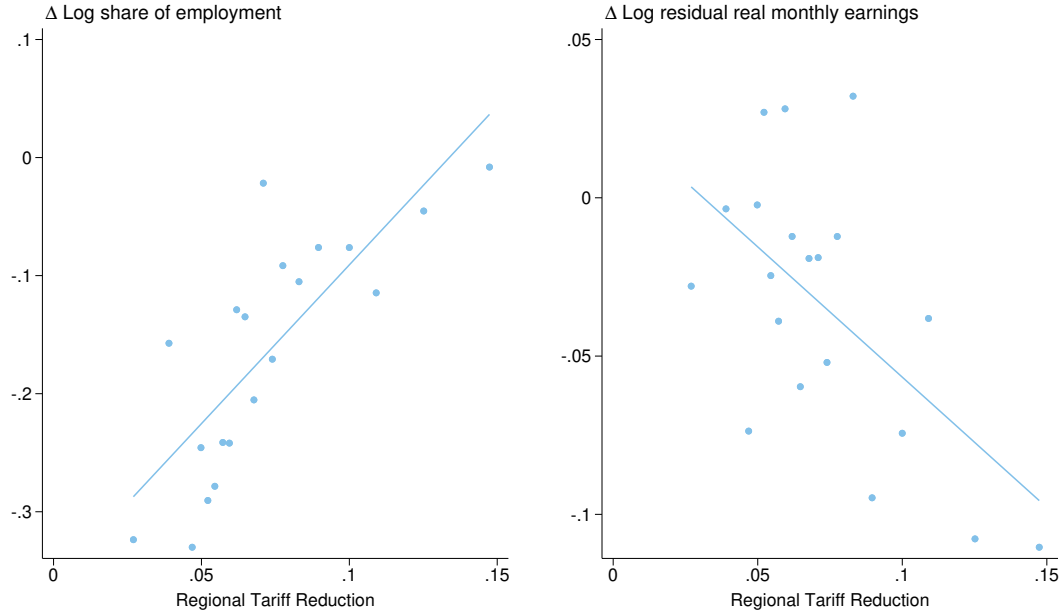
Notes: This figure plots the distribution of raw and residual formal real monthly earnings, informal real monthly earnings, and self-employment real monthly earnings, across 486 Brazilian microregions. Residual real monthly earnings are conditional on a fully saturated vector of dummies for gender, age groups, and education groups. Real earnings are expressed in 2000 reais.

Figure F.3: Effect of Dix-Carneiro and Kovak (2017)'s RTR on wage work and self-employment

Panel A: Wage work



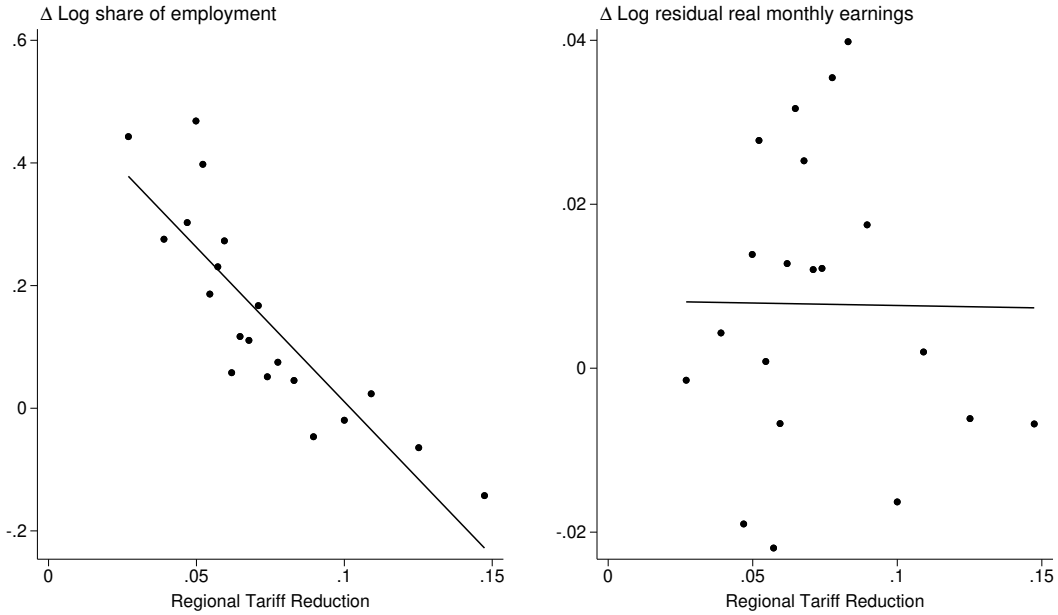
Panel B: Self-employment



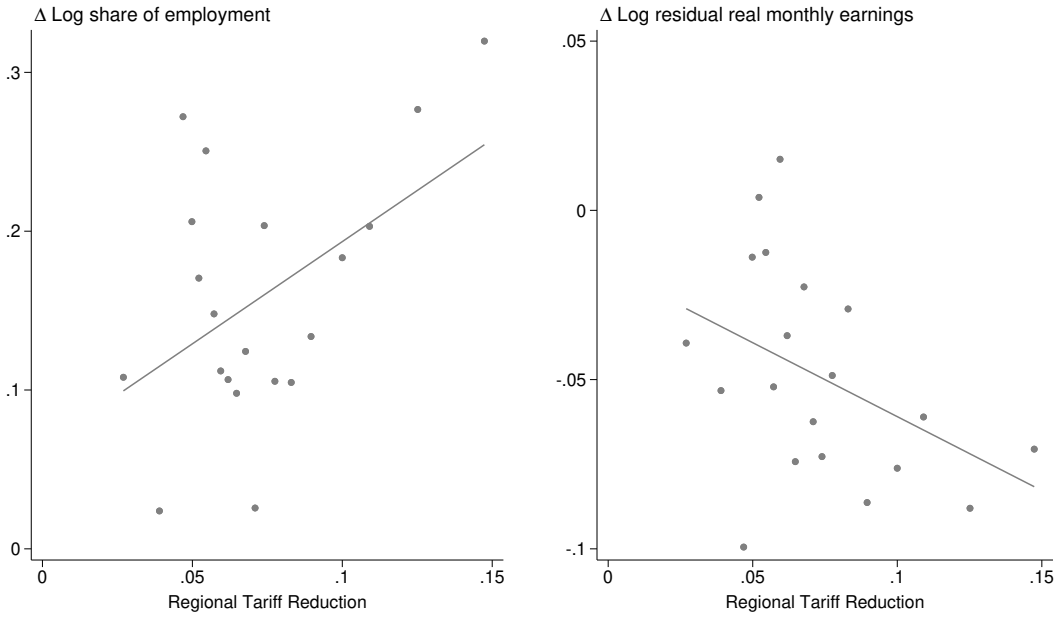
Notes: This figure shows the correlation between Dix-Carneiro and Kovak (2017)'s exposure measure, RTR (regional tariff reductions, i.e., microregion-level exposure to 1990–1994 import tariff reductions), changes in employment shares (left panel), and changes in log residual real monthly earnings (right panel) between the 1991 and 2000 census, for individuals employed in wage work (formal or informal) (Panel A) versus in self-employment (Panel B) across 486 Brazilian microregions. Log residual real monthly earnings are conditional on flexible controls for gender, age, and education. Real earnings are computed from nominal earnings by expressing all values in 2000 reais.

Figure F.4: Effect of import competition on formal versus informal wage work

Panel A: Formal Wage Work



Panel B: Informal Wage Work



Notes: Each panel shows the correlation between Dix-Carneiro and Kovak (2017)'s exposure measure, RTR (regional tariff reductions, i.e., microregion-level exposure to 1990–1994 import tariff reductions), changes in employment shares (left) and log residual real monthly earnings (right) between 1991 and 2000 across 486 Brazilian microregions. Log residual real monthly earnings are conditional on a fully saturated vector of dummies for gender, age groups, and education groups. Real earnings are expressed in 2000 reais.