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TRADE, LABOR MARKET CONCENTRATION, AND WAGES

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ABSTRACT

I estimate the effect of trade on local labor market concentration and its implications for wages using employer-employee linked data and tariff shocks from Brazil's trade liberalization. Trade increased concentration by 7%, an effect driven by firm exit and worker flows to surviving import-competing firms. Increased concentration reduced wage take-home shares—estimated at 50 cents on the dollar pre-shock—enough to offset small wage gains from reallocation, but did not meaningfully reduce wages on net. Most of the wage declines attributed to Brazil's trade liberalization resulted instead from reductions in the marginal revenue product of labor. Incorporating informality reveals substantial regional heterogeneity.

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An online appendix is available at http://www.nber.org/data-appendix/w35018

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), and the U.S. (Autor, Dorn and Hanson, 2013). Why? This paper tests one potential mechanism: 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 kicking out 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 for several reasons. First, formal workers experienced the most negative effects of import competition on wages (Kovak, 2013), and these effects have persisted for decades (Dix-Carneiro and Kovak, 2017). 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., forthcoming).

Using a sufficient statistics approach, employer-employee linked data, and import tariff reduction shocks, I test whether trade increased labor market concentration among formal employers in more exposed markets, and estimate the consequences for formal sector wages. I then quantify how much of trade liberalization's net-negative effect on local formal sector wages is accounted for by increased labor market power. Trade increased concentration by 7%, driven by direct worker reallocation from shrinking to expanding non-exporting tradable sector firms, consistent with the pro-competitive effects of import competition predicted by Melitz and Ottaviano (2008). Higher concentration raised average wage markdowns—estimated at 50 cents on the dollar pre-shock—by enough to offset small reallocation gains, but increased markdowns account for only 6% of trade's overall negative effect on average wages. Within-firm reductions in the marginal revenue product of labor, consistent with reduced output prices, account for the remaining 94%. These conclusions—which concern average effects—are largely unaffected by informal wage work and self-employment, but mask substantial heterogeneity in formal firms' wage markdowns across regions and their response to trade.

I develop a parsimonious model of imperfectly competitive labor markets where firm size matters for wage setting. The model provides the link between local labor market concentration and wage markdowns, a standard measure of firm labor market power defined as the ratio of the marginal revenue product of labor to the wage. On the supply side, workers have idiosyncratic preferences over jobs, yielding labor supply decisions that, when aggregated, follow a nested CES structure, as in Berger, Herkenhoff and Mongey (2022), henceforth BHM. On the demand side, firms maximize profits holding constant local employers' labor demand—namely, competing for workers à la Cournot. In this environment, labor reallocation in response to shocks is governed by two key elasticities: a *cross-market* elasticity of substitution, and a *within-market* cross-firm elasticity of substitution. Along with a firm's payroll share in its local labor market, these key elasticities of substitution determine the firm's wage markdown.

This paper's first theoretical result concerns the link between local labor market concentration and the average wage markdown at a specific local labor market. Taking a weighted average of firms' markdowns across all firms in a market, I show that the market-level average markdown is determined by the same two key elasticities of substitution, along with the market's payroll Herfindahl index, defined as the sum of its firms' squared payroll shares. This result holds regardless of the shape of firms' production functions or the competition structure in product markets, on which I remain agnostic. Overall, the more concentrated a market is, and the more inelastic the elasticities of substitution are, the larger is the market's average markdown.

The second result concerns the effect of import competition exposure on a local labor market's average wage markdown. A direct implication of my model's expression for a market's average wage markdown is that its response to an increase in import competition can be quantified by just two sufficient statistics: the effect of import competition exposure on local labor market concentration, and the *gap* between workers' cross-market vs. the within-market cross-firm inverse elasticities of substitution. To see the intuition for why the *gap* in elasticities is what matters for *changes* in markdowns, consider the following. Trade fundamentally changes firms' relative size. But, if it is just as easy for workers to substitute locally (i.e., within markets) as it is for them to substitute globally (i.e., across markets), then firms effectively operate in a single national market, where their relative size is negligible and inconsequential to market power. Overall, the larger the gap between the key elasticities, and the larger the effect on concentration, the larger the effect on markdowns.

With clear guidance on the key sufficient statistics needed to quantify the effect of trade on firm labor market power, I proceed to estimate them using employer-employee linked data and Brazil's trade liberalization. In 1990, Brazil announced an import tariff reduction reform, to be completed by 1994, whereby import tariffs on all sectors would be reduced from a pre-reform average of 33% to a post-reform average of 13%. As sectors differed in their pre-reform levels of protection, the reform generated substantial cross-sector variation in import tariff changes. This cross-sector variation in 1990-1994 changes in import tariffs is the policy-induced variation I exploit to estimate my model's sufficient statistics.

My empirical analysis begins by estimating the effect of import competition exposure on local labor market concentration. 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 throughout the entire period, and report results using microregion-only boundaries as robustness. My identification strategy leverages local labor markets' differential exposure to import tariff reductions depending on each market's pre-liberalization sectoral composition, similar to the approach in Dix-Carneiro and Kovak (2017). I estimate a difference-in-differences regression of the change in a local market's payroll Herfindahl on the change in its "import competition exposure," a shift-share treatment intensity measure whose "shift" is the set of tariff reductions experienced by each firm in the local labor market, and whose "share" is each firm's contribution to its market's baseline year payroll Herfindahl. This particular functional form is guided by the model outlined above, though I also consider alternative measures as robustness checks.

I find that a 10 percent increase in import competition exposure increased local labor markets' payroll Herfindahl by 0.02 points relative to less affected markets, with no evidence of pre-trends. This effect is quite large: it corresponds to a 7% increase relative to the 0.28 pre-reform unweighted mean, or a 27% increase relative to the 0.08 payroll-share-weighted mean. The effect is robust to alternative measures of import competition exposure and concentration, to defining labor markets solely as microregions, and to weighting by baseline size, showing that it is not driven by a handful of small markets.

To examine the source of increased concentration, I trace direct worker flows between three firm types—exporters, non-exporting tradable firms, and non-tradable firms—within each local labor market. Import competition primarily reduced employment among non-exporting tradable firms, consistent with Melitz and Ottaviano (2008), with little detectable effect on exporters or non-tradable firms. The key driver of increased concentration is direct within-market cross-firm reallocation of workers from shrinking to expanding import-competing firms (i.e., non-exporting tradable sector firms). Consistent with Melitz and Ottaviano (2008), total employment at exporters did not meaningfully respond to unilateral import tariff reductions, contributing only modestly to the increase in concentration as a larger share of the surviving total employment is located in exporting firms, who are much larger, but not via direct worker reallocation towards exporters.

The next step towards quantifying the effect of import competition on firm labor market power is to estimate workers' two key elasticities of substitution. My model provides the regression specifications, and my setting the quasi-exogenous variation. The availability of trade shocks that vary across firms within markets allows 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 the within-market cross-firm elasticity of substitution using within-market cross-firm variation in tariff reductions as shocks to firm wage premia and employment, and the cross-market elasticity using cross-market variation in changes to import competition exposure as shocks to indices of market wage premia and employment.

I estimate a within-market cross-firm inverse elasticity of substitution of 0.990, and a cross-market inverse elasticity of substitution of 1.448. Both point estimates are robust to alternative tariff shocks and to relevant alternative samples, and are not driven by unobservable worker characteristics or by changes in workforce composition, both of which I can control for in estimating firms' wage premia. These elasticities—along with the pre-liberalization level of labor market concentration—imply that prior to liberalization, Brazilian workers took home only 50 cents for every marginal dollar they generated for the firm.

This suggests substantial levels of firm labor market power in Brazil's formal sector—much higher than, for example, estimates for the US, which range from 65 to 80 cents on the dollar, though consistent with estimates for other developing countries.[1] Comparing to estimates by BHM for the US, the key difference between the two contexts is that Brazil's within-market cross-firm elasticity of substitution is *seven* times as inelastic as the US', suggesting that Brazilian workers have a much tougher time making within-market cross-firm substitutions than US workers do.[2] Combined with the estimates for the effect of import competition on labor market concentration, the 0.459 gap between these elasticities implies that a 10% increase in import competition exposure reduced local labor markets' average wage take-home share by 0.24 cents on the dollar, via a statistically significant increase in wage markdowns. However, increased wage markdowns account for only about 6% of the overall negative effect of trade on average wages (Table 6); the remaining 94% was driven by within-firm reductions in the other component of the wage, the marginal revenue product of labor.

These baseline results focus on formal firms alone. However, Brazil's labor market is dual: formal firms co-exist with a large informal sector and substantial self-employment. Section 8 extends the model to incorporate self-employment as an alternative to wage work, informal wage work as a consequence of involuntary separation from formality, and preference heterogeneity across demographics and regions. Self-employment substantially curbs formal firms' market power by providing an attractive outside option, while the threat of involuntary separation into informal wage work slightly increases it. On average, these margins bring the wage take-home share to 51 cents on the dollar—close to the baseline's 50 cents, because most formal sector employment is located in predominantly formal microregions (Figure 1), and most of these microregions are in the Southeast, where the within-market cross-firm elasticity is most inelastic (Appendix Table C.2)—but the extension reveals substantial heterogeneity: 46 to 73 cents depending on local conditions outside the formal sector and workforce composition (Table 7). The effect of trade on markdowns through the extended model is also muted on average, but more heterogeneous across regions, with the North and Northeast experiencing a 1.2 p.p. *increase* in wage take-home shares.

Overall, this paper offers three new take-aways from Brazil's trade liberalization episode concerning the interaction between trade, labor market concentration, and wages: (i) Brazilian firms command substantial firm labor market power, primarily driven by difficult within-market cross-firm substitution relative to contexts such as the US; (ii) Opening to trade increased that labor market power a bit further as it raised concentration, by enough to offset wage gains from cross-firm reallocation, but (iii) on net the magnitude of the market power effect was small, and cannot explain most of the relative wage decline due to trade. Combined with evidence that trade raised firm productivity in Brazil (Muendler, 2004) and that the effects of import

[1] Amodio et al. (2025) apply a production function approach to estimate wage markdowns in a harmonized global panel of firms from 82 low- and middle-income countries. While Brazil is not in their sample, consistent with this paper they find a median wage take-home share of 43 cents on the dollar and that markdowns are increasing in firm size.

[2] These preference parameters might be microfounded in institutional and cultural factors that differ between the two countries and that shape worker preferences. For example, Brazil's inelastic cross-firm substitution might be a result of high valuations for formal sector job amenities combined with high search frictions. Using discrete choice experiments, Felix et al. (2026) find that Brazilian workers, and especially those formally employed, have extraordinarily high willingness to pay for formal sector attributes (i.e., 34% of wages for unemployment insurance, 24% for parental leave, 18% for termination notice). They also find that, conditional on searching for jobs, workers primarily search for formal sector jobs. For comparison, recent estimates of willingness to pay for typical job attributes in the US from Maestas et al. (2023) are substantially smaller. US workers are willing to pay less than 10% for 8 of the 12 attributes they evaluate, the most comparable across the two studies being training opportunities, valued at 5.4% by US workers in Maestas et al. (2023) but at 28.5% by Brazilian workers in Felix et al. (2026).

competition are very similar whether measured using output tariffs or effective rates of protection, within-firm MRPL reductions likely reflect reduced output prices rather than productivity losses—consistent with the pro-competitive effects predicted by Melitz and Ottaviano (2008).

Literature. The analysis speaks to large literatures on the regional incidence of trade, on labor market power in developing countries, and on methods for estimating wage markdowns. Brazil's liberalization has been widely studied for various outcomes (e.g., Muendler (2004), Gonzaga, Menezes Filho and Terra (2006), Krishna, Poole and Senses (2012), Dix-Carneiro and Kovak (2017), Dix-Carneiro, Soares and Ulyssea (2018)). More recently, Dix-Carneiro et al. (forthcoming) study how size-dependent distortions—taxes, regulations, and labor market frictions—change the aggregate gains from trade liberalization in Brazil.

My focus is on mechanisms: I zoom into the wage effects on formal sector workers to test whether trade-induced increases in labor market concentration reduce 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 comprehensive study of the relationship between trade, labor market concentration, and wages. I provide the theoretical link between concentration and average markdowns, derive the sufficient statistics needed to quantify the effect of shocks to the former on the latter, estimate both the negative (via markdowns) and positive (via reallocation) wage effects of trade through concentration, and extend the model to show how formal sector markdown levels and their response to trade change when self-employment, informal wage work, and demographic heterogeneity are incorporated.

The estimates here are among the first of wage markdowns in a developing country, documenting a take-home share of 50 cents on the dollar. Since an earlier version of this paper was first circulated, this literature has greatly advanced. Amodio et al. (2025) use a production function approach to estimate wage markdowns in a panel of firms across 82 low- and middle-income countries, finding a median take-home share of 43 cents on the dollar and that markdowns increase in firm size, consistent with my findings for Brazil, though Brazil is not in their sample. Amodio, Medina and Morlacco (2025) use a labor supply approach to estimate wage markdowns in Peru, finding that self-employment curbs oligopsony power, with comparative advantage driving worker sorting across sectors. Evidence of strategic competition for workers continues to emerge: Sharma (2024) detects collusion among Indian textile firms.

On methods, the labor supply approach to estimating wage markdowns builds on work in international trade (Goldberg (1995), Atkeson and Burstein (2008)) that leverages nested preference structures to approximate market-level behavior originating from richer preferences. Grigolon and Verboven (2014) show that nested structures can approximate key substitution patterns of richer random coefficients models while remaining tractable. Random utility maximization has been increasingly adopted in labor economics to model and

[3] Segerstrom and Sugita (2015) show that extending Melitz (2003) to a multi-industry setting reverses the model's difference-in-differences prediction for bilateral liberalization: productivity rises more in nonliberalized industries, as cross-industry labor reallocation and wage adjustments offset within-industry selection. This critique does not apply to unilateral liberalization settings such as Brazil's, where the relevant framework is Melitz and Ottaviano (2008) and wages are pinned down by a freely traded outside good, shutting down the cross-industry wage channel driving the reversal.

estimate worker preferences (Mas and Pallais, 2017; Azar, Berry and Marinescu, 2022; Maestas et al., 2023; Felix et al., 2026). Market-level analyses (e.g., Berger, Herkenhoff and Mongey (2022)) follow the micro-to-macro approach now common in other fields. By showing when and how the key elasticities of substitution in models of oligopsony can be identified with IV, as opposed to indirect inference, I also contribute to a broader 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)).

The analysis extension in Section 8 advances this work by embedding 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 formal-sector 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 serve as the source of formal firms' market power. Felix et al. (2026) conducts discrete choice experiments in Brazil's largest slum complex, finding that formal workers value formal amenities most, the self-employed least, and informal workers have mixed valuations, consistent with preference-based sorting across sectors. The extension shows that the threat of involuntary separation into informality or unemployment increases formal firms' power, and provides a bias formula for market power estimates when only employer-employee linked data is available. It also delivers heterogeneous markdown estimates by gender, age, education, and region, contributing to the literature on markdowns and gender (e.g., Sharma (2023); Hoang, Mitra and Pham (2024)).

2 Concentration and markdowns: An empirical model

In this section I introduce an empirical model of Brazilian labor markets that provides the relationship between labor market concentration and wage markdowns. As in BHM, labor supply is nested CES, firms compete for workers à la Cournot, and there is a large number of labor markets.[4] Combined, these assumptions imply that the impact of trade on firm labor market power can be quantified by two key sufficient statistics only: the effect of trade 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.

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.

[4] 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 25). 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 economy consists of a continuum of homogenous 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_{zm}^j 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}:

$$\min_{zm} V_{zm}^j = \ln l_{zm}^j + \ln \xi_m + \ln \xi_{zm} - \xi_{zm}^j$$
$$\text{s.t. } l_{zm}^j w_{zm} \geq y^j$$

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_{zm}^j is an idiosyncratic worker taste shifter with a General Extreme Value (GEV) distribution:[5]

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

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,[6] whose nesting structure is shown in Figure 2 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_{zm}^j 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 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

[5]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.

[6]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.

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 a similar intuition as 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,[7] 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

[7] 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).

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

As in BHM, a nice 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.[8] 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.[9] 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 on changes in the marginal revenue product of labor (via changes in productivity) and on wage markdowns (via relative sizes). At the local labor market level, Corollary 1 below isolates the effect on markdowns.

[8] 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.

[9] Around 3% of firms are multi-establishment (operating in different microregions) in this context. Appendix Figure A.2 shows near non-existent within-firm cross-region movement in the data.

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:*

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

where $\bar{r}_m$ and $\bar{w}_m$ are market m's (employment-weighted) average marginal revenue product of 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 directly proportional to its level of concentration, and more specifically to the 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.[10]

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.

[10] 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.

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. 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. RAIS covers the universe of Brazilian formal sector workers. I focus on the sample of private sector workers aged 18 to 65. 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. 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 were provided via 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 Section 8, downloaded from the replication data for Dix-Carneiro and Kovak 2017.

3.2 Brazil's formal sector labor market structure

In 1991, 42% of employed individuals held formal sector jobs, 28% informal jobs (i.e., worked for wages but without a formal contract that guarantees formal sector benefits like unemployment insurance), and 30% were self-employed (Appendix Table A.1). By 2000, the formal wage work share had fallen to 36% of 55.6 million employed, while informal wage work rose to 38% and self-employment fell to 25%.

Most formal employment is located in regions where employment is primarily formal (Figure 1). In terms of flows, 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 C.1 and Appendix Table A.4). Flows also reveal that Brazilian formal labor markets have geographic and occupational components. 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.2 and A.3). In contrast, most firm switchers change sectors (Appendix Figure A.4).

Based on these patterns, I define a formal sector local labor market as a microregion × occupational group cell,[11] and present effects with local labor markets defined by microregions only as robustness.[12] Appendix Table A.2 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 0.28, 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 (Appendix Table A.3). Formal sector wages and employment decline with formal sector employment concentration (Appendix Figure A.1), consistent with labor market oligopsony.

While these patterns suggest that an analysis of formal sector wage markdowns can be reasonably conducted using employer-employee linked data on formal firms alone, Section 8 extends the paper's model to provide a comprehensive analysis that explicitly incorporates the informal sector.

[11] 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 (2021)) or region × sector (e.g., BHM, Lamadon, Mogstad and Setzler (2022), and Alfaro Urena, Manelici and Vasquez (2021)), with few studies using region only boundaries (e.g., Pham (2023)). Section 6 shows that finer boundaries yield similar but more precise estimates of elasticities of substitution.

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

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%,[13] 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 striking correlation between pre-liberalization tariff levels and reform-induced tariff cuts, as documented in Figure 3, is precisely the biggest 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.

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 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 on these outcomes. Reassuringly, Appendix Figure A.5 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 on local labor market concentration

My first step towards quantifying the effect of trade 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 on local labor markets' payroll Herfindahl indices. As discussed in Section 3, I define a local labor market as a microregion $\times$ occupational group cell, and present robustness to market boundaries defined by microregion only.

4.1 Empirical strategy

My identification strategy for estimating the effect of trade on local labor market concentration follows the shift-share treatment intensity approach adopted by other papers on the regional incidence of trade (e.g.,

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

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,[14] $s_{zm,1991}$ 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.

In other words, ΔICE_m is a weighted average of the firm-level shocks experienced by tradable sector firms,[15] 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, $HHI_{m,1991}^T \equiv \sum_{j\in\Theta_m^T} s_{jm,1991}^2$.[16] 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, 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. I then present robustness checks to alternative definitions of ΔICE and to alternative measures of tariff shocks.[17]

Figure 4 displays the variation in ΔICE_m across geography for two example occupations, while Appendix Table A.2 provides the mean and key percentiles of the distribution of ΔICE_m across local labor markets. The mean change in import competition exposure was 12%, 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

[14] 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.

[15] A small number of microregion × occupation markets have no tradable sector firms in 1991. I set ΔICE_m to zero in those markets, which serve as pure controls in market-level regressions. When markets are defined by microregion only, all markets have at least one tradable sector firm, but results are similar.

[16] By construction, the κ_{zm} weights sum to one, constituting "complete shares" (Borusyak, Hull and Jaravel (2022)).

[17] My measure of import competition exposure serves as a shift-share shock for identification but does not have an independent structural interpretation as in Kovak (2013), which would require assumptions on production functions, product market structure, and the equilibrium entry game.

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,[18] and δ_m and δ_k are local labor market and year fixed effects. As the specification is in stacked differences, 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.[19]

Since equation 12 is a difference-in-differences regression with shift-share treatment intensity, causal interpretation of ζ_k coefficients depends on two assumptions: a) that the import tariff "shifts" composing ΔICE_m are as good as randomly assigned (i.e., shock-driven identification, see Borusyak, Hull and Jaravel (2022)), an assumption discussed in Section 3.3 and which relies on the reform-driven nature of the tariff reductions; and b) that absent trade liberalization, the potential outcomes of markets more exposed to import competition would have followed the same trend as those of least exposed markets, an untestable assumption whose reasonableness can be argued by the lack of pre-trends, to which I turn next when discussing my findings.

4.2 Estimates of effect of trade on concentration

Figure 5 and Table 1 present my main estimates of the effect of trade on local labor market concentration. Column (1) shows the main specification: a 10 percent increase in import competition exposure increased local labor markets' (wage premium) payroll Herfindahls by 0.02 points (SE of 0.002). This is a 7% increase relative to the pre-liberalization 0.28 unweighted average, or a 27% increase relative to the 0.08 payroll-share-weighted average.

This effect is large in magnitude and it is robust to various alternative specifications, several of which are presented in the remaining columns of Table 1. Within column (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. The effect on concentration is also present, and is about half as large, when effective rates of protection—much noisier measures of tariff shocks—are used to construct ΔICE_m (column (2) of Table 1).[20] The effect is also present even when labor markets are defined more broadly by microregions only (column (3)), and when dropping markets with no change in import competition exposure (column (4)).[21]

[18] Following Dix-Carneiro and Kovak (2017): long differences use 1991 as the base year; for pre-treatment years, ΔY_{mt} is the long difference from 1991 back to year t to keep the timing convention consistent.

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

[20] Effective rates of protection are output tariffs netted out of input tariffs, constructed using Brazil's 1995 input-output table at broader sector levels (43 sectors) than output tariffs (CNAE95, 285 sectors). Smaller treatment effects are expected due to attenuation bias from the greater noise. See Appendix A.

[21] The effect is also robust to alternative weights for ΔICE_m and to weighting by baseline employment, confirming it is not driven by small markets. Consistent with the labor supply framework, $s^2_{zm,1991}$ weights yield the least noisy estimates. Statistical significance is also robust to two-way clustering by occupation and region, and to spatial-correlation-adjusted standard errors following Adao, Kolesár and Morales (2019). See Appendix Tables A.8–A.10.

I also estimate equation 12 for local labor market employment and wage premia, presented in Appendix Table A.5. My estimates for the effect of import competition on employment and wage premia are in line with patterns documented by Dix-Carneiro and Kovak (2017): trade liberalization reduced employment and wages in local labor markets more exposed to import competition relative to less exposed markets, although the effect on wages exhibited positive pre-trends. Given the evidence of pre-trends, I also present effects on wage premia relative to trend (see Appendix A). 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 sectors used to construct ΔICE_m by an order of magnitude, to 285, relative to those currently used in the literature—[22] 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.9 shows that, in this context, inference results are similar.

The effect of trade on concentration is pervasive across informality levels. Appendix Table A.6 shows that the effect is present in both above- and below-median informality markets, with somewhat larger point estimates where informal employment is a larger share of total employment. While this confirms that the concentration effect is not confined to one segment of the informality distribution, understanding how informality interacts with wage markdowns requires the full model extension in Section 8, which incorporates self-employment, informal wage work, and preference heterogeneity into labor supply decisions.

4.3 Source of increased concentration

What drives the increase in concentration shown in Figure 5? I leverage the employer-employee linked data to compute direct worker flows between firm types within each local labor market. I classify firms into three groups: exporters (those exporting at any point between 1991 and 1994), non-exporting tradable firms, and non-tradable firms. I then estimate the effect of import competition exposure on total employment by firm type and on within-market (microregion × occupation) worker flows between each pair of firm types.

Figure 6 presents results by firm type. Import competition primarily reduced employment among non-exporting tradable firms in the most affected markets, with little detectable effect on exporters or non-tradable firms. This differential incidence is consistent with Melitz and Ottaviano (2008): under unilateral liberalization, import competition hits non-exporting tradable firms hardest, as they depend entirely on the domestic market.

Where do the displaced workers go? Appendix Figures A.6 and A.7–A.8 trace within-market worker flows between firm types.[23] They reveal that the key source of increased concentration is direct reallocation from shrinking to expanding non-exporting tradable sector firms. Appendix Figure A.6 shows that, in more affected markets, hirings into non-exporting tradable firms increased from all three firm types (Panel a), while separations from non-exporting tradable firms rose differentially towards other non-exporting tradable firms

[22] Previous papers use 20 sectors (Nível 50) or (Nível 80). See Appendix A.

[23] Appendix Figures A.9–A.11 replicate this analysis including flows into and out of the formal sector. Conclusions are unchanged: differential flows into or out of the formal sector cannot explain the increase in concentration.

(Panel b). On net, non-exporting tradables experienced a reallocation of employment towards surviving firms within the group—consistent with import competition generating a pro-competitive push among import-competing firms—while also losing workers to non-tradable firms.[24]

This direct reallocation from shrinking to expanding import-competing firms is labor market evidence of the pro-competitive effects of trade predicted by Melitz and Ottaviano (2008). Topalova and Khandelwal (2011) find that tariff reductions raised firm productivity in India's 1991 unilateral liberalization through both pro-competitive effects and reduced input costs, the latter driving most of the effect. Muendler (2004) documents similar productivity gains among Brazilian manufacturers in response to the same tariff reductions studied here. Combined with my finding that most of the wage reductions attributed to trade operate through within-firm declines in the marginal revenue product of labor, this suggests that the wage losses in more exposed regions were likely driven entirely by reduced output prices rather than productivity losses. Consistent with this interpretation, the effects of import competition on concentration, wages, and employment are very similar whether measured using output tariffs or effective rates of protection (Tables 1 and 2, column (2)), as in Dix-Carneiro and Kovak (2017).

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

[24] Appendix Figure A.7 reveals a different pattern for exporters: both hirings and separations declined, suggesting a near-complete halt in dynamism, possibly reflecting the 1994 Real Plan dollar peg dampening export demand. Appendix Figure A.8 shows non-tradable firms' dominant pattern is non-tradable to non-tradable reallocation.

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 × 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 than attempting to directly estimate ε_{zm}^{-1}—a general equilibrium object—with regression, which cannot be done for Cournot competition, as shown by BHM.[25]

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.[26] 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.).[27]

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 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.[28] This is equivalent to assuming that each worker provides one "effective monthly unit" of labor, whereas the model allows l_{zm}^{j} to be more generally pinned down by worker j's exogenous reservation earnings y^{j}.[29]

I measure w_{zmt} as the firm z's wage *premium* in market m for the month of December of year t. That is, the total compensation w_{zm}^{j} received by worker j for all labor j provided in December *conditional on worker j's characteristics*.[30] It is important here to clarify that my use of the term "wage premia" follows papers

[25] 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).

[26] 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.

[27] 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 (2010)).

[28] Workers employed as of December 31 who were hired on or before December 1. This is the standard measure of firm-level employment in RAIS (e.g., Kovak (2013); Dix-Carneiro and Kovak (2017)).

[29] 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) shows incorporating hours does not affect estimates of trade's effect on wages in later years.

[30] For each year, I estimate firm wage premia as firm × market fixed effects in a regression of worker log December earnings on firm × market fixed effects plus controls for age, education, and gender (see Appendix A). These differ from Abowd, Kramarz and

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).[31]

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

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

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). I also report results using effective rates of protection, which include 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. Appendix Figure A.12 plots this identifying variation. I estimate equations 15 and 16 clustering standard errors at the firm level, and weighting the regression by the firm's base year employment to focus on variation coming from firms where most workers were located at baseline. I then present robustness estimates to alternative clustering schemes, weighting 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 $\Delta\delta_m$:

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

Margolis (1999) firm fixed effects, which condition on worker fixed effects and may be biased under limited mobility (Bonhomme et al., 2023). Appendix Table A.11 shows that conditioning on worker fixed effects does not substantially change elasticity estimates.

[31] 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 (2021); Pham (2023); Tortarolo and Zarate (2018)). To check whether unobservable worker characteristics matter, I present results conditioning on worker fixed effects, as in Abowd, Kramarz and Margolis (1999). To check whether differential sorting might confound estimates, I restrict to the sub-sample of stayers within firm-market pairs.

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.

The two identifying assumptions are that there is a first stage (i.e., $\lambda \neq 0$), and 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 $\frac{1}{\eta}$.[32] I estimate equation 19 clustering standard errors at local labor market level, and present robustness to alternative clustering and wage measurements.

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 binned scatters of its identifying variation are plotted in Appendix Figure A.12. 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 0.556 percent (SE 0.044). This is a strongly identified first stage, with an F-statistic of 156.771. Panel B shows that the proportional effect on firms' wage premia was roughly of the same magnitude, at a 0.550 percent reduction (SE 0.024). Combined, these effects imply a within-market cross-firm inverse elasticity of substitution of 0.990 (SE 0.089).

A within-market cross-firm inverse elasticity of substitution of 0.990 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 1 percent. This is a large estimate, nearly seven times larger than BHM's corresponding estimate of 0.14 for the US,[33] suggesting that Brazilian workers substitute a lot less swiftly across firms in response to wage changes than US workers do. This rather inelastic preference parameter places an upper bound of $1/(1+0.990) \approx 50\%$ on firms' wage take-home shares. In other words, the slow change in firm choice in response to wage changes implies that in the 1990s Brazilian workers were paid at most 50 cents of every marginal dollar they generated.

This point estimate is robust to important alternative specifications, several of which are presented in the remaining columns of Table 2. A first concern is that such 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 (4) of Table 2 shows that defining local labor markets more broadly by microregion only yields a very similar within-market cross-firm inverse elasticity of substitution, of 0.969, albeit with a slightly larger standard error than the baseline estimate. I cannot reject that this is equal to my baseline estimate of 0.990. Since the latter 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 similar elasticity estimates when using effective rates of protection as opposed to import tariffs as shocks (column (2) of Table 2), with smaller first stage and reduced form effects consistent with the greater noise in effective rates of protection.[34] Column (3) of Table 2 shows that controlling for worker fixed effects

[32] 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 $\ln \xi_{zmt} = \epsilon_{zmt}/(1+\eta) = \nu_{zmt}/(1+\eta)$.

[33] BHM reports an η of 6.96, whose inverse is 0.14, based on local labor markets defined as a commuting zone × NAICS3 pairs.

[34] Effective rates of protection are output tariffs netted out of input tariffs. Smaller effects are expected due to attenuation bias.

and demographic-by-year controls—thereby netting out unobservable worker characteristics and changes in workforce composition—yields a somewhat more elastic inverse elasticity of 0.811 (SE 0.080), placing the implied upper bound on wage take-home shares at 55 cents on the dollar. These results suggest that, at least in the Brazilian context, the labor market power exerted by local firms is rather invariant to worker sorting across firms.[35] 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. Section 8 addresses this by extending the model to allow for demographic heterogeneity in worker preferences.

Finally, while $\frac{1}{\eta}$ is a preference parameter, and thus assumed stable over the period of study, Appendix Figure A.14 presents estimates of $\frac{1}{\eta}$ estimated one year at a time during the post-liberalization period. Appendix Figure A.13 plots the year by year identifying variation. Estimates are stable, though slightly increasing—meaning labor supply becomes more inelastic—consistent with the estimation sample capturing a growing share of within-firm stayers over time, who are likely more inelastic than recent switchers.[36]

6.2 Cross-market inverse elasticity of substitution

Table 3 presents my estimate of $\frac{1}{\theta}$ based on equation 19. The first stage in Panel A shows that a 1 percent increase in a market's import competition exposure reduced employment by 0.261 percent (SE 0.032), whereas Panel B shows that the proportional effect on markets' wage premia indices was roughly half as large, at a 0.120 percent reduction (SE 0.043). Combined, the first stage and reduced form produce an IV estimate of 0.459 (SE 0.190) for the difference between $\frac{1}{\theta}$ and $\frac{1}{\eta}$, which implies a cross-market inverse elasticity of substitution of 1.448 (SE 0.168) given the estimate for $\frac{1}{\eta}$ from Section 6.1.

There are three important take-aways 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 < 0.02), 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 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 1.448 for Brazil's cross-market inverse elasticity to other contexts. A cross-market inverse elasticity of substitution of 1.448 means that a market's wage premium index (i.e., the taste-adjusted wage premium) would have to increase by 1.448 percent before one percent more workers were

See Appendix A.

[35] Using average wages changes the implied upper bound to 51 cents (Appendix Table A.11, column (4)); restricting to stayers within firm-market pairs yields 55 cents (column (3)). Restricting to markets whose tradable firms are the only local producers of their sector yields similar elasticities but smaller first stage and reduced form effects (Appendix Table A.12, column (2)), suggesting sectoral agglomeration exacerbates trade's effects, consistent with Dix-Carneiro and Kovak (2017). See also Appendix Tables A.12 and A.13.

[36] Adding exiting and entering firms to the sample (coding employment and wages as zero outside their active period) yields similar estimates (Appendix Table A.12, column (4)).

attracted from other markets. While relatively inelastic, this point estimate is about two-thirds of BHM's corresponding estimate of 2.2 for the US, suggesting Brazilian workers substitute more swiftly across local markets than US workers do.[37] Overall, the main difference in substitution patterns between Brazilian vs. US workers seems to be that US workers substitute a lot more swiftly across firms within markets relative to Brazilian workers. On net, this relatively inelastic cross-market elasticity of substitution places a lower bound of $1/(1+1.448) \approx 41\%$ on wage take-home shares. That is, during the 1990s Brazilian workers were paid at least 41 cents of every marginal dollar they generated.

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.

Appendix Table A.3 shows that in the baseline year of 1991, this weighted average concentration was 0.08 on a scale that ranges from zero (infinitely tiny firms) to one (one firm). This is equivalent to saying that on average Brazilian workers were in labor markets whose equilibria were pinned down as if only $12.5 = 1/0.08$ equally-sized firms operated them. Because most workers work in larger labor markets, note that the payroll-share-weighted average concentration is much smaller, less than one third, of its 0.28 unweighted counterpart,[38] a fact that is taken into account in the country-level average wage markdown.

Combined with Section 6 estimates for $\frac{1}{\theta}$ and $\frac{1}{\eta}$, a 0.08 level of labor market concentration implies per equation 9 that Brazil's formal sector pre-liberalization average wage markdown was approximately 2, whose inverse gives an average wage take-home share of approximately 50 percent.[39] That is, Brazilian formal sector workers took home 50 cents of every dollar of marginal revenue product of labor they generated. This places Brazil's formal sector as substantially less competitive than US labor markets, where wage take-home shares are estimated at 65% for manufacturing (Yeh, Macaluso and Hershbein, 2022) and 73% for tradables by BHM,[40] 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 substantially curbs formal firms' market power, while Pham (2023) report 47% for Chinese manufacturing.[41] Across these studies, markdowns are consistently found to increase with

[37] BHM reports $\theta = 0.45$, whose inverse is 2.2, based on local labor markets defined as a commuting zone × NAICS3 pairs.

[38] The payroll-share-weighted HHI was also smaller than the median of 0.21: many local labor markets are highly concentrated, but most workers are in less concentrated ones. See Appendix A.2.

[39] 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 47 percent if evaluated at the (unweighted) average payroll Herfindahl of 0.28.

[40] $73\% \approx 1/(1+(1/0.45)*0.11+(1-0.11)*(1/6.96))$ using $\theta = 0.45$, $\eta = 6.96$, and payroll-share-weighted Herfindahl of 0.11 from BHM. See Appendix B.2.4 and B.2.6.

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

firm size.[42]

7 Effect of trade on local wage premia: Markdowns vs MRPL

What does the trade-induced increase in labor market concentration documented in Section 4 imply for average wage markdowns, marginal revenue product of labor, and wage premia? This section answers this question by combining the theoretical results from Section 2 with the estimates of workers' key elasticities of substitution from Section 6 and putting these together into an accounting exercise.

7.1 Effect decomposition

We can decompose the effect of trade on average wages via concentration into its effect on markdowns versus its effect on MRPL by applying the product rule to a simple average wage decomposition. Start by noting that the average wage in a market is the product of the average marginal revenue product of labor in that market and the market's average wage take-home share. The model allows us to estimate the latter, while wages are measured directly in the data, so the residual wage variation corresponds to the market's average MRPL. This already allows us to decompose the effect of trade on average wages into its effect on the average wage markdown and the average MRPL. But we can go one step further: because we observe all firms in the market, we can also decompose the effect on average MRPL into within-firm changes over time versus other changes, which stem from labor being reallocated across firms within markets.

To be precise, recall from equation 9 that market m's average wage in year t is given by $\bar{w}_{mt} = \mu_{mt}^{-1}\bar{r}_{mt}$, where μ_{mt}^{-1} is the market average wage take-home share and $\bar{r}_{mt}$ is the market average marginal revenue product of labor. Therefore, the effect of trade on average wages can be decomposed as:

$$\frac{d\bar{w}_{mt}}{dICE_m} = \frac{d\mu_{mt}^{-1}}{dICE_m}\bar{r}_{mt} + \frac{d\bar{r}_{mt}}{dICE_m}\mu_{mt}^{-1} = -\underbrace{\frac{d\mu_{mt}}{dICE_m}}_{\gamma_t}\frac{1}{\mu_{mt}^2}\bar{r}_{mt} + \frac{d\bar{r}_{mt}}{dICE_m}\mu_{mt}^{-1} \tag{21}$$

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 effect of trade on the average marginal revenue product

[42] Appendix Table A.14 confirms this: regressing $\ln(\mu_{zm})$ on log employment yields +0.003, robust across specifications. Amodio et al. (2025) find a similar positive relationship (+0.025 with firm FE, Table A.3; +0.079 without, Table 1). I also find a near-zero positive correlation between log markdowns and log wages. This result is different from Amodio et al. (2025)'s negative correlation, which are negative by construction, since the production function method calculates log markdowns as log estimated MRPL minus log wage. In contrast, I estimate markdowns directly from elasticities of substitution and wage bill shares. The near-zero correlation I find suggest that cross-firm wage differentials in Brazil are primarily driven by the marginal revenue product of labor, not by market power. The fact that these correlations are positive also suggests that firms with more market power pay higher wages, not lower, on average, such that these are also firms with higher MRPL.

of labor $\bar{r}_{mt}$ can be further decomposed as:

$$\frac{d\bar{r}_{mt}}{dICE_m} = \sum_{z\in\Theta_{mt}} s^e_{zmt}\frac{dr_{zmt}}{dICE_m} + \sum_{z\in\Theta_{mt}} r_{zmt}\frac{ds^e_{zmt}}{dICE_m} = \underbrace{\frac{d\left(\bar{r}_{mt}|_{s^e_{jm0}}\right)}{dICE_m}}_{\text{Within-firm effect}} + \underbrace{\frac{d\left(\bar{s}^e_{mt}|_{r_{jm0}}\right)}{dICE_m}}_{\text{Cross-firm reallocation}} \tag{22}$$

where Θ_{mt} is the set of firms operating in market m in year t, $\bar{r}_m|_{s^e_{jm0}}$ is market m's average marginal revenue product of labor at time t using firms' baseline employment shares as weights for aggregation, and $\bar{s}_{mt}|_{r_{jm0}}$ is market m's average employment share using firms' baseline marginal revenue product as weights.[43] Plugging equation 22 into equation 21 and expressing changes relative to the baseline year ($t = 0$) gives:

$$\overbrace{\frac{d\bar{w}_{mt}}{dICE_m}}^{\text{Effect on avg wage}} = \underbrace{-\frac{\gamma_t}{\mu_0^2}\times\bar{r}_0}_{\text{Via average markdown}} + \overbrace{\left[\underbrace{\frac{d\left(\bar{r}_{mt}|_{s^e_{jm0}}\right)}{dICE_m}}_{\text{Within-firm}} + \underbrace{\frac{d\left(\bar{s}^e_{mt}|_{r_{jm0}}\right)}{dICE_m}}_{\text{Cross-firm reallocation}}\right]\times\mu_0^{-1}}^{\text{Via average MRPL}} \tag{23}$$

where $\bar{r}_0$ and μ_0^{-1} are the baseline average MRPL and wage take-home share, respectively.

7.2 Estimates

Figure 7 presents my estimates of γ_t for all sample years, summarized in Table 4 as the post-reform midpoint estimate. The estimated 0.459 gap between the within-market and cross-market inverse elasticities of substitution implies that a 10% increase in import competition exposure increased the average wage markdown by 0.010 (SE of 0.004) points, an effect driven by the 0.02 point average increase in markets' payroll Herfindahls. This is equivalent to a reduction of the pre-liberalization average wage take-home share of approximately 50 cents on the dollar by 0.24 cents ($\frac{1}{2.027+0.010} - \frac{1}{2.027} \approx 0.24\%$).

This rather muted effect of increased labor market concentration on local labor markets' average wage take-home shares reflects the fact that the 0.459 gap between the inverse elasticities of substitution, while statistically significant, is still moderate in magnitude: Brazilian workers find it somewhat harder to substitute globally (across markets) than locally (within markets, across firms), but the difference is not large. This is in contrast to the 2.08 gap estimate for US local labor markets from BHM—about five times larger— indicating that US workers find it much easier to substitute within markets than across markets. If Brazilian workers had the same within- and cross-market elasticities of substitution as US workers, the average wage take-home share would have declined by 2.3 cents on the dollar—as opposed to 0.24 cents—,[44] an effect about 10 times as large.

[43] Changes in concentration do not feature into the within-firm effect because $\bar{r}_{mt}|_{s^e_{jm0}}$ holds firms' relative size constant, both when computing each firm's r_{zmt} and when weighting to obtain $\bar{r}_{mt}$.

[44] Calculated as $2.3\% \approx \left[\mu_{0,US} + (1/0.45 - 1/6.96) * 0.02\right]^{-1} - \mu^{-1}_{0,US}$, where $\mu_{0,US}$ is the implied pre-liberalization average wage markdown evaluated at Brazil's average Herfindahl-weighted labor market concentration of 0.08 but US elasticities of substitution from BHM ($\theta = 0.45, \eta = 6.96$): $\mu_{0,US} = 1 + (1/0.45) * 0.08 + (1 - 0.08) * (1/6.96)$.

Tables 5 and 6 summarize the implication of this effect to average wages per equation 23. Table 5 presents estimates of the overall effect of import competition exposure on the average wage premium and its components, showing that a 10% increase in import competition exposure reduced the average wage premium by 0.163 multiples of the minimum wage.[45] It also shows that a 10% increase in import competition exposure reduced the average MRPL by 0.361 (SE of 0.067) multiples of the minimum wage. This large negative effect is entirely driven by a within-firm MRPL reduction of 0.369 (SE of 0.095) multiples of the minimum wage, and attenuated slightly by a cross-firm employment reallocation positive average MRPL effect of 0.012 (SE of 0.002) multiples of the minimum wage. Since firm productivity rose during Brazil's liberalization (Muendler, 2004), and results are very similar using output tariffs or effective rates of protection (Tables 1 and 2, column (2)), these within-firm MRPL reductions likely reflect reduced output prices rather than productivity losses—consistent with the pro-competitive effects of import competition predicted by Melitz and Ottaviano (2008) and with productivity gains documented in India's comparable episode by Topalova and Khandelwal (2011).

Table 6 puts these effects in perspective relative to the overall effect of trade on average wages. A 10% increase in import competition exposure reduced the average wage premium by 0.184 multiples of the minimum wage, or roughly 7.40% of the pre-liberalization average of 2.48. Of this total reduction, roughly 6% is accounted for by the decline in the average wage take-home share—i.e., by increased firm labor market power—and roughly 94% is accounted for by the decline in the average marginal revenue product of labor. The within-firm MRPL reduction (−0.178 multiples of the minimum wage, or −7.17% of the baseline average premium) accounts for nearly the entire MRPL effect, with a small positive cross-firm reallocation effect (+0.006, or +0.24%).

8 Self-employment, informal wage work, and heterogeneity

While this paper focuses on the formal sector, a natural question is whether Brazil's large informal sector—where informal wage work and self-employment represent roughly 40%–50% of total employment—affects estimates of formal firms' wage markdowns derived from a model that does not explicitly incorporate informality. Appendix C extends the model on three dimensions: adding self-employment as an alternative to wage work, modeling informal wage work as a possible outcome of involuntary separation from formality (and quantifying how unemployment insurance may curb the market power that involuntary separation concerns confer on formal firms), and allowing elasticities of substitution to vary by age, education, gender, and region. I summarize the extensions and main findings here.

8.1 Extension margins

Self-employment. The extension splits each local labor market into wage work and self-employment, introducing an elasticity of substitution $\tilde{\rho}$ between the two. This embeds the relationship between self-employment and wage markdowns studied in Amodio, Medina and Morlacco (2025) into an oligopsony model where firms are heterogeneous from workers' perspective. Self-employment enters the labor supply

[45] Wage effects are reported relative to trend to account for pre-trends in wage premia, also documented in Dix-Carneiro and Kovak (2017).

curves faced by formal firms through $(L_{\bar{g}m}/L_m)^{1/\tilde{\rho}}$ in equation C.5: the larger the wage work share in a market, the more labor supplied to each formal firm, all else equal. A more elastic $\tilde{\rho}$ means self-employment is a closer substitute for wage work, curbing formal firms' market power. The extended markdown formula (equation C.23) shows how $\tilde{\rho}$ interacts with concentration: the wage work share of total market employment, s_m, mediates the contribution of $\tilde{\rho}$ to markdowns.

Informal wage work. Incorporating informal wage work poses different challenges, because it can occur both inside and outside formal firms (Ulyssea, 2018) and cannot be distinguished in the data with the same granularity as formal employment. Rather than adding informality as a separate nest, the extension models it as a possible outcome of involuntary separation from formality. Workers seeking formal jobs consider both wages and the probability of being fired and, until finding another formal job, working in the local informal sector. The relevant wage for attracting labor supply to firm z becomes the *expected* wage $\bar{w}_{z\bar{g}m} = p_{z\bar{g}m} w^o_m + (1 - p_{z\bar{g}m}) w_{z\bar{g}m}$, where $p_{z\bar{g}m}$ is workers' prior about involuntary separation probability and w^o_m are expected local informal wages. This preserves the aggregation properties of the discrete choice problem while allowing informal sector conditions to influence formal sector labor supply.

Preference heterogeneity. The extension allows elasticities of substitution to vary by age, education, gender, and region. Labor markets are partitioned into demographic cells—defined by gender (2), education (3), and age (3) groups—with cell-specific or group-specific parameters. This contributes the first labor market oligopsony framework inclusive of self-employment and informal wage work to the literature on demographic wage differentials, where evidence linking wage markdowns to gender wage gaps has begun to emerge (Sharma, 2023; Hoang, Mitra and Pham, 2024).

Bias formula. A key result is the decomposition of the within-market cross-firm inverse elasticity into two components (equation C.17): $1/\tilde{\eta} = 1/\eta + \Omega$, where $1/\eta$ is the labor supply response to formal wage changes (as estimated in Section 5), and Ω captures the additional response to changes in expected formal-informal wage gaps. If $\Omega > 0$, workers supply *more* labor to firms with higher expected wage gaps relative to the informal sector—preferring formal employment despite potentially higher informal pay—so the original model *understates* firms' market power. If $\Omega < 0$, the informal sector serves as an attractive outside option that curbs formal firms' power. In either case, Ω directly measures how much the within-market cross-firm elasticity would change if informal wage work conditions were incorporated, and its sign reveals whether informal wage work amplifies or mitigates formal firms' market power.

Unemployment insurance. Since involuntary separation concerns may amplify formal firms' market power, the extension also quantifies whether unemployment insurance helps curb it by cushioning expected post-separation earnings. Cross-country evidence suggests that markdowns are increasing in self-employment shares in countries with unemployment insurance systems (Amodio et al., 2025), and micro evidence from Brazil shows that workers have high willingness to pay for unemployment insurance (Felix et al., 2026), a potential microfoundation for formal firms' market power. In Brazil, formally separated workers were entitled to roughly four months of salary-based benefits during the 1990s (see Appendix C for details). The extension re-estimates the bias Ω with and without incorporating these benefits (Appendix Table C.5).

8.2 Empirical strategy

Measurement. Estimation combines RAIS with Brazil's 1991 and 2000 censuses, which provide informal wage work and self-employment earnings at the microregion × demographic cell level. Appendix C provides all estimating equations and data mappings. The within-market cross-firm elasticity $1/\tilde{\eta}$ is recovered by combining the IV estimate for $1/\eta$ with an estimate of the bias Ω from ignoring informal wage work. Ω is estimated using within-firm cross-demographic-cell variation in expected formal-informal wage gaps, where firm-level separation probabilities are directly observed in RAIS. The wage work–self-employment elasticity $1/\tilde{\rho}$ and cross-market elasticity $1/\tilde{\theta}$ are estimated jointly using census data at the microregion × demographic cell level.

Trade shocks. Identification differs from the baseline in important ways. While $1/\eta$ uses firm-level tariff variation as in Section 5, the elasticities $1/\tilde{\rho}$ and $1/\tilde{\theta}$ are identified using cell-specific Regional Tariff Reductions (RTR) from Dix-Carneiro and Kovak (2017)—shift-share measures of import competition at the microregion level whose shares include *all* workers, formally or informally employed, in tradable industries, making them well-suited for census-based estimation. The within-market cross-sector variation identifying $\tilde{\rho}$ comes from demographic cells with their own wage work and self-employment sectors and differential trade exposure; cross-market variation identifying $\tilde{\theta}$ comes from cross-microregion RTR differences. The bias term Ω is identified from the interaction of firm-level tariff reductions and demographic-specific RTRs, capturing how trade shocks jointly affect separation probabilities and informal sector conditions.

8.3 Findings

Elasticities of substitution. I find that the bias from ignoring informal wage work is positive but small ($\Omega \approx 0.033$; Appendix Table C.4), implying that the original model's $1/\eta$ estimates are affected by only about 3%. Incorporating unemployment insurance benefits into expected post-separation earnings reduces Ω from 0.033 to 0.022 (Appendix Table C.5), indicating that unemployment insurance partially curbs the market power that involuntary separation concerns confer on formal firms, consistent with cross-country evidence from Amodio et al. (2025). Self-employment substantially curbs formal firms' market power ($1/\tilde{\rho} \approx 0.47$; Appendix Table C.6), consistent with Amodio, Medina and Morlacco (2025) for Peru and with Amodio et al. (2025) using markdowns estimated with the production function approach. This curbing is very heterogeneous by demographics: women and less educated workers substitute more elastically between formal wage work and self-employment ($1/\tilde{\rho}$), making self-employment a stronger outside option that limits formal firms' wage setting for these groups. There is little demographic heterogeneity in within-market cross-firm substitution, even when taking into account informal wage work ($1/\tilde{\eta}$). There is substantial regional heterogeneity in the within-market cross-firm elasticity of substitution, however: Southeast estimates (1.029) are nearly twice those for the Northeast (0.458) (Appendix Table C.2), yielding more market power in the former.[46]

Wage markdowns. On average, Brazilian formal sector workers were paid 51 cents of their marginal revenue product once all margins are accounted for, compared to 50 cents in the baseline model (Table 8; Figures 9

[46] The muted cross-firm substitution heterogeneity may reflect wage compression from union bargaining; Appendix Figure A.16 shows a mild positive correlation between union presence and wages.

and 8). This average masks substantial heterogeneity: take-home shares range from 46 to 73 cents depending on local labor market conditions outside the formal sector and workforce composition. Table 7 presents summary statistics for the distribution of wage take-home shares across microregions, separately by year and demographic groups. The dispersion is driven by regional heterogeneity in the within-market cross-firm elasticity $1/\eta$: most microregions cluster around 50 cents on the dollar, while those in the North and Northeast—where cross-firm elasticities are almost half those of the rest of the country—feature take-home shares reaching up to 70 cents on the dollar; the dispersion across occupations is centered between 50 and 55 cents (Appendix Figure C.9). These estimates combine the extended model's elasticities—including the wage work–self-employment elasticity ($1/\tilde{\rho}$) and the informal wage work bias (Ω)—with concentration measures that account for expected informal wages by demographic type and the share of employment in wage work versus self-employment. Most of the demographic heterogeneity in markdowns is driven by the elasticity of substitution between wage work and self-employment, $1/\tilde{\rho}$ (Appendix Table C.8): women substitute more elastically than men ($1/\tilde{\rho}$ of 0.390 vs. 0.660), while young workers (0.850) and the highly educated (0.935) are the least elastic, suggesting self-employment is a weaker outside option for these groups. These differences translate into modest variation in wage take-home shares (Table 7): median take-home shares in 1991 range from 0.480 for tertiary-educated workers to 0.509 for primary-educated workers, from 0.487 for young workers to 0.541 for older workers, and are nearly identical across gender (0.498 for men, 0.500 for women). Regional heterogeneity is considerably larger: workers in the Southeast—where formal labor markets are thicker and self-employment less prevalent—face higher markdowns than those in the Northeast, driven by the substantially more inelastic within-market cross-firm elasticity in the Southeast.

Markdowns and concentration. In the baseline model, equation 9 implies a linear relationship between formal sector concentration and the average wage markdown, with slope $(1/\theta - 1/\eta)$. Incorporating informal wage work, self-employment, and demographic preference heterogeneity introduces non-linearities: the extended markdown formula (equation C.23) depends not only on formal sector concentration but also on its interaction with the local wage work share s_m and on region- and demographic-specific elasticities. Appendix Figures C.9, C.11, and C.12 plot the resulting distributions and the relationship between the formal sector's Herfindahl-Hirschman Index (HHI) and estimated take-home shares. The relationship is non-linear, likely driven by cross-regional differences in elasticities: take-home shares increase with concentration for markets with many similarly sized firms (HHI below 0.2), but past that threshold, higher concentration typically reduces take-home shares.

Effect of trade on markdowns. Combining these estimates with reduced-form effects of RTR on concentration and the wage work share, two channels operate in opposite directions: a *within-market net substitution channel* reduces markdowns by shifting weight toward the more elastic self-employment margin, while a *cross-market substitution channel* increases markdowns by shifting weight toward the least elastic cross-market margin. At the country average, these channels largely offset ($\Delta\mu \approx 0$ percentage points). In low-informality regions the take-home share is essentially unchanged (+0.04 pp), while in high-informality regions it falls modestly (−0.54 pp). Across regions, the North and Northeast see take-home shares rise by +1.2 pp while the Southeast, South, and Center-West see essentially no change (−0.1 pp). Across demographics, differences are small: women and older workers see take-home shares fall by −0.1 pp; tertiary-educated and young workers see slight increases (+0.1 to +0.2 pp) (Table 8; Appendix Section C.3.3). Appendix Table A.6 confirms that trade affects concentration in both high- and low-informality markets. The key demographic heterogeneity takeaway is that the within-market net substitution and cross-market substitution effects nearly cancel for all groups, despite substantial demographic heterogeneity in markdown levels, resulting in similar and muted effects of trade's effect on markdowns across groups at the country level.

Overall, these findings add richness to the baseline analysis, but the qualitative conclusions are largely unchanged: formal sector firms in Brazil command high wage markdowns, and the sufficient statistics framework in Section 2 captures the essential forces. This alignment arises because most formal employment is concentrated in predominantly formal microregions, and most of these microregions are in the Southeast, where the within-market cross-firm elasticity is most inelastic ($1/\eta = 1.029$; Appendix Table C.2) and self-employment and informal wage work represent a small share of total employment, so the additional margins have a modest effect on estimated markdowns. Appendix Figure A.16 further shows that unions are unlikely to be a primary driver of the wage dynamics documented here. The full model, estimation strategy, and detailed results are in Appendix C.

9 Conclusion

This paper is an empirical study of the relationship between trade, local labor market concentration, and wages in the context of Brazil's 1990s unilateral trade liberalization. I showed that the effect of trade on wage markdowns across local labor markets can be quantified by two sufficient statistics: the effect of trade on local labor market concentration, and the gap between workers' cross-market vs. within-market cross-firm inverse elasticities of substitution. I then leveraged Brazil's rich employer-employee linked data and the variation in import tariff reductions to estimate these sufficient statistics.

The findings in Sections 4, 6, and 7 can be summarized into three take-aways: (i) In the 1990s, formal sector firms in Brazil commanded substantial labor market power, with workers taking home only 50 cents of every marginal dollar generated—primarily driven by workers' very inelastic within-market cross-firm substitution, nearly seven times as inelastic as in the US; (ii) Opening to trade increased that labor market power a bit further as it raised local labor market concentration—by enough to offset wage gains from cross-firm reallocation—but (iii) on net, increased wage markdowns account for only about 6% of the overall negative effect of trade on average wages (Table 6); the remaining 94% was driven by within-firm reductions in the marginal revenue product of labor. Since firm productivity rose during Brazil's liberalization (Muendler, 2004)—as in India's comparable episode (Topalova and Khandelwal, 2011)—and the effects of import competition are nearly identical using output tariffs or effective rates of protection, these MRPL reductions likely reflect reduced output prices, consistent with the pro-competitive channel in Melitz and Ottaviano (2008).

Section 8 extends these results to incorporate self-employment, informal wage work, and preference heterogeneity by demographics and regions. These additional margins bring average wage take-home shares to 51 cents on the dollar—close to the baseline—but reveal substantial heterogeneity: take-home shares range from 46 to 73 cents depending on local labor market conditions outside the formal sector and workforce composition (Table 7). Self-employment curbs formal firms' power by providing an attractive outside option, while the threat of involuntary separation into informal wage work slightly increases it—though unemployment insurance partially offsets this effect by cushioning expected post-separation earnings. The effect of trade on markdowns through the extended model is near zero at the country average, with modest heterogeneity across regions and demographics.

Future research should further disentangle the components of the marginal revenue product of labor—price markups, productivity, and production technology—to confirm and refine this output-price interpretation across contexts.

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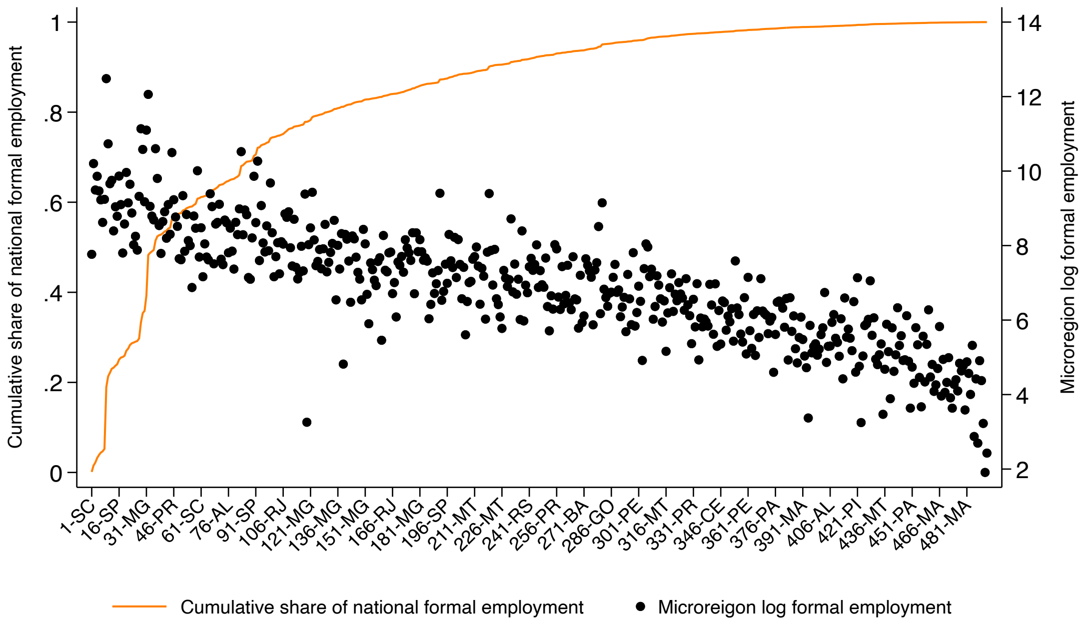
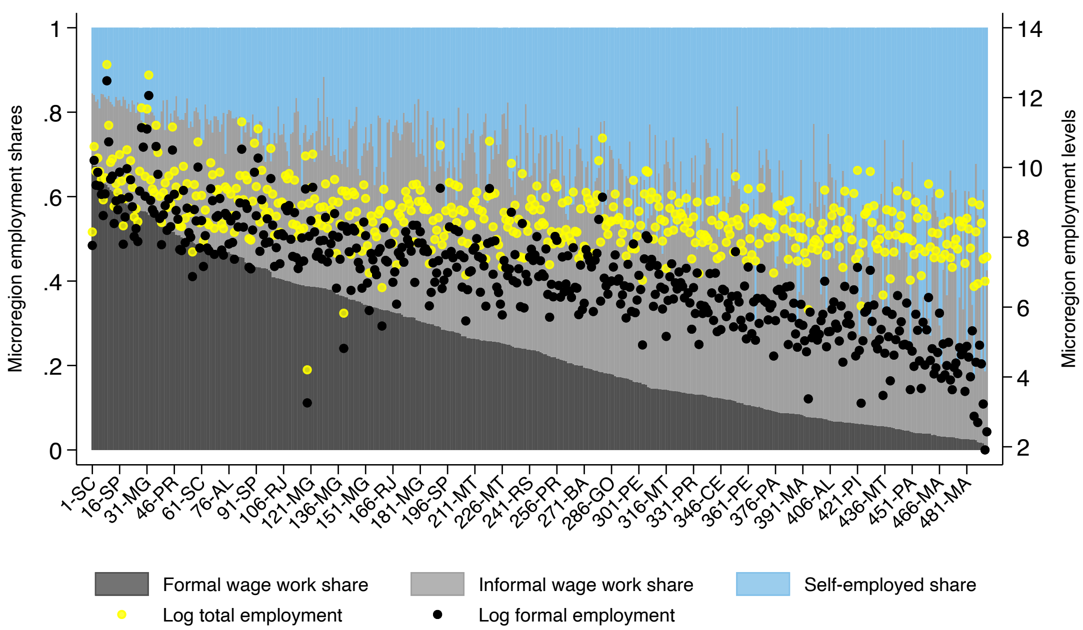
Figure 1: 1991 distribution of formal employment across microregions

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

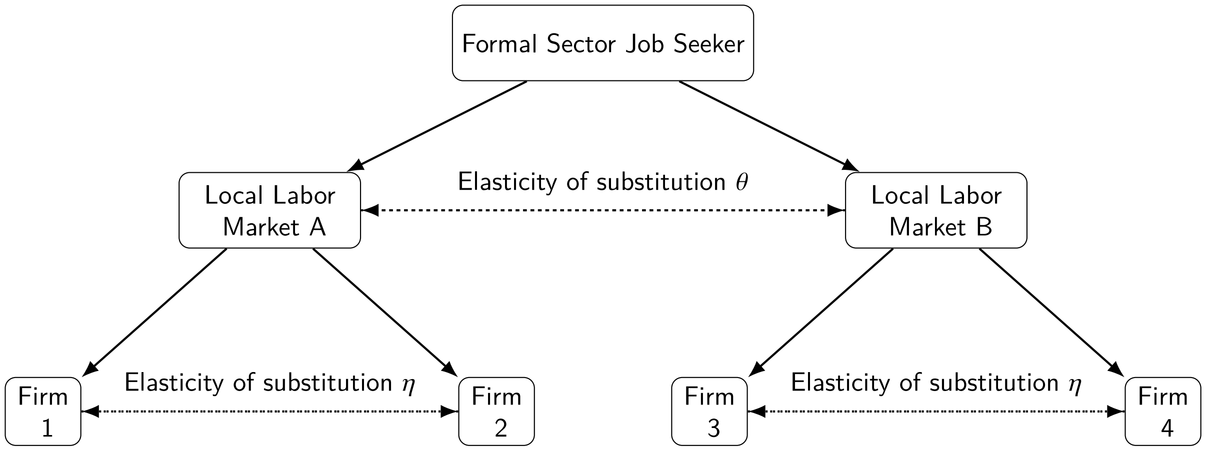


(b) Formal wage work, informal wage work, and self-employment



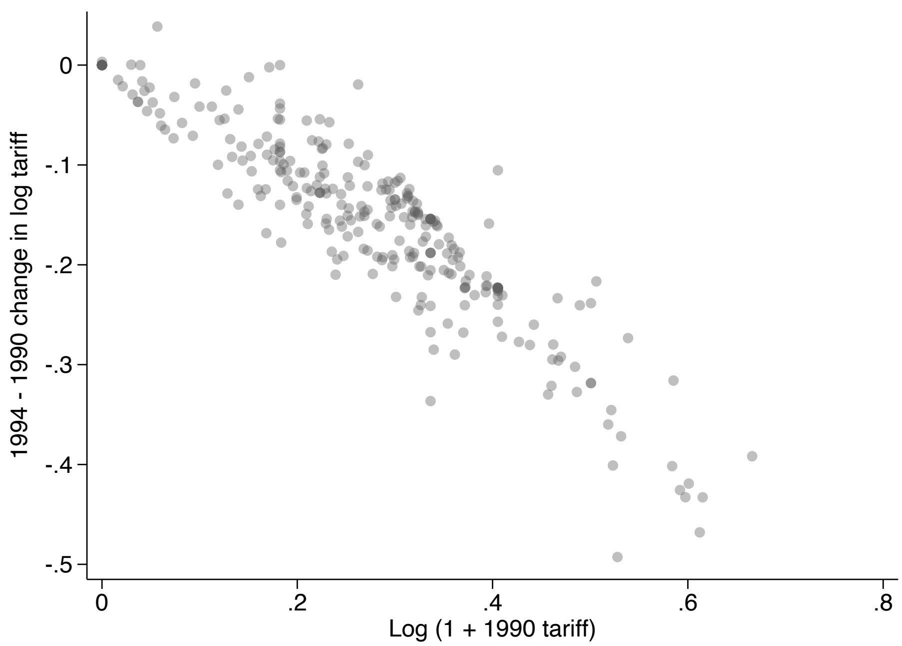
Notes: This figure plots the distribution of formal employment across 486 microregions per the 1991 Population Census. Panel (a) shows the cumulative distribution of formal employment (red line, left y-axis) and log formal employment (black scatter, right y-axis). Labels in the x-axis correspond to different microregions. Panel (a) shows the distribution of formal wage work (black bins, left y-axis), informal wage work (gray bins, left y-axis), self-employment (blue bins, left y-axis), log total employment (yellow scatter, right y-axis), and log formal employment (black scatter, right y-axis). The letters indicate the microregion's state, and the number indicates the microregion's rank in the national sorting (e.g., the country's most formal microregion, labeled 1-SC, is a microregion in the state of Santa Catarina).

Figure 2: Discrete choice microfoundation of nested CES preferences over formal sector jobs



Notes: This figure displays a diagram of workers' labor supply decision, microfounded in discrete choice, as presented in Section 2. In the empirical exercise, a local labor market is defined as a microregion x 2-digit occupation pair based on firm-to-firm flow data in Figures A.2 through A.4. Appendix C extends this setup in three ways: 1) Adds a within-market cross-sector elasticity of substitution into self-employment; 2) Adds informal wage work as a potential outcome of involuntary separation from a formal sector job; and 3) Adds heterogeneity by gender, education, and age groups to all elasticities of substitution, and heterogeneity by regions to the within-market cross-firm elasticity. These extra margins are estimated by combining employer-employee linked data with census data and group-specific trade shocks. See Appendix C.

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

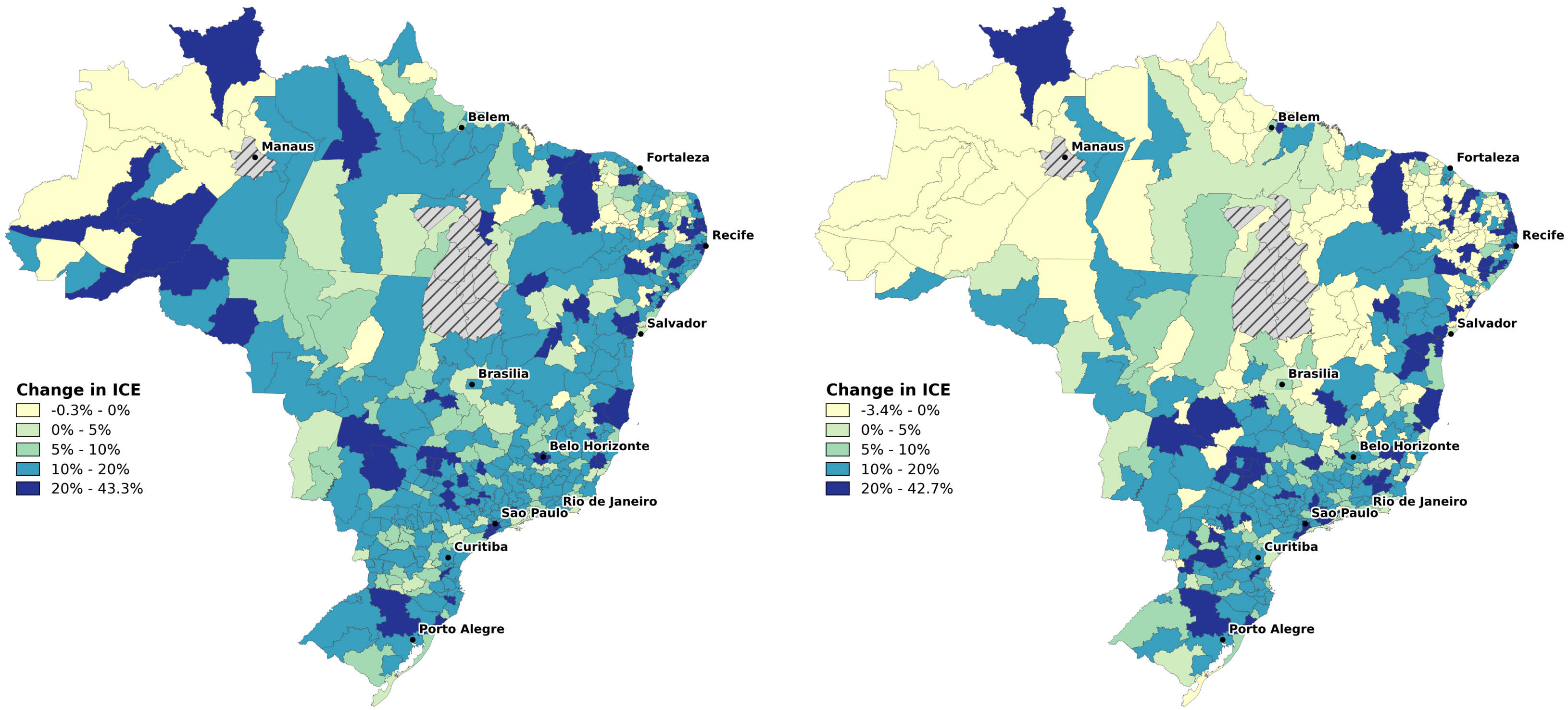


Note: This figure plots import tariff reductions from Brazil's 1990–1994 import tariff reform across RAIS' 5-digit sector variable "CNAE-95", including 285 tradable sectors and 280 non-tradable sectors. Sector-level tariffs are simple averages of product-level tariffs for the products produced in each sector, and are constructed by mapping 6-digit product-level tariffs from UNCTAD TRAINS to CNAE-95 using Brazil's product-to-sector mappings from IBGE. See Section 3.3 for details.

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

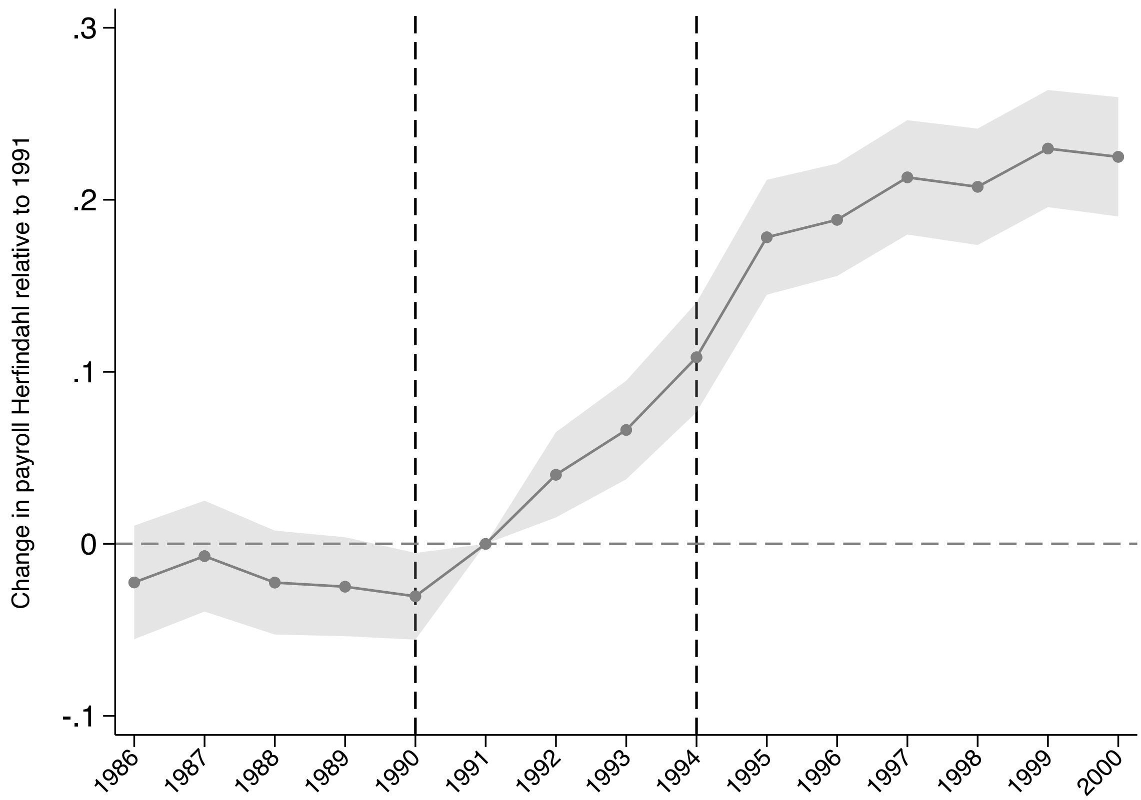
(a) Office administration workers

(b) Managers and supervisors of industrial workers



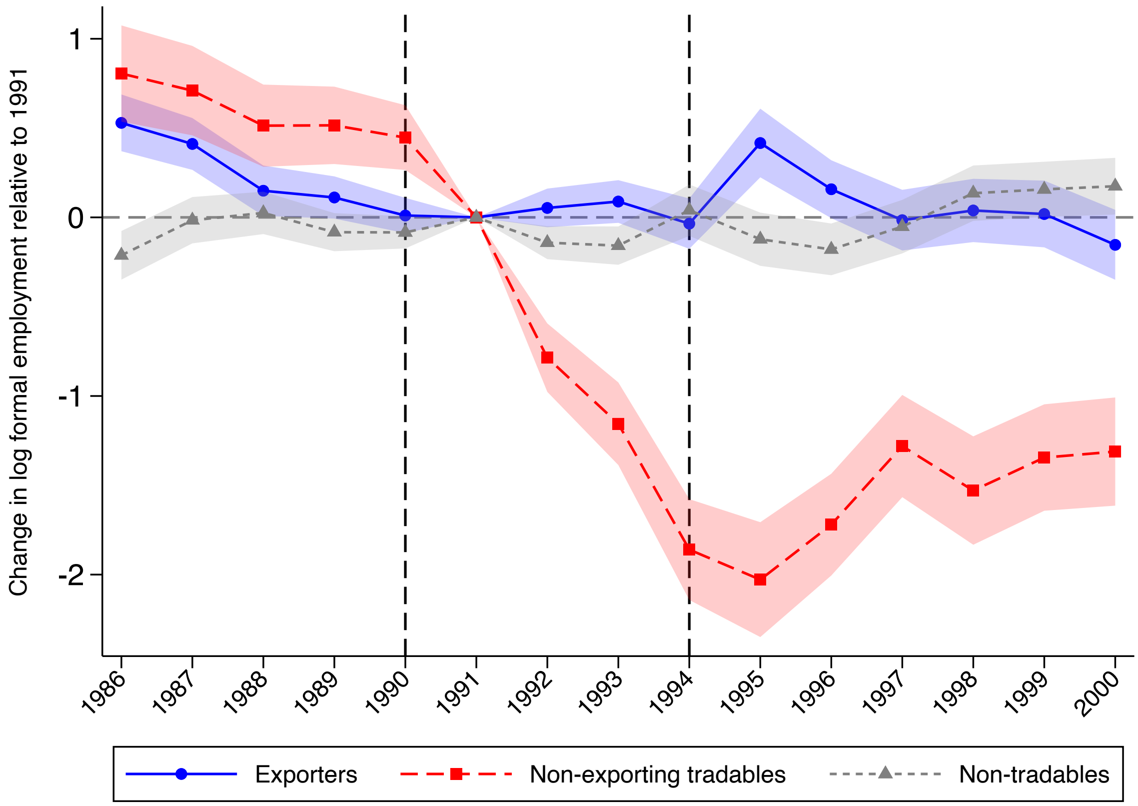
Note: This figure displays variation in ΔICE_m: the change in import competition exposure across local labor markets for two occupation groups. Produced using QGIS with microregion boundaries from Dix-Carneiro and Kovak (2017).

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



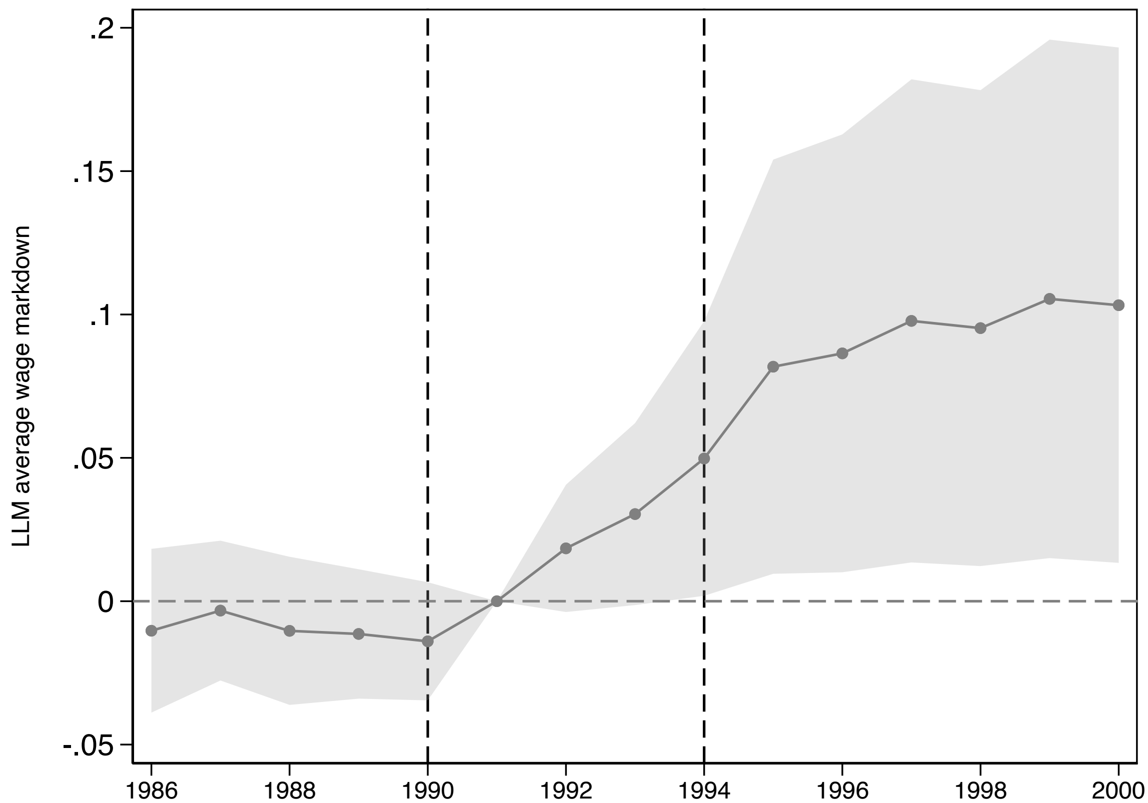
Notes: This figure plots regression coefficients ζ_k on regressor ΔICE_m from equation 12, where the outcome is the change in payroll Herfindahl relative to 1991. Since ΔICE_m is a weighted average log change in import tariffs, note that this is a units-on-logs regression, such that a 10% increase in import competition exposure changed the outcome by $(\zeta_k/100) \times 10$ units. Shaded areas report the 95% confidence interval based on clustered standard errors at the local labor market level.

Figure 6: Effect of import competition exposure on employment of non-exporting tradable sector firms, exporting firms, and non-tradable sector firms



Note: This figure plots coefficients of three regressions about the cumulative effect of the change in import competition exposure: on changes in log employment of exporters; on changes in log employment of non-exporting tradables; and on changes in log employment of non-tradables. To keep the composition of firms fixed, firms exporting any time between 1991 and 1994 are tagged as exporters for the whole period. Each point is a ζ_k coefficient from equation 12. Dotted lines indicate the beginning and end of the tariff reductions reform. So that all differences reflect a change from a future year to a past year, for the pre-liberalization years the outcome is the 1991 log employment minus each respective year's log employment, whereas for the post-reform years the outcome is each respective year's log employment minus the 1991 log employment. All regressions are weighted by 1991 employment. Standard errors are two-way clustered by microregion and occupation group.

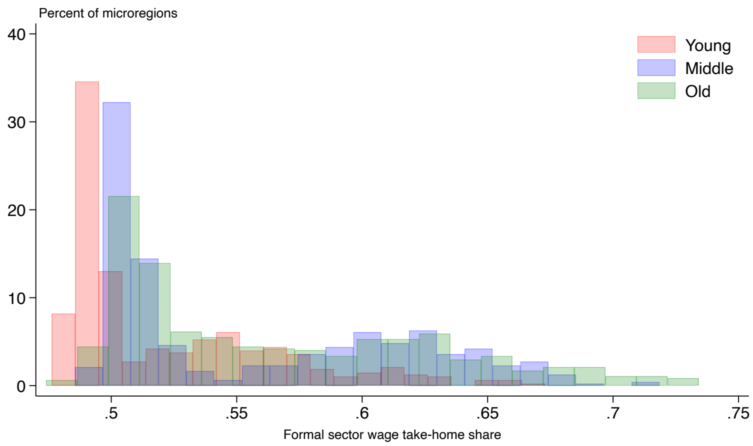
Figure 7: Effect of import competition on local labor market average wage markdowns



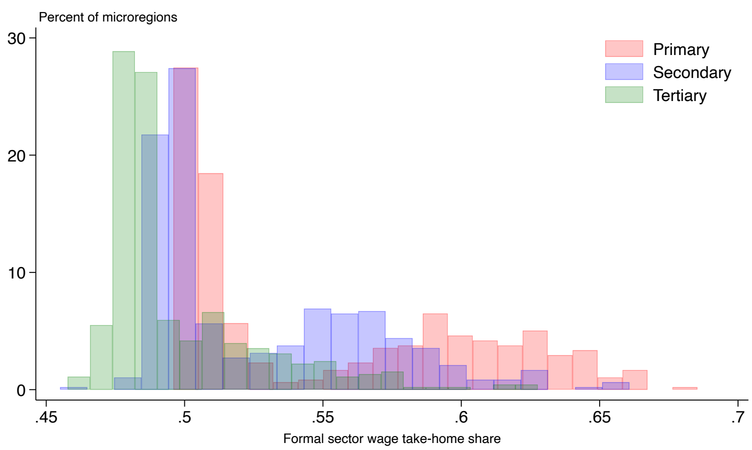
Notes: This figure plots γ_t, the effect of import competition on local labor markets' average wage markdown share at year t, derived in equation 10. The two components of γ_k are $\left(\frac{1}{\theta} - \frac{1}{\eta}\right)$, whose estimates are presented in Table 3, and the β_t coefficients that estimate the effect of import competition on labor market concentration, presented in Figure 5. Standard errors are estimated assuming β_t and $\left(\frac{1}{\theta} - \frac{1}{\eta}\right)$ are independent (see Appendix B for details).

Figure 8: Extended model: Formal sector wage take-home share heterogeneity by demographics

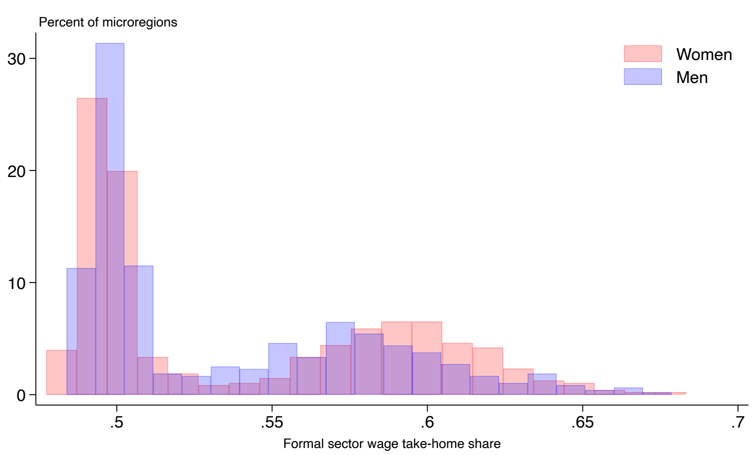
Panel A: By age



Panel B: By education



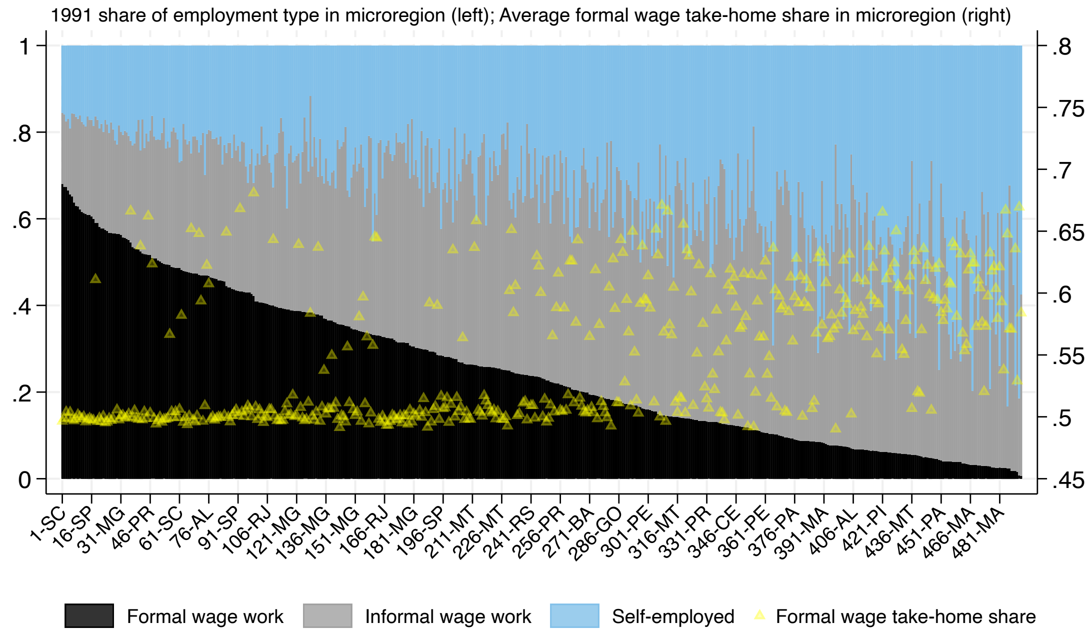
Panel C: By gender



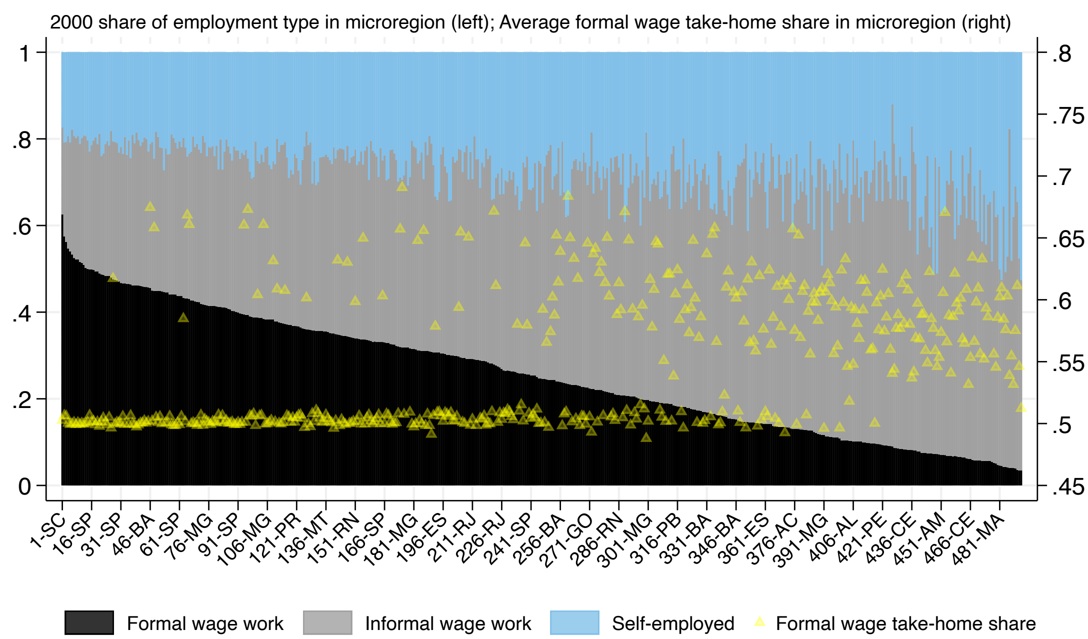
Notes: This figure shows the distribution of average wage take-home shares based on the extended model (Appendix C) across microregions, separately by demographic groups and conditional on year fixed effects. The extended model features heterogeneous (by demographics) substitution to self-employment ($1/\tilde{\rho}$) and formal sector labor supply sensitivity to involuntary separation into informal wage work (Ω). Average wage take-home shares are calculated at the local labor market level (microregion x occupation) using region-specific estimates of $1/\tilde{\eta}$, demographic-specific estimates of $1/\tilde{\rho}$, and the pooled estimate for $1/\tilde{\theta}$.

Figure 9: Extended model: Formal sector wage take-home share heterogeneity by microregion

Panel A: 1991



Panel B: 2000



Notes: This figure plots the distribution of formal wage work (black bins, left y-axis), informal wage work (gray bins, left y-axis), self-employment (blue bins, left y-axis), and estimated average wage take-home shares per the extended model in Appendix C across 486 microregions (yellow scatter, right y-axis). Labels in the x-axis correspond to different microregions. The first two letters indicate the microregion's state, and the number indicates the microregion's rank in the national sorting (e.g., the country's most formal microregion, labeled 1-SC, is a microregion in the state of Santa Catarina).

Table 1: Effect of import competition exposure on local labor market concentration

	Main specification	Alternative shocks, boundaries, sample: ICE constructed with effective rates of protection	Alternative shocks, boundaries, sample: Local labor market is microregion	Alternative shocks, boundaries, sample: Drop markets where $\Delta ICE = 0$	Effect per 10% increase in ICE
	(1)	(2)	(3)	(4)	(5)
Δ Payroll Herfindahl (based on wage premium)	0.213	0.119	0.102	0.036	0.021
	(0.017)	(0.011)	(0.046)	(0.018)	(0.002)
Mean (unweighted), 1991	0.287	0.287	0.038	0.248	0.287
Mean (weighted), 1991	0.079	0.079	0.006	0.076	0.079
Δ Payroll Herfindahl	0.213	0.121	0.110	0.040	0.021
	(0.017)	(0.012)	(0.064)	(0.019)	(0.002)
Mean (unweighted), 1991	0.292	0.292	0.113	0.253	0.292
Mean (weighted), 1991	0.082	0.082	0.026	0.080	0.082
Δ Employment Herfindahl	0.247	0.141	0.058	0.030	0.025
	(0.016)	(0.011)	(0.056)	(0.017)	(0.002)
Mean (unweighted), 1991	0.239	0.239	0.071	0.200	0.239
Mean (weighted), 1991	0.056	0.056	0.011	0.054	0.056
Observations	289,680	289,680	7,124	251,220	289,680
Local labor markets	19,759	19,759	475	16,748	19,759

Notes: Standard errors in parentheses, clustered at the local labor market level. Each cell reports the 1997 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) presents the main specification. Column (2) constructs ICE using effective rates of protection instead of import tariffs. Column (3) defines local labor markets as microregions only. Column (4) drops markets with $\Delta ICE = 0$. Column (5) reports the effect per 10% increase in ICE. HHI is on a 0–1 scale: the coefficient per unit of ICE is multiplied by 10 and divided by 100 to obtain the effect on the 0–1 HHI scale per 10% ICE increase (i.e., $\zeta_{1997}/10$). Weighted means weight by total formal employment in the mmc×cbo942d pair.

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

	Main specification	Shock is effective rate of protection	Premia has worker FE and demo × year FEs	Local labor market is microregion
	(1)	(2)	(3)	(4)
	Panel A: First stage			
Δ Firm log employment in LLM	-0.556	-0.359	-0.604	-0.417
	(0.044)	(0.035)	(0.053)	(0.037)
First stage F	156.771	106.426	130.168	124.666
	Panel B: Reduced form			
Δ Firm wage premium in LLM	-0.550	-0.354	-0.489	-0.404
	(0.024)	(0.019)	(0.030)	(0.017)
	Panel C: 2SLS			
Labor supply within-market cross-firm	0.990	0.984	0.811	0.969
inverse elasticity of substitution	(0.089)	(0.109)	(0.080)	(0.092)
Implied upper bound on wage take-home share	50%	50%	55%	51%
Local labor market (LLM) FE	Yes	Yes	Yes	Yes
Observations	855,104	852,702	463,138	440,966
Firms	344,534	344,027	213,704	420,246
Local labor markets	15,730	15,679	13,627	474

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 mmc × occupation, using December wage premium conditional on observables. Column (2): uses effective rate of protection as the instrument instead of TRAINS import tariff. Column (3): wage premium conditional on worker fixed effects and demographic-by-year controls. Column (4): defines LLM as microregion only (no occupation dimension).

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

	Δ Import Competition Exposure
	(1)
Panel A: First stage	
Δ LLM employment index	-0.261
	(0.032)
First stage F	66.253
Panel B: Reduced form	
Δ LLM wage premium index	-0.120
	(0.043)
Panel C: 2SLS	
$1/\theta - 1/\eta$	0.459
	(0.190)
Panel D: Cross-market inverse elasticity of substitution	
$1/\theta$	1.448
	(0.168)
Panel E: OLS	
Regression of Δ LLM employment index on Δ LLM wage premium index	-0.390
	(0.015)
Implied lower bound on wage take-home share	0.408
Observations (Local labor markets)	15,730

Notes: This table presents first stage, reduced form, and two-stage least squares estimates of $\frac{1}{\theta} - \frac{1}{\eta}$, and implied $\frac{1}{\theta}$, based on equations 19 and 20. Panel E reports the OLS estimate of a regression of the market employment index on the market wage premium index. Implied lower bound on wage take-home share is calculated as $\left(1 + \frac{1}{\theta}\right)^{-1}$ per equation 9 under the limiting assumption that each local labor market is composed of one firm (i.e. $HHI_m = 1$ for all m). Standard errors shown in parentheses are clustered at the local labor market level.

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	0.0098 (0.0043)
β Effect of Δ Import Competition Exposure on payroll Herfindahl	0.021 (0.003)
$\frac{1}{\theta} - \frac{1}{\eta}$ Difference between key inverse elasticities of labor supply	0.459 (0.190)
Effect of Δ Import Competition Exposure on market average wage take-home share	-0.0022 (0.0010)
Local labor markets	19,759

Notes: This table presents estimates of γ_{1997} per equation 9, listing its two components: β_{1997} taken from Table 1, and $(1/\theta - 1/\eta)$ from Table 3. Standard errors are estimated assuming ζ_{1997} and $(1/\theta - 1/\eta)$ are independent (see Appendix B for details).

Table 5: Effect of import competition on the average wage and its sub-components

	Effect per 10% increase in ICE (1)
Δ Average wage premium	-0.163 (0.030)
Δ Average wage premium take-home share	-0.0023 (0.0010)
Δ Average marginal revenue product of labor	-0.361 (0.067)
Δ Within-firm	-0.369 (0.095)
Δ Cross-firm	0.012 (0.002)
Observations	243,750
Local labor markets	16,250

Notes: This table presents estimates of $\tilde{\zeta}_{1997}$, the de-trended specification coefficient equivalent to ζ_{1997} from equation 12, separately estimated for the change in average wage premium, the change in average marginal revenue product of labor, and its subcomponents. The coefficient for the change in average wage premium wage take-home share is the same as in Table 4. Column (1) presents regression estimates, whereas column (2) presents the effect per 10% increase in import competition exposure to facilitate interpretation.

Table 6: Effect of import competition on local average wages: Accounting

	Pre-reform level	Effect of 10% increase in ICE	Change in multiples of min wage per equation 23	Percent change from baseline average wage premium	Effect as percent of total effect on average wage premium
	(1)	(2)	(3)	(4)	(5)
Average wage premium	2.48	-0.163	-0.184	- 7.40%	100%
Average wage take-home share	0.48	-0.0023	-0.012	- 0.47%	6%
Average marginal revenue product of labor	5.15	-0.361	-0.172	- 6.93%	94%
Δ Within-firm	–	-0.369	-0.178	- 7.17%	–
Δ Cross-firm	–	0.012	0.006	+ 0.24%	–

Notes: Column (1) displays the 1991 unweighted average December wage premium across local labor markets from Table A.2, the corresponding unweighted average wage take-home share across these markets, and the implied unweighted average MRPL, computed as their ratio. Note that the unweighted average wage take-home share is slightly different from the country-level average take-home share of 50 cents/dollar, but it is the correct baseline level to align with the effects in column (2), which are based on unweighted market-level regressions. Column (2) repeats the point estimates from Table 5. Column (3) reports the effect of a 10% increase in ICE on multiples of the minimum wage calculated as the sum of the scaled effects on each sub-component, per equation 23, and thus differs slightly from the regression-based estimate column (4), but allows for the accounting decomposition in columns (5)-(6).

Table 7: Extended Model: Distribution of wage take-home shares across microregions

	Min	p10	p25	p50	p75	p90	Max
Panel A.1: All (1991)							
Wage take-home share μ	0.483	0.486	0.489	0.504	0.587	0.616	0.658
N (microregions)				478			
Panel A.2: By age (1991)							
Young (18–29)	0.473	0.479	0.481	0.487	0.542	0.570	0.649
Middle (30–49)	0.483	0.489	0.494	0.513	0.601	0.629	0.718
Old (50–64)	0.474	0.494	0.503	0.541	0.608	0.647	0.731
Panel A.3: By education (1991)							
Primary	0.485	0.490	0.494	0.509	0.588	0.616	0.661
Secondary	0.460	0.483	0.486	0.498	0.552	0.575	0.641
Tertiary	0.462	0.473	0.476	0.480	0.508	0.535	0.610
Panel A.4: By gender (1991)							
Men	0.480	0.484	0.486	0.498	0.567	0.593	0.655
Women	0.475	0.483	0.487	0.500	0.581	0.604	0.670
Panel B.1: All (2000)							
Wage take-home share μ	0.484	0.487	0.490	0.498	0.584	0.609	0.665
N (microregions)				478			
Panel B.2: By age (2000)							
Young (18–29)	0.470	0.479	0.481	0.486	0.558	0.587	0.657
Middle (30–49)	0.485	0.490	0.494	0.506	0.591	0.615	0.666
Old (50–64)	0.488	0.497	0.505	0.523	0.586	0.616	0.668
Panel B.3: By education (2000)							
Primary	0.486	0.489	0.493	0.503	0.585	0.606	0.658
Secondary	0.469	0.485	0.487	0.494	0.562	0.587	0.655
Tertiary	0.459	0.474	0.476	0.480	0.516	0.543	0.622
Panel B.4: By gender (2000)							
Men	0.482	0.486	0.488	0.495	0.572	0.599	0.661
Women	0.475	0.484	0.486	0.494	0.567	0.587	0.647

Notes: This table reports percentiles of the distribution of the wage take-home share μ across microregions, computed from the 1991 and 2000 Brazilian population censuses. Each panel shows the distribution for all workers and by demographic subgroups (age, education, gender). N reports the number of microregions. See Section A for details on sample construction.

Table 8: Extended model: Effect of 10% increase in Regional Tariff Reductions (RTR) on wage markdowns

	Average	Informality		Gender		Education			Age			Region	
	effect	High	Low	Men	Women	Primary	Second.	Tertiary	Young	Middle	Old	N+NE	SE+S+CW
Panel A: Parameter Estimates													
$1/\eta$ (within-market inverse elast.)	0.990	0.330	0.779	0.990	0.990	0.990	0.990	0.990	0.990	0.990	0.990	0.385	0.819
	(0.181)												
Ω (separation bias)	0.033	0.033	0.033	0.033	0.035	0.033	0.033	0.033	0.034	0.035	0.033	0.033	0.033
	(0.013)												
$1/\tilde{\eta} = 1/\eta + \Omega$	1.023	0.701	0.991	1.023	1.025	1.023	1.023	1.023	1.024	1.025	1.023	0.484	1.057
$1/\tilde{\rho}$ (wage work vs self-emp)	0.482	0.482	0.482	0.660	0.390	0.482	0.460	0.935	0.850	0.355	0.303	0.482	0.482
	(0.059)												
$1/\tilde{\theta}$ (cross-market)	1.191	1.191	1.191	1.191	1.191	1.191	1.191	1.191	1.191	1.191	1.191	1.191	1.191
	(0.619)												
Panel B: Market Characteristics (1991, wagebill-weighted)													
Average HHI	0.167	0.425	0.156	0.167	0.167	0.167	0.167	0.167	0.167	0.167	0.167	0.247	0.155
Average s_r	0.785	0.620	0.792	0.785	0.785	0.785	0.785	0.785	0.785	0.785	0.785	0.716	0.796
Average take-home share μ	0.506	0.562	0.504	0.492	0.494	0.494	0.494	0.490	0.490	0.494	0.495	0.627	0.487
Panel C: Reduced-Form Effects of RTR													
β_{HHI} $(\partial HHI/\partial RTR)$	-0.722	-0.910	-0.663	-0.722	-0.722	-0.722	-0.722	-0.722	-0.722	-0.722	-0.722	-0.794	-0.699
	(0.075)	(0.373)	(0.089)	(0.075)	(0.075)	(0.075)	(0.075)	(0.075)	(0.075)	(0.075)	(0.075)	(0.128)	(0.091)
β_{s_m} $(\partial s_m/\partial RTR)$	0.477	1.585	0.291	0.477	0.477	0.477	0.477	0.477	0.477	0.477	0.477	0.869	0.329
	(0.069)	(0.496)	(0.072)	(0.069)	(0.069)	(0.069)	(0.069)	(0.069)	(0.069)	(0.069)	(0.069)	(0.183)	(0.063)
Panel D: Policy Effect of 10% Increase in RTR													
$\partial \bar{\varepsilon}^{-1}/\partial RTR$	0.005	0.028	-0.000	0.000	0.007	0.005	0.005	-0.006	-0.004	0.008	0.009	-0.025	0.004
$\Delta\mu$ (change in take-home share)	0.000	0.001	-0.000	0.000	0.000	0.000	0.000	-0.000	-0.000	0.000	0.000	-0.001	0.000
Change in take-home share (pp)	-0.051	-0.537	0.042	0.036	-0.100	-0.051	-0.062	0.169	0.126	-0.117	-0.139	1.177	-0.064

Notes: Each column reports the estimated effect of a 10% increase in Regional Tariff Reductions on average wage markdowns for the indicated subgroup. Panel A parameter sources: $1/\eta$ from Table C.2 (region-specific estimates for region columns; pooled for others); Ω from Table C.4 (demographic-specific estimates); $1/\tilde{\eta} = 1/\eta + \Omega$; $1/\tilde{\rho}$ from Table C.6 (demographic-specific estimates where significant; pooled otherwise); $1/\tilde{\theta}$ from Table C.6 (pooled estimate). Standard errors in parentheses. Panel B reports 1991 formal wagebill-weighted averages. Panel C reports reduced-form regressions of market-level changes (1991–2000) on Regional Tariff Reductions, controlling for region fixed effects and clustering standard errors at the microregion level; β_{HHI} and β_{s_m} are from subgroup-specific regressions for informality and region columns, and pooled estimates for demographic columns. Panel D applies the policy effect formula from Section 2. Informality classified by 1991 microregion-level share of non-formal employment above/below the median. Region groups follow the $1/\eta$ heterogeneity: North + Northeast vs Southeast + South + Center West.