Oligopsony in Dual Labor Markets

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Abstract

I present a model of labor market oligopsony where finitely many formal-sector firms operate in dual local labor markets. Local labor markets feature formal wage work, informal wage work within or outside of formal firms, and self-employment opportunities that vary by worker age, education, and gender. Workers seeking formal sector employment make labor supply decisions taking into account the probability of involuntary separation into informal wage work, wages in the local informal wage work sector, and expected returns to self-employment opportunities for their type. Formal sector firms face upward-sloping labor supply curves originating from workers' heterogeneous preferences over different employment modes and over job characteristics, and set formal sector wages accordingly, holding fixed competitors' labor demand. I then estimate workers' preference parameters and implied endogenous wage markdowns combining employer-employee linked data, Census data, and exogenous labor demand shocks from Brazil's trade liberalization. Workers are paid 52 cents on the dollar on average, but this figure varies substantially, between 45 and 70 cents on the dollar, depending on demographic-specific and region-specific preferences and depending on the labor market conditions outside the formal sector. Overall, the estimates suggest that self-employment curbs formal firms' labor market power, while the threat of involuntary separation into informal wage work increases it.

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1 Introduction

As with most emerging economies, Brazil's labor market is of dual nature: formal or registered firms co-exist with informal or unregistered firms and with a large share of the working age population in self-employment. This paper extends the labor market oligopsony model in Felix (2022) to embed informal wage work and self-employment as into a model of labor market oligopsony where formal sector firms exploit heterogeneity in worker preferences over job attributes to set wages below the marginal revenue product of labor. Informal wage work might occur within or outside formal firms as in Ulyssea (2018), while self-employment opportunities are an alternative to wage work, as in Amodio, Medina and Morlacco (2022).

I model self-employment as an alternative to wage work (formal or informal) within each region. This follows from the premise that job seekers may decide to start a small business when local wage work options are hard to secure or unsatisfactory, and embeds the relationship between self-employment and wage markdowns studied in Amodio, Medina and Morlacco (2022) into a model of labor market oligopsony where firms are heterogeneous from workers' perspective. The relationship between self-employment and wage markdowns differs depending not only on firm decisions, but also on regional labor market conditions. The extension adds a new key parameter to the model, the elasticity of substitution between wage work and self-employment, which I estimate using census data and differential shocks to wage work demand across regions and across demographic groups.

The question of how formal firms' wage markdowns relate to informal wage work is more challenging. Informal wage work might occur inside or outside formal firms, and both margins are empirically relevant in the Brazilian context (Ulyssea, 2018). The existence of within-firm informality means that incorporating informal wage work as *nest* within wage work is neither conceptually sound nor empirically feasible, as the two margins of informality can't be distinguished in the data with the same granurality as firm-level formal employment. I overcome these challenges by modeling informal wage work as a possible outcome of involuntary separation from formality and allowing workers to take that into account when choosing between formal sector firms. I assume that people seeking formal sector jobs take into account not only the wages paid by different firms, but also the probability that they might be fired from that firm and, until they find another formal sector job, work in the local informal sector.

This approach also allows me to test how the value of unemployment insurance alters workers' sensitivity to the probability of involuntary separation, a plausible hypothesis in light of recent micro and macro evidence. At the micro-level, recent evidence shows that unemployment insurance increases the appeal of formal sector jobs, as documented in Felix et al. (2026) via survey experiments on Brazilian slum residents, potentially increasing formal firms' wage-setting power. At the macro-level, cross-country evidence from Amodio et al. (2024) suggests that unemployment insurance mediates the relationship between wage markdowns and self-employment. Markdowns increase with self-employment when UI is available, decrease otherwise.

There are several attractive features to this approach. In terms of theory, while the introduction of priors and expectations is a conceptually significant departure from the original model,[1] the aggregation properties of workers' discrete choice problem are still preserved by modeling choices as a function of *expected*, rather than *de facto*, formal sector wages. The expectation takes into account informal sector wages and workers' priors about the probability of involuntary separation from each firm into the local informal wage work sector. The approach's preservation of aggregation properties allows us to derive an *expression for the bias* in estimates of the within-market cross-firm elasticity of substitution η if the extended model is true but data from the informal sector are not incorporated to adjust the estimation.

In terms of empirical advantages, involuntary separations are observed in Brazil's employer-employee linked data, which allows us to proxy for workers' priors over separation probabilities with firm-level separation rates over any prespecified time horizon. In addition, the Census contains information on earnings for informal wage work, separately for different regions and demographic groups, which allows us to obtain reasonable estimates of expectations about regional informal sector pay for different demographic groups and locations.

Literature. Significant advances have been made in estimating wage markdowns in developing countries. Amodio et al. (2024) estimate wage markdowns in a multi-country panel of manufacturing firms in low- and middle-income countries, following the markup estimation method pioneered by De Loecker and Warzynski (2012) and subsequently improved and extended to recover markdowns, as in Yeh, Macaluso and Hershbein (2022) and Ackerberg, Caves and Frazer (2015). They find a median wage markdown of 2, implying a median cross-country wage take-home share of 50 cents on the dollar. While Brazil is not in their sample, this is remarkably similar to the estimate of average wage markdowns for Brazil reported in Felix (2022), based on estimating workers' elasticities of substitution across jobs and the implied wage markdown through the lens of a model of labor market oligopsony. Evidence of strategic competition for workers is also increasing: Sharma (2024) provide evidence of collusion in the labor market among Indian textile firms.

Progress has also been made on understanding the relationship between self-employment and strategic wage setting (i.e., taking other firms' decisions into account). For example, Amodio, Medina and Morlacco (2022) study the relationship between oligopsony and self-employment in Peru, finding that the latter curbs the former, with implications for industrial policy. Relative to their

[1] As I elaborate below, this departure will necessitate additional simplifying assumptions on firm behavior, timing of worker decisions, and dynamics in order for the problem to remain tractable and useful to the exercise at hand.

model—in which comparative advantage is the key driver of worker sorting across sectors—the model I develop features worker preferences as the key driver of sorting and as the source of formal firms' market power, consistent with evidence from survey experiments in Brazil's largest favela complex in Felix et al. (2026). In my model, while comparative advantage indirectly affects each worker's expectations of what they might earn if they chose self-employment, the decision between self-employment and wage work is driven by worker preferences over these job modes, given the set of available wage work options and expected earnings from self-employment.

Importantly, advances have also been made in attempting to understand which regulatory features of labor markets are more strongly correlated with high wage markdowns. Amodio et al. (2024) document that markdowns follow an inverted-U shape with respect to a country's self-employment share, with an inflection point at around 40%, and that unemployment protection mediates this relationship. Countries with unemployment protection, such as those in the former Soviet Union, fall primarily to the left of the inverted-U, with wage markdowns increasing with self-employment. Countries without unemployment protection, such as most countries in Africa, fall primarily to the right of the inverted-U, with markdowns decreasing with self-employment. While Brazil is not in their sample, the estimates from Felix (2022)—of an average wage take-home share of 50 cents on the dollar—combined with roughly 30% of Brazil's employed population in self-employment in the 2000 census, place Brazil to the left of this inverted-U figure's inflection point (of roughly 40% self-employment), along with other countries with unemployment protection.

Last but not least, the literature has begun to document heterogeneity in wage markdowns across demographic groups. Sharma (2023) estimates that over half of the 18 p.p. gender wage gap in Brazil's formal sector in the mid-2000s can be explained by women having stronger preferences for their specific employer (in other words, more inelastic substitution across formal-sector firms than men). Using a similar preferences-based approach to estimate wage markdowns, and the end of the Multifibre Agreement in 2005 as a shock to labor demand in Brazil's formal-sector textile industry, she finds that men earned on average 73 cents on the marginal dollar, while women earned on average 55 cents. Heterogeneous preferences by demographics are a key feature that firms may exploit when setting wages, even if regulatory frameworks prohibit it. Felix et al. (2026) document that women in Brazil's largest favela have a high willingness to pay for parental leave—an amenity only provided by formal firms—whereas men do not value this amenity at all. At the same time, open-text responses reveal that women highly prioritize flexibility, an amenity more commonly available in self-employment, which curbs formal firms' market power.

2 Model

I model the labor supply preferences of *people seeking formal employment*. Firms exploit these preferences to set wages below workers' marginal revenue product when maximizing profits while holding local competitors' labor demand decisions constant. The starting point is the labor supply preference structure in Felix (2022), which I extend on three dimensions. First, I split each local labor market into two sectors, wage work and self-employment. Second, I allow workers to take into account the probability of involuntary separation into informal wage work—which varies by firm—when making formal sector labor supply decisions. Finally, when bringing the model to the data, whenever possible I allow the model's key elasticities of substitution to vary by age, education, gender, and region.

2.1 Discrete choice formal labor supply in dual labor markets

There is a continuum of homogenous workers j, each choosing where to supply l^j effective units of labor.[2] Each worker chooses a single market m from a continuum of local labor markets. Within each market, workers choose a sector s, which is either wage work $\overline{g}$ or self-employment $\underline{g}$. When considering wage work, workers take into account expected earnings from a finite number a formal sector firms z, taking into account their perceived probability of separation into informal wage work o in market m. When considering self-employment, workers take into account expected earnings from starting a small business in market m.

Let l^j_{zgm} denote the effective units of labor that worker j supplies to employment option zsm.[3] Worker j chooses the employment option zsm that minimizes their dis-utility of labor given their preference parameters ξ, expected earnings $\mathbb{E}[l^j_{zgm} w_{zms} | X_{zgm}]$, and their minimum earnings requirement y^j:

$$\min_{zgm} V^j_{zgm} \equiv \ln l^j_{zgm} + \ln \xi_m + \ln \xi_{zgm} - \xi^i_{zgm} \tag{1}$$

$$\text{s.t.} \qquad l^j_{zgm} \bar{w}_{zgm} \geq y^j \tag{2}$$

where $\bar{w}_{zgm} = \mathbb{E}[w_{zms} | X_{zgm}]$ are expected wages from employment option zsm given firm, sector,

[2] I assume that workers are homogenous in both preferences and productivity to center the model discussion around wage differentials driven by worker preferences—the key underlying source of market power—which facilitates exposition and delivers clearer insights. However, both assumptions are easy to relax empirically. I relax the assumption of homogeneous elasticities of substitution in Felix (2022), allowing them to vary by gender, education, and age, after presenting the regression equations that I use to estimate model's key elasticities of substitution. As in Felix (2022)'s original model, the assumption of homogenous productivity is straightforwardly mapped to Brazil's rich microdata by focusing on wages conditional on flexible controls for worker gender, education, age, and worker fixed effects when possible.

[3] For formal sector options (i.e., $s = \overline{g}$), $z\overline{g}m$ represents a formal sector firm z in a market m's wage work sector. For self-employment options (i.e., $s = \underline{g}$), the z subscript refers to the worker (as each worker is their own employer).

and local labor market characteristics. The terms ξ_m, ξ_{gm}, and ξ_{zgm} are market, sector-market, and option-market taste shifters, respectively, while ξ^j_{zgm} is an idiosyncratic worker taste shock for option zsm, drawn from the following nested Generalized Extreme Value (GEV) distribution:

$$G(\xi) = \exp\Big(-\sum_m \Big\{\sum_s \Big(\sum_z e^{-(1+\tilde{\eta})\,\xi^j_{zgm}}\Big)^{\frac{1+\tilde{\rho}}{1+\tilde{\eta}}}\Big\}^{\frac{1+\tilde{\theta}}{1+\tilde{\rho}}}\Big]^{\frac{1}{1+\tilde{\theta}}}\Big). \tag{3}$$

Equations 1-3 preserve the structure of the original labor supply decision problem in Felix (2022), but differ from it in two ways. First, it adds self-employment as an alternative to wage work within each market, with elasticity of substitution $\tilde{\rho}$ between the two sectors, as shown in equation 3. Second, it centers workers' labor supply decisions around *expected earnings*.

Introducing earnings uncertainty is a natural way to incorporate both self-employment and informal wage work into the labor supply decisions of people seeking formal sector employment. For self-employment, workers don't know with certainty the returns to starting a small business ex-ante. For wage work, workers might care about the risk of involuntary separation from any formal sector job into informality (and related expected earnings losses)—whether it is inside or outside firms—when making labor supply decisions. I next derive the implications of these extensions to the labor supply curves facing each formal sector firm and, consequently, markdowns.

2.2 Labor supply curves faced by individual formal firms

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

$$l_{z\overline{g}m} = L\left(\frac{\bar{w}_{z\overline{g}m}}{W_{\overline{g}m}}\right)^{\tilde{\eta}}\left(\frac{W_{\overline{g}m}}{W_m}\right)^{\tilde{\rho}}\left(\frac{W_m}{W}\right)^{\tilde{\theta}} \times \left[\xi_{zgm}^{1+\tilde{\eta}}\xi_{gm}^{1+\tilde{\rho}}\xi_m^{1+\tilde{\theta}}\right]^{-1}. \tag{4}$$

whose inverse is:

$$\bar{w}_{z\overline{g}m} = W\left(\frac{l_{z\overline{g}m}}{L_{\overline{g}m}}\right)^{\frac{1}{\tilde{\eta}}}\left(\frac{L_{\overline{g}m}}{L_m}\right)^{\frac{1}{\tilde{\rho}}}\left(\frac{L_m}{L}\right)^{\frac{1}{\tilde{\theta}}}\xi_{z\overline{g}m}^{1+\frac{1}{\tilde{\eta}}}\xi_{\overline{g}m}^{1+\frac{1}{\tilde{\rho}}}\xi_m^{1+\frac{1}{\tilde{\theta}}}. \tag{5}$$

These are the labor supply equations for formal sector firms, and—as in the original model—equation 5 is the key equation whose derivative with respect to formal sector firms' own employment decisions will determine wage markdowns. Importantly, the relevant wage for attracting labor supply to any one formal sector firm is the *expected* wage $\bar{w}_{z\overline{g}m}$ from taking that formal sector job, which takes into account probabilities of separation into informality and local informal sector conditions:

$$\bar{w}_{z\bar{g}m} = p_{z\bar{g}m} w_m^o + (1 - p_{z\bar{g}m}) w_{z\bar{g}m} \tag{6}$$

where $p_{z\bar{g}m}$ is workers' prior about probability of separation from firm z into informal wage work, w_m^o are expected local informal sector wages, and $w_{z\bar{g}m}$ is firm z's posted formal wage.

In addition, because the elasticity of substitution between self-employment and wage work must necessarily be estimated using census data—which does not include firm boundaries but does include regions—it is useful to introduce the market-level aggregate of equation 5, the inverse labor supply curve to wage work in each market:

$$W_{\bar{g}m} = W \left(\frac{L_{\bar{g}m}}{L_m}\right)^{\frac{1}{\tilde{\rho}}} \left(\frac{L_m}{L}\right)^{\frac{1}{\tilde{\theta}}} \xi_{\bar{g}m}^{1+\frac{1}{\tilde{\rho}}} \xi_m^{1+\frac{1}{\tilde{\theta}}}. \tag{7}$$

Finally, the wage and labor supply indices in equations 4-7 follow the standard nested CES structure:

$$W_{\bar{g}m} = \left[\sum_z \left(\frac{\bar{w}_{z\bar{g}m}}{\xi_{z\bar{g}m}}\right)^{1+\tilde{\eta}}\right]^{\frac{1}{1+\tilde{\eta}}}, \qquad L_{\bar{g}m} = \left[\sum_z \left(\xi_{z\bar{g}m} l_{z\bar{g}m}\right)^{\frac{1+\tilde{\eta}}{\tilde{\eta}}}\right]^{\frac{\tilde{\eta}}{1+\tilde{\eta}}} \tag{8}$$

$$W_m = \left[\left(\frac{W_{\underline{g}m}}{\xi_{\underline{g}m}}\right)^{1+\tilde{\rho}} + \left(\frac{W_{\bar{g}m}}{\xi_{\bar{g}m}}\right)^{1+\tilde{\rho}}\right]^{\frac{1}{1+\tilde{\rho}}}, \qquad L_m = \left[\left(\xi_{\underline{g}m} l_{\underline{g}m}\right)^{\frac{1+\tilde{\rho}}{\tilde{\rho}}} + \left(\xi_{\bar{g}m} L_{\bar{g}m}\right)^{\frac{1+\tilde{\rho}}{\tilde{\rho}}}\right]^{\frac{\tilde{\rho}}{1+\tilde{\rho}}} \tag{9}$$

$$W = \left[\sum_m \left(\frac{W_m}{\xi_m}\right)^{1+\tilde{\theta}}\right]^{\frac{1}{1+\tilde{\theta}}}, \qquad L = \left[\sum_m \left(\xi_m L_m\right)^{\frac{1+\tilde{\theta}}{\tilde{\theta}}}\right]^{\frac{\tilde{\theta}}{1+\tilde{\theta}}} \tag{10}$$

where $W_{\underline{g}m}$ are expected earnings from self-employment in market m.[4]

2.3 Firm-specific wage markdowns

Taking logs of Equation 5 and differentiating it with respect to $l_{z\bar{g}m}$—holding constant labor demand from competing formal firms, priors $p_{z\bar{g}m}$ about the probability of separation, expected returns to

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

self-employment $W_{\underline{g}}$, and expected informal sector wages $\bar{w}^o_m$—gives:

$$\varepsilon^{-1}_{z\overline{g}m} \equiv \frac{\partial \ln \bar{w}_{z\overline{g}m}}{\partial \ln l_{z\overline{g}m}} = \frac{1}{\tilde{\eta}} + \left(\frac{1}{\tilde{\rho}} - \frac{1}{\tilde{\eta}}\right) \underbrace{\frac{\partial \ln L_{gm}}{\partial \ln l_{z\overline{g}m}}}_{s_{z\overline{g}m}} + \left(\frac{1}{\tilde{\theta}} - \frac{1}{\tilde{\rho}}\right) \underbrace{\frac{\partial \ln L_m}{\partial \ln l_{z\overline{g}m}}}_{s_{zm}} \tag{11}$$

where

$$s_{z\overline{g}m} \equiv \frac{\bar{w}_{z\overline{g}m} \cdot l_{z\overline{g}m}}{\sum_k \left(\bar{w}_{k\overline{g}m} \cdot l_{k\overline{g}m}\right)} \quad \text{and} \quad s_{zm} \equiv \frac{\bar{w}_{z\overline{g}m} \cdot l_{z\overline{g}m}}{\sum_s \sum_k \left(\bar{w}_{kgm} \cdot l_{kgm}\right)}$$

are formal sector firm z's *expected* wage bill as a fraction of the formal sector's *expected* wage bill (aka, taking into account separation probabilities into informality and local informal sector wages), and as a fraction of each market's overall *expected* wage bill (including self-employment as an alternative to wage work), respectively.[5] Rearranging:

$$\varepsilon^{-1}_{z\overline{g}m} = \left[\frac{1}{\tilde{\rho}} s_{z\overline{g}m} + \frac{1}{\tilde{\eta}}(1 - s_{z\overline{g}m})\right] + \left(\frac{1}{\tilde{\theta}} - \frac{1}{\tilde{\rho}}\right) s_{zm}. \tag{12}$$

Equation 12 differs from its counterpart in Felix (2022) in three ways. First, it states that the elasticity of labor supply to any one formal sector depends not only on the ease with which workers can substitute between firms within markets $\tilde{\eta}$, and between markets $\tilde{\theta}$, but also on the ease with which they can substitute between wage work and self-employment within markets $\tilde{\rho}$. Second, the relevant wage bill shares for the markdowns are based on expected wages—taking into account probabilities of separation into wage work and local informal sector wages.

Local average wage markdown. As in the original model, the average wage markdown is $\mu_{\overline{g}m} \equiv 1+\varepsilon^{-1}_{\overline{g}m}$ and the wage take-home share is $\mu^{-1}_{\overline{g}m}$. Taking a weighted average of the firm-specific inverse elasticity of substitution in Equation 12 using $s_{z\overline{g}m}$ as weights gives the expression for the inverse elasticity of residual labor supplied to each firm:

$$\bar{\varepsilon}^{-1}_{\overline{g}m} = \frac{1}{\tilde{\rho}} HHI_{\overline{g}m} + \frac{1}{\tilde{\eta}}(1 - HHI_{\overline{g}m}) + \left(\frac{1}{\tilde{\theta}} - \frac{1}{\tilde{\rho}}\right) HHI_{\overline{g}m} \cdot s_m. \tag{13}$$

where $HHI_{\overline{g}m} \equiv \sum_{z\in\Gamma_{\overline{g}m}} s^2_{z\overline{g}m}$ is the Herfindahl-Hirschman Index for wage work sector $\overline{g}$ in market m, measuring the wage bill concentration in sector $\overline{g}$ in market m, and

$$s_m \equiv \frac{\sum_{z\in\Gamma_{\overline{g}m}} s_{z\overline{g}m}}{\sum_{z\in\Gamma_m} s_{z\overline{g}m}}$$

[5] The equivalence between the partial derivatives and market shares can be shown by computing the partial derivatives given the definition of labor supply indices, and then contrasting this result with what we obtain when plugging the inverse labor supply equation into the definition of wage bill market shares.

is the wage bill share of wage work sector $\overline{g}$ in market m.

No dynamic gains from influencing workers' priors. Importantly, these theoretical results assume that firms do not *internalize* that their hiring and firing decisions at time t might influence workers' priors about the probability of separation from the firm at $t+1$ or after. This assumption of *no dynamic gains from formation of priors* simplifies firms' wage setting problem by shutting off $p_{z\overline{g}m}$ as an additional lever for (future) wage setting. As a result, it preserves the original model's static structure while allowing priors to enter labor supply decisions. While workers may form priors based on past firm-level involuntary separations (observed in the administrative data) and act accordingly, neither workers nor firms consider *future* benefits or costs of labor market decisions made at t. The possibility of dynamic gains to labor market power form firms' ability to influence worker priors about the possibility of separation from formation of priors (and about other within-firm dynamic considerations, such as promotion timelines, tenure benefits, etc) merits further research.

3 Estimation strategy

3.1 Substitution Within markets, across firms

I next show how $1/\tilde{\eta}$ can be estimated by combining the IV estimate for $1/\eta$ in Felix (2022) with an estimate of the bias introduced into that estimation strategy if the extended model is true instead. That is, under the assumption that workers take into account the probability of involuntary separation into informality when making formal sector labor supply decisions. First, re-arrange the expected wage equation we have:

$$\bar{w}_{z\overline{g}m} = p_{z\overline{g}m} w^o_m + (1 - p_{z\overline{g}m}) w_{z\overline{g}m} = w_{z\overline{g}m} \left(1 + p_{z\overline{g}m}\sigma_{z\overline{g}m}\right) \tag{14}$$

where $\sigma_{z\overline{g}m} \equiv (w^o_m - w_{z\overline{g}m})/w_{z\overline{g}m}$ is the wage gap between the local informal sector wage w^o_m (for observationally equivalent workers) and firm z's formal wage. Note that this expected wage gap may be negative (informal sector pays less) or positive (informal sector pays more). Taking logs:

$$\ln \bar{w}_{z\overline{g}m} = \underbrace{\ln w_{z\overline{g}m}}_{\text{Formal wage}} + \underbrace{\ln\left(1 + p_{z\overline{g}m}\sigma_{z\overline{g}m}\right)}_{\text{Expected wage gap}} \tag{15}$$

Next, take logs of Equation 5 and let δ_m denote a market fixed effect to write:

$$\ln \bar{w}_{z\overline{g}m} = \frac{1}{\tilde{\eta}} \ln l_{z\overline{g}m} + \delta_m + \ln \xi^{1+\tilde{\eta}}_{z\overline{g}m} \tag{16}$$

Bias formula. Given equation 16, the within-market cross-firm inverse elasticity of substitution $\frac{1}{\tilde{\eta}}$ in the extended model can estimated as the second-stage coefficient from an IV regression where $\ln l_{z\bar{g}m}$ is instrumented by an exogenous labor demand shock $X_{z\bar{g}m}$. Expanding expected wages per equation 15 gives and letting $(*)$ denote partialled-out variables to apply the Frisch-Waugh-Lovell theorem, we have:

$$\frac{1}{\tilde{\eta}} = \frac{\mathrm{Cov}(\ln \bar{w}^*_{z\bar{g}m}, X_{z\bar{g}m})}{\mathrm{Cov}(\ln l^*_{z\bar{g}m}, X_{z\bar{g}m})} = \underbrace{\frac{\mathrm{Cov}(\ln w^*_{z\bar{g}m}, X_{z\bar{g}m})}{\mathrm{Cov}(\ln l^*_{z\bar{g}m}, X_{z\bar{g}m})}}_{\frac{1}{\eta}\text{: Supply response to formal wage}} + \underbrace{\frac{\mathrm{Cov}[\ln(1 + p_{z\bar{g}m}\sigma_{z\bar{g}m})^*, X_{z\bar{g}m}]}{\mathrm{Cov}(\ln l^*_{z\bar{g}m}, X_{z\bar{g}m})}}_{\Omega\text{: Supply response to expected wage gap}} \tag{17}$$

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

The first term in equation 17—the labor supply response to a firm's change in its formal wage—equals the IV estimate for the within-market cross-firm elasticity of substitution $1/\eta$ in the original model in Felix (2022). If the original model were true, workers do not take into account the possibility of separation into informality when making labor supply decisions and, as a result, the second term in equation is zero and the elasticities in the extended and original models are the same.

However, *if the extended model is true*, then the relevant elasticity of substitution for wage setting is $1/\tilde{\eta}$. The term Ω in equation 17 is thus the misspecification bias in the original model if the extended model is true. This bias is zero if there is no within-market cross-firm variation in perceived probabilities of involuntary separation into the informal sector at the time workers make decisions. If the bias in non-zero, its sign depends not only on local informal sector wages, but also on the joint distribution (across formal sector firms in the same market) between perceived probabilities of separation and formal sector wages.

Interpretation. If Ω is *negative*, *less* labor is supplied to firms with higher expected wage gaps relative to the same market's informal sector wage. That is, workers dislike firms that in expectation pay less than the local informal sector.[6]. Conversely, if Ω is *positive*, *more* labor is supplied to firms with higher expected wage gaps relative to the same market's informal sector wage. That is, workers prefer to supply labor to a formal firm *despite* higher expected wages in the local informal sector. In this case, the inverse elasticity of substitution $1/\eta$ estimated based on the original model is *too small* relative to the true inverse elasticity $1/\tilde{\eta}$, meaning that *firms have more labor market power* than originally estimated.

[6]The possibility that informal sector options might pay more than formal sector options should not be discarded. Appendix Figure 3 shows that a non-trivial share of regions exhibit larger monthly earnings for workers of the same demographic composition than monthly earnings at formal sector firms

Regression. I estimate Ω, the labor supply response to changes in expected formal-informal wage gaps *conditional on firm-level changes formal wages*, by running the following second stage regression:

$$\Delta \ln(1 + p_{z\bar{g}r}\sigma^{c}_{z\bar{g}r}) = \Omega\Delta \ln l^{c}_{z\bar{g}r} + \delta_z + \delta_r + \delta_c + \epsilon_{z\bar{g}r} \tag{18}$$

where the left-hand side are 1991-2000 changes in expected wage gaps for each demographic cell, the right-hand side are 1991-2000 changes in employment at firm z in microregion r for demographic cell c, and the δ terms are firm, region-cell fixed effects. Including a firm fixed effect ensures that Ω captures the supply response to changes in expected wage gaps relative to the informal sector conditional on firm-level changes in wage levels, the identifying variation for η in both the original and extended models.

Measurement. Brazil's employer-employee linked dataset RAIS includes information on separation reason for all separations, discerning between firings and quits. I use this information to proxy for workers' prior about the *probability of involuntary separation* from firm z in microregion r using:

$$p_{z\bar{g}r} = \frac{\text{Fired workers from firm } z\bar{g}r}{\text{Fired workers from + Stayers in firm } z\bar{g}r} \tag{19}$$

Since firm employment is measured as of December 31 of each calendar year, I measure priors for year t as of December 31 of year t using separations during that elapsed calendar year. For example, priors as of December 31, 1991 are based on separations between January 1, 1991 and December 30, 1991. Finally, to ensure that the formal-informal wage gaps I measure are not confounded by productivity differences across demographics, I calculate $\sigma_{z\bar{g}m}$ separately for each demographic cell, following the definitions as in the self-employment analysis. I merge RAIS' data for 1991 and 2000 with the census those years at the microregion and demographic cell level, matching on gender, education group, and age group dummies.

Panel A of Figure 3 plots a binned scatter of firm-level probabilities of involuntary separation distribution against informal-formal wage gaps. The probability is on average 3%. It is highest in local labor markets where the informal sector pays less than the formal sector (for the same demographic cell). This means that workers who are at higher risk of separation would also face lower-paying informal sector jobs. Panel B of Figure 3 shows that a substantial number of jobs are in local labor markets where the informal sector pays less, but there is a sizable number of jobs for which informal wage work pays higher wages than the formal sector. This is consistent with the residual wage distributions plotted in Figure 2.

Instruments. Since Ω is a labor supply parameter governing within-market cross-firm substitution in response to formal-informal wage gaps, its identification requires labor demand shocks that vary: (a) across firms within markets, as in the identification of η in the original model; and (b) *within* firm relative to the informal sector, since Ω captures labor supply responses to changes in wage gaps conditional on changes in formal wage levels, captured by η instead. I leverage demographic heterogeneity within firm to construct such an instrument as the interaction between firm-level changes in import tariff reductions $\Delta \ln(1 + \tau_{i(z)})$—the labor demand used to identify η in the original model—and demographic-specific regional tariff reductions ΔRTR_m^c, the labor demand shock used to identify the elasticities of substitution between wage work and self-employment in the extended model.

The key idea is that, conditional on firm-level changes in formal wages, which identify η, the remaining within-market cross-firm labor supply reallocation is driven by cross-firm differences in expected wage gaps relative to the informal sector. Expected wage gaps are, in turn, affected both by the direct import tariff reduction shock to the firm—which change workers' perceived probability of separation from any one firm into informality over the decade—and by regional tariff reductions, which change equilibrium informal sector wages for each demographic group. The exclusion restriction assumes that, conditional on firm, region, and demographic group fixed effects, the only labor supply is affected by tariff reductions is by changing separation probabilities and equilibrium wages, as opposed to workers' distaste for individual firms or for formal versus informal wage work.

Unemployment insurance. Finally, to test how the prospect of receiving unemployment insurance in the event of involuntary separations affect workers' formals ector labor supply decisions, I consider an alternative measure of the expected wage. I do this by measuring the expected wage after separation as 4/12 the same wage as paid by the formal firm (i.e., the typical unemployment insurance benefit at the time) and 8/12 the local informal sector wage. [7]

3.2 Substitution to self-employment and across markets

To arrive at the regression equation to estimate the elasticity of substitution between wage work and self-employment, take logs of equation 7, express it in long-differences, add a constant to absorb

[7]While Brazil's regulation of unemployment insurance benefits in the 1990s were quite complicated—sometimes featuring earnings brackets to limit benefit amounts, and being contingent on proof of time of service—the typical benefit amount was four-months salary, paid monthly. See Brasil (1990) for the 1990 law that instituted the unemployment insurance system. Details on benefit brackets and how the rules have changed over time are available at https://www.debit.com.br/tabelas/seguro-desemprego?utm_source=chatgpt.com.

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

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

A challenge in estimating equation 20 in the context of trade liberalization is that each labor market has a single wage work sector, and import tariff reductions were primarily a labor demand shock to wage work. As a result, $\tilde{\rho}$ cannot be separately identified from $\tilde{\theta}$ unless markets can be partitioned into multiple sub-markets,[8] each with their own wage work and self-employment sectors and their own differential shock to wage work labor demand (the key within-market, cross-sector variation needed to identify $\tilde{\rho}$),[9] but all subject to the same market-level general equilibrium effects of trade liberalization (the key cross-market variation needed to identify $\tilde{\theta}$).

Regression. I overcome this challenge by partitioning labor markets into worker demographic cells, each with their own wage work and self-employment sectors. Let g denote demographic groups defined by gender (men; women), education (primary; secondary; tertiary), and age (young: ages 18-29; middle: ages 30-49; old: ages 50-64) and c denote the fully saturated demographic cells defined by these three groups (e.g., young men with at most primary education). I then estimate:

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

where Δ denotes long-differences within cell by microregion pairs and δ_r is a region fixed effect (denoting one of Brazil's major 5 regions) to absorb regional general equilibrium effects of trade liberalization.

Measurement. I estimate regression equation 22 using data from Brazil's 1991 and 2000 censuses and defining markets as microregions.[10] While $W^c_{\overline{g}m}$, $L^c_{\overline{g}m}$, and L^c_m are technically wage and labor supply indices that reflect lower-nest elasticities and taste shifters, to keep the analysis entirely executable with census data I use the corresponding directly observable measures of these objects.

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

[9] An alternative approach is to introduce additional cross-market variation in import tariff reductions, such as those generated by interactions between regional tariff reductions and baseline market characteristics. I explore this approach in Appendix Table A.2, where I instrument equation 20 with regional tariff reductions and its interaction with: (a) each microreigon's log maximum distance to the nearest labor law enforcement office, borrowing the estimation strategy from Ponczek and Ulyssea (2022); and (b) each microregion's 1991 share of employment that was formal. The approach lacks power but yields similar point estimates, especially for $\tilde{\theta}$, relative to my preferred specification, which adds within-market cross-group variation to better identify $\tilde{\rho}$.

[10] While the 1991 and 2000 censuses contain occupation codes, they are missing for a large share of observations, and the two years also follow different definitions from each other (and from RAIS). As a result, I define labor markets more broadly as microregions only for the purposes of estimating the extended model's elasticities of substitution between self-employment and wage work and across markets.

Specifically, I measure L^c_{gm}/L^c_m as demographic cell c's wage work employment share in microregion m, and $\ln W^c_{gm}$ as demographic cell's c's log residual real monthly earnings in microregion m. Log real monthly earnings are residualized conditional on a fully saturated vector of the gender, education, and age variables, so that the variation in wages used to estimate the model's preference parameters are not driven by productivity differences across demographic cells.

Instruments. I instrument the two endogenous variables in equation 22 with cell-specific regional tariff reductions ΔRTR^c_m and their interaction with: (a) each cell's 1991 formal sector employment share, as tariffs hit formal firms most directly; and (b) demographic group dummies, to soak up variation and allow for heterogeneous effects by groups. The primary instrument ΔRTR^c_m is calculated following the shift-share methodology in Dix-Carneiro and Kovak (2017), featuring exposure shares that are cell-industry-microregion-specific and include *all* workers—formally or informally employed—in tradable sector industries. The exclusion restriction assumption is all cell-specific regional tariff reductions and their interactions with baseline and demographic characteristics are orthogonal to changes in workers' tastes-shifters ξ for wage work and for the market.

Figure A.2 shows that regional tariff reductions reduced *both* employment and earnings in wage work. This same-sign effect is consistent with tariff reductions tracing out labor supply to wage work. The same shock had effects of opposite signs for self-employment: regional tariff reductions *increased* the number of workers in self-employment in more affected regions but, on average, *reduced* their earnings. These opposite-sign effects are consistent with the release of workers from wage work constituting a labor supply shock to the self-employment sector.

Appendix Table A.2 shows a similar pattern for informal wage work as for self-employment, accompanied by a near-zero effect on formal sector residual real monthly earnings. The latter finding differs from the significantly *negative* effects of regional tariff exposure documented based on RAIS-reported earnings in prior literature (e.g., Dix-Carneiro and Kovak (2017), Felix (2022)). Informality within firms, as studied by Ulyssea (2018) and, more recently, Derenoncourt et al. (2025), is a strong candidate explanation for this near-zero effect effect. If higher-wage formal sector workers are moved "off the books," they disappear from RAIS,[11] but they might still report that they are formally employed in the Census, as their job is the same as before. This induces a positive selection into the set of workers marked as formally employed in the Census, even if they are in fact informal. In the absence of data that can discern the intensive margin and extensive margin of informality during this period, it is hard to make empirical progress on elasticities of substitution between informal and formal wage works.

[11] Moving a worker off the books–potentially tagging their leaving as a quit so that no firing cots are paid—saves the firm nearly 100% the worker's wage in the labor costs that are now evaded, even if the worker keeps the same take-home earnings. See Debaere (2003) for a table listing non-wage labor costs to the firm.

3.3 Heterogeneous preferences by demographics

An advantage of partitioning markets by demographic cells—and in constructing cell-specific regional tariff reductions— is that it is the first step in estimating demographics-based heterogeneity in elasticities of substitution. I estimate heterogeneity in the elasticity of substitution between wage work and self-employment by demographic groups as follows:

$$\Delta \ln W^{c}_{\bar{g}m} = \frac{1}{\tilde{\rho}^{g}} \Delta \ln\left(\frac{L^{c}_{\bar{g}m}}{L^{c}_{m}}\right) + \delta_m + \epsilon^{c}_{\bar{g}m} \tag{22}$$

where δ_m is a market fixed effect, absorbing market-level general equilibrium effects of trade common to all demographic cells. I use the same cell-level instruments as for equation 22. The exclusion restriction assumption is that the shocks to wage work labor demand are orthogonal to workers' tastes-shifters ξ for wage work in that market. Unfortunately, there is not enough within-group cross-cell variation in tariff reductions to estimate equation 22 separately by demographic groups, which would yield estimates of heterogeneity in the cross-market elasticity of substitution $\tilde{\theta}$.

If exploited by firms when setting wages, such heterogeneity could generate differential wage markdowns by demographics. Thus, partitioning labor markets by demographic groups effectively extends the original model in Felix (2022) in a third direction, contributing the first labor market oligopsony framework inclusive of self-employment and informal wage work to a vast literature on wage differentials by demographics, where evidence on the link between wage markdowns and gender wage gaps has begun to emerge (e.g., see Sharma (2023), Hoang, Mitra and Pham (2024)).

3.4 Wage markdowns and wage take-home shares

I combine estimates of the extended model's parameters with direct measures of labor market equilibrium wage bill shares to compute local labor market m's average wage markdown according to the extended model. The markdown is given by:

$$\mu_{\bar{g}m} \equiv 1 + \varepsilon^{-1}_{\bar{g}m} = \left[\frac{1}{\tilde{\rho}} HHI_{\bar{g}m} + \frac{1}{\tilde{\eta}}(1 - HHI_{\bar{g}m})\right] + \left(\frac{1}{\tilde{\theta}} - \frac{1}{\tilde{\rho}}\right) HHI_{\bar{g}m} \times s_m. \tag{23}$$

where $1/\tilde{\eta}$ is calculated as $1/\tilde{\eta} = 1/\eta - \Omega$ using estimates of the bias Ω introduced by expected involuntary separation into informality, $s_{z\bar{g}m}$ and s_{zm} are formal sector firm z's *expected* wage bill as a fraction of the formal sector's *expected* wage bill, and as a fraction of each market's overall *expected* wage bill (including self-employment as an alternative to wage work), respectively. Table **??** summarizes the model parameters and key endogenous objects—which act as weights of these preference parameters in computing workers inverse elasticity of residual labor supply—along with their main estimates and ranges based on detected heterogeneity.

4 Results

4.1 Elasticities of substitution

Substitution within markets, across firms. Table 3 reports IV estimates of the within-market cross-firm inverse elasticity of substitution, $1/\eta$, and its regional heterogeneity. Column (1) replicates the 0.0985 estimate in Felix (2022), which is the relevant within-market cross-firm elasticity of substitution in Section 2 if workers are indifferent to the threat of involuntary separation to informal wage work (namely, if $\Omega = 0$). The near-unit value of this elasticity is nearly 7 times higher than its corresponding value for the United States, and remains—as in Felix (2022)—the leading driver of Brazil's high markdown levels.

This near-unit and relatively inelastic value is not homogeneous across Brazil, however. Columns (2)-(5) report heterogeneity by Brazil's major regions. Column (2) shoes that the pooled elasticity is driven by the Southeast— the largest region by employment and where the formal sector made up more than half of total employment share in 1991, as shown in the microregions to the left of Figure 1—where the within-market cross-firm elasticity of substitution is estimated at 1.022.

The picture is very different in the Northeast, Brazil's second largest region by population and employment. Column (2) shows that the within-market cross-firm elasticity of substitution in the Northeast is nearly half of that in the Southeast, at 0.462. Since nearly all employment in the Northeast is informal—these are primarily the microregions to the right of Figure 1, with the exception of a few metropolitan areas (e.g., PE-3 and AL-3), where formality is high—more elastic within-market cross-firm elasticity of substitution in this region suggests raises the interesting possibility of congestion—or larger employer-employee matching frictions within the formal sector—in markets where many more firms are formal and small. Columns (4) and (5) report coefficients pooling with other regions where the first stage of Brazil's firm-level tariff shocks is too weak to yield estimates restricted to the smaller regions of the North, South, and Center West. I pool the North with the Northeast given their similarity in levels of development, and the South and Center West with the Southeast, for similar reasons. Both estimates are attenuated when pooled.

Bias from separation into informal wage work. Panel A of Figure 5 plots the reduced form and first stage relationships that identify Ω. They are both negative: employment and expected informal-formal wage gaps declined in firms and markets more exposed to import competition. Combined these imply that Ω is positive, such that the threat of involuntary separation from a formal job increases firms' market power. For ease of interpretation, Panel B of Figure shows the OLS relationship between the two variables. The OLS says that firms that grow are also those that distance themselves–in terms of earnings–from local labor market conditions. Some of these firms pay more than local informal jobs, but many pay less, as shown in Panel B of Figure 2. Table 4 reports the implied IV coefficients and its heterogeneity by demographic groups.

Table 4 examines the bias term Ω in the within-market cross-firm elasticity of susbtitution presented in Table 3 if the extended model is true, namely, if workers take into account the probability of involuntary separation into informal wage work when making labor supply decisions. The most important take-away is that the magnitude of the bias is very small. At 0.0326 on average per Column (1), Ω shows that the within-market cross-firm elasticity of substitution is roughly 3% *smaller* in Felix (2022)'s model than in the extended model where involuntary separations are considered. In other words, the threat of involuntary separation into informal wage work slightly increases formal firms' labor market power. Columns (2)-(4) then show some heterogeneity in this estimate by demographics, though the magnitudes are not significantly different from each other.

Unemployment insurance. An interesting question is whether the value of unemployment insurance alters workers' attitude towards being fired from a formal sector job. Appendix Table 4 tests this hypothesis. Consistent with the cross-country evidence Amodio et al. (2024), I find that unemployment insurance makes workers less responsive to the threat of involuntary separation, curbing labor market power. This results from contrasting the magnitude of the Ω bias introduced by involuntary separations if it is estimated with or without the value of unemployment insurance benefits, available in Appendix Table A.1. When Ω is positive, the threat of involuntary separation to the informal sector increases firms' labor market power. When it is negative, the informal sector curbs that market power. While all estimates I find for Ω are positive, adding unemployment insurance to the estimation reduces Ω from 0.0326 to 0.0217. This suggests that unemployment insurance, on average, operates like a substitute to self-employment at Brazil's level of economic development. This is consistent with the findings in Amodio and de Roux (2021). The elasticity of substitution to self-employment is however a much stronger force in curbing labor market power than the contribution of unemployment insurance to its relevant elasticity ($1/\eta$).

Substitution to self-employment. Tables 5 and 6 report estimates of inverse elasticities governing substitution between wage work and self-employment, $1/\tilde{\rho}$, as well as cross-market substitution, $1/\tilde{\theta}$. The magnitudes are generally small, indicating highly elastic reallocation between wage work and self-employment and across markets, which limits equilibrium labor market power. Heterogeneity results show that substitution is particularly elastic for demographic groups with stronger outside options, reinforcing the role of self-employment and informal work as key competitive constraints on formal-sector wage setting.

Appendix Figure A.1 plots the first stage of equation 22—namely, the effect of demographic-specific regional tariff reductions on log share of wage work employment—separately by demographic groups. These figures are demographic-specific versions of Figure 4, which plots the pooled, across groups, relationship between wage work and regional tariff reductions. The variation in Appendix Figure A.1 is the one I use to estimate the elasticities of substitution, reported in Tables 5 and 6.

4.2 Wage markdowns and key take-aways

On average, Brazilian formal sector workers were paid 52 cents of their marginal revenue product of labor in 1991. However, this average masks substantial heterogeneity across regions and across demographics. Take-home shares are highest for less educated workers, for whom the minimum wage is a substantial increase relative to their earnings in either informal wage work or self-employment, and for women, for whom self-employment earnings are higher than either formal sector earnings or informal sector earnings at any point of the formal wage take-home share distribution (see, e.g., Figure A.6).

Figure 6 shows the distributions of average wage take-home shares—the inverse of wage markdowns—across microregions and across occupations, since a local labor market is defined as a microregion x 2-digit occupation pair. The dispersion is driven by regional heterogeneity in the within-market cross-firm inverse elasticity of susbtitution $1/\eta$. Most microregions have wage markdowns around 50 cents on the dollar, with that mass increasing between 1991 and 2000. Microregions in the North and Northeast of Brazil, for which within-market cross-firm elasticities of substitution are almost half those of the rest of the country, feature higher wage take-home shares. The dispersion across occupations is centered between 50 and 55 cents on the dollar.

Figure 7 displays the microregion-level average markdowns by microregion, along with each microregion's share of employment in formal wage work, informal wage work, and self-employment. A clear pattern is that wage take-home shares are typically higher, though not always, in places where a larger share of the workforce is self-employed. Regions that are primarily formal, such as microregions in the states of São Paulo or Santa Catarina, show little dispersion of average wage take-home shares, though variation still exists as places differ in demographic composition, and demographic groups have different elasticities of substitution to self-employment and respond differently to the threat of separation into informal wage work. Appendix Figures A.4 through A.6 replicate these figures but showing demographic-specific local wage markdowns.

It should also be noted that a substantial share of the earnings gaps between the informal, informal, and self-employment sector is accounted for by the demographics of the workers, as shown in Figure 2. In fact, the 1991 residual earnings distribution shows that observationally-equivalent workers were paid on average higher earnings in either the informal wage work sector or in self-employment. By 2000, the residual earnings gap between the formal and informal wage work sectors closes, whereas self-employment–primarily composed of women—remained a higher return activity on average.

Wage markdowns and concentration. Since the model in Section 2 features ample heterogeneity in preference parameters and implies that local labor market conditions outside of the formal sector affect firms' wage setting, the relationship between local labor market concentration and wage markdowns is no longer a linear function of formal sector concentration, as in Felix (2022)'s framework. The relationship is now tied to many more variables, including cross-market variation in demographic composition.

Figure 10 and its demographic-specific versions (Appendix Figures A.3) plot the relationship between the formal sector's local Herfindahl-Hirschmann Index (HHI)–measured with respect to expected wages—and wage take-home share estimates. It shows a non-linear relationship likely driven by cross-regional differences in elasticities of substitution. Wage take-home shares increase in labor market concentration for markets with numerous similarly sized firms (i.e., HHI below 0.2). Past that threshold, labor market concentration among formal sector firms typically reduces formal sector wage take-home shares.

These findings reveal three important lessons about oligopsony in dual labor markets: (1) Self-employment and informal wage work materially shape wage markdowns by altering workers' outside options; (2) The threat of involuntary separation into informal work increases formal firms' labor market power, though this effect is attenuated by unemployment insurance; and (3) Self-employment curbs formal firms' labor market power, an effect that is especially pronounced for women.

5 Discussion

This paper shows that oligopsony in Brazil's dual labor market is shaped not only by concentration among formal firms, but by the structure of workers' outside options. Extending a standard preferences-based framework to incorporate self-employment and the risk of involuntary separation into informal wage work, I recover the elasticities that discipline firms' wage-setting behavior and map them into wage markdowns. The estimates indicate that formal firms pay workers, on average, roughly half of their marginal product, with meaningful variation across regions and demographic groups. Self-employment emerges as an important competitive force limiting firms' market power, while separation risk modestly amplifies it. Together, the results suggest that policies affecting the returns to self-employment, the functioning of the informal sector, or the insurance value of formal jobs can all shape equilibrium wage markdowns.

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Table 1: Brazilian workforce during 1990s trade liberalization

	Population Census Year	
	1991	2000
Number of employed people	41,916,200	55,562,018
Formal wage work share	0.42	0.36
Informal wage work share	0.28	0.38
Self-employment share	0.30	0.25
Hours worked per week	44.19 (11.69)	44.18 (14.77)
Monthly earnings (R$ 2000)	304.28 (380.81)	539.05 (692.15)
Monthly earnings (R$ 2000) - Formal	382.65 (403.00)	589.40 (644.85)
Monthly earnings (R$ 2000) - Informal	161.20 (248.60)	457.60 (693.29)
Monthly earnings (R$ 2000) - Self-employed	329.97 (413.88)	570.84 (746.29)
Monthly minimum wage (R$ 2000)	126.51	151.00
Female share	0.29	0.38
Age	34.08 (11.53)	35.28 (11.47)
Years of education	5.35 (4.28)	6.91 (4.44)

Source: This table shows descriptive statistics from the Brazilian 1991 and 2000 population censuses, calculated using the 1991 and 2000 census sample extract files from the supplementary materials for Dix-Carneiro and Kovak (2017). Earnings and hours are for main job. All statistics are calculated using individual-level sample extract weights. Earnings statistics exclude top 1% and bottom 1% of earnings distribution. The sample includes all workers employed in the private sector (that is, excludes public administration). Formal wage work follows the standard definition for Brazilian labor markets, which is to have a signed *Carteira de Trabalho* (variables v350 and v0447 in 1991 and 2000 censuses, respecively). Real monthly earnings are based on the the IPCA deflator and are expressed in 2000 reais. Minimum wage reports the federal monthly minimum wage for July 1991 and for 2000. Brazil's minimum wage are regulated as minimum monthly earnings for a 44-hour work week.

Table 2: Wage markdowns in dual labor markets: Worker preference parameters and weights

Parameter (1)	Definition (2)	Estimation: Data (3)	Tables (4)	IV estimate (5)	Statistically signif. heterogeneity: Range (6)	Detected by (7)
$1/\eta$	Within-market cross-firm inverse elasticity of substitution	RAIS + Tariffs	3	0.985	0.462 - 1.022	Region
Ω	Bias in $1/\eta$ due to expected involuntary separation into informal wage work	RAIS + Census + Tariffs	4	0.0326	0.0273 - 0.0346	Age, gender
$1/\tilde{\rho}$	Inverse elasticity of substitution between wage work and self-employment	Census + Tariffs	5 and 6	0.482	0.303 - 0.935	Age, educ, gender
$1/\tilde{\theta}$	Cross-market inverse elasticity of substitution	Census + Tariffs	5	1.215	-	-
Directly measured labor market equilibrium objects ("weights")						
$HHI_{\bar{g}m}$	*Expected* wage bill concentration among formal firms in market m.	RAIS + Census				
s_m	Wage work sector wage bill share as fraction of overall wage bill in market m.	RAIS + Census				

Notes: Column (3) lists the datasets used; column (6) lists the heterogeneity ranges; column (7) lists the heterogeneity dimensions. A market is a microregion × 2-digit occupation pair.

Table 3: IV estimates of $1/\eta$ and its heterogeneity across major regions

	$\Delta \ln w_{z\bar{g}m}$				
	(1)	(2)	(3)	(4)	(5)
	Main	NE	SE	N + NE	SE + S + CW
$\Delta \ln l_{z\bar{g}m}$	0.985***	0.462**	1.022***	0.406*	1.000***
	(0.173)	(0.231)	(0.142)	(0.206)	(0.161)
First-stage F	68.85	22.49	88.13	23.98	71.96
Observations	846365	99639	529409	114717	731648

Notes: This table shows IV estimates of $1/\eta$, the within-market cross-firm inverse elasticity of residual labor supply in Felix (2022) (and in Section 2 if $\Omega = 0$) and its heterogeneity by Brazil's major regions. The sample includes all formal sector firms in RAIS in 1991 and 1997. Each observation is a firm X local labor market cell. A local labor market is a microregion x 2-digit occupation pair. All regressions include local labor market fixed effects. The instruments are firm-level import reductions from Felix (2022). Column (1) replicates the average elasticity from Felix (2022). Column (2) re-estimates Column (1) with a microregion fixed effect instead of a microregion x 2-digit occupation pair fixed effect. Column (3) re-estimate Column (1) with a 2-digit occupation fixed effect instead of a microregion fixed effect. Columns (4)-(7) re-estimates Column (1) limiting the sample to microregions within major regions: NE (Northeast), SE (Southeast), N (North), S (South), and CW (Center West). Limiting the sample to the N, S, or CW alone yields weak first stages and statistically insignificant results, hence why Columns (4)-(5) present results pooling them with Brazil's largest regions in terms of population (Northeast and Southeast). Standard errors are clustered two-ways by microregion and 2-digit occupation. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 4: IV estimates of Ω and its heterogeneity by demographic groups

	$\Delta \ln(1 + p_{z\bar{s}m}\sigma^c_{z\bar{s}m})$			
	All	By gender	By education	By age
$\Delta \ln l^c_{z\bar{s}m}$	(1)	(2)	(3)	(4)
All	0.0326**			
	(0.0129)			
Men		0.0326**		
		(0.0131)		
Women		0.0343**		
		(0.0160)		
Primary education			0.0120	
			(0.0256)	
Secondary education			-0.0412	
			(0.0564)	
Tertiary education			0.0426	
			(0.0387)	
Young (18-29)				0.0339**
				(0.0144)
Middle (20-49)				0.0346***
				(0.0134)
Old (50-64)				0.0273*
				(0.0164)
First-stage F	1.526	3.243	2.727	5.835
Anderson-Rubin F		4.199	25.95	3.993
Observations	394800	394800	394800	394800

Notes: This table shows IV estimates of Ω, the bias term in IV estimates of the within-market cross firm elasticity of substitution in the model of Felix (2022) if the extended model is true, and its heterogeneity by worker characteristics. The sample includes all formal sector firms in RAIS in 1991 and 2000, merged at the microregion X demographic cell X year level to the corresponding census data. Each observation is a firm X microregion X demographic cell (gender X education group X age group dummies). All regressions include firm, microregion, and demographic cell fixed effects. Expected informal-formal wage gap is calculated separately by firm, microregion, and demographic cell as $p_{z\bar{s}r}\sigma^c_{z\bar{s}r}$. The first term, $p_{z\bar{s}m}$, is the within-year probability of *involuntary* separation from firm z. The second, $\sigma^c_{z\bar{s}m} \equiv (w^o_m - w_{z\bar{s}m})/w_{z\bar{s}m}$, is the monthly earnings gap between the local informal sector and firm z's formal wage for that demographic cell. Monthly earnings gaps are calculated by first converging all reported earnings to 2000 constant Reais. $p^c_{z\bar{s}m}$ is calculated using separations between between January 1 and December 30 of each year. Firm total formal employment is measured as of December 31 of each year. The instrument for the IV estimate is the interaction between firm-level import reductions and regional tariff reductions, interacted with demographic group dummies and firm baseline formal employment. Regional tariff reductions are microregion-level exposure to 1990-1994 import tariff reductions from Dix-Carneiro and Kovak (2017) and differ by demographic group. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 5: IV Estimates of $1/\tilde{\rho}$ and $1/\tilde{\theta}$

	$\Delta \ln W^c_{\bar{s}m}$ (1)
$\Delta \ln(L^c_{\bar{s}m}/L^c_m)$	0.458^{***}
	(0.0767)
$\Delta \ln(L^c_m/L^c)$	1.215^{*}
	(0.625)
First-stage F ($1/\rho$)	66.57
Anderson-Rubin F ($1/\rho$)	60.46
First-stage F ($1/\theta$)	3.099
Anderson-Rubin F ($1/\theta$)	37.47
Observations	8055

Notes: This table shows second stage estimates from an instrumental variables regression that estimates the inverse elasticity of substitution between wage work and self-employment $1/\tilde{\rho}$ as the coefficient on $\Delta \ln(L^c_{\bar{s}m}/L^c_m)$, and the cross-market elasticity of substitution $1/\tilde{\theta}$ as the coefficient on $\Delta \ln(L^c_{\bar{s}m}/L^c_m)$. The sample includes 486 microregions and 18 demographic cells defined by 8 major demographic groups (2 by gender) x (3 by education) x (3 by age). The outcome variable is the 1991-2000 change in log residual real monthly earnings among individuals employed in wage work, either formally or informally, for each demographic group. The dependent variables are (1) the 1991-2000 change in log share of wage work employment in a microregion for each demographic cell; and (2) the 1991-2000 change in the log share of microregion total employment relative to national employment for each demographic cell. The instruments are group-specific Regional Tariff Reductions interacted with demographic group dummies and with the 1991 formal share of employment in the microregion. The regression includes region fixed effects for each of Brazil's five major regions and is weighted by each cell's 1991 microregion population. Log residual real monthly earnings are condition flexible controls for gender, education, and age. Real monthly earnings are based on the the IPCA deflator and are expressed in 2000 reais. Standard errors in parentheses are clustered by microregion. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

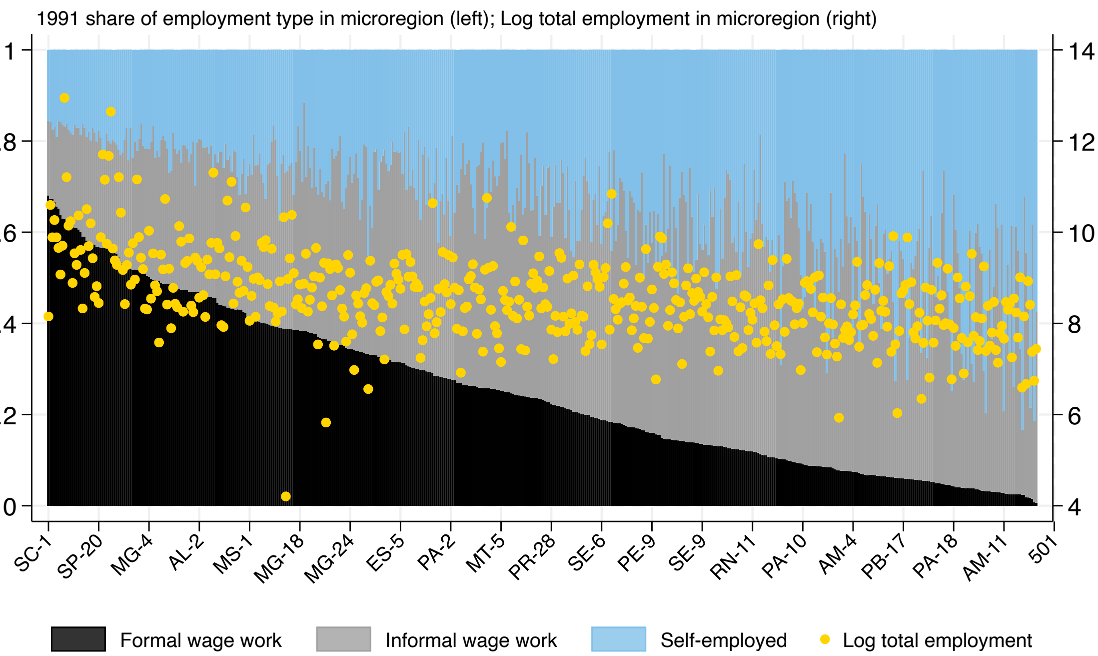
Table 6: IV Estimates of $1/\tilde{\rho}$ and its heterogeneity by demographic groups

	$\Delta \ln W^c_{\bar{s}m}$			
	All	By gender	By education	By age
$\Delta \ln(L^c_{\bar{s}m}/L^c_m)$	(1)	(2)	(3)	(4)
All	0.482***			
	(0.0594)			
Men		0.660***		
		(0.186)		
Women		0.390***		
		(0.122)		
Primary education			0.0686	
			(0.0668)	
Secondary education			0.460***	
			(0.177)	
Tertiary education			0.935***	
			(0.307)	
Young (18-29)				0.850***
				(0.157)
Middle (20-49)				0.355***
				(0.100)
Old (50-64)				0.303***
				(0.104)
First-stage F	52.64	90.64	33.53	30.18
Anderson-Rubin F	52.64	105.1	44.43	44.89
Observations	8055	8055	8055	8055

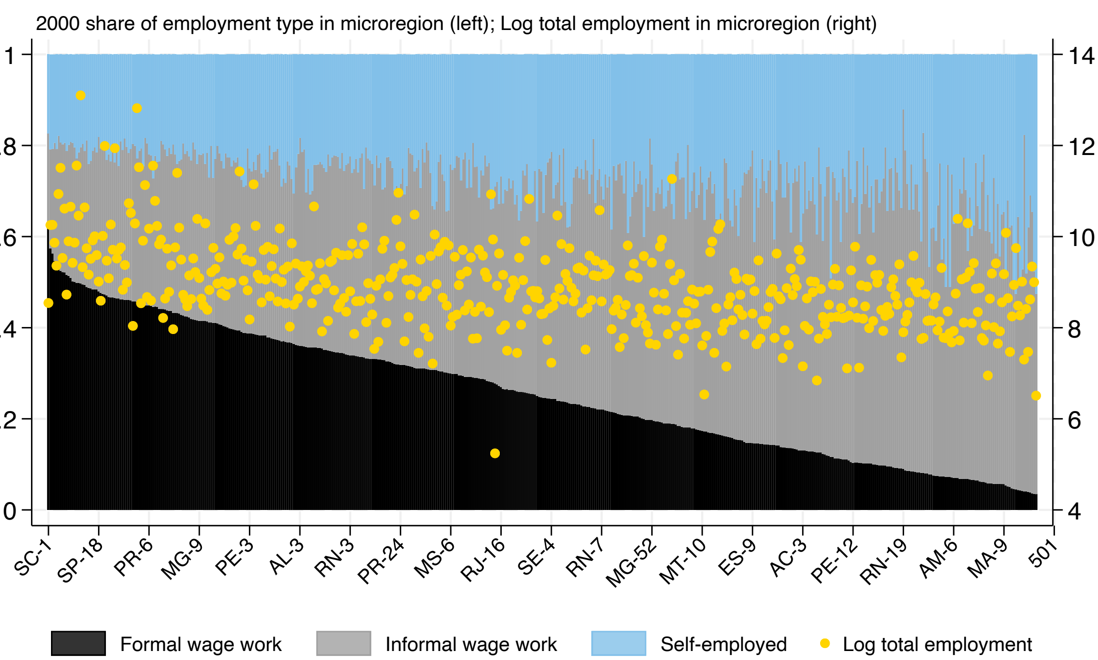
Notes: This table shows second stage estimates from instrumental variables regressions that estimate the inverse elasticity of substitution between wage work and self-employment $1/\tilde{\rho}$ jointly or separately by demographic groups. The sample includes 486 microregions and 18 demographic cells defined by 8 demographic groups (2 by gender) x (3 by education) x (3 by age). The outcome variable is the 1991-2000 change in log residual real monthly earnings among individuals employed in wage work, either formally or informally, for each demographic cell. The dependent variable is the 1991-2000 change in log share of wage work employment in a microregion for each cell. The instruments are cell-specific Regional Tariff Reductions interacted with group dummies and with each cell's 1991 formal share of employment in the microregion. All regressions include microregion fixed effects. Column (1) reports a pooled regression coefficient. Columns (2)-(4) report coefficients on the dependent variable interacted with dummies for each of the 8 major demographic groups. Columns (1) and (2) are weighted by each cell's 1991 microregion population. Columns (3) and (4) are unweighted due to substantial heterogeneity in workforce composition by education and age across microregions, resulting in larger and noiser estimates. Log residual real monthly earnings are conditional on flexible controls for gender, education, and age. Real monthly earnings are based on the the IPCA deflator and are expressed in 2000 reais. Standard errors in parentheses are clustered by microregion. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Figure 1: Wage work and self-employment by microregion, 1991 and 2000

Panel A: 1991



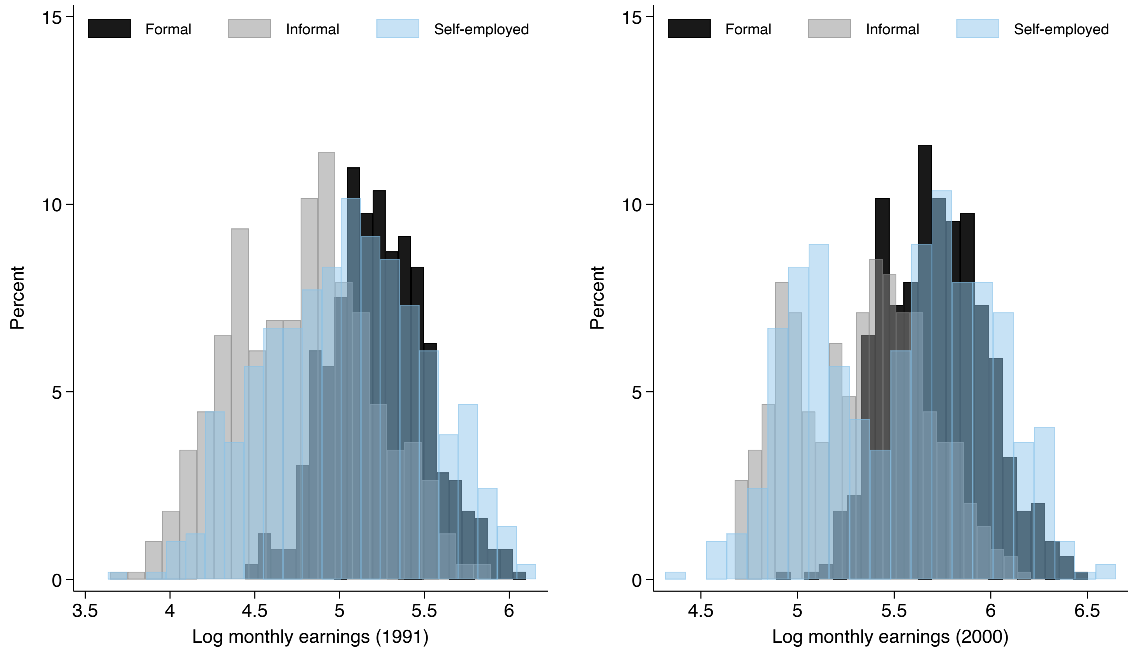
Panel B: 2000



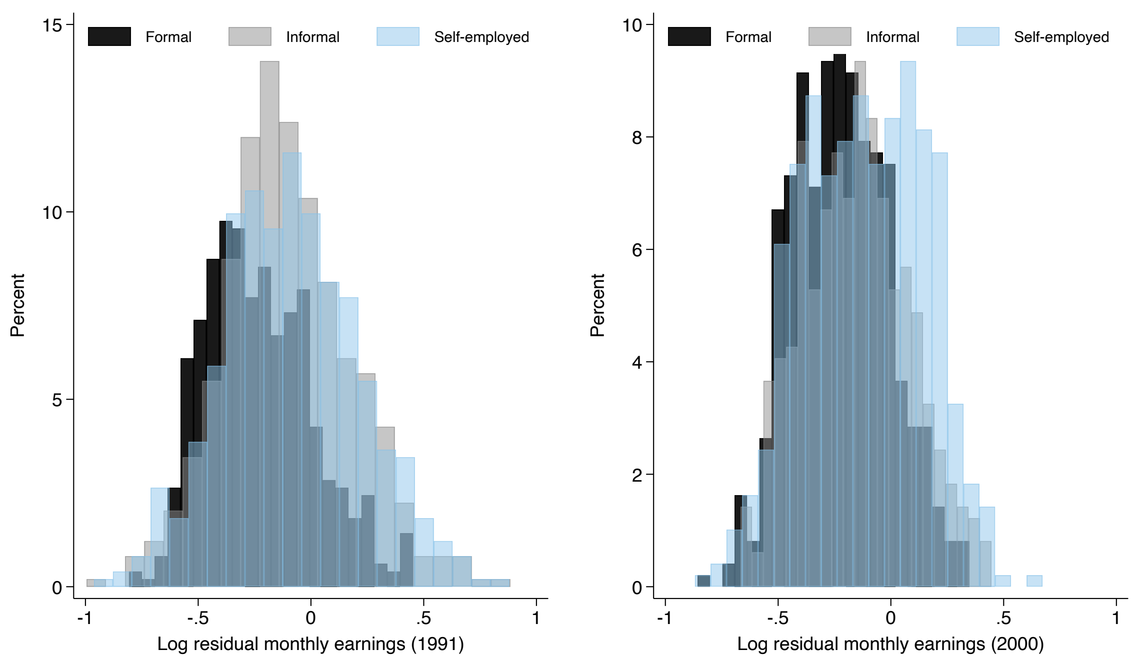
Notes: This figure plots the distribution of formal wage work (black bins, left y-axis), informal wage work (gray bins, left y-axis), self-employment (blue bins, left y-axis), and log total employment (yellow scatter, right y-axis) across 486 microregions per the 1991 and 2000 Population Census. Labels in the x-axis correspond to different microregions. The first two letters indicate the micoregion's state, and the number indicates the microregion's rank in the state by formal sector share (e.g., "SC-1" = top-ranked microregion in Santa Catarina).

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

Panel A: Raw earnings



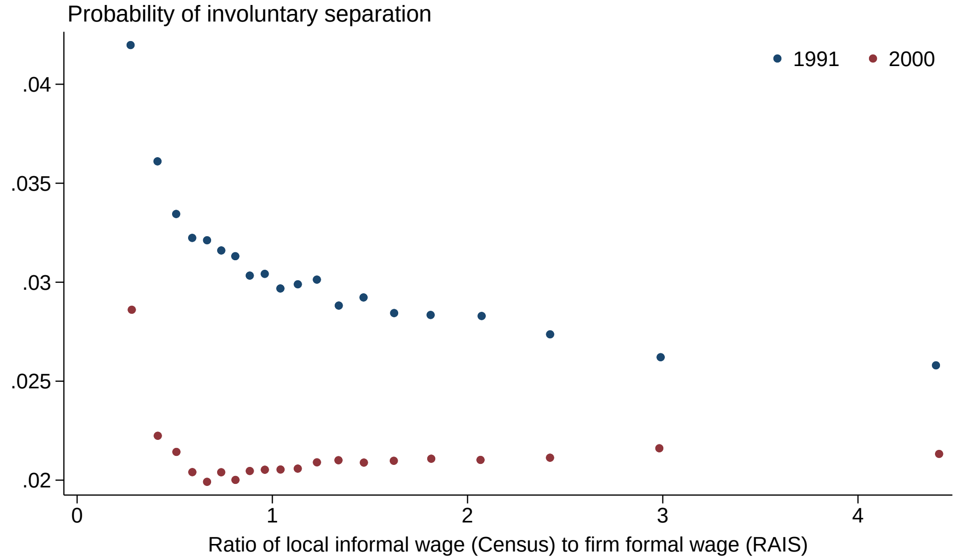
Panel B: Residual earnings



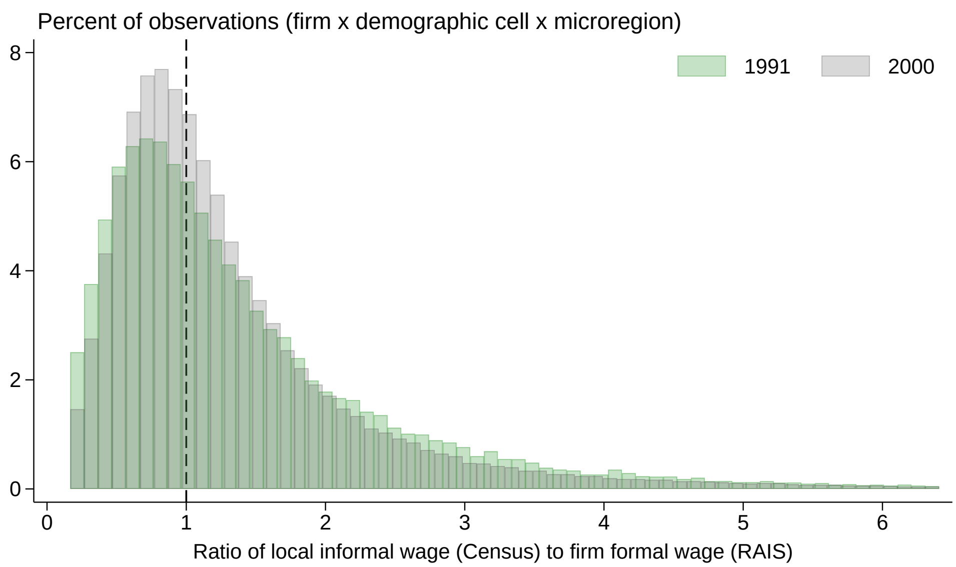
Notes: This figure plots the distribution of raw and residual formal real monthly earnings, informal real monthly earnings, and self-employment real monthly earnings, across 486 Brazilian microregions. Residual real monthly earnings are conditional on a fully saturated vector of dummies for gender, age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

Figure 3: Probability of involuntary separation from formal firm and informal-formal wage gaps

Panel A: Probability of involuntary separation and wage gap



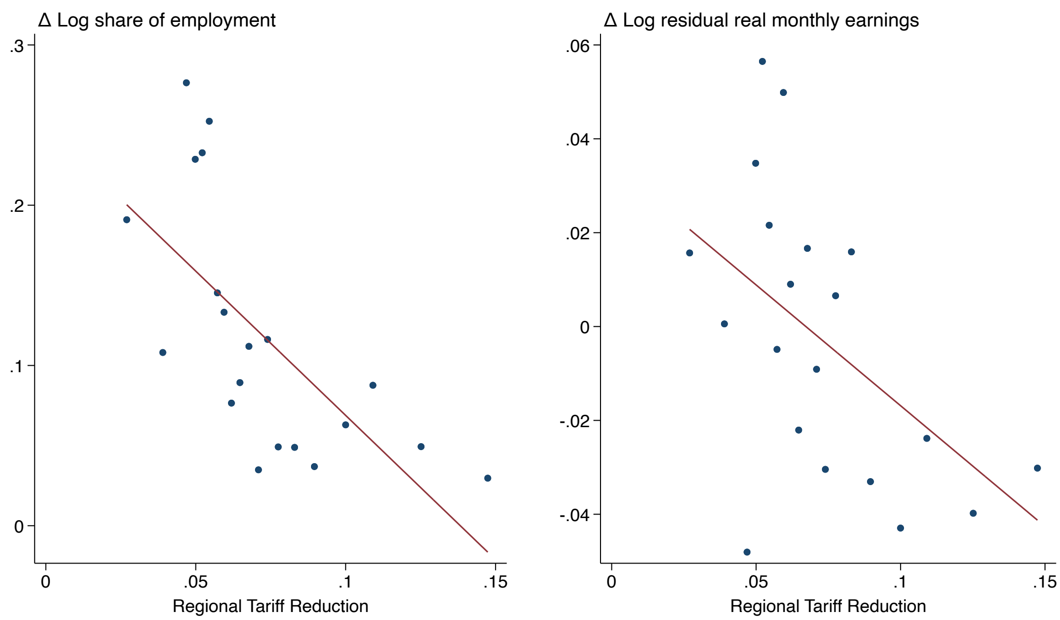
Panel B: Wage gap distribution



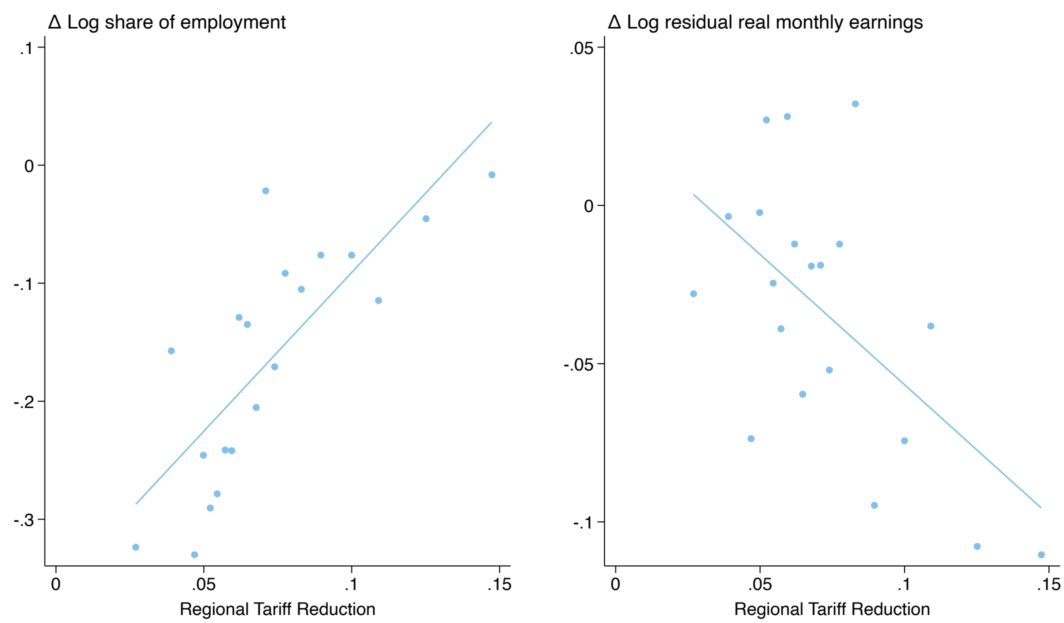
Notes: This figure shows the relationship between probabilities of involuntary separation from a formal sector firm in RAIS and the gap between local informal sector monthly earnings and the firm's formal monthly earnings, separately computed for each demographic cell. The sample includes all formal sector firms in RAIS in 1991 and 2000, merged at the microregion X demographic cell X year level to the corresponding census data. Each observation is a firm X microregion X demographic cell (gender X education group X age group dummies). Monthly earnings gaps are calculated by first converging all reported earnings to 2000 constant Reais. The probability of involuntary separation is calculated using separations between between January 1 and December 30 of each year. Firm total formal employment is measured as of December 31 of each year.

Figure 4: Effect of Regional Tariff Reductions on wage work and self-employment

Panel A: Wage work



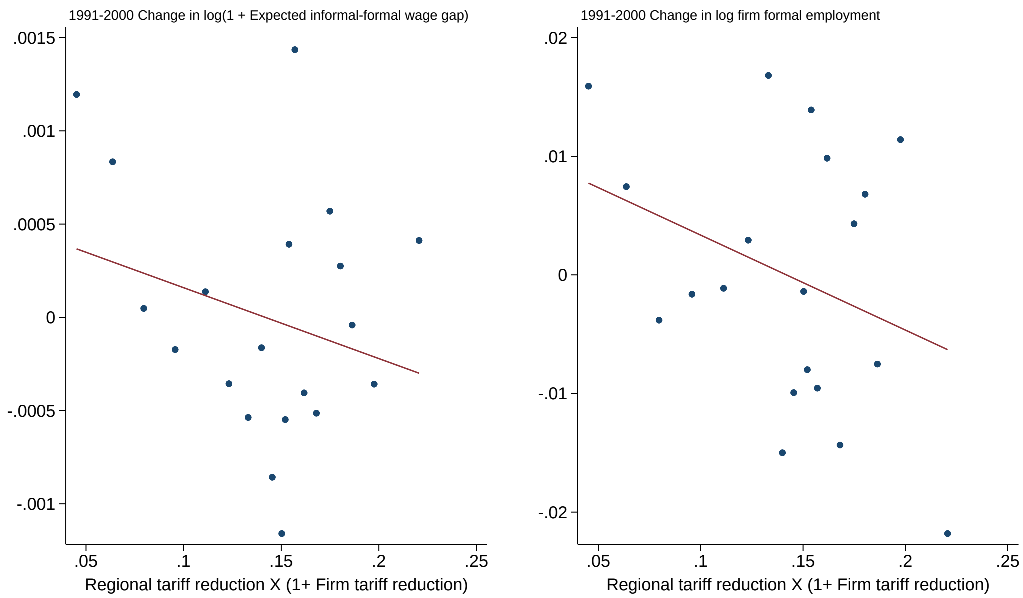
Panel B: Self-employment



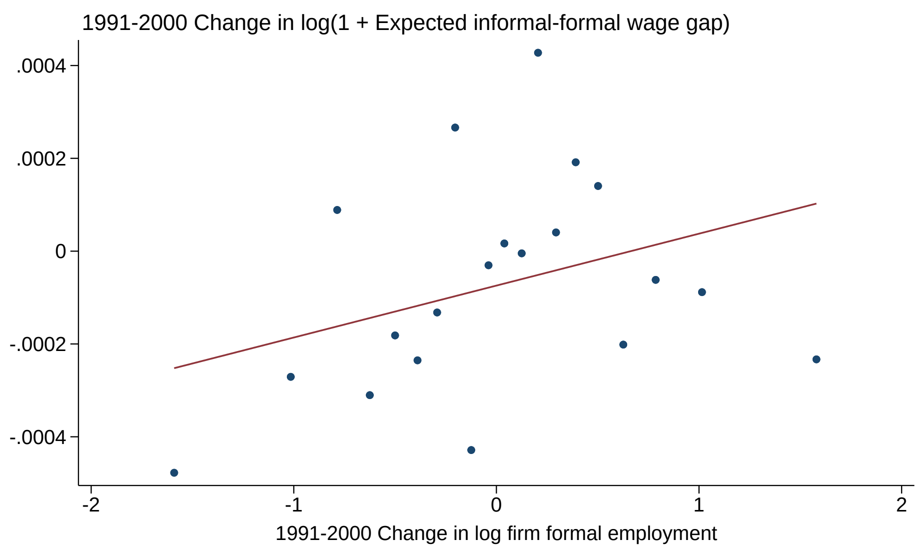
Notes: This figures shows the correlation between Regional Tariff Reductions (i.e., microregion-level exposure to 1990-1994 import tariff reductions from Dix-Carneiro and Kovak (2017)), changes in employment shares (left panel), and changes log residual real monthly earnings (right panel) between the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

Figure 5: Firm-specific labor supply response Ω to expected informal-formal wage gaps

Panel A: Visual IV: Reduced form (left) and First Stage (right)



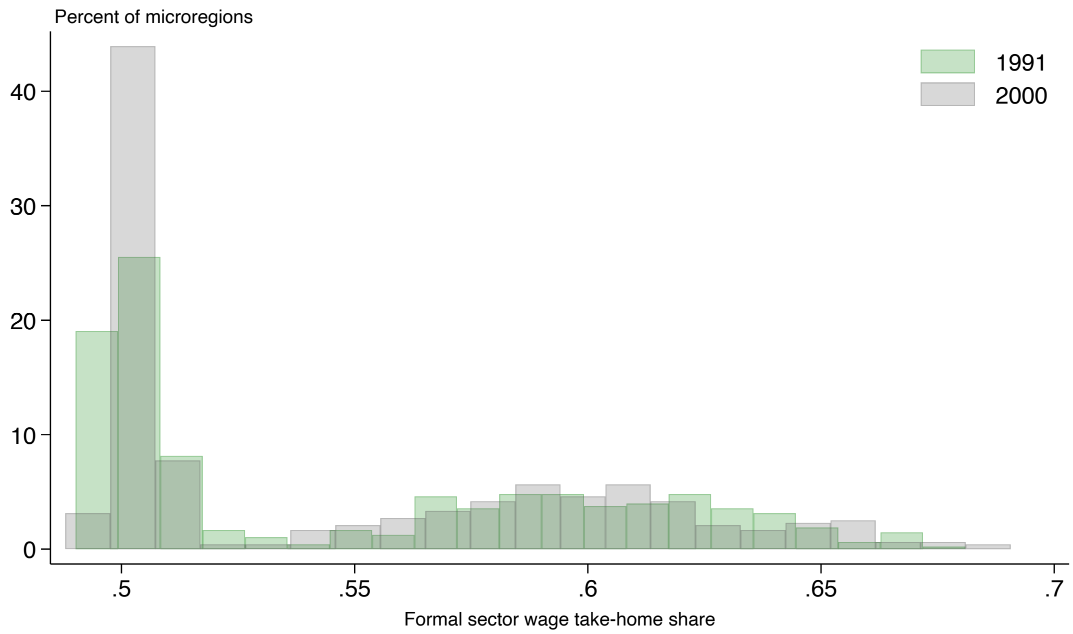
Panel B: Visual OLS



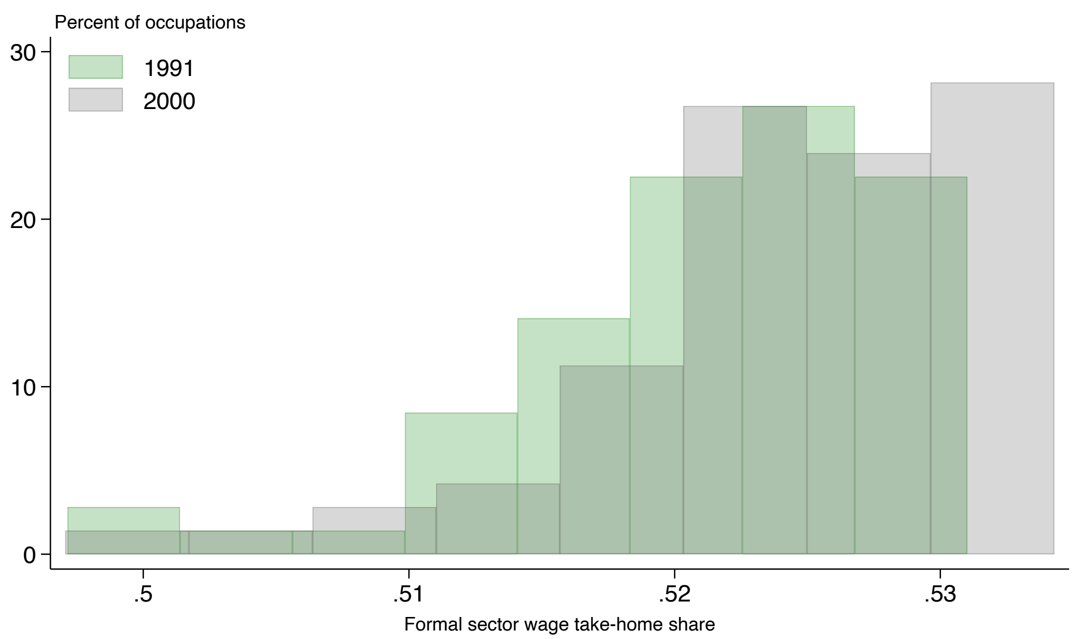
Notes: This figure shows binned scatters of 1991-2000 changes in expected informal-formal wage gaps, firm formal employment, and import tariff reductions. The sample includes all formal sector firms in RAIS in 1991 and 2000, merged at the microregion X demographic cell X year level to the corresponding census data. Each observation is a firm X microregion X demographic cell (gender X education group X age group dummies). All binned scatters show residual variation conditional on firm, microregion, and demographic cell fixed effects. Expected informal-formal wage gap is calculated separately by firm, microregion, and demographic cell as $p_{z\bar{s}r}\sigma^c_{z\bar{s}r}$. The first term, $p_{z\bar{s}m}$, is the probability of *involuntary* separation from firm z , measured as the ratio of firings to firings plus stayers in firm z. The second, $\sigma^c_{z\bar{s}m} \equiv (w^o_m - w^c_{z\bar{s}m})/w^c_{z\bar{s}m}$, is the monthly earnings gap between the local informal sector and firm z's formal wage for that demographic cell. Monthly earnings gaps are calculated by first converging all reported earnings to 2000 constant reais. $p_{z\bar{s}m}$ is calculated using separations between between January 1 and December 30 of each year. Firm total formal employment is measured as of December 31 of each year. Regional tariff reductions are microregion-level exposure to 1990-1994 import tariff reductions from Dix-Carneiro and Kovak (2017) and differ by demographic group.

Figure 6: Formal sector wage take-home share in dual labor markets

Panel A: Distribution across microregions



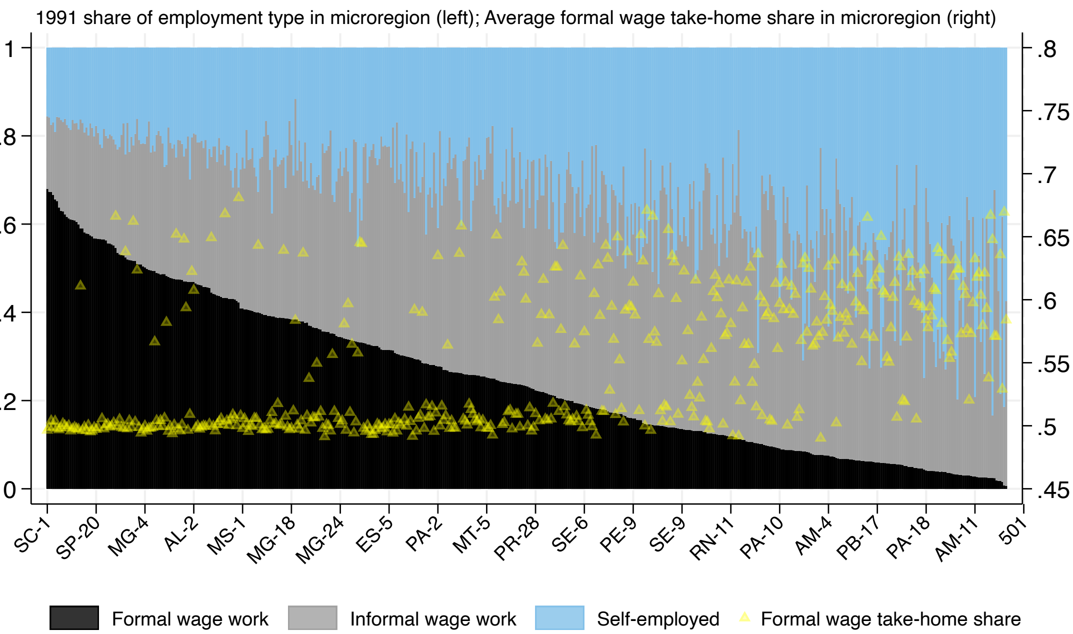
Panel B: Distribution across occupations



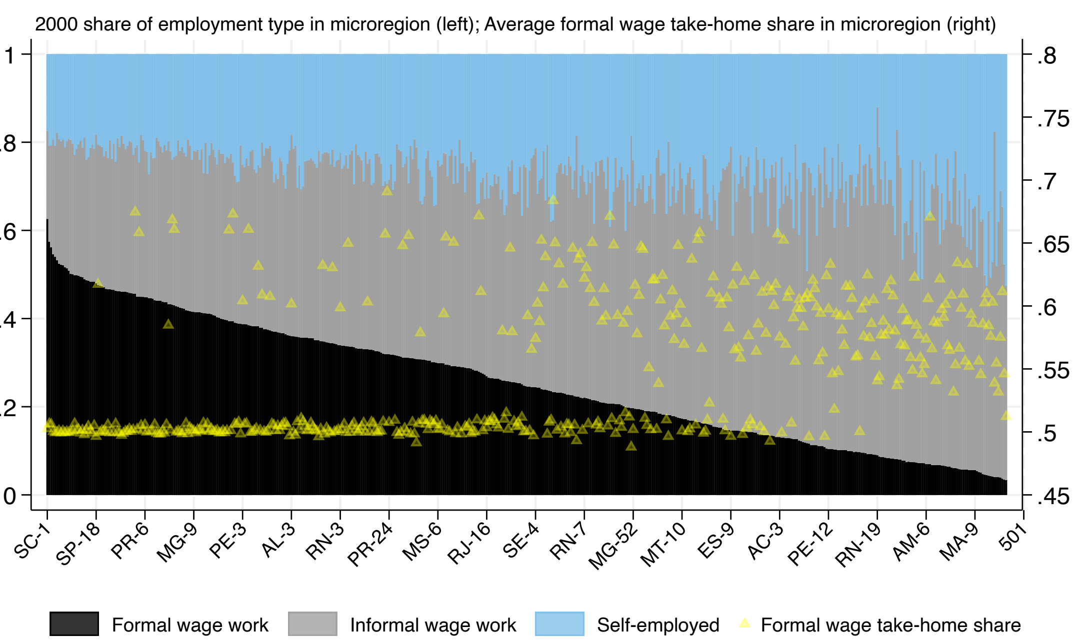
Notes: This figure show the distribution of average wage take-home shares (the inverse of average wage markdown) across microregions (Panel A) and occupations (Panel B), separately by year. Average wage take-home shares are calculated at the local labor market level (microregion x occupation) following equation 23, using region-specific estimates of $1/\tilde{\eta}$ and pooled estimates of $1/\tilde{\rho}$ and $1/\tilde{\theta}$. Panels A and B aggregate these estimates by microregion and 2-digit occupation, respectively, weighing observations by total formal wage bill.

Figure 7: Formal sector wage take-home shares and the composition of regional employment

Panel A: 1991



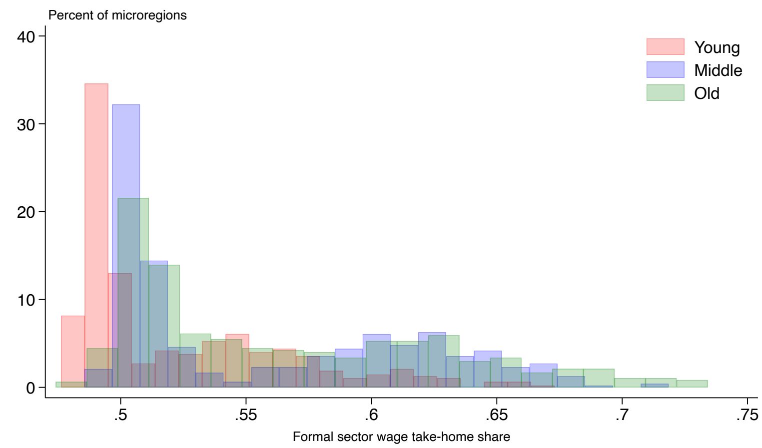
Panel B: 2000



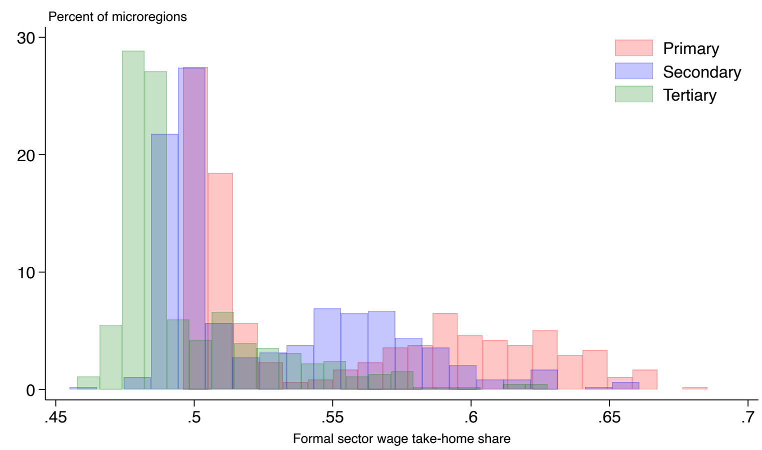
Notes: This figure plots the distribution of formal wage work (black bins, left y-axis), informal wage work (gray bins, left y-axis), self-employment (blue bins, left y-axis), average wage take-home shares across 486 microregions (yellow scatter, right y-axis). Labels in the x-axis correspond to different microregions. The first two letters indicate the micoregion's state, and the number indicates the microregion's rank in the state by formal sector share (e.g., "SC-1" = top-ranked microregion in Santa Catarina). See notes to Figure 6 for details and Appendix Figures A.4 through A.6 for demographic-specific distributions.

Figure 8: Formal sector wage take-home share in dual labor markets: Demographic heterogeneity

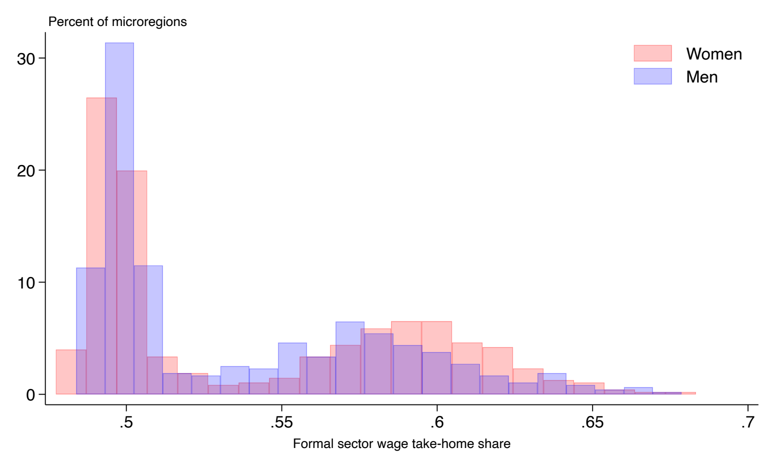
Panel A: By age



Panel B: By education

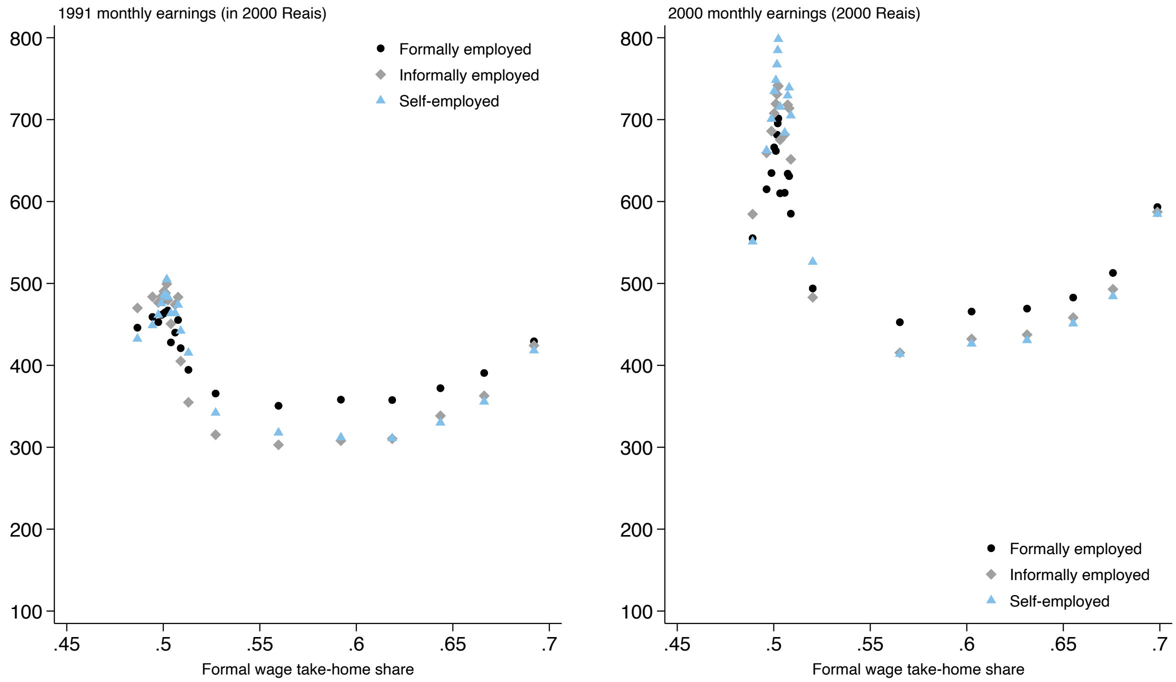


Panel C: By gender



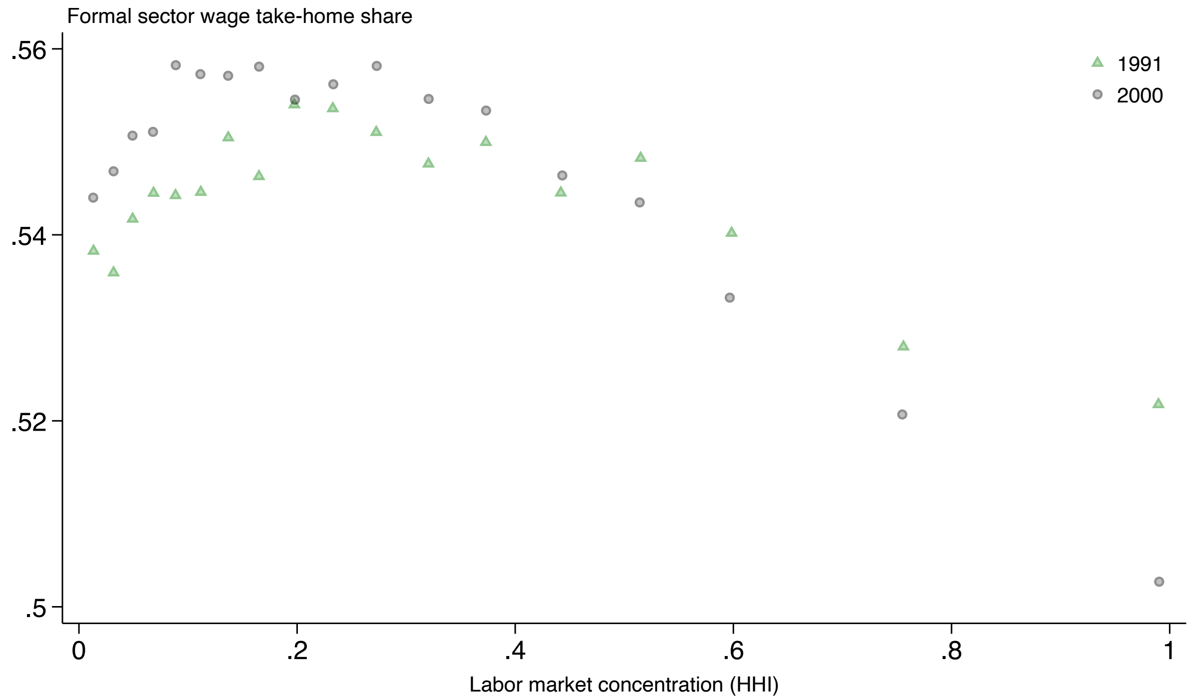
Notes: This figure show the distribution of average wage take-home shares (the inverse of average wage markdown) across microregions, separately by demographic groups and conditional on year fixed effects. Average wage take-home shares are calculated at the local labor market level (microregion x occupation) following equation 23, 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: Wages and formal sector wage take-home shares



Notes: This figure plots binned scatters between real monthly earnings (separately by employment type) and average formal sector wage take-home shares across microregions. See Figure 6 for details and Appendix Figures A.3 for demographic-specific binned scatters.

Figure 10: Concentration and formal sector wage take-home shares



Notes: This figure plots binned scatters between average formal sector wage take-home shares and formal firms' local labor market concentration (HHI), measured in terms of *expected* wage bill shares, which take into account probabilities of separation into informal wage work, across microregion x occupation pairs, separately by year. See Figure 6 for details and Appendix Figures A.4 through A.6 for demographic-specific distributions.

6 Additional Tables and Figures

Table A.1: Estimates of Ω and its sensitivity to unemployment insurance

	$\Delta \ln(1 + p_{z\bar{s}m}\sigma^c_{z\bar{s}m})$		
		IV	
	OLS	Without UI	With UI
	(1)	(2)	(3)
$\Delta \ln l^c_{z\bar{s}m}$	0.000443***	0.0326**	0.0217**
	(0.000135)	(0.0129)	(0.00853)
First-stage F		1.526	1.526
Observations	407862	394813	394813

Notes: This table shows OLS and IV estimates of Ω, the bias term in IV estimates of the within-market cross firm elasticity of substitution in the model of Felix (2022) if the extended model is true, and its heterogeneity by worker characteristics. The sample includes all formal sector firms in RAIS in 1991 and 2000, merged at the microregion X demographic cell X year level to the corresponding census data. Each observation is a firm X microregion X demographic cell (gender X education group X age group dummies). All regressions include firm, microregion, and demographic cell fixed effects. Expected informal-formal wage gap is calculated separately by firm, microregion, and demographic cell as $p_{z\bar{s}r}\,\sigma^c_{z\bar{s}r}$. The first term, $p_{z\bar{s}m}$, is the within-year probability of *involuntary* separation from firm z. The second, $\sigma^c_{z\bar{s}m} \equiv (w^o_r - w_{z\bar{s}m})/w_{z\bar{s}m}$, is the monthly earnings gap between the local informal sector w^o_r and firm z's formal wage for that market and demographic cell. Column (3) substitutes $w^o_{r,UI} = (4 * w_{z\bar{s}m} + 8 * w^o_r)/12$ for the informal sector wage to account for unemployment insurance benefits of 4-months salary. Monthly earnings gaps are calculated by first converging all reported earnings to 2000 constant Reais. $p_{z\bar{s}m}$ is calculated using separations between between January 1 and December 30 of each year. Firm total formal employment is measured as of December 31 of each year. The instrument for the IV estimate is the interaction between firm-level import reductions and regional tariff reductions, interacted with demographic group dummies and firm baseline formal employment. Regional tariff reductions are microregion-level exposure to 1990-1994 import tariff reductions from Dix-Carneiro and Kovak (2017) and differ by demographic group. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

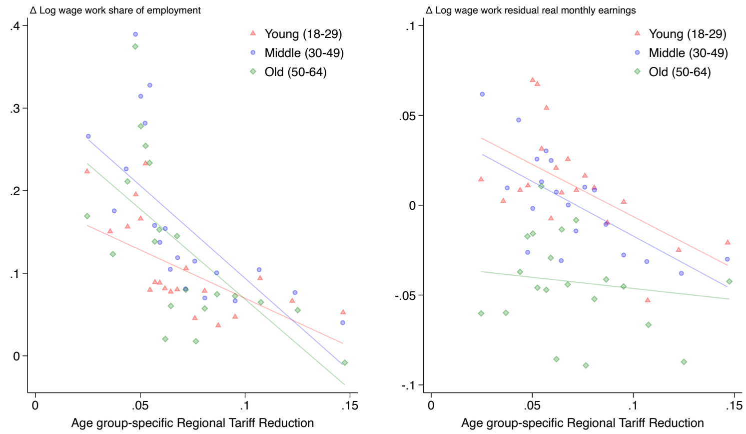
Table A.2: Robustness estimates of $1/\tilde{\rho}$ and $1/\tilde{\theta}$

	$\Delta \ln W_{\bar{s}m}$
	(1)
$\Delta \ln(L_{\bar{s}m}/L_m)$	0.120
	(0.133)
$\Delta \ln(L_m/L)$	1.111***
	(0.405)
First-stage F for $\Delta \ln(L_{\bar{s}m}/L_m)$	38.19
Anderson-Rubin F	49.92
First-stage F for $\Delta \ln(L_m/L)$	1.965
Anderson-Rubin F	2.454
Observations	478

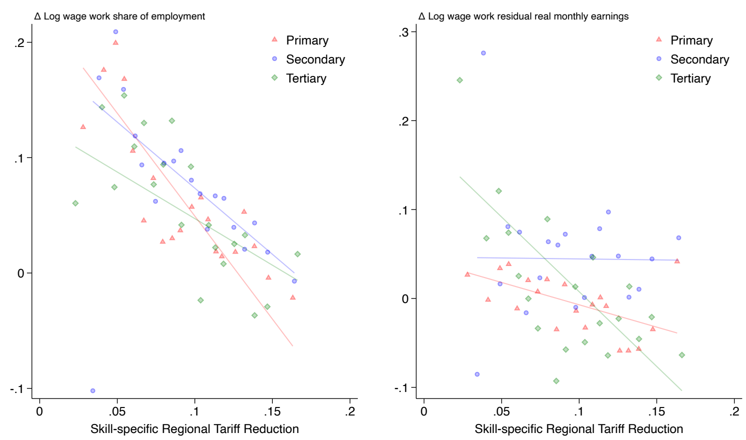
Notes: This table shows second stage estimates from an instrumental variables regression that estimates the inverse elasticity of substitution between wage work and self-employment $1/\tilde{\rho}$ as the coefficient on $\Delta \ln(L_{\bar{s}m}/L_m)$, and the cross-market elasticity of substitution $1/\tilde{\theta}$ as the coefficient on $\Delta \ln(L_{\bar{s}m}/L_m)$. The sample includes 486 microregions. The outcome variable is the 1991-2000 change in log residual real monthly earnings among individuals employed in wage work, either formally or informally, in each microregion. The dependent variables are (1) the 1991-2000 change in log share of wage work employment in a microregion; and (2) the 1991-2000 change in the log share of microregion total employment relative to national employment. The instruments are Regional Tariff Reductions interacted with (a) each microreigon's log maximum distance to the nearest labor law enforcement office, borrowing the estimation strategy from Ponczek and Ulyssea (2022); and (b) each microregion's 1991 share of employment that was formal. The regression includes region fixed effects for each of Brazil's five major regions and is weighted by each cell's 1991 microregion population. Log residual real monthly earnings are condition flexible controls for gender, education, and age. Real monthly earnings are based on the the IPCA deflator and are expressed in 2000 reais. Standard errors in parentheses are clustered by microregion. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Figure A.1: Effect of Regional Tariff Reductions on wage work by demographics

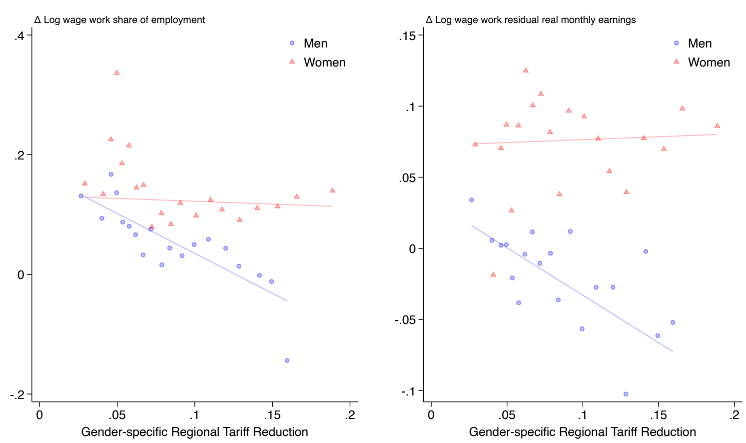
Panel A: By age



Panel B: By education



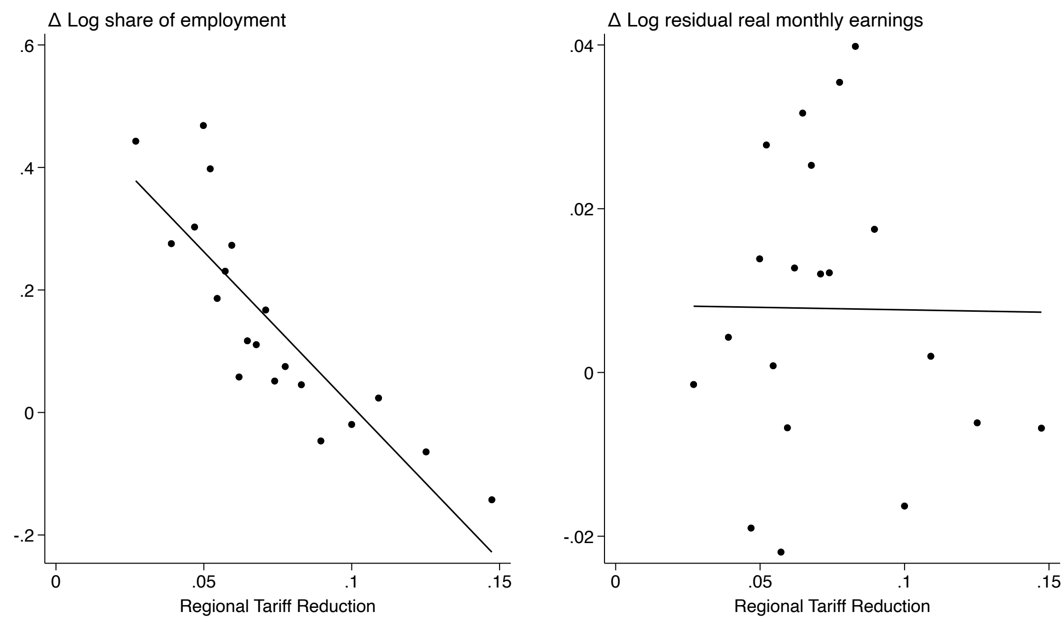
Panel C: By gender



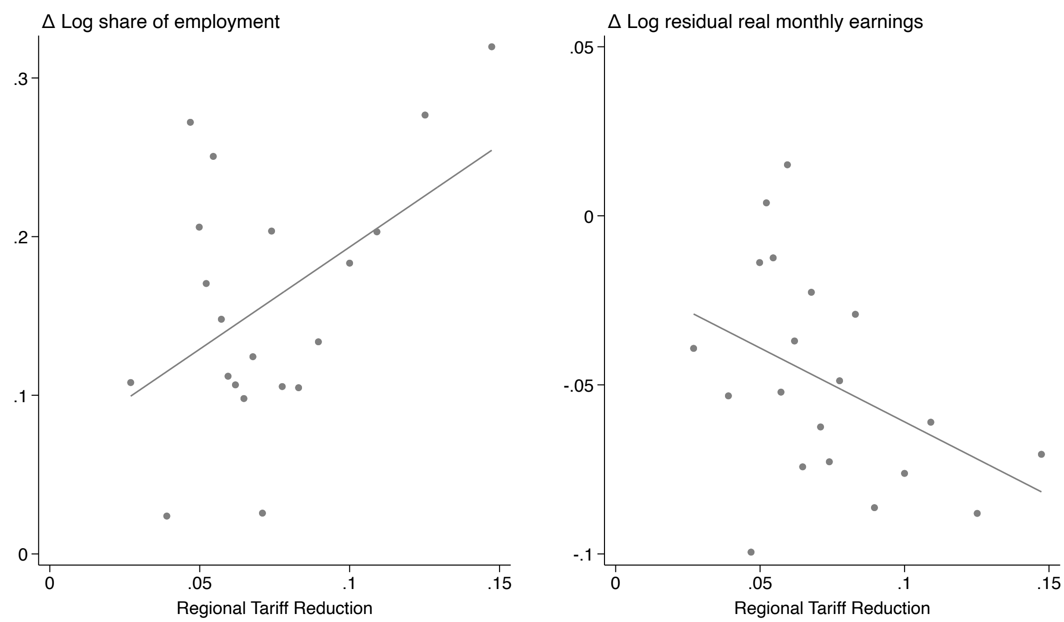
Notes: is figures shows the correlation between Regional Tariff Reductions (i.e., microregion-level exposure to 1990-1994 import tariff reductions from Dix-Carneiro and Kovak (2017)), changes in employment shares (left panel), and changes log residual real monthly earnings (right panel) between the 1991 and 2000 census, for self-employed individuals across 486 Brazilian microregions, separately for different demographic groups. Log residual real monthly earnings are conditional on a fully saturated vector of dummies for gender, age groups, and education groups. Real earnings are computed from nominal earnings by expressing all values in 2000 reais.

Figure A.2: Effect of Import Tariff Reductions on formal versus informal wage work

Panel A: Formal Wage Work



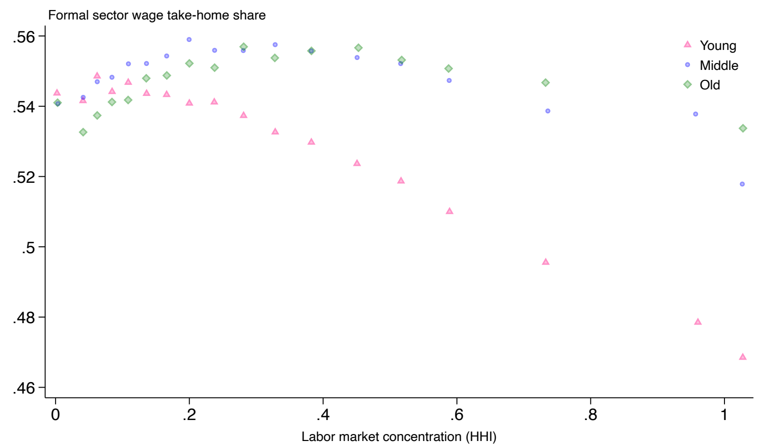
Panel B: Informal Wage Work



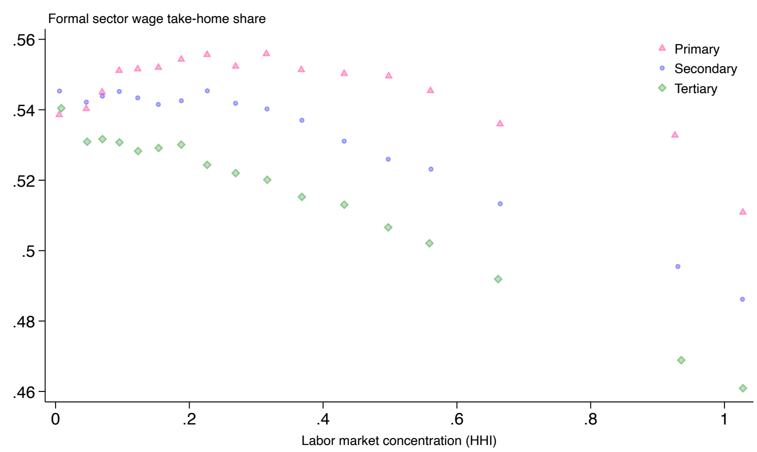
Notes: Each panel shows the correlation between Regional Tariff Reductions (microregion-level exposure to 1990-1994 import tariff reductions from Dix-Carneiro and Kovak (2017)), changes in employment shares (left) and log residual real monthly earnings (right) between 1991 and 2000 across 486 Brazilian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

Figure A.3: Formal sector wage take-home share and labor market concentration

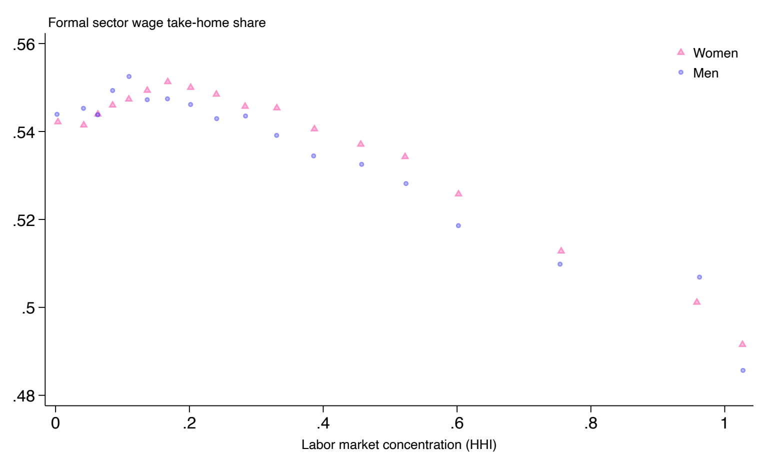
Panel A: By age



Panel B: By education



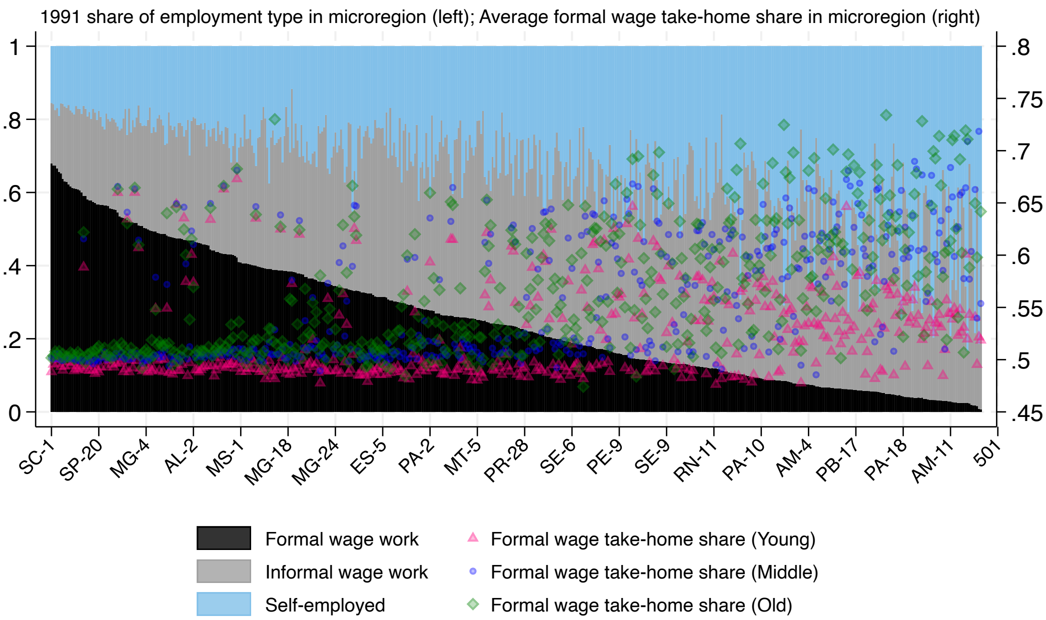
Panel C: By gender



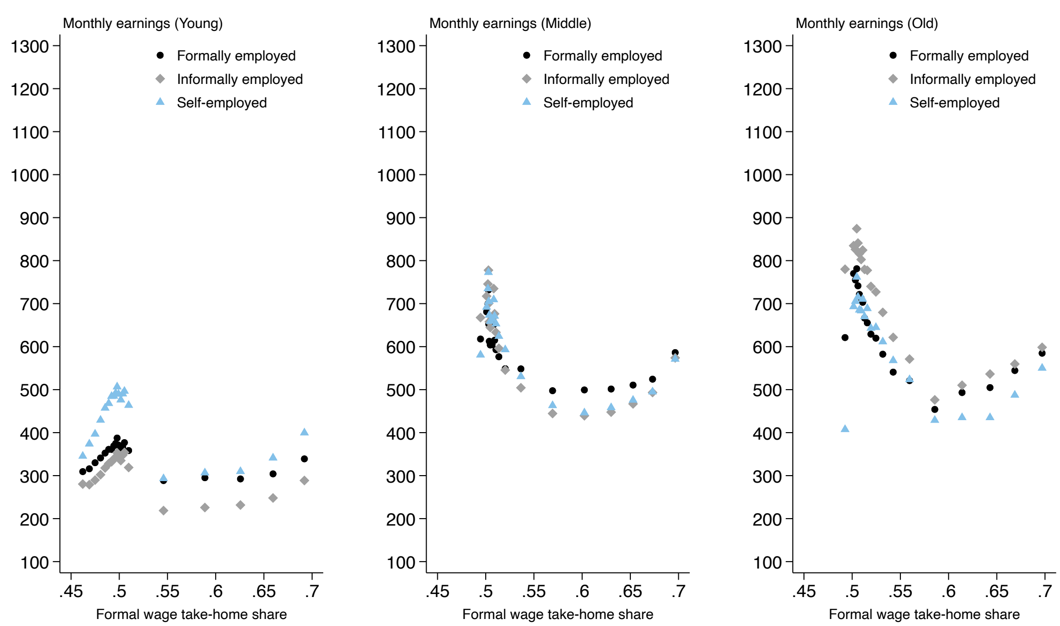
Notes: This figure plots binned scatters between demographic-specific average formal sector wage take-home shares and formal firms' local labor market concentration (HHI). See notes to Figure 10.

Figure A.4: Formal wage take-home share in dual labor markets: Heterogeneity by age

Panel A: 1991 distribution across microregions



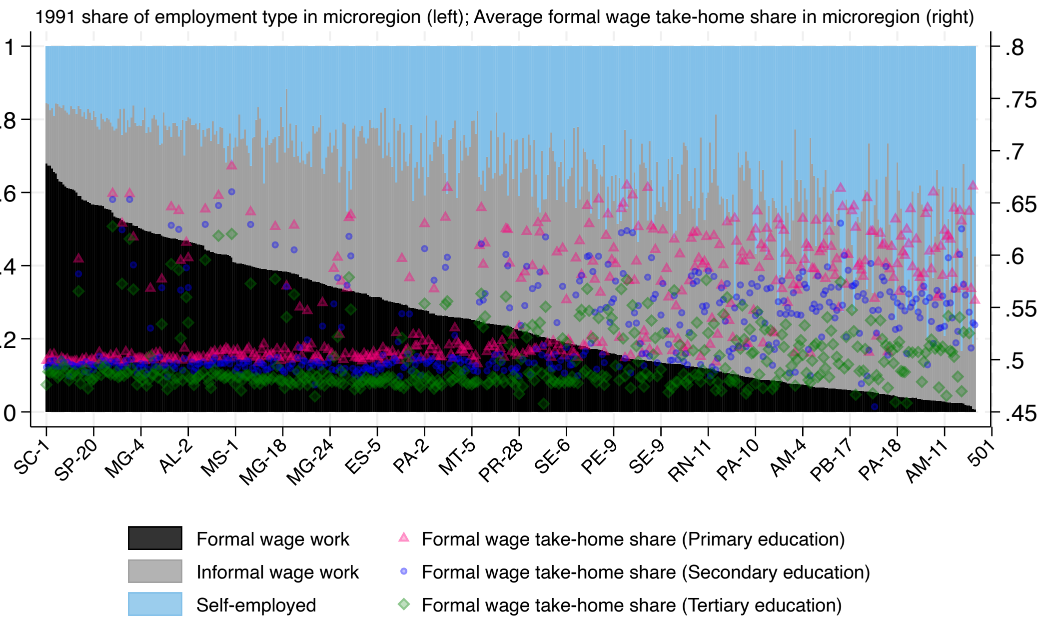
Panel B: Wages and wage take home-shares



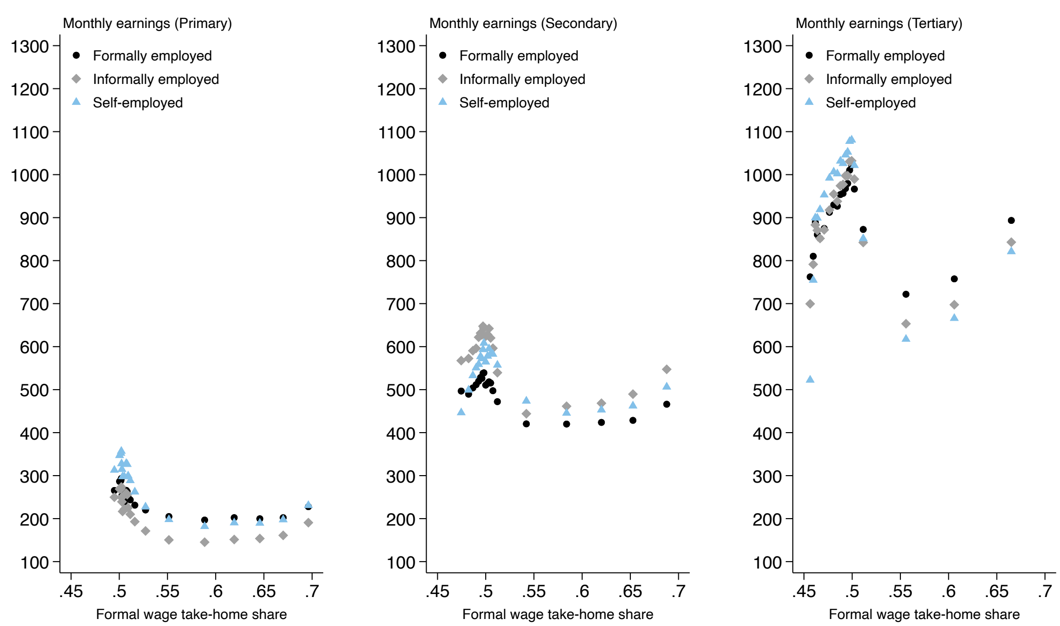
Notes: This figures age-specific versions of Figures 7 and 9. Average wage take-home shares are calculated at the local labor market level (microregion x occupation) following equation 23, using region-specific estimates of $1/\tilde{\eta}$, demographic-specific estimates of $1/\tilde{\rho}$, and the pooled estimate for $1/\tilde{\theta}$. For Panel A, average take-home shares are aggregated at the microregion level using formal sector wage bill as weights.

Figure A.5: Formal wage take-home share in dual labor markets: Heterogeneity by education

Panel A: 1991 distribution across microregions



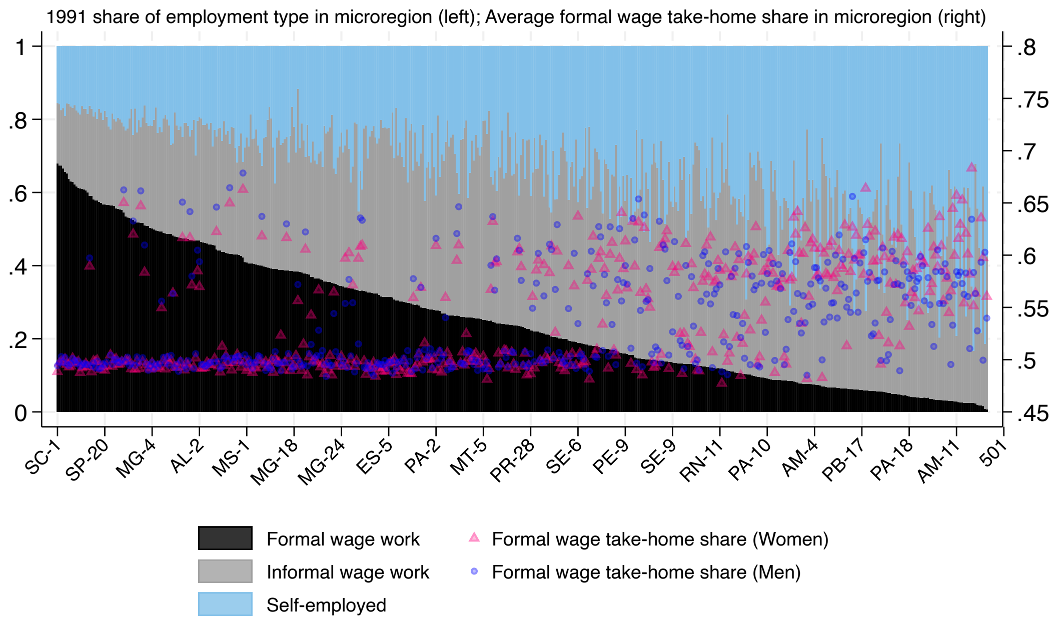
Panel B: Wages and wage take home-shares



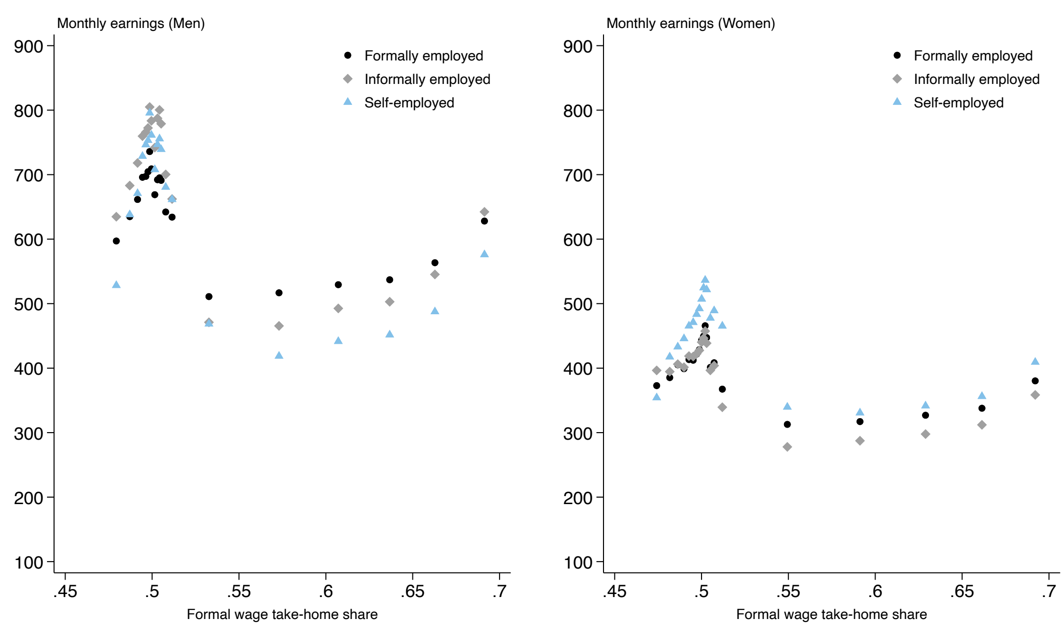
Notes: This figures education-specific versions of Figures 7 and 9. Average wage take-home shares are calculated at the local labor market level (microregion x occupation) following equation 23, using region-specific estimates of $1/\tilde{\eta}$, demographic-specific estimates of $1/\tilde{\rho}$, and the pooled estimate for $1/\tilde{\theta}$. For Panel A, average take-home shares are aggregated at the microregion level using formal sector wage bill as weights.

Figure A.6: Formal wage take-home share in dual labor markets: Heterogeneity by gender

Panel A: 1991 distribution across microregions



Panel B: Wages and wage take home-shares



Notes: This figures gender-specific versions of Figures 7 and 9. Average wage take-home shares are calculated at the local labor market level (microregion x occupation) following equation 23, using region-specific estimates of $1/\tilde{\eta}$, demographic-specific estimates of $1/\tilde{\rho}$, and the pooled estimate for $1/\tilde{\theta}$. For Panel A, average take-home shares are aggregated at the microregion level using formal sector wage bill as weights.