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Between sticky floors and glass ceilings: the effect of trade liberalization on double discrimination in Brazil

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Between sticky floors and glass ceilings: the effect of trade liberalization on double discrimination in Brazil

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Abstract

This article investigates how trade liberalization affects gender and racial pay inequalities in the short run. Guided by an intersectional perspective, we consider overlapping effects across gender, race, and wage levels. We exploit Brazil's trade liberalization process (1988–95) as a natural experiment. On average, liberalization increased wages of nonwhite women relative to men and white women. However, this average effect masks substantial heterogeneity. When we decompose pay gaps along the wage distribution, we find that liberalization reduced racial and gender discrimination at low wages, which mitigated preexisting 'sticky floors' by gender. In contrast, at the top of the distribution, liberalization increased racial discrimination, which reinforced existing 'glass ceilings' by race.

JEL-Classification: F13, F14, J15, J71.

Keywords: trade liberalization, wage inequality, intersectionality, gender, race

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1 Introduction

In developing countries, where segregated labor markets and diffused gains from trade are widespread, there is a growing interest in the implications of international trade for economic inequality. Whereas most studies have focused on the average effects of trade across countries, sectors, and skill levels,¹ increasing attention has been given to the effects of international trade for gender equality.² Yet, little is known about the heterogeneous effects of opening up to international trade along the wage distribution and from an intersectional perspective—that is, one that takes into account how identities of gender, race, class (and others) intersect and overlap to create complex patterns of social advantage and disadvantage (Crenshawt, 1989).

In this article, we investigate the consequences of Brazil's trade liberalization for gender and racial inequalities both on average and along the wage distribution. Between 1988 and 1995, Brazil drastically opened up to international trade. The main policy objective was to reduce and equalize import tariffs across sectors. As a result, cross-sectoral variation in tariff reduction is almost perfectly predicted by initial sectoral tariff levels. In short, conditional on the initial tariff level, tariff reductions between 1988 and 1995 were exogenous to local labor market conditions. Due to this feature, this episode of trade liberalization has been considered a close-to-ideal natural experiment and has been widely studied in the literature.³

We focus particularly on the effects for nonwhite women, a group which, so far, has received little attention in the literature. In her seminal article on intersectionality, legal scholar Kimberle Crenshawt posits, for the United States, that "[b]ecause the intersectional experience is greater than the sum of racism and sexism, any analysis that does not take

¹E.g., Wood (1998); Dollar and Kraay (2004); Ferreira *et al.* (2007); Egger and Kreickemeier (2009, 2012); Topalova (2010); Autor *et al.* (2013); Kovak (2013); Helpman *et al.* (2017).

²See, among others, Black and Brainerd (2004) and Autor *et al.* (2019) for the US; Berik *et al.* (2004) for Korea and Taiwan; Juhn *et al.* (2014) and Ben Yahmed and Bombarda (2020) for Mexico; Anukriti and Kumler (2019) for India; Kis-Katos *et al.* (2018) for Indonesia; and Gaddis and Pieters (2017) for Brazil.

³A non-exhaustive list includes Castilho *et al.* (2012); Kovak (2013); Dix-Carneiro and Kovak (2015, 2017); Gaddis and Pieters (2017); Braga (2018); Costa *et al.* (2018); Dix-Carneiro *et al.* (2018); Hirata and Soares (2020).

intersectionality into account cannot sufficiently address the particular manner in which Black women are subordinated" (Crenshawt, 1989, p. 140). However, very few papers on applied economics take intersectionality seriously (Lovell, 1994; Brewer *et al.*, 2002; Ruwanpura, 2008; Elu and Loubert, 2013; Weichselbaumer, 2020). Most articles study gender or racial inequalities in isolation. In Brazil, as in other former colonies engaged in the Atlantic slave trade, nonwhite women have been, throughout history, particularly disadvantaged in the labor market. In comparison to other gender-race groups, nonwhite women earn the lowest average wage, are less likely to be formally employed, and are over-represented in jobs with poor working conditions (Lovell, 1994; Soares, 2000; Ipea, 2011).

Theoretically, international trade is expected to affect gender and racial inequalities in the labor market through two channels. First, according to Gary Becker's theory of taste-based discrimination, increased competition forces firms to abandon discriminatory practices, because discriminating individuals of similar productivity is costly (Becker, 1957). Once competition increases, firms reduce discrimination to minimize costs, at the risk of being driven out of the market. Second, in the framework of the Hecksher-Ohlin model, trade will benefit sectors that employ the country's relatively most abundant factor. In labor markets with high levels of sectoral segregation, trade liberalization will have distributional consequences along gender and racial lines, thus creating winners and losers. These differential impacts of trade liberalization are further mediated by labor market conditions, such as the size of the informal market (Ben Yahmed and Bombarda, 2020), mobility across sectors, and wage rigidity (Pavcnik, 2017). While the Beckerian channel unambiguously predicts that exposure to liberalization reduces discrimination, the second channel is ambiguous, depending on the patterns of labor market segregation and on which sectors are more exposed to import competition.

The empirical evidence on gendered effects of liberalization is mixed. For the United States, Black and Brainerd (2004) show that exposure to competition in manufacturing industries reduces the gender wage gap. In Indonesia, Kis-Katos *et al.* (2018) find that exposure to tariff reductions on intermediate goods increases female labor force participation

and reduces the share of women primarily occupied with domestic tasks. For Mexico, Juhn *et al.* (2014) show that trade integration following the North American Free Trade Agreement (NAFTA) increased female productivity in blue collar jobs, mostly through technology diffusion. On the other hand, Ben Yahmed and Bombarda (2020) document that, following liberalization in Mexico, women became more likely to enter the informal service sector, while men were more likely to work formally in the manufacturing sector. Exploiting China's tariff liberalization, between 1990 and 2005, Wang *et al.* (2020) find that exposure to import competition reduces the gender employment gap.

For Brazil, studies on trade liberalization and gender and racial inequality mostly investigate medium to long-term average effects, exploiting the long difference between the 1991 and 2000 censuses. Gaddis and Pieters (2017) find that exposure to liberalization reduces both male and female average employment rates. Since the negative effects are larger for males, the gender gap in employment reduces in absolute terms. However, because men had higher initial employment rates, there was no reduction in the *relative* gender gap. In one of the few studies on the racial effects of liberalization, Hirata and Soares (2020) test Becker's (1957) model of taste-based discrimination among men. Consistent with Becker's hypothesis, the authors find that in regions more exposed to liberalization there was a reduction in the unexplained wage gap (i.e., discrimination) between white and nonwhite male workers. Their preferred estimate suggests that a tariff cut of 9.7 percentage points (the sample average in 1990–1995) reduces the racial wage gap among men by 18% between 1991 and 2000, with the effect persisting with a similar magnitude until 2010 (the year of the last available census). Whether similar effects occurred in the short run for the pay gap across gender-race groups and throughout the wage distribution remain open questions that we tackle in this article.

We complement the existing literature by focusing on short-term dynamics and by considering heterogeneity for all gender and racial groups, both on average and along the wage distribution. For identification, we follow the standard strategy in the literature (e.g., Topalova, 2010; Kovak, 2013; Dix-Carneiro and Kovak, 2017) and construct a regional measure of trade exposure based on pre-liberalization sectoral employment shares and

exogenous cuts in sectoral import tariffs over time. By covering the exact years of liberalization (1988–1995) and measuring labor market outcomes yearly, we capture the short-term effects of liberalization for men and women of different races. The identification strategy exploiting yearly variation in exposure to liberalization instead of long differences is similar to that of Erten *et al.* (2019), who study the labor market adjustments of liberalization in South Africa.

Because gender and racial pay gaps vary along the wage distribution, we decompose the impact of trade liberalization at different wage quantiles, using the method developed in Firpo *et al.* (2018). For gender, many studies from different contexts find 'glass ceilings'—larger pay gaps at the top of the distribution than at the median—or 'sticky floors'—larger pay gaps at the bottom of the distribution than at the median (Albrecht *et al.*, 2003; Arulampalam *et al.*, 2007; Chi and Li, 2008; Salardi, 2012; Carrillo *et al.*, 2014; Bertrand, 2018; Deshpande *et al.*, 2018). A major contribution of our paper is to connect the literature on gendered and racial effects of trade liberalization to the literature on pay gaps over the wage distribution.

Another contribution is to examine the dynamics of the liberalization period itself, 1988–1995. Most of the literature focuses on the period after 1996, once the process of tariff reduction was concluded. After the mid-1990s, poverty and inequality reduced sharply in Brazil (Ferreira *et al.*, 2008; Alvarez *et al.*, 2018), as in many Latin American countries (Gasparini and Lustig, 2011). Racial and gender occupational segregation have also declined substantially since 1996 in Brazil (Salardi, 2016). Yet, the macroeconomic environment before 1996 was remarkably different: poverty and inequality rose in the 1980s and early 1990s (Ferreira *et al.*, 2008), and occupational segregation by gender and race did not fall between 1987 and 1995 (Salardi, 2016). So far, the short-run effects of trade liberalization under the distinct pre-1996 economic regime remain largely overlooked.

We find that, on average, trade liberalization causes a immediate decrease in hourly wages for all gender-race groups. However, the effect on female wages becomes positive with one and two-year lags. More interestingly, this positive effect is significantly larger for nonwhite women. Among men, in contrast, the mean racial pay gap remains unaffected by tariff reductions. Because the initial negative effect was stronger for males and the succeeding positive effect was larger for nonwhite females, trade liberalization was, on average, mostly beneficial to nonwhite women. Overall, liberalization contributed to a reduction of 17.7% in the mean racial wage gap among women and a reduction of 7.25% in the mean gender wage gap between white (nonwhite) men and white (nonwhite) women.

We then decompose wage gaps between gender-race groups at several quantiles of the wage distribution. We find that trade liberalization reduces racial wage discrimination at the bottom half of the wage distribution, but increases discrimination at the upper half. These distributional effects suggest that, overall, liberalization reinforced existing 'glass ceilings' for nonwhite men and women alike. In contrast, trade liberalization reduced gender wage discrimination among whites and nonwhites alike, but only at the bottom quartile. Thus, liberalization mitigated preexisting 'sticky floors' for women, but had no effect for higher quantiles of the wage distribution. In comparison, the contribution of trade liberalization to the explained portion of the wage gap was quantitatively negligible for all gender-race pairs.

To study the short-term effects of trade liberalization by gender and race, we are restricted to the yearly national household survey (PNAD).⁴ The smallest geographical units available in the PNAD are the 26 federal states plus the federal district. We define a local labor market as a state-urban or state-rural cell, which leaves us with substantially less spatial variation when compared to the literature using microregions from the decennial censuses. However, the analysis at the state-urban-rural level has the advantage of including a large enough number of observations per gender-race group in each sector and regional cell, which is crucial for the estimation of intersectional effects.

We confirm that the results are robust to several sensitivity tests. A clear concern is selection into employment, a widely discussed issue in labor economics (e.g., Blau and Kahn, 2017; Machado, 2017). We take a closer look at employment by gender and race.

⁴More generally, the literature on the labor market effects of liberalization in Brazil uses a range of different data sources. Bosch *et al.* (2012) use the Monthly Employment Survey (Pesquisa Mensal de Emprego) and the PNAD to assess the effects of liberalization and labor market reforms on informality in Brazil. Krishna *et al.* (2014) exploit administrative linked employer-employee data from the Relação Anual de Informações Sociais (RAIS) to assess the wage-effects of liberalization.

Although liberalization differentially affects the employment rates of men and women of different races, our results are robust to the inclusion of a fine-grained measure of average employment for different demographic groups. Altogether, the evidence suggests that our results are not entirely driven by selection into employment.

The paper is structured as follows. Section 2 introduces the process of liberalization and facts about gender and racial inequality in Brazil. In section 3, we present the data and, in section 4, discuss the identification strategy and empirical specifications. In section 5, we report and discuss the results. Section 6 concludes.

2 Background

Trade liberalization Starting in 1988, Brazil initiated a comprehensive process of trade liberalization involving extensive reductions in import tariffs, elimination of discretionary controls, and overall reduction of non-tariff barriers (Kume *et al.*, 2003; Abreu, 2004). As in many other Latin American countries, Brazil's liberalization agenda occurred in a larger context of economic liberalization in the region, related to the Washington Consensus and the advancements of the negotiations of the Mercosur agreement (Castilho *et al.*, 2012; Gaddis and Pieters, 2017). Between 1988 and 1995, import tariffs decreased substantially across economic sectors, albeit at varying speeds. Appendix Figure A1 plots the effective tariff levels over time across sectors.⁵ Except for the automotive sector, whose tariffs oscillated throughout the period, the figure shows a general trend of rapidly declining import tariffs throughout the period.

The main objectives of the liberalization process were to reduce distortions in production and to equalize tariff levels across sectors. Following this logic, sectors that were initially highly protected experienced larger tariff cuts (Kovak, 2013). Figure 1 plots the change in effective tariff between 1987 and 1995 against the pre-liberalization tariff level for each sector at the 2-digit level.⁶ The strong negative correlation between the two measures and

 $^{^5\}mathrm{We}$ describe the data in detail in the next section.

⁶As discussed in further detail in section 4, to assess short-run effects of liberalization, our identification strategy relies on yearly variation in tariffs, rather than on long differences. Figures A2 and A3 plot the biannual changes in tariffs and the initial tariff levels with and without the automotive sector. Similarly



Figure (1) Changes in effective tariffs between 1987 and 1995 and pre-liberalization tariff levels

close to perfect fit confirm that tariff cuts were mostly determined by initial tariff levels. Kovak (2013) and Gaddis and Pieters (2017) emphasize, in addition, that the federal government was able to restrain protectionist interests and to put forward the liberalization process. Overall, the fact that the process of liberalization was mostly determined by policy, rather than by sectoral performance or economic interests, explains its suitability as a natural experiment.

Gender and race inequalities The Brazilian labor market exhibits substantial levels of segmentation and wage inequality by gender and race (Lovell, 1994; Salardi, 2016). Because historical race and gender inequalities accumulate and reinforce each other, labor market outcomes are particularly disadvantageous for nonwhite women. In addition to being overrepresented in marginalized sectors, nonwhite women work on average longer hours and receive lower wages (Lovell, 1994). Whereas these differences are partially explained by observable characteristics (e.g., education, experience), a substantial residual remains, which is often interpreted in the literature as the effect of discrimination. In our sample from the PNAD, nonwhite women earn the lowest average log hourly wage and are most

to Figure 1, the plots show that, with the exception of the automotive sector, tariff cuts were larger for sectors that were more protected before liberalization. As the liberalization process advanced, tariff cuts converged to zero.



Figure (2) Employment trends by gender and race, 1987–1996 Notes: Authors' calculation from PNAD, 1987–1996. Survey weights are used. The figure considers individuals aged 25–64.

likely to be working informally, as compared to all other gender-race groups (see Online Appendix Table A3). Using survey data for 2010, Layton and Smith (2017) show that nonwhite women are more likely to report having suffered gender discrimination than white women of similar socioeconomic backgrounds. Moreover, perceptions of discrimination by class, gender, and race are closely interlinked, with a measure of skin color coded by the interviewer being more predictive of discrimination by class than household wealth or educational attainment. As put by the authors, "race underlies discrimination even when respondents fail to perceived it as race-based" (Layton and Smith, 2017, p. 54).

What were the broader trends in labor market outcomes by gender and race during the trade liberalization period considered in this article? The first key trend is that women were rapidly joining the labor market. Between 1987 and 1996, the gender gap in employment rates fell by 10 percentage points (Figure 2), with very similar trends across racial groups. At the same time, however, occupational segregation by gender and race did not fall (Salardi, 2016). Overall, this was a period of rising poverty and inequality (Ferreira *et al.*, 2008).

In 1987, the gender wage gap among whites—conditional on age, education, number of children, state of residence and 2-digit employment sector—was 54%. The conditional racial wage gap among men was 14%. On top of both these effects, nonwhite women



Figure (3) Evolution of conditional pay gaps across gender-race groups. Reference group: white men

suffered an additional pay penalty of 5% relative to other groups. Figure 3 shows how these conditional pay gaps evolved over time. The gender wage gap falls, in absolute terms, to 41% in 1996, whereas the racial wage gap actually increases over time, reaching 17% in the last sample year. Lastly, the wage penalty for being nonwhite *and* woman shrinks after 1989, and actually turns positive in 1995 and 1996. In sum, gender pay gaps are much larger than racial pay gaps, but while the former decline over the liberalization period, the latter slightly increase. Nonwhite women's experience is not fully captured by the additive effects of being female and nonwhite; there is an additional intersectional effect, which is negative until 1992 and becomes positive afterwards.

Importantly, the *average* wage gap by gender or race hides substantial heterogeneity along the wage distribution. Studying the 1987–2006 period, Salardi (2012) finds that gender pay gaps exhibit both sticky floors and glass ceilings, although both phenomena become smaller over the period. Racial wage gaps reveal a persistent glass ceiling for nonwhite workers. By the early 2000s, urban Brazil had the highest gendered glass ceiling, conditional on education and experience, among the urban areas of 12 Latin American countries (Carrillo *et al.*, 2014). Figure 4 shows the unconditional wage gap across the

Notes: Authors' calculation from PNAD, 1987–1996. OLS estimates of women, nonwhite, and women \times nonwhite dummies with 95% confidence intervals from Mincerian regressions of log hourly wage on age (quadratic), education, number of children, state dummies, and 21 sectoral dummies. Regressions are estimated separately for each year. Survey weights from PNAD are used. Includes all workers aged 25–64 with positive earnings.



(a) Racial gap among men (white vs. nonwhite)

(b) Racial gap among women (white vs. nonwhite)





Figure (4) Raw wage gaps: log hourly wage difference between social groups Notes: Kernel-weighted local polynomial smooth plots with 95% confidence bands. Common Y-axis for all subfigures. Years

are 1987 and 1996. Survey weights from PNAD are used. Includes all workers aged 25-64 with positive earnings.

distribution for 1987 and 1996. Overall, the figure reveals clear glass ceilings by race and large sticky floors for nonwhite women. Between 1987 and 1996, sticky floors and glass ceilings by gender became less pronounced; racial glass ceilings, on the other hand, did not improve substantially.

The remainder of the paper will attempt to rigorously estimate if and how trade liberalization contributed to changing wage differentials across demographic groups over time, both on average and along the wage distribution.

3 Data

We combine two data sources in our empirical analysis. Data on import tariffs by economic sector come originally from Kume *et al.* (2003) and are compiled in Abreu (2004). We have yearly information on import tariffs for 20 2-digit sectors between 1987 and 1996. We make use of the information on the effective tariff rate, which considers both tariffs on final as well as on intermediate goods.⁷ All in all, data on import tariffs reflect the exposure of different sectors to import competition.

Individual-level data on labor market outcomes and socio-demographics come from the PNAD, a nationally representative yearly household survey.⁸ We use eight survey rounds, covering the period 1987 to $1996.^9$ For our purposes, the PNAD offers three main advantages. First, since it is a household survey, it includes a large sample of individuals irrespective of their employment status, including both formal and informal workers, self-employed, unemployed and inactive individuals. Second, the survey contains a large enough number of observations for each gender-race group by sector and state. This is critical, because our main objective is to estimate the heterogeneous effect of trade opening across gender and race. Third, by being conducted annually, the data allow us to cover the exact years of liberalization and to estimate its short-run effects on labor market outcomes. Despite its comprehensiveness, the PNAD has one main drawback: it is only representative at the state level, which is a relatively large unit of analysis as compared to other administrative units that have been used in the literature, such as microregions or municipalities. We further divide each state into rural and urban cells, as those cells are more homogeneous in terms of sectoral composition and labor market characteristics and increase the spatial variation in our analysis.

Two alternative data sources which have been previously used in the literature are the Demographic Censuses and the administrative records of the Relação Anual de Informações

⁷For methodological details on how effective tariff rates are calculated, see Kume *et al.* (2003).

⁸Before 2003, the PNAD did not include the rural areas of the Northern states of Acre, Amapá, Amazonas, Pará, Rondônia and Roraima.

⁹The PNAD was not conducted in 1991, which was a census year, and in 1994, due to budgetary reasons.

Sociais (RAIS). The main disadvantage of the census is its decennial time span 1991, 2000. Liberalization of import tariffs occurred mainly between 1990 and 1995, with a small reversion of the process after 1995 (see Online Appendix Figure A1). Measuring outcomes using the 2000 census, therefore, captures the net effect of liberalization and could be potentially confounded by other policy changes happening at the same period. Between 1995 and 1998, the Brazilian government started to register current account imbalances (Kume *et al.*, 2003). This was related to rising imports that resulted from tariff reductions and exchange rate appreciation following the monetary stabilization plan (Plano Real). Additionally, capital flight following the Mexican crisis of 1994 made it more difficult for the government to finance current account deficits. In this context, between 1995 and 1998, the Brazilian government increased import tariff rates in some sectors that were driving the increase in overall imports (Kume *et al.*, 2003, p. 18).

The other alternative data source would be the RAIS, a yearly administrative census, covering all workers employed in the formal sector. Although very comprehensive, the database excludes self-employed workers, informal workers and unemployed individuals. Additionally, information on employees' race was only introduced in the 2000s, so it is not available for the period analyzed in this paper. For the reasons discussed above, we believe the PNAD is the most suitable source of microdata for our empirical analysis. Next, we discuss our measure of local trade exposure as well as our empirical model.

4 Empirical strategy

Our identification strategy relies on a shift-share design, which is the standard design in the literature for estimating the causal effects of aggregate shocks on local labor markets (e.g., Topalova, 2010; Autor *et al.*, 2013; Kovak, 2013). Our measure of exposure to trade liberalization varies depending on pre-liberalization sectoral composition of employment across state-rural-urban cells and changes in sectoral tariffs over time. Intuitively, although sectoral tariff cuts occur at the national level, their differential impact across regions depends on pre-existing local sectoral shares (Castilho *et al.*, 2012; Gaddis and Pieters, 2017; Dix-Carneiro and Kovak, 2017). To assess the short-run effects of liberalization, we exploit yearly variation in trade exposure—a strategy used, among others, by Erten *et al.* (2019) for South Africa. We measure trade protection as:

$$TP_{dut} = \sum_{s=1}^{20} \frac{L_{dus}^0}{L_{du}^0} \times \pi_{st}$$
(1)

where d denotes state, u is a urban/rural indicator, t denotes year, and s denotes sector. L_{dus}^{0} is the sectoral employment in a state-urban-rural cell in 1987, before liberalization started. L_{du}^{0} is overall employment in a state urban-rural area, also in 1987. π_{st} is the effective tariff level in sector s and year t. When computing sectoral employment shares, we follow Gaddis and Pieters (2017) and exclude the nontradable sector, because its implicit tariff is always zero. We assume, however, that prices in the tradable sector are transmitted to the nontradable sector (Kovak, 2013; Gaddis and Pieters, 2017). Thus, while excluded from the variable TP, workers in the nontradable sector are included in the regression analysis. The higher the value of TP_{dut} , the higher the level of trade protection in a local labor market. Accordingly, tariff reduction corresponds to a fall of TP_{dut} over time.

The validity of the shift-share strategy for causal identification relies on the assumption that either the shifts—here, changes in sectoral tariffs—or the shares—here, sectoral employment shares—are exogenous (Borusyak *et al.*, 2018; Goldsmith-Pinkham *et al.*, 2020). While assuming exogeneity of employment shares is unrealistic, prior literature has convincingly argued that the 1988–1995 cuts in import tariffs in Brazil are exogenous to local labor market conditions.¹⁰

In the Online Appendix, we plot yearly tariff changes across sectors, with and without the automotive sector (Figures A2 and A3).¹¹ Initially, tariff cuts were larger for sectors that were heavily protected in 1987, with a stronger negative correlation between 1987 and 1992. Afterwards, yearly tariff changes slowly approached zero, indicating the completion

¹⁰An incomplete list of papers using this identification assumption for Brazil's trade liberalization includes: Castilho *et al.* (2012); Kovak (2013); Dix-Carneiro and Kovak (2017); Gaddis and Pieters (2017); Braga (2018); Costa *et al.* (2018); Dix-Carneiro *et al.* (2018).

¹¹We show later that the results are robust to excluding the automative sector altogether.

of the process of liberalization. Overall, the figure with yearly variation also corroborates the hypothesis that the main objective of the liberalization process was to equalize tariff levels, without being susceptible to substantial protectionist interests by certain groups.

We estimate the reduced-form model:

$$ln(Y_{idsut}) = \beta_1 F_{idsut} + \beta_2 N_{idsut} + \beta_3 F_{idsut} \times N_{idsut} + \beta_4 T P_{dut} + \beta_5 T P_{dut} \times F_{idsut} + \beta_6 T P_{dut} \times N_{idsut} + \beta_7 T P_{dut} \times F_{idsut} \times N_{idsut} + \lambda X_{idut} + \delta_t + \gamma_{du} + \phi_s +$$
(2)
$$\eta_d t + \epsilon_{idsut}$$

where the dependent variable is the logarithm of the deflated hourly wage of individual iliving in state d, rural or urban area u, employed in sector s at time t. TP is the trade protection measure for each state-urban-rural area regressed separately with lags up to 5 years. F is a female dummy; N is a nonwhite dummy equal to one if the individual self-declares as black (*preto*) or brown (*pardo*). X is a vector of individual controls: age, age squared, educational attainment and number of children. We also include year fixed effects δ_t , state-urban-rural fixed effects γ_{du} , sector fixed effects ϕ_s , and state trends $\eta_d t$. These fixed effects and state trends account for yearly shocks that commonly affected the Brazilian labor market, historical factors that are constant over time for each state-urbanrural area or 2-digit sector, and differential wage trends for each state. Standard errors are clustered at the state-urban-rural level.

We restrict our sample to working individuals aged 25–64, surveyed between 1987 and 1996. The number of observations ranges from 424,252 to 549,716, depending on the lag structure of the trade protection variable. See Online Appendix A.1 for variable definitions and Table A2 for summary statistics of the estimation sample. The average log hourly wage in our sample is around 1.55 BRL; 45% of the individuals self-identify as nonwhites, and 34% are women. Tables A4-A7 in the Online Appendix show the sectoral occupation for the gender-race groups in our sample.

The main coefficients of interest are β_4 —the effect of trade liberalization on white male wages; β_5 —the differential effect of liberalization on female wages; β_6 —the differential effect of liberalization on nonwhite wages; and β_7 —the additional effect of liberalization on the wages of nonwhite women.

The time lag between a tariff cut, increased market competition, and firm decisions will depend on many unobservable factors, such as labor market rigidities and international trade frictions. Because, beforehand, we are unsure about the most appropriate time lag for the tariff protection variable, we transparently run regressions with up to five lags of TP. Because the lagged variables are highly correlated from one year to the next, we do not include the full lag structure simultaneously as our main specification.¹² Instead, we introduce time lags separately. As such, the lagged coefficients of TP can be understood as cumulative effects over previous periods. Overall, this model specification allows us to assess the short-term dynamics of liberalization for the gender-race groups.

5 Results

5.1 Average effects

We start by estimating the average effect of trade protection on individual wages. The dependent variable is the logarithm of deflated hourly wages. By normalizing with respect to hours worked, this variable comes closest to the concept of pay discrimination.¹³ The point estimates are presented in Table 1, with column (1) showing the contemporaneous effect of trade protection, and columns (2) to (6) showing the lagged effects up to five years before the wage measurement.

In line with the stylized facts presented in section 2, the estimated conditional wage gaps by gender and race are large. Women earn, on average, 43%–47% less than comparable males in the same sector of employment. Nonwhites earn 15%–18% less than comparable whites. For nonwhite women, the interaction coefficient (Nonwhite × Women) is always statistically insignificant. This suggests that the disadvantage of this group, which obtains

 $^{{}^{12}}$ cor $(TP_t, TP_{t-1})=0.9267$ and cor $(TP_t, TP_{t-2})=0.7809$. Additional results including all lagged variables simultaneously are shown in the Online Appendix.

 $^{^{13}{\}rm We}$ acknowledge, however, that not accounting for individual productivity or ability in the wage regression makes the interpretation of discrimination questionable.

the lowest hourly wage, arises entirely from the cumulative effect of two identities—being female and being nonwhite—rather than through their interaction. Later on, however, when decomposing pay gaps across the wage distribution, we show that nonwhite women suffer a large unexplained pay loss at the bottom of the distribution, suggesting a negative interaction effect of being nonwhite and women at low-paying jobs. These nuanced findings highlight the importance of understanding gender and race inequalities from an intersectional perspective.

Turning to the effects of trade liberalization, Figure 5 plots the marginal TP coefficients by gender-race group across different time lags. At first, a decline in trade protection reduces hourly wages of all gender-race groups (Figure 5a). With a lag of one and two years, however, tariff cuts increase wages of women (Figure 5b). For nonwhite women, there is an additional increase in wages that remains significant and sizable up to three years (Figure 5d). The estimates for nonwhite men are always insignificant (Figure 5c).

As a whole, in the short run, liberalization reduced average gender and racial inequality in Brazil. At first, the contemporaneous negative wage effect was larger for males than for women. A one SD decline in TP_t (≈ 1.19) reduces male hourly wages by approximately 2.15%, whereas female hourly wages fall by only 1.49% (Table 1, column (1)). Afterwards, lagged tariff cuts do not significantly affect wages for men, but significantly increase wages for women, and especially so for nonwhite women. A one SD decline in trade exposure TP_{t-1} (TP_{t-2}) increases hourly wages of white women by 1.24% (2.25%) and hourly wages of nonwhite women by 2.71% (4.7%) (Table 1, columns (2) and (3)).¹⁴ The larger increase in wages of nonwhite women one and two years after liberalization contributes to a reduction of 17.7% in the mean racial wage gap among women and a reduction of 7.25% in the mean gender wage gap between white (nonwhite) men and white (nonwhite) women.

¹⁴The standard deviation (SD) of TP_{t-1} (TP_{t-2}) is 1.349 (2.125); see Table A2.

		Log(hourly wage)						
	(1)	(2)	(3)	(4)	(5)	(6)		
	au = t	$\tau = t - 1$	au = t - 2	au = t - 3	$\tau = t - 4$	$\tau = t - 5$		
Women	-0.4709^{***}	-0.4533^{***}	-0.4369^{***}	-0.4494^{***}	-0.4408***	-0.4422^{***}		
	(0.0174)	(0.0196)	(0.0141)	(0.0151)	(0.0216)	(0.0189)		
Nonwhite	-0.1531^{***}	-0.1528^{***}	-0.1523^{***}	-0.1635^{***}	-0.1845^{***}	-0.1874^{***}		
	(0.0123)	(0.0123)	(0.0111)	(0.0104)	(0.0157)	(0.0159)		
Nonwhite \times Women	-0.0034	0.0078	0.0186	0.0229	0.0099	0.0219		
	(0.0198)	(0.0206)	(0.0164)	(0.0156)	(0.0230)	(0.0204)		
TP_{τ}	0.0181^{**}	0.0120	-0.0104	-0.0245	-0.0387	0.0022		
	(0.0090)	(0.0091)	(0.0178)	(0.0180)	(0.0240)	(0.0221)		
$TP_{\tau} \times Women$	-0.0056**	-0.0092^{***}	-0.0106^{***}	-0.0016	-0.0010	-0.0004		
	(0.0025)	(0.0032)	(0.0025)	(0.0020)	(0.0052)	(0.0033)		
$TP_{\tau} \times Nonwhite$	-0.0018	-0.0021	-0.0034	0.0006	0.0078	0.0060^{*}		
	(0.0039)	(0.0032)	(0.0032)	(0.0024)	(0.0047)	(0.0032)		
$TP_{\tau} \times Nonwhite \times Women$	-0.0075	-0.0109^{*}	-0.0115^{***}	-0.0079^{**}	-0.0002	-0.0031		
	(0.0054)	(0.0062)	(0.0040)	(0.0036)	(0.0066)	(0.0042)		
N	549716	487338	424252	359430	292574	292574		
Year FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		
State-Urban FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		
State trends	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		
Sector FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		
Individual controls	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		

Table (1) Trade protection and hourly wage

Notes: * p < 0.10, ** p < 0.05, *** p < 0.01. Ordinary least squares (OLS) estimates reported with robust standard errors clustered at the state-urban level. Tariff protection measured at time t - 5, t - 4, t - 3, t - 2, t - 1, and t. The sample includes individuals aged 25 to 64. Period is 1987–1996. Survey weights from PNAD are used. Control variables include age, squared age, educational attainment and number of children.



Figure (5) Trade protection and hourly wage. TP coefficients

5.2 Selection into employment

A well-known issue in the gender wage gap literature is selection into employment (Gronau, 1974; Heckman, 1979). Because wages are only observed for the employed, and women's labor force participation rates are much lower than men's, selection on unobservables in the participation decision will bias the estimated coefficients of wage gap regressions. In our case, the effect of trade liberalization on wage gaps across social groups could be driven (to an unknown extent) by changes in selection into employment over time.

Unfortunately, there is no consensual econometric fix for the selection problem, with different correction methods producing disparate results (e.g., Blau and Kahn, 2017; Machado, 2017). Most of the existing correction methods rely on stringent assumptions on positive or negative selection patterns, whereas, in reality, the selection process is unknown

Notes: The figure plots the marginal TP coefficients by gender-race group across different time lags. Point estimates shown with 95% confidence intervals.

and may even be heterogeneous across population groups (Neal, 2004).

In the absence of an econometric fix, we gather two pieces of evidence that seem inconsistent with a fully selection-driven story. First, we directly estimate the effects of trade liberalization on employment for the different gender and racial groups. Second, we re-estimate the baseline models, but now controlling for gender-race-cohort employment rates that vary by state-urban-rural area and year.

In Table 2, we estimate the effects of trade liberalization on the employment probability of the different population groups. At first, liberalization decreases employment. After two and three years, however, less tariff protection is associated with higher employment probability. The magnitude of these effects differs between the groups with negative effects of liberalization being strongest for males, followed by nonwhites females and, finally, by white females. These short-run estimates are consistent with the longer term effects by gender reported by Gaddis and Pieters (2017). While they estimate that one SD decline in TP reduces female employment by 1.03 pp and male employment by 2.92 pp, we estimate a drop of 2.19 pp in nonwhite female employment, 1.24 in white female employment, and 2.45 pp in male employment (column (1)). Thus, while the short-run effects are generally larger, the relative effects by group are very similar to those of Gaddis and Pieters (2017). After two and three years, however, we find evidence of an increase in employment particularly strong for white females. From columns (3) and (4), liberalization increases employment of white females by 2.21 pp (4.28 pp) and employment of nonwhite females by 0.62 pp (2.35 pp) with a lag of two and three years.

Under the simplifying neoclassical assumption that the first jobs to be cut are those with the lowest marginal productivity, we would expect that more jobs are lost among the lowest paid men than among the lowest paid women. This process would upward bias the gender wage gap estimate. In contrast, we find that trade liberalization reduces the gender wage gap, which is at odds with a purely trade-induced selection effect. Among women, however, if more jobs are lost among the lowest paid nonwhites than among whites, the racial wage gap estimate is downward biased. Indeed, we find that trade liberalization increases nonwhite women wages relative to those of white women. Selection

		Worked in the ref week					
	(1)	(2)	(3)	(4)	(5)	(6)	
	$\tau = t$	$\tau = t - 1$	$\tau = t - 2$	$\tau = t - 3$	$\tau = t - 4$	$\tau = t - 5$	
Women	-0.4102^{***}	-0.4015***	-0.3915***	-0.3822***	-0.3870***	-0.3855***	
	(0.0177)	(0.0165)	(0.0162)	(0.0159)	(0.0248)	(0.0227)	
Nonwhite	0.0236^{***}	0.0227^{***}	0.0219^{***}	0.0219^{***}	0.0227^{**}	0.0231^{***}	
	(0.0066)	(0.0060)	(0.0060)	(0.0061)	(0.0090)	(0.0085)	
Nonwhite \times Women	-0.0225^{*}	-0.0220^{*}	-0.0245^{*}	-0.0297^{**}	-0.0412^{*}	-0.0347^{*}	
	(0.0133)	(0.0130)	(0.0134)	(0.0142)	(0.0214)	(0.0194)	
$\mathrm{TP}_{ au}$	0.0206^{***}	0.0168^{***}	-0.0028	-0.0094^{***}	-0.0113^{**}	0.0012	
	(0.0038)	(0.0040)	(0.0036)	(0.0026)	(0.0051)	(0.0040)	
$TP_{\tau} \times Women$	-0.0102^{***}	-0.0106^{***}	-0.0104^{***}	-0.0099***	-0.0051	-0.0039	
	(0.0034)	(0.0032)	(0.0027)	(0.0026)	(0.0066)	(0.0041)	
$TP_{\tau} \times Nonwhite$	0.0005	0.0007	0.0009	0.0004	-0.0010	-0.0006	
	(0.0012)	(0.0010)	(0.0009)	(0.0010)	(0.0020)	(0.0013)	
$TP_{\tau} \times Nonwhite \times Women$	0.0080**	0.0075^{***}	0.0075^{***}	0.0087***	0.0142^{**}	0.0077**	
	(0.0031)	(0.0025)	(0.0024)	(0.0027)	(0.0055)	(0.0036)	
N	863041	765278	666731	565745	461497	461497	
Year FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	
State-Urban FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	
State trends	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	
Individual controls	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	

 Table (2)
 Trade protection and employment

Notes: * p < 0.10, ** p < 0.05, *** p < 0.01. Ordinary least squares (OLS) estimates reported with robust standard errors clustered at the state-urban level. Tariff protection measured at time t - 5, t - 4, t - 3, t - 2, t - 1, and t. The sample includes individuals aged 25 to 64. Period is 1987–1996. Survey weights from PNAD are used. Control variables include age, squared age, educational attainment and number of children.

into employment could be driving this effect, although it is unlikely to drive the reduction in average wage gaps between nonwhite women and men.

We then try to explicitly model changing employment rates for the different social groups. We assign to each individual the average employment share of her/his gender, race, and 5 year age cohort, in the state and urban-rural area of residence and survey year. By controlling for this variable, we purge the variation in wages that is systematically related to the evolution of employment rates by different demographic groups over time and across cohorts.¹⁵ As shown in Table 3, the coefficient of the employment share is positive and significant: on average, a 10 percentage point increase in the predicted employment rate, is associated with a 0.011 to 0.031 log point increase in hourly wages. Reassuringly, the estimated effects of trade liberalization remain qualitatively similar.

In sum, selection into employment is an important caveat of this article, but, from a quantitative perspective, our evidence suggests that it is not a first-order concern. However, we cannot exclude the possibility that part of the wage-effect of trade liberalization is operating via the employment margin.

One question that arises from the results presented so far is whether liberalization had differential effects beyond the mean. In what follows, we decompose the effect of tariff reductions at different points of the wage distribution to investigate if trade liberalization affected inequality and discrimination between gender-race groups.

5.3 Decompositions

We decompose the gap in log hourly wages (Y_g) between two mutually exclusive groups, g = A, B, at the τ th quantile of the unconditional wage distribution $(Q_{g,\tau})$ as¹⁶

$$\Delta_O^\tau \equiv Q_{B,\tau} - Q_{A,\tau} = \Delta_X^\tau + \Delta_U^\tau \tag{3}$$

 $^{^{15}}$ In total, there are 13,436 cells defined by gender, race, 5-year age cohort, state-urban-rural area, and year. The employment share across cells has mean 0.658, standard deviation of 0.291, and ranges between the whole unit interval, 0–1. In the most restricted estimation sample, there are, on average, 32 individuals per cell.

¹⁶With a few exceptions, we follow the notation in Fortin *et al.* (2011).

		Log(hourly wage)						
	(1)	(2)	(3)	(4)	(5)	(6)		
	$\tau = t$	$\tau = t - 1$	$\tau = t - 2$	$\tau = t - 3$	$\tau = t - 4$	$\tau = t - 5$		
Women	-0.3622***	-0.3562***	-0.3584***	-0.3902***	-0.4029***	-0.4021***		
	(0.0343)	(0.0365)	(0.0291)	(0.0296)	(0.0349)	(0.0329)		
Nonwhite	-0.1513^{***}	-0.1509^{***}	-0.1503^{***}	-0.1619^{***}	-0.1839^{***}	-0.1867^{***}		
	(0.0120)	(0.0123)	(0.0111)	(0.0105)	(0.0157)	(0.0159)		
Nonwhite \times Women	0.0036	0.0131	0.0228	0.0269^{*}	0.0137	0.0252		
	(0.0183)	(0.0195)	(0.0157)	(0.0146)	(0.0215)	(0.0192)		
TP_{τ}	0.0144^{*}	0.0096	-0.0097	-0.0231	-0.0369	0.0019		
	(0.0084)	(0.0091)	(0.0175)	(0.0178)	(0.0240)	(0.0222)		
$TP_{\tau} \times Women$	-0.0019	-0.0061^{**}	-0.0082^{***}	0.0001	-0.0004	0.0001		
	(0.0026)	(0.0027)	(0.0027)	(0.0018)	(0.0049)	(0.0032)		
$TP_{\tau} \times Nonwhite$	-0.0016	-0.0020	-0.0034	0.0007	0.0081^{*}	0.0062^{*}		
	(0.0038)	(0.0032)	(0.0032)	(0.0024)	(0.0046)	(0.0032)		
$TP_{\tau} \times Nonwhite \times Women$	-0.0107^{**}	-0.0132^{**}	-0.0132^{***}	-0.0094^{***}	-0.0017	-0.0040		
	(0.0047)	(0.0056)	(0.0037)	(0.0031)	(0.0060)	(0.0039)		
Employment share	0.3099^{***}	0.2806^{***}	0.2321^{***}	0.1784^{***}	0.1123^{*}	0.1194^{*}		
	(0.0816)	(0.0776)	(0.0690)	(0.0661)	(0.0598)	(0.0606)		
N	549716	487338	424252	359430	292574	292574		
Year FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		
State-Urban FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		
State trends	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		
Sector FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		
Individual controls	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		

Table (3) Trade protection and hourly wage. Selection into employment

Notes: * p < 0.10, ** p < 0.05, *** p < 0.01. Ordinary least squares (OLS) estimates reported with robust standard errors clustered at the state-urban level. Tariff protection measured at time t - 5, t - 4, t - 3, t - 2, t - 1, and t. The sample includes individuals aged 25 to 64. Period is 1987–1996. Survey weights from PNAD are used. Control variables include age, squared age, educational attainment, number of children and the average employment share of the individual's gender, race and 5 year age cohort, in the state and urban-rural area of residence and survey year.

where Δ_X^{τ} is the composition effect (or explained term), which is the part of the gap explained by differences in the distribution of covariates X between the two groups; and Δ_U^{τ} is the unexplained term, which is the part of the gap explained by differences in the returns to covariates and unobservables between the two groups. In our setting, because the features defining group membership—gender, race—are (mostly) fixed from birth, the unexplained term is usually associated with labor market discrimination.

Method We use the decomposition method proposed in Firpo *et al.* (2018). This method has three main advantages. First, it can be used to decompose any general distributional statistic. In our context, the relevant distributional statistics are different quantiles of the wage distribution. Second, the method in Firpo *et al.* (2018) provides a detailed decomposition of each variable's contribution to the explained and unexplained components of the wage gap, allowing us to isolate the contribution of trade liberalization. Third, because the method uses recentered influence function (RIF)-regressions (Firpo *et al.*, 2009), it follows the same logic and computational attractiveness of other regression-based methods, such as the Oaxaca-Blinder mean-decomposition method (Oaxaca, 1973; Blinder, 1973). In sum, it is a convenient tool to estimate how trade liberalization affected wage gaps between different demographic groups, at different points of the wage distribution.

The RIF for quantile τ th is given by

$$RIF(y_g; Q_{g,\tau}) = Q_{g,\tau} + \frac{\tau - \mathbb{1}\{y_g \le Q_{g,\tau}\}}{f_{Y_g}(Q_{g,\tau})}, \qquad g = A, B$$
(4)

where $\mathbb{1}\{.\}$ is an indicator function and $f_{Y_g}(.)$ is the density of the marginal distribution of Y for group g. In equation (4), we then plug in the estimated sample quantile, $\hat{Q}_{g,\tau}$, and the density $\hat{f}_{Y_g}(\hat{Q}_{g,\tau})$ and run OLS regressions of $\widehat{RIF}(y_g; \hat{Q}_{g,\tau})$ on covariates X_g .¹⁷ The OLS coefficients $(\hat{\gamma}_{g,\tau})$ play a role similar to the coefficients in a Oaxaca-Blinder mean

 $^{{}^{17}\}widehat{f}_{Y_g}(\widehat{Q}_{g,\tau})$ is estimated with a kernel density estimator.

decomposition. The empirical counterpart of equation (3) becomes

$$\widehat{\Delta}_{O}^{\tau} = \widehat{\Delta}_{X}^{\tau} + \widehat{\Delta}_{U}^{\tau}
= (\overline{X}_{B} - \overline{X}_{A})\widehat{\gamma}_{A,\tau} + \overline{X}_{B}(\widehat{\gamma}_{B,\tau} - \widehat{\gamma}_{A,\tau})$$
(5)

The individual contribution of a covariate k, X_k , to the composition effect is $(\overline{X}_{Bk} - \overline{X}_{Ak})\widehat{\gamma}_{Ak,\tau}$. Its contribution to the unexplained term is $\overline{X}_{Bk}(\widehat{\gamma}_{Bk,\tau} - \widehat{\gamma}_{Ak,\tau})$.

We decompose wage gaps at five different percentiles: 10th, 25th, 50th, 75th, and 90th. The dependent variable is, as before, the log hourly wage. The covariates mimic the model in column 3 of Table 3, which uses tariff protection with a two-year lag, sector fixed effects, and controls for the group's employment share. As a result, we are decomposing wage differentials between two groups *within* sector, year, and state-urban/rural cells, holding their employment propensities constant.

We decompose racial gaps for each gender (male, female), and gender gaps for each racial group (white, nonwhite). For each decomposition, the reference group—i.e., group A in equation (5), whose coefficients, $\hat{\gamma}_{A,\tau}$, weigh the composition effect—is the group with highest hourly wage: white in the racial gap decompositions, and men in the gender gap decompositions. As suggested in Fortin *et al.* (2011), standard errors are obtained by bootstrapping the whole procedure (100 replications).

Findings Before considering the impact of trade liberalization, it is worth noting a few interesting decomposition patterns. Figure 6 plots the composition effect and unexplained term of wage differentials across the distribution for gender-race pairs. The two terms sum up to the observed wage differential. Overall, wage gaps by race are much larger across all quantiles than wage gaps by gender. However, discrimination (proxied by the unexplained term) matters much more by gender than by race.

Racial wage gaps increase over the wage distribution, with the gradient being steeper for men than women (Figure 6). Among men, the racial wage gap at the 90th percentile is 1.7 times higher than the gap at the 10th percentile (Table 4). Among women, the racial gap at the top is 1.2 times higher than at the bottom (Table 5). Moreover, discrimination by race matters more at higher quantiles of the wage distribution. This pattern is particularly strong among men: at the 10th percentile, discrimination accounts for 13% of the wage differential; at the 90th percentile, it accounts for 54%. These patterns suggest strong glass ceiling effects by race, both for nonwhite men and nonwhite women. Among women, Figure 6 reveals a racial sticky floor, which is almost entirely explained by differences in observable characteristics.

The gender gaps within racial groups exhibit different patterns. Among whites, gender gaps increase from 0.18 log points, at the 10th percentile, up to 0.30 log points, at the 90th percentile (Table 6). Among nonwhites, however, the gender wage gap is largest at the 10th percentile: 0.39 log points (Table 7), revealing a large sticky floor for nonwhite women. As one moves up the wage distribution, the magnitude of the gender wage gap becomes similar among whites and nonwhites. With respect to composition vs. discrimination effects, the gender gap is entirely due to discrimination. Composition effects alone would predict a reversal of the gender gap across the whole distribution for nonwhites. For whites, composition effects are also negative at the bottom and top of the distribution, and statistically indistinguishable from zero at the 25th and 50th percentiles. Discrimination effects are particularly large at the top of the wage distribution and, among nonwhites, at the bottom as well. Absent discrimination, white women would out-earn white men by 0.24 log points at the 90th percentile; nonwhite women would out-earn nonwhite men by 0.17 log points at the 10th percentile and by 0.19 log points at the 90th (Figure 6).

We now turn to the contribution of trade liberalization.¹⁸ For racial wage gaps, trade liberalization reduces discrimination at the bottom of the wage distribution, between the 10th and 50th percentiles, as shown by the negative and significant contributions of tariff protection to the unexplained term among men (Table 4) and among women (Table 5). However, between the 50th and 90th percentiles, trade liberalization contributes positively to the unexplained term—that is, it increases discrimination at the top. At the median, trade liberalization significantly reduces unexplained racial gaps among men, whereas it

 $^{^{18}}$ Tables 4-7 only show the contribution of trade liberalization to the decomposition terms. The full tables with the contributions of all explanatory variables are shown in the Online Appendix (Tables A8-A11).

has no significant effect for racial gaps among women.

For gender wage gaps, trade liberalization reduces the unexplained term at the 10th percentile among whites, but has no statistically significant effect for higher quantiles (Table 6). Among nonwhites, tariff reduction reduced discrimination at the 10th and 25th percentiles (Table 7). Similar to the white population, the effect becomes insignificant at higher quantiles. For an overview, Figure 7 plots the contribution of trade liberalization to the unexplained term across quantiles for each gender-race pair.

For all decompositions, the contribution of trade liberalization to the unexplained term is much larger (in absolute terms) than its contribution to the composition effect. Therefore, while gender-race groups were differentially exposed to tariff reduction, the contribution of this differential exposure was small, when compared to the difference in the *impact* of tariff reduction experienced by each group. This is particularly true at the tails of the distribution. For example, the estimates suggest that, in the absence of trade liberalization, racial pay discrimination among men would be 0.39 log points larger at the 10th percentile and 0.21 smaller at the 90th percentile (Table 4). Among women, no liberalization would imply a 0.46 log point larger unexplained racial gap at the bottom and a 0.24 smaller gap at the top of the wage distribution (Table 5).

Overall, in the short term, trade liberalization reduced existing sticky floors for nonwhite women, but reinforced racial glass ceilings in the Brazilian labor market.

5.4 Robustness Checks

In what follows, we briefly report on several robustness checks. For a detailed presentation, see the Online Appendix. First, as discussed in section 4, there is a concern that tariff cuts were smaller in the automotive sector due to protectionist interests. We show that our findings are not driven by individuals employed in this sector (see Table A12). Second, we control for part-time workers, by adding a dummy variable for whether an individual reports working less than 40 hours in the survey's reference week. Although part-time status is highly endogenous to individual unobservables, we find it reassuring that, overall,



Figure (6) RIF-decompositions: composition effect and unexplained term over selected quantiles of the wage distribution

trade protection coefficients remain similar (see Table A15). Third, we regress a model with all TP variables (up to two years) jointly (Table A16). As noted before, these results are very likely to suffer from multicollinearity, because the trade protection variables are highly correlated over time. As before, we find statistically significant positive wage effects of liberalization on women and, especially, nonwhite women after a two-year lag. The coefficients of the contemporaneous effect and one-year lag are, as expected, less robust. Finally, we explore alternative estimators for standard errors (see Table A17) and remove potential outliers in the wage distribution by winsorizing (Table A13) and trimming (Table A14) the dependent variable. The results are robust throughout.

We also check the robustness of the detailed decompositions to several specification choices. The contribution of trade liberalization to the unexplained term does not depend on the inclusion of the employment correction term (Figure A5). In addition, we performed all decompositions for two alternative lag choices of tariff protection: contemporaneous effects (TP_t) , and with a one year lag (TP_{t-1}) . Figure A6 plots the contribution of trade liberalization to the unexplained term of the wage gap for the different lag choices. For most pairs, the direction of the contribution across the wage distribution is similar across the time lags. However, the contributions are much larger for the two-year lag (TP_{t-2})

Notes: RIF-decomposition (Firpo *et al.*, 2018) estimates. Period is 1987–1996. Survey weights from PNAD are used. For bootstrap standard errors and more details, see Tables 4-7.

	Log wage at percentile:						
	(1)	(2)	(3)	(4)	(5)		
	10th	25th	50th	75th	90th		
Differential							
Log wage White men	0.5992^{***}	1.1666^{***}	1.8425^{***}	2.5927^{***}	3.3435^{***}		
	(0.0050)	(0.0041)	(0.0028)	(0.0036)	(0.0046)		
Log wage Nonwhite men	0.1176^{***}	0.6182^{***}	1.2086^{***}	1.8717^{***}	2.5267^{***}		
	(0.0021)	(0.0035)	(0.0036)	(0.0067)	(0.0092)		
Difference	0.4816^{***}	0.5484^{***}	0.6338^{***}	0.7210^{***}	0.8168^{***}		
	(0.0051)	(0.0049)	(0.0043)	(0.0077)	(0.0102)		
Explained							
Tariff protection	0.0219^{***}	-0.0092	-0.0378***	-0.0324^{***}	-0.0305***		
	(0.0065)	(0.0065)	(0.0051)	(0.0061)	(0.0075)		
Total	0.4199^{***}	0.5411^{***}	0.4786^{***}	0.4612^{***}	0.3788^{***}		
	(0.0161)	(0.0127)	(0.0089)	(0.0094)	(0.0125)		
Unexplained							
Tariff protection	-0.3889***	-0.1524^{***}	-0.1080***	0.1398^{***}	0.2129^{***}		
	(0.0462)	(0.0372)	(0.0382)	(0.0477)	(0.0599)		
Total	0.0617^{***}	0.0073	0.1552^{***}	0.2598^{***}	0.4380***		
	(0.0201)	(0.0152)	(0.0095)	(0.0116)	(0.0138)		

 Table (4)
 RIF-decompositions: racial gap among men

Notes: N = 278744; N white men = 152014; N nonwhite men = 126730. * p < 0.10, ** p < 0.05, *** p < 0.01. RIF-decomposition (Firpo *et al.*, 2018) estimates with bootstrap standard errors in parentheses (100 replications). Tariff protection measured at time t - 2. Period is 1987–1996. Survey weights from PNAD are used. For the contribution of remaining explanatory variables, see Online Appendix Table A8.

and, in fact, most of the effects for TP_t and TP_{t-1} are statistically insignificant. Lastly, we obtain very similar results when we re-estimate the decompositions based on a model without sector fixed effects (Figure A7).

6 Conclusion

This article revisits Brazil's trade liberalization process (1987–1995), a natural experiment that has been widely studied in the literature. By combining a local labor market approach with decomposition methods, we paint a rich picture of the short-run effect of trade liberalization for gender and racial inequality, both at the mean and along the wage distribution. After an initial adjustment period, trade liberalization caused an increase of approximately 1.24%–2.25% in hourly wages of white women and 2.71%–4.7% in hourly wages of nonwhite women, which contributed to a reduction both in *average* gender wage

	Log wage at percentile:						
	(1) 10th	(2) 25th	(3) 50th	(4) 75th	(5) 90th		
Differential							
Log wage White women	0.4207^{***}	0.8880^{***}	1.5478^{***}	2.3374^{***}	3.0473^{***}		
	(0.0038)	(0.0074)	(0.0046)	(0.0091)	(0.0073)		
Log wage Nonwhite women	-0.2711^{***}	0.3774^{***}	0.9301^{***}	1.5684^{***}	2.2406^{***}		
	(0.0032)	(0.0055)	(0.0033)	(0.0061)	(0.0088)		
Difference	0.6918^{***}	0.5106^{***}	0.6177^{***}	0.7690^{***}	0.8067^{***}		
	(0.0051)	(0.0089)	(0.0052)	(0.0111)	(0.0109)		
Explained							
Tariff protection	0.0454^{***}	-0.0170^{**}	-0.0362^{***}	-0.0365^{***}	-0.0376^{***}		
	(0.0116)	(0.0071)	(0.0068)	(0.0086)	(0.0102)		
Total	0.6703^{***}	0.5075^{***}	0.4926^{***}	0.6016^{***}	0.4540^{***}		
	(0.0194)	(0.0128)	(0.0152)	(0.0184)	(0.0162)		
Unexplained							
Tariff protection	-0.4560^{***}	-0.3529^{***}	-0.0565	0.1607^{**}	0.2387^{***}		
	(0.1289)	(0.0685)	(0.0583)	(0.0751)	(0.0923)		
Total	0.0215	0.0031	0.1251^{***}	0.1675^{***}	0.3527^{***}		
	(0.0212)	(0.0144)	(0.0137)	(0.0173)	(0.0188)		

Table (5) RIF-decompositions: racial gap among women

Notes: N = 145508; N white women = 81173; N nonwhite women = 64335. * p < 0.10, *** p < 0.05, *** p < 0.01. RIF-decomposition (Firpo *et al.*, 2018) estimates with bootstrap standard errors in parentheses (100 replications). Tariff protection measured at time t - 2. Period is 1987–1996. Survey weights from PNAD are used. For the contribution of remaining explanatory variables, see Online Appendix Table A9.

inequality and *average* racial wage inequality among women. However, trade liberalization had heterogeneous consequences along the wage distribution. Liberalization reduced racial and gender discrimination at low wages, mitigating preexisting 'sticky floors' by gender, but reinforced existing 'glass ceilings' by race.

Although in slight decline, those sticky floors and glass ceilings reflect the persistence of discrimination in the Brazilian labor market. In terms of public policy, it is important to understand how economic shocks and policies shape those patterns of labor market inequality, even if unintentionally. Our results show that even though liberalization contributed to a reduction in *average* gender and racial wage inequality and discrimination at the *bottom* of the distribution—consistent with Becker (1957)— it contributed to an increase in racial discrimination in the upper part of the wage distribution. As discussed by Borrowman and Klasen (2020), there is a great amount of persistence in labor market inequalities. Breaking these patterns requires concatenated efforts. In particular, the

	Log wage at percentile:						
	(1) 10th	(2) 25th	(3) 50th	(4) 75th	(5) 90th		
Differential							
Log wage White men	0.5992^{***}	1.1666^{***}	1.8425^{***}	2.5927^{***}	3.3435^{***}		
	(0.0051)	(0.0040)	(0.0033)	(0.0036)	(0.0040)		
Log wage White women	0.4207^{***}	0.8880^{***}	1.5478^{***}	2.3374^{***}	3.0473^{***}		
	(0.0036)	(0.0064)	(0.0046)	(0.0102)	(0.0066)		
Difference	0.1786^{***}	0.2786^{***}	0.2947^{***}	0.2552^{***}	0.2962^{***}		
	(0.0065)	(0.0078)	(0.0056)	(0.0107)	(0.0077)		
Explained							
Tariff protection	-0.0012^{*}	0.0005	0.0022^{**}	0.0018^{**}	0.0017^{**}		
	(0.0007)	(0.0004)	(0.0008)	(0.0009)	(0.0009)		
Total	-0.0384	0.0442	0.0083	-0.1584^{***}	-0.2387***		
	(0.0337)	(0.0323)	(0.0255)	(0.0257)	(0.0312)		
Unexplained							
Tariff protection	-0.1829^{**}	0.0620	0.0211	0.0565	0.0755		
	(0.0889)	(0.0646)	(0.0504)	(0.0679)	(0.0799)		
Total	0.2170^{***}	0.2345^{***}	0.2864^{***}	0.4136^{***}	0.5349^{***}		
	(0.0347)	(0.0332)	(0.0262)	(0.0272)	(0.0323)		

Table (6) RIF-decompositions: gender gap among whites

Notes: N = 233187; N white men = 152014; N white women = 81173. * p < 0.10, ** p < 0.05, *** p < 0.01. RIF-decomposition (Firpo *et al.*, 2018) estimates with bootstrap standard errors in parentheses (100 replications). Tariff protection measured at time t - 2. Period is 1987–1996. Survey weights from PNAD are used. For the contribution of remaining explanatory variables, see Online Appendix Table A10.

double burden of discrimination experienced by nonwhite women, which is often invisible to society, should be given more focus in future research and policy agendas.

	Log wage at percentile:						
	(1)	(2)	(3)	(4)	(5)		
	10th	25th	50th	75th	90th		
Differential							
Log wage Nonwhite men	0.1176^{***}	0.6182^{***}	1.2086^{***}	1.8717^{***}	2.5267^{***}		
	(0.0024)	(0.0042)	(0.0032)	(0.0069)	(0.0102)		
Log wage Nonwhite women	-0.2711^{***}	0.3774^{***}	0.9301^{***}	1.5684^{***}	2.2406^{***}		
	(0.0037)	(0.0061)	(0.0036)	(0.0063)	(0.0091)		
Difference	0.3887^{***}	0.2408^{***}	0.2786^{***}	0.3033^{***}	0.2861^{***}		
	(0.0045)	(0.0076)	(0.0048)	(0.0098)	(0.0141)		
Explained							
Tariff protection	-0.0177^{***}	-0.0037^{***}	0.0030***	0.0108^{***}	0.0130^{***}		
	(0.0025)	(0.0010)	(0.0010)	(0.0015)	(0.0024)		
Total	-0.1706^{***}	-0.0326	-0.0171	-0.0771^{***}	-0.1926^{***}		
	(0.0311)	(0.0199)	(0.0186)	(0.0261)	(0.0377)		
Unexplained							
Tariff protection	-0.2099^{*}	-0.1421^{***}	0.0734	0.0644	0.0829		
	(0.1099)	(0.0512)	(0.0546)	(0.0659)	(0.0857)		
Total	0.5593^{***}	0.2735^{***}	0.2957^{***}	0.3803^{***}	0.4786^{***}		
	(0.0313)	(0.0215)	(0.0196)	(0.0269)	(0.0395)		

 Table (7)
 RIF-decompositions: gender gap among nonwhites

Notes: N = 191065; N nonwhite men = 126730; N nonwhite women = 64335. * p < 0.10, ** p < 0.05, *** p < 0.01. RIF-decomposition (Firpo *et al.*, 2018) estimates with bootstrap standard errors in parentheses (100 replications). Tariff protection measured at time t - 2. Period is 1987–1996. Survey weights from PNAD are used. For the contribution of remaining explanatory variables, see Online Appendix Table A11.





Figure (7) Contribution of trade liberalization to unexplained term over selected quantiles of the wage distribution

Notes: Common Y-axis for all subfigures. RIF-decomposition (Firpo *et al.*, 2018) estimates. Tariff protection measured at time t - 2. Period is 1987–1996. Survey weights from PNAD are used. For bootstrap standard errors and more details, see Tables 4-7.

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Online Appendix

Between sticky floors and glass ceilings: the effect of trade liberalization on double discrimination in Brazil

Variable	Description
Ln(Hourly wage)	Logarithm of deflated hourly wages from main occupation (BRL).
Worked in the ref week	Worked in the reference week
Nonwhite	Individual self-declared as black $(preto)$ or brown $(pardo)$
Women	Female respondent
Age	Age in completed years
Squared Age	Squared age
Education 1	Respondent completed no years of education
Education 2	Respondent completed between 1 and 3 years of education
Education 3	Respondent completed between 4 and 7 years of education
Education 4	Respondent completed between 8 and 10 years of education
Education 5	Respondent completed between 11 and 14 years of education
Education 6	Respondent completed more than 14 years of education
Number of Children	Number of children in the household
TP_t	Trade exposure at the state-rural/urban cell
TP_{t-1}	One year lagged trade exposure at the state-rural/urban cell
TP_{t-2}	Two-years lagged trade exposure at the state-rural/urban cell
TP_{t-3}	Three-years lagged trade exposure at the state-rural/urban cell
TP_{t-4}	Four-years lagged trade exposure at the state-rural/urban cell
TP_{t-5}	Five-years lagged trade exposure at the state-rural/urban cell
Part time	Worked less than 40 hours per week
Employment share	Average employment share of respondent's gender, race and 5 year age cohort in the state and urban-rural area of residence

A.1 Variable definition

A.2 Industry concordance

The construction of our trade protection measure, specified in equation 1, requires the concordance between the PNAD industry classification—which we use to construct our employment shares—with the sectoral tariff data by Kume *et al.* (2003)—which we use as our shifter. We follow the concordance methodology in Ferreira *et al.* (2007) with three alterations proposed by Gaddis and Pieters (2017): combining sectors 'processing of vegetal products' with 'meat packing, dairy industry, vegetal and other food products'; 'leather and skins' with 'footwear'; and 'manufacturing of synthetic materials' with 'unclassified

manufacturing'.

A.3 Robustness Checks

Here, we present in more detail the robustness checks of section 5.4. First, we show that our preferred estimates (Table 1) are not driven by workers in the automotive sector. In this sector, tariffs did not decrease monotonically, but oscillated considerably over time (see Figure A1). This raises the concern that automotive tariffs could be endogenous to protectionist interests or sectoral performance. To alleviate this concern, we exclude workers in the automotive sector from the analysis (Table A12). Reassuringly, the point estimates of trade liberalization remain very similar, both in magnitude and statistical significance. Additionally, when running regressions separately for each of the 21 sectors of employment (Figure A8), we find no evidence that our results by gender and race are driven by the automotive sector. For nonwhite men, there is little heterogeneity in the effects of liberalization across sectors, while for white and nonwhite women heterogeneity is more pronounced.

Second, we remove potential outliers in the wage data by winsorizing and trimming the dependent variable at the 1st and 99th and at the 5th and 95th percentiles, as shown in Tables A13 and A14. Interestingly, winsorizing or trimming the dependent variable at the 5% yield considerably smaller gender and race pay gaps at baseline. Whereas the baseline gender gap ranges around 44%–47% and the race gap around 15%–19% (Table 1), estimates of Tables A13 and A14 reveal gender and race gaps around 33%–42% and 12%–14%, respectively. This reduction in gender and wage inequality, as a result of winsorizing or trimming, is consistent with the existence of sticky floors and glass ceilings in the wage distribution, as discussed in section 2. The effects of trade liberalization across gender-race groups remain qualitatively unchanged to either trimming or winsorizing.

Third, we control for part-time workers, by adding a dummy variable for whether an individual worked less than 40 hours in the survey's reference week. Although parttime status is highly endogenous to individual unobservables, we find it reassuring that, overall, the trade protection estimates remain comparable, although slightly smaller and marginally less significant for the triple interaction coefficient (Table A15). We also note that controlling for part-time status considerably widens the conditional gender gap and slightly reduces the racial gap.

Fourth, we regress all trade protection variables (lagged up to two years) jointly (Table A16). Because the TP variables are highly correlated over time, these results should be interpreted with caution, as they are very likely plagued by multicollinearity. As before, we find statistically significant positive wage effects of liberalization on women and, especially, nonwhite women after a two-year lag. The coefficients of the contemporaneous effect and one-year lag are, as expected, less robust.

Fifth, we consider alternative estimators for standard errors, by allowing for correlation across individual standard errors at different levels. Table A17 presents the estimates. Standard errors clustered at the state level are shown in parenthesis, while standard errors clustered at the state-urban-sector level appear in brackets. All in all, the significance levels are very similar to our baseline results.

A.4 Additional Figures and Tables



Figure (A1) Tariffs across sectors and over time



Figure (A2) Changes in tariffs and initial tariff level across sectors



Figure (A3) Changes in tariffs and initial tariff level across sectors; excluding automotive sector



Figure (A4) RIF-decompositions: composition effect and unexplained term over selected quantiles of the wage distribution

	mean	sd	\min	max
Ln(Hourly wage)	1.556	1.095	-9	8
Worked in the ref week	0.982	0.132	0	1
Nonwhite	0.450	0.498	0	1
Women	0.343	0.475	0	1
Nonwhite \times Women	0.152	0.359	0	1
Age	40.090	9.787	25	64
Squared Age	1702.992	836.885	625	4096
Education 1	0.161	0.368	0	1
Education 2	0.173	0.379	0	1
Education 3	0.303	0.459	0	1
Education 4	0.116	0.321	0	1
Education 5	0.163	0.369	0	1
Education 6	0.083	0.276	0	1
Number of Children	2.193	1.709	0	14
TP_t	1.857	1.190	0	5
$\mathrm{TP}_t \times \mathrm{Women}$	0.659	1.136	0	5
$\mathrm{TP}_t \times \mathrm{Nonwhite}$	0.754	1.112	0	5
$\mathrm{TP}_t \times \mathrm{Nonwhite} \times \mathrm{Women}$	0.265	0.757	0	5
TP_{t-1}	2.091	1.349	0	6
$\mathrm{TP}_{t-1} \times \mathrm{Women}$	0.738	1.281	0	6
$TP_{t-1} \times Nonwhite$	0.859	1.278	0	6
$TP_{t-1} \times Nonwhite \times Women$	0.300	0.864	0	6
TP_{t-2}	2.888	2.125	0	8
$\mathrm{TP}_{t-2} \times \mathrm{Women}$	1.004	1.855	0	8
$TP_{t-2} \times Nonwhite$	1.212	1.925	0	8
$TP_{t-2} \times Nonwhite \times Women$	0.417	1.267	0	8
TP_{t-3}	2.998	2.218	0	8
$\mathrm{TP}_{t-3} \times \mathrm{Women}$	1.057	1.938	0	8
$TP_{t-3} \times Nonwhite$	1.248	2.004	0	8
$\mathrm{TP}_{t-3} \times \mathrm{Nonwhite} \times \mathrm{Women}$	0.435	1.326	0	8
TP_{t-4}	2.670	1.456	0	6
$TP_{t-4} \times Women$	0.968	1.561	0	6
$TP_{t-4} \times Nonwhite$	1.077	1.522	0	6
$TP_{t-4} \times Nonwhite \times Women$	0.382	1.048	0	6
TP_{t-5}	3.964	2.062	0	8
$\text{TP}_{t-5} \times \text{Women}$	1.421	2.270	0	8
$TP_{t-5} \times Nonwhite$	1.633	2.274	0	8
$\mathrm{TP}_{t-5} \times \mathrm{Nonwhite} \times \mathrm{Women}$	0.571	1.557	0	8
Observations	424252			

 Table (A2)
 Summary statistics

Panel A: Women						
	White women		Nonwhite	e women	Differe	ence
	mean	sd	mean	sd	b	t
Ln(Hourly wage)	1.66	1.08	1.03	1.01	0.63^{***}	113.27
Age	39.21	9.18	39.58	9.33	-0.37***	-7.61
Squared Age	1621.35	770.07	1653.37	788.18	-32.02^{***}	-7.80
Education 1	0.08	0.26	0.20	0.40	-0.13^{***}	-73.58
Education 2	0.12	0.33	0.19	0.39	-0.07***	-37.12
Education 3	0.28	0.45	0.29	0.46	-0.01***	-4.79
Education 4	0.12	0.32	0.11	0.31	0.01^{***}	8.54
Education 5	0.24	0.43	0.16	0.37	0.08^{***}	38.78
Education 6	0.16	0.37	0.04	0.21	0.11^{***}	71.21
Number of Children	1.87	1.43	2.47	1.85	-0.59***	-69.03
Observations	81173		64335		145508	

Table (A3) Group differences

Panel B: Women and men

Difference MenWomen sd mean \mathbf{b} \mathbf{t} mean sd 0.26^{***} Ln(Hourly wage) 1.651.081.381.0975.26 0.01^{***} Nonwhite 0.450.500.440.507.77 1.10^{***} Age 40.4710.0439.37 9.2534.66102.71*** Squared Age 1738.22863.811635.51778.2938.01Education 1 0.180.380.130.34 0.04^{***} 36.88Education 2 0.03^{***} 0.180.390.150.3624.09 0.02^{***} Education 30.310.460.290.4516.64 Education 4 0.320.32 0.00^{***} 4.440.120.11Education 5 0.140.350.210.40 -0.07^{***} -54.66-0.04*** Education 6 0.070.260.110.31-42.15Number of Children 1.73 0.09^{***} 15.732.222.141.66Observations 278744145508424252

	b	pct	cumpct
Agricultural products	4416.00	6.86	7
Mining Products	79.00	0.12	7
Oil and Coal extraction	14.00	0.02	7
Non-metallic minerals	129.00	0.20	7
Steel, non-ferrous and other metal products	110.00	0.17	7
Machinery and Tractors	88.00	0.14	8
Electrical and electronic equipment	202.00	0.31	8
Automobiles, trucks and buses; parts, comp. and other vehicles	58.00	0.09	8
Wood products and furniture	367.00	0.57	8
Cellulose, paper and printing	158.00	0.25	9
Rubber products	20.00	0.03	9
Chemical elements and products	101.00	0.16	9
Oil refining and petrochemicals	19.00	0.03	9
Pharmaceutical and perfumery products	68.00	0.11	9
Plastic products	107.00	0.17	9
Textile products	642.00	1.00	10
Apparel	1142.00	1.78	12
Footwear	194.00	0.30	12
Meat packing, dairy industry, vegetable and other food products	1666.00	2.59	15
Nontradables	54536.00	84.77	100
Unclassified manufacturing	219.00	0.34	100
Total	64335.00	100.00	
Observations	64335		

Table (A4) Sectoral occupation: Nonwhite women

Table (A5) Sectoral occupation: White women

	b	pct	cumpct
Agricultural products	2457.00	3.03	3
Mining Products	45.00	0.06	3
Oil and Coal extraction	22.00	0.03	3
Non-metallic minerals	186.00	0.23	3
Steel, non-ferrous and other metal products	380.00	0.47	4
Machinery and Tractors	191.00	0.24	4
Electrical and electronic equipment	269.00	0.33	4
Automobiles, trucks and buses; parts, comp. and other vehicles	181.00	0.22	5
Wood products and furniture	348.00	0.43	5
Cellulose, paper and printing	408.00	0.50	6
Rubber products	50.00	0.06	6
Chemical elements and products	218.00	0.27	6
Oil refining and petrochemicals	58.00	0.07	6
Pharmaceutical and perfumery products	153.00	0.19	6
Plastic products	207.00	0.26	6
Textile products	765.00	0.94	7
Apparel	1789.00	2.20	10
Footwear	1099.00	1.35	11
Meat packing, dairy industry, vegetable and other food products	1987.00	2.45	13
Nontradables	69946.00	86.17	99
Unclassified manufacturing	414.00	0.51	100
Total	81173.00	100.00	
Observations	81173		

	b	pct	cumpct
Agricultural products	29722.00	23.45	23
Mining Products	1259.00	0.99	24
Oil and Coal extraction	200.00	0.16	25
Non-metallic minerals	1676.00	1.32	26
Steel, non-ferrous and other metal products	2441.00	1.93	28
Machinery and Tractors	638.00	0.50	28
Electrical and electronic equipment	471.00	0.37	29
Automobiles, trucks and buses; parts, comp. and other vehicles	841.00	0.66	29
Wood products and furniture	2960.00	2.34	32
Cellulose, paper and printing	831.00	0.66	32
Rubber products	168.00	0.13	33
Chemical elements and products	699.00	0.55	33
Oil refining and petrochemicals	241.00	0.19	33
Pharmaceutical and perfumery products	192.00	0.15	33
Plastic products	270.00	0.21	34
Textile products	630.00	0.50	34
Apparel	245.00	0.19	34
Footwear	366.00	0.29	35
Meat packing, dairy industry, vegetable and other food products	5282.00	4.17	39
Nontradables	77270.00	60.97	100
Unclassified manufacturing	328.00	0.26	100
Total	126730.00	100.00	
Observations	126730		

Table (A6) Sectoral occupation: Nonwhite men

Table (A7)Sectoral occupation: White men

	b	pct	cumpct
Agricultural products	25809.00	16.98	17
Mining Products	881.00	0.58	18
Oil and Coal extraction	244.00	0.16	18
Non-metallic minerals	1755.00	1.15	19
Steel, non-ferrous and other metal products	4042.00	2.66	22
Machinery and Tractors	1692.00	1.11	23
Electrical and electronic equipment	1086.00	0.71	23
Automobiles, trucks and buses; parts, comp. and other vehicles	1830.00	1.20	25
Wood products and furniture	3484.00	2.29	27
Cellulose, paper and printing	1645.00	1.08	28
Rubber products	343.00	0.23	28
Chemical elements and products	1184.00	0.78	29
Oil refining and petrochemicals	363.00	0.24	29
Pharmaceutical and perfumery products	389.00	0.26	29
Plastic products	559.00	0.37	30
Textile products	1015.00	0.67	30
Apparel	529.00	0.35	31
Footwear	1338.00	0.88	32
Meat packing, dairy industry, vegetable and other food products	5600.00	3.68	35
Nontradables	97629.00	64.22	100
Unclassified manufacturing	597.00	0.39	100
Total	152014.00	100.00	
Observations	152014		

		Log	wage at perce	entile:	
	(1)	(2)	(3)	(4)	(5)
	10th	25th	50th	75th	90th
Differential	0 5000***	1 1000***	1 0405***	0 5007***	0.0405***
Log wage white men	0.5992^{-11}	1.1000	1.8425	2.592(3.3435
T NT 1.4	(0.0050)	(0.0041)	(0.0028)	(0.0030)	(0.0046)
Log wage Nonwhite men	(0,0021)	(0.0182)	1.2080	1.8(1)	2.5207
D:#	(0.0021)	(0.0035)	(0.0030)	(0.0007)	(0.0092)
Difference	(0.0051)	(0.0040)	(0.0042)	(0.0077)	(0.0102)
	(0.0051)	(0.0049)	(0.0045)	(0.0077)	(0.0102)
Explained					
Tariff protection	0.0219^{***}	-0.0092	-0.0378^{***}	-0.0324^{***}	-0.0305***
	(0.0065)	(0.0065)	(0.0051)	(0.0061)	(0.0075)
Demographic	-0.0057^{***}	-0.0006	0.0033^{**}	0.0051^{***}	0.0034^{*}
	(0.0017)	(0.0018)	(0.0014)	(0.0016)	(0.0018)
Education	0.1735^{***}	0.2816^{***}	0.3232^{***}	0.3871^{***}	0.3564^{***}
	(0.0073)	(0.0065)	(0.0058)	(0.0079)	(0.0109)
Employment share	0.0012^{***}	0.0025^{***}	0.0019^{***}	0.0005^{*}	-0.0004
	(0.0003)	(0.0004)	(0.0003)	(0.0003)	(0.0003)
Sector	0.0650^{***}	0.0707^{***}	0.0377^{***}	0.0125^{***}	0.0003
	(0.0030)	(0.0022)	(0.0014)	(0.0014)	(0.0014)
Region	0.1654^{***}	0.1949^{***}	0.1457^{***}	0.0841^{***}	0.0448^{***}
	(0.0098)	(0.0086)	(0.0076)	(0.0072)	(0.0095)
Year	-0.0014^{**}	0.0013^{**}	0.0046^{***}	0.0043^{***}	0.0049^{***}
	(0.0007)	(0.0006)	(0.0011)	(0.0014)	(0.0016)
Total	0.4199^{***}	0.5411^{***}	0.4786^{***}	0.4612^{***}	0.3788^{***}
	(0.0161)	(0.0127)	(0.0089)	(0.0094)	(0.0125)
Unexplained					
Tariff protection	-0.3889***	-0.1524^{***}	-0.1080***	0.1398^{***}	0.2129^{***}
	(0.0462)	(0.0372)	(0.0382)	(0.0477)	(0.0599)
Demographic	-0.0551	0.6361***	0.4478***	0.2766^{**}	-0.2465*
	(0.1030)	(0.0968)	(0.0865)	(0.1159)	(0.1473)
Education	0.2610^{***}	0.3703^{***}	0.0944^{***}	-0.0667***	-0.2460***
	(0.0237)	(0.0178)	(0.0147)	(0.0232)	(0.0175)
Employment share	0.3620^{***}	0.3512^{***}	0.0344	-0.1460	-0.1411
	(0.1050)	(0.0895)	(0.0781)	(0.0909)	(0.1231)
Sector	0.0843^{***}	0.1659^{***}	-0.1339^{***}	-0.1579^{***}	-0.1064^{***}
	(0.0274)	(0.0185)	(0.0135)	(0.0156)	(0.0136)
Region	-0.0018	-0.0941^{**}	-0.0064	-0.0178	0.2345^{**}
	(0.0406)	(0.0399)	(0.0516)	(0.0601)	(0.1031)
Year	-0.3552^{***}	-0.1333^{**}	-0.2900^{***}	0.0823	-0.0055
	(0.0658)	(0.0649)	(0.0836)	(0.1115)	(0.1514)
Constant	0.1555	-1.1365^{***}	0.1168	0.1496	0.7361^{***}
	(0.1641)	(0.1391)	(0.1546)	(0.2058)	(0.2708)
Total	0.0617^{***}	0.0073	0.1552^{***}	0.2598^{***}	0.4380^{***}
	(0.0201)	(0.0152)	(0.0095)	(0.0116)	(0.0138)
N	278744	278744	278744	278744	278744
N White men	152014	152014	152014	152014	152014
N Nonwhite men	126730	126730	126730	126730	126730

Table (A8)RIF-decompositions: racial gap among men

Notes: * p < 0.10, ** p < 0.05, *** p < 0.01. Bootstrap standard errors in parentheses (100 replications). Tariff protection measured at time t - 2. Period is X987–1996. Survey weights from PNAD are used.

	Log wage at percentile:					
	(1)	(2)	(3)	(4)	(5)	
	10th	25th	50th	75th	90th	
Differential	0 4007***	0.0000***	1 5470***	0.0074***	9.0479***	
Log wage white women	(0.4207)	0.8880	1.54(8)	$2.33(4^{-1})$	3.0473^{+++}	
	(0.0038)	(0.0074)	(0.0046)	(0.0091)	(0.0073)	
Log wage Nonwhite women	-0.2711	0.3774	(0.0022)	1.3084	2.2400	
D:#	(0.0032)	(0.0055)	(0.0033)	(0.0001)	(0.0088)	
Difference	(0.0918)	(0.0080)	(0.0177)	(0.0111)	(0.0100)	
	(0.0031)	(0.0089)	(0.0052)	(0.0111)	(0.0109)	
Explained						
Tariff protection	0.0454^{***}	-0.0170^{**}	-0.0362^{***}	-0.0365***	-0.0376^{***}	
	(0.0116)	(0.0071)	(0.0068)	(0.0086)	(0.0102)	
Demographic	0.0162^{***}	0.0125^{***}	0.0072^{***}	-0.0015	-0.0067**	
	(0.0034)	(0.0023)	(0.0019)	(0.0024)	(0.0029)	
Education	0.2691^{***}	0.2872^{***}	0.3555^{***}	0.5008***	0.4218^{***}	
	(0.0091)	(0.0076)	(0.0110)	(0.0147)	(0.0140)	
Employment share	0.0001	0.0001	0.0001	0.0000	0.0000	
~	(0.0001)	(0.0002)	(0.0002)	(0.0001)	(0.0001)	
Sector	0.0539***	0.0286***	0.0069***	-0.0055***	-0.0066***	
	(0.0033)	(0.0020)	(0.0014)	(0.0013)	(0.0019)	
Region	0.2813***	0.1967***	0.1619***	0.1451***	0.0833***	
	(0.0157)	(0.0097)	(0.0095)	(0.0121)	(0.0138)	
Year	0.0043**	-0.0006	-0.0028	-0.0008	-0.0003	
	(0.0018)	(0.0015)	(0.0018)	(0.0025)	(0.0028)	
Total	0.6703^{***}	0.5075^{***}	0.4926***	0.6016^{***}	0.4540^{***}	
	(0.0194)	(0.0128)	(0.0152)	(0.0184)	(0.0162)	
Unexplained						
Tariff protection	-0.4560^{***}	-0.3529^{***}	-0.0565	0.1607^{**}	0.2387^{***}	
	(0.1289)	(0.0685)	(0.0583)	(0.0751)	(0.0923)	
Demographic	-0.0157	-0.3420^{**}	-0.1374	0.4392^{**}	0.2834	
	(0.1706)	(0.1501)	(0.1319)	(0.1907)	(0.2180)	
Education	0.2979^{***}	0.1031^{***}	0.0425^{*}	0.0046	-0.1707^{***}	
	(0.0386)	(0.0254)	(0.0225)	(0.0228)	(0.0253)	
Employment share	0.0518	0.1039^{*}	0.1184**	0.0100	-0.0339	
-	(0.0931)	(0.0624)	(0.0521)	(0.0607)	(0.0684)	
Sector	0.4026***	-0.0319	-0.1307***	-0.1653***	-0.1276***	
	(0.0664)	(0.0380)	(0.0244)	(0.0237)	(0.0344)	
Region	-0.0107	0.1385**	0.1434**	0.0095	0.2281*	
	(0.0787)	(0.0667)	(0.0676)	(0.0975)	(0.1254)	
Year	-0.5662***	-0.4590***	-0.1301	0.2442	0.1060	
	(0.1551)	(0.1278)	(0.1171)	(0.1778)	(0.2223)	
Constant	0.3178	0.8433^{***}	0.2757	-0.5354*	-0.1714	
σ Γ (1	(0.3046)	(0.2127)	(0.1815)	(0.2933)	(0.3272)	
Iotal	0.0215	0.0031	0.1251^{***}	0.1675^{***}	0.3527^{***}	
	(0.0212)	(0.0144)	(0.0137)	(0.0173)	(0.0188)	
N	145508	145508	145508	145508	145508	
N White women	81173	81173	81173	81173	81173	
N Nonwhite women	64335	64335	64335	64335	64335	

 Table (A9)
 RIF-decompositions: racial gap among women

Notes: * p < 0.10, ** p < 0.05, *** p < 0.01. Bootstrap standard errors in parentheses (100 replications). Tariff protection measured at time t - 2. Period is **1337**–1996. Survey weights from PNAD are used.

		Log wage at percentile:					
	(1) 10th	$\begin{array}{c} (2) \\ 25 \mathrm{th} \end{array}$	(3) 50th	(4) 75th	(5) 90th		
Differential							
Log wage White men	0.5992^{***}	1.1666^{***}	1.8425^{***}	2.5927^{***}	3.3435^{***}		
	(0.0051)	(0.0040)	(0.0033)	(0.0036)	(0.0040)		
Log wage White women	0.4207^{***}	0.8880^{***}	1.5478^{***}	2.3374^{***}	3.0473^{***}		
	(0.0036)	(0.0064)	(0.0046)	(0.0102)	(0.0066)		
Difference	0.1786^{***}	0.2786^{***}	0.2947^{***}	0.2552^{***}	0.2962^{***}		
	(0.0065)	(0.0078)	(0.0056)	(0.0107)	(0.0077)		
Explained							
Tariff protection	-0.0012^{*}	0.0005	0.0022^{**}	0.0018^{**}	0.0017^{**}		
	(0.0007)	(0.0004)	(0.0008)	(0.0009)	(0.0009)		
Total	-0.0384	0.0442	0.0083	-0.1584^{***}	-0.2387^{***}		
	(0.0337)	(0.0323)	(0.0255)	(0.0257)	(0.0312)		
Unexplained							
Tariff protection	-0.1829^{**}	0.0620	0.0211	0.0565	0.0755		
	(0.0889)	(0.0646)	(0.0504)	(0.0679)	(0.0799)		
Total	0.2170^{***}	0.2345^{***}	0.2864^{***}	0.4136^{***}	0.5349^{***}		
	(0.0347)	(0.0332)	(0.0262)	(0.0272)	(0.0323)		
Ν	233187	233187	233187	233187	233187		
N White men	152014	152014	152014	152014	152014		
N White women	81173	81173	81173	81173	81173		

Table (A10) RIF-decompositions: gender gap among whites

Notes: * p < 0.10, ** p < 0.05, *** p < 0.01. Bootstrap standard errors in parentheses (100 replications). Tariff protection measured at time t - 2. Period is 1987–1996. Survey weights from PNAD are used.

	Log wage at percentile:					
	(1)	(2)	(3)	(4)	(5)	
	10th	25th	50th	75th	90th	
Differential	0 1170***	0.0100***	1 0000***	1 0717***	0 5007***	
Log wage Nonwhite men	(0,0024)	(0.0182)	1.2080	1.8(1)	2.5207	
T NI 1.4	(0.0024)	(0.0042)	(0.0032)	(0.0069)	(0.0102)	
Log wage Nonwhite women	-0.2(11)	0.3774	(0.9301)	1.5084	2.2400	
D:#	(0.0037)	(0.0001)	(0.0030)	(0.0003)	(0.0091)	
Difference	(0.0045)	(0.0076)	(0.0048)	(0.0008)	(0.0141)	
	(0.0045)	(0.0070)	(0.0048)	(0.0098)	(0.0141)	
Explained						
Tariff protection	-0.0177^{***}	-0.0037***	0.0030***	0.0108^{***}	0.0130^{***}	
	(0.0025)	(0.0010)	(0.0010)	(0.0015)	(0.0024)	
Demographic	-0.0040***	0.0005	0.0014^{**}	0.0017^{*}	0.0022^{**}	
	(0.0011)	(0.0006)	(0.0007)	(0.0010)	(0.0011)	
Education	-0.0281^{***}	-0.0370***	-0.0837***	-0.1374^{***}	-0.2033***	
	(0.0013)	(0.0016)	(0.0026)	(0.0079)	(0.0080)	
Employment share	-0.0260	0.1052^{***}	0.1767^{***}	0.1094^{***}	0.0144	
	(0.0305)	(0.0198)	(0.0195)	(0.0275)	(0.0375)	
Sector	-0.1029^{***}	-0.0837***	-0.0856***	-0.0292^{***}	0.0036	
	(0.0040)	(0.0032)	(0.0032)	(0.0039)	(0.0043)	
Region	0.0195^{***}	-0.0120***	-0.0304^{***}	-0.0402^{***}	-0.0290***	
	(0.0031)	(0.0021)	(0.0025)	(0.0038)	(0.0053)	
Year	-0.0114^{***}	-0.0019^{**}	0.0015	0.0079^{***}	0.0065^{*}	
	(0.0018)	(0.0009)	(0.0013)	(0.0023)	(0.0035)	
Total	-0.1706^{***}	-0.0326	-0.0171	-0.0771^{***}	-0.1926^{***}	
	(0.0311)	(0.0199)	(0.0186)	(0.0261)	(0.0377)	
Unexplained						
Tariff protection	-0.2099^{*}	-0.1421^{***}	0.0734	0.0644	0.0829	
	(0.1099)	(0.0512)	(0.0546)	(0.0659)	(0.0857)	
Demographic	-0.2554	-0.4563^{***}	0.2089^{**}	0.6915^{***}	0.8007***	
	(0.1966)	(0.1132)	(0.0999)	(0.1552)	(0.2172)	
Education	-0.1989***	-0.2135^{***}	0.0152	0.1055^{***}	0.2104^{***}	
	(0.0265)	(0.0207)	(0.0164)	(0.0404)	(0.0363)	
Employment share	-0.0900	0.1286^{**}	0.2473^{***}	0.0911	-0.0804	
	(0.0912)	(0.0511)	(0.0428)	(0.0565)	(0.0742)	
Sector	0.1163^{*}	-0.0223	0.2058^{***}	0.1506^{***}	0.1196^{***}	
	(0.0600)	(0.0320)	(0.0207)	(0.0208)	(0.0207)	
Region	0.2598^{***}	0.1817^{***}	0.1474^{***}	0.0914	-0.0574	
	(0.0581)	(0.0431)	(0.0560)	(0.0832)	(0.1332)	
Year	-0.3646^{***}	-0.2166^{***}	0.0679	0.0107	-0.1526	
	(0.1297)	(0.0802)	(0.0995)	(0.1382)	(0.1911)	
Constant	1.3020^{***}	1.0140^{***}	-0.6702^{***}	-0.8248^{***}	-0.4446	
	(0.2541)	(0.1575)	(0.1900)	(0.2427)	(0.3175)	
Total	0.5593^{***}	0.2735^{***}	0.2957^{***}	0.3803^{***}	0.4786^{***}	
	(0.0313)	(0.0215)	(0.0196)	(0.0269)	(0.0395)	
N	191065	191065	191065	191065	191065	
N Nonwhite men	126730	126730	126730	126730	126730	
N Nonwhite women	64335	64335	64335	64335	64335	

 Table (A11)
 RIF-decompositions: gender gap among nonwhites

Notes: * p < 0.10, ** p < 0.05, *** p < 0.01. Bootstrap standard errors in parentheses (100 replications). Tariff protection measured at time t - 2. Period is **1987**–1996. Survey weights from PNAD are used.

		Log(hourly wage)					
	(1)	(2)	(3)	(4)	(5)	(6)	
	$\tau = t$	$\tau = t-1$	$\tau = t - 2$	$\tau = t - 3$	$\tau = t - 4$	$\tau = t - 5$	
Women	-0.4708***	-0.4539^{***}	-0.4371^{***}	-0.4496***	-0.4427^{***}	-0.4434***	
	(0.0173)	(0.0191)	(0.0138)	(0.0147)	(0.0203)	(0.0181)	
Nonwhite	-0.1527^{***}	-0.1534^{***}	-0.1522^{***}	-0.1636^{***}	-0.1846^{***}	-0.1870^{***}	
	(0.0124)	(0.0120)	(0.0111)	(0.0103)	(0.0158)	(0.0160)	
Nonwhite \times Women	-0.0035	0.0085	0.0185	0.0230	0.0105	0.0221	
	(0.0197)	(0.0202)	(0.0163)	(0.0153)	(0.0224)	(0.0199)	
TP_{τ}	0.0195^{**}	0.0126	-0.0119	-0.0261	-0.0391	0.0025	
	(0.0096)	(0.0091)	(0.0186)	(0.0186)	(0.0240)	(0.0222)	
$TP_{\tau} \times Women$	-0.0057^{**}	-0.0092^{***}	-0.0107^{***}	-0.0017	-0.0005	-0.0002	
	(0.0025)	(0.0031)	(0.0025)	(0.0020)	(0.0050)	(0.0033)	
$TP_{\tau} \times Nonwhite$	-0.0020	-0.0019	-0.0036	0.0005	0.0077	0.0059^{*}	
	(0.0039)	(0.0032)	(0.0032)	(0.0025)	(0.0047)	(0.0032)	
$TP_{\tau} \times Nonwhite \times Women$	-0.0074	-0.0110^{*}	-0.0113^{***}	-0.0078**	-0.0002	-0.0031	
	(0.0053)	(0.0061)	(0.0040)	(0.0035)	(0.0065)	(0.0042)	
N	545787	483915	421342	357034	290686	290686	
Year FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	
State-Urban FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	
State trends	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	
Sector FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	
Individual controls	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	

Table (A12) Trade protection and hourly wage. Excluding automotive sector

Notes: * p < 0.10, ** p < 0.05, *** p < 0.01. Ordinary least squares (OLS) estimates reported with robust standard errors clustered at the state-urban level. Tariff protection measured at time t - 5, t - 4, t - 3, t - 2, t - 1, and t. The sample includes individuals aged 25 to 64 employed in sectors other than "Automobiles, trucks and buses; parts, comp. and other vehicles". Period is 1987–1996. Survey weights from PNAD are used. Control variables include age, squared age, educational attainment and number of children.

		Log(hourly wage)				
	1%	5%	1%	5%	1%	5%
	TP_t	TP_t	TP_{t-1}	TP_{t-1}	TP_{t-2}	TP_{t-2}
Women	-0.4603^{***}	-0.4152^{***}	-0.4427^{***}	-0.3996***	-0.4271^{***}	-0.3860***
	(0.0165)	(0.0134)	(0.0187)	(0.0162)	(0.0132)	(0.0104)
Nonwhite	-0.1518^{***}	-0.1444^{***}	-0.1506^{***}	-0.1420^{***}	-0.1499^{***}	-0.1409^{***}
	(0.0115)	(0.0099)	(0.0118)	(0.0108)	(0.0105)	(0.0091)
Nonwhite \times Women	0.0018	0.0187	0.0115	0.0247	0.0222	0.0324^{**}
	(0.0180)	(0.0150)	(0.0191)	(0.0171)	(0.0149)	(0.0126)
TP_{τ}	0.0194^{**}	0.0209^{**}	0.0127	0.0130	-0.0100	-0.0105
	(0.0088)	(0.0085)	(0.0092)	(0.0086)	(0.0176)	(0.0167)
$TP_{\tau} \times Women$	-0.0056**	-0.0055^{**}	-0.0093***	-0.0090**	-0.0105^{***}	-0.0099***
	(0.0025)	(0.0024)	(0.0032)	(0.0034)	(0.0024)	(0.0021)
$TP_{\tau} \times Nonwhite$	-0.0017	-0.0011	-0.0023	-0.0021	-0.0036	-0.0034
	(0.0037)	(0.0032)	(0.0031)	(0.0026)	(0.0031)	(0.0028)
$TP_{\tau} \times Nonwhite \times Women$	-0.0086^{*}	-0.0115^{***}	-0.0112^{*}	-0.0128^{***}	-0.0120^{***}	-0.0128^{***}
	(0.0049)	(0.0036)	(0.0057)	(0.0045)	(0.0037)	(0.0025)
N	549716	549716	487338	487338	424252	424252
Year FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark
State-Urban FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark
State trends	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark
Sector FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark
Individual controls	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark

Table (A13) Trade protection and hourly wage. Winsorizing

Notes: p < 0.10, p < 0.05, p < 0.05, p < 0.01. Ordinary least squares (OLS) estimates reported with robust standard errors clustered at the state level in parenthesis and state-urban-rural sector level. Tariff protection measured at time t - 3, t - 2, t - 1, and t. The sample includes individuals aged 25 to 64. Period is 1987–1996. Survey weights from PNAD are used. Control variables include age, squared age, educational attainment and number of children.

		Log(hourly wage)					
	1%	5%	1%	5%	1%	5%	
	TP_t	TP_t	TP_{t-1}	TP_{t-1}	TP_{t-2}	TP_{t-2}	
Women	-0.4329***	-0.3526***	-0.4166***	-0.3431***	-0.4016***	-0.3347***	
	(0.0137)	(0.0081)	(0.0156)	(0.0114)	(0.0105)	(0.0053)	
Nonwhite	-0.1463^{***}	-0.1238^{***}	-0.1451^{***}	-0.1245^{***}	-0.1433^{***}	-0.1226^{***}	
	(0.0100)	(0.0075)	(0.0105)	(0.0088)	(0.0093)	(0.0064)	
Nonwhite \times Women	0.0130	0.0336^{***}	0.0203	0.0370***	0.0288**	0.0436^{***}	
	(0.0155)	(0.0098)	(0.0157)	(0.0131)	(0.0116)	(0.0090)	
TP_{τ}	0.0203**	0.0187**	0.0127	0.0065	-0.0105	-0.0170	
	(0.0086)	(0.0084)	(0.0091)	(0.0073)	(0.0168)	(0.0152)	
$TP_{\tau} \times Women$	-0.0058**	-0.0058**	-0.0092***	-0.0076*	-0.0104^{***}	-0.0077***	
	(0.0024)	(0.0025)	(0.0031)	(0.0039)	(0.0022)	(0.0019)	
$TP_{\tau} \times Nonwhite$	-0.0017	-0.0022	-0.0023	-0.0019	-0.0040	-0.0033	
	(0.0034)	(0.0023)	(0.0028)	(0.0018)	(0.0029)	(0.0021)	
$TP_{\tau} \times Nonwhite \times Women$	-0.0101**	-0.0101***	-0.0116**	-0.0107***	-0.0116***	-0.0111***	
	(0.0041)	(0.0026)	(0.0044)	(0.0037)	(0.0026)	(0.0018)	
N	538775	494032	477537	437961	415737	381341	
Year FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	
State-Urban FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	
State trends	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	
Sector FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	
Individual controls	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	

 Table (A14)
 Trade protection and hourly wage. Trimming

Notes: * p < 0.10, ** p < 0.05, *** p < 0.01. Ordinary least squares (OLS) estimates reported with robust standard errors clustered at the state level in parenthesis and state-urban-rural sector level. Tariff protection measured at time t - 3, t - 2, t - 1, and t. The sample includes individuals aged 25 to 64. Period is 1987–1996. Survey weights from PNAD are used. Control variables include age, squared age, educational attainment and number of children.

			Log(hour	rly wage)		
	(1)	(2)	(3)	(4)	(5)	(6)
	$\tau = t$	$\tau = t - 1$	$\tau = t-2$	$\tau = t - 3$	$\tau = t - 4$	$\tau = t - 5$
Women	-0.5565***	-0.5387***	-0.5255***	-0.5416^{***}	-0.5411^{***}	-0.5400^{***}
	(0.0188)	(0.0213)	(0.0152)	(0.0164)	(0.0227)	(0.0196)
Nonwhite	-0.1490^{***}	-0.1494^{***}	-0.1479^{***}	-0.1586^{***}	-0.1792^{***}	-0.1834^{***}
	(0.0124)	(0.0122)	(0.0109)	(0.0101)	(0.0152)	(0.0153)
Nonwhite \times Women	-0.0145	-0.0044	0.0055	0.0087	-0.0107	0.0030
	(0.0216)	(0.0221)	(0.0176)	(0.0166)	(0.0240)	(0.0216)
TP_{τ}	0.0111	0.0052	-0.0113	-0.0220	-0.0371	0.0084
	(0.0093)	(0.0088)	(0.0176)	(0.0181)	(0.0241)	(0.0207)
$TP_{\tau} \times Women$	-0.0053**	-0.0094^{***}	-0.0104^{***}	-0.0015	0.0012	0.0006
	(0.0026)	(0.0028)	(0.0028)	(0.0020)	(0.0049)	(0.0032)
$TP_{\tau} \times Nonwhite$	-0.0018	-0.0018	-0.0034	0.0006	0.0077^{*}	0.0064^{**}
	(0.0039)	(0.0032)	(0.0032)	(0.0023)	(0.0045)	(0.0031)
$TP_{\tau} \times Nonwhite \times Women$	-0.0060	-0.0092	-0.0099**	-0.0063*	0.0044	-0.0006
	(0.0059)	(0.0065)	(0.0043)	(0.0037)	(0.0069)	(0.0045)
N	549716	487338	424252	359430	292574	292574
Year FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark
State-Urban FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark
State trends	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark
Sector FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark
Individual controls	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark

Table (A15) Trade protection and hourly wage. Controlling for part-time status

Notes: * p < 0.10, ** p < 0.05, *** p < 0.01. Ordinary least squares (OLS) estimates reported with robust standard errors clustered at the state-urban level. Tariff protection measured at time t - 5, t - 4, t - 3, t - 2, t - 1, and t. The sample includes individuals aged 25 to 64. Period is 1987–1996. Survey weights from PNAD are used. Control variables include age, squared age, educational attainment, number of children and part-time status.

	Log(hourly wage)
	(1)
Women	-0.4691***
	(0.0172)
Nonwhite	-0.1507***
	(0.0143)
Nonwhite \times Women	0.0072
	(0.0195)
TP_t	-0.0374^{*}
	(0.0213)
$\mathrm{TP}_t \times \mathrm{Women}$	-0.0032
	(0.0090)
$TP_t \times Nonwhite$	-0.0175
	(0.0132)
$TP_t \times Nonwhite \times Women$	0.0224^{*}
	(0.0130)
TP_{t-1}	0.0699*
	(0.0356)
$TP_{t-1} \times Women$	0.0754^{***}
	(0.0164)
$TP_{t-1} \times Nonwhite$	0.0171
	(0.0165)
$TP_{t-1} \times Nonwhite \times Women$	0.0143
	(0.0269)
TP_{t-2}	-0.0481*
	(0.0272)
$TP_{t-2} \times Women$	-0.0550***
	(0.0075)
$TP_{t-2} \times Nonwhite$	-0.0066
	(0.0058)
$TP_{t-2} \times Nonwhite \times Women$	-0.0288**
	(0.0113)
Observations	424252
Year FE	\checkmark
State-Urban FE	\checkmark
State trends	\checkmark
Individual controls	\checkmark
Sector FE	\checkmark

Table (A16) Trade protection and hourly wage. Lagged variables jointly

Notes: * p < 0.10, ** p < 0.05, *** p < 0.01. Ordinary least squares (OLS) estimates reported with robust standard errors clustered at the state level in parenthesis and stateurban-rural sector level. Tariff protection measured at time t-2, t-1, and t regressed jointly. The sample includes individuals aged 25 to 64. Period is 1987–1996. Survey weights from PNAD are used. Control variables include age, squared age, educational attainment and number of children.

		Log(hourly wage)						
	(1)	(2)	(3)	(4)	(5)	(6)		
	$\tau = t$	$\tau = t - 1$	$\tau = t - 2$	$\tau = t - 3$	$\tau = t - 4$	$\tau = t - 5$		
Women	-0.4709	-0.4533	-0.4369	-0.4494	-0.4408	-0.4422		
	$(0.0194)^{***}$	$(0.0215)^{***}$	$(0.0154)^{***}$	$(0.0170)^{***}$	$(0.0262)^{***}$	$(0.0238)^{***}$		
	$[0.0143]^{***}$	$[0.0128]^{***}$	[0.0086]***	$[0.0108]^{***}$	$[0.0252]^{***}$	$[0.0249]^{***}$		
Nonwhite	-0.1531	-0.1528	-0.1523	-0.1635	-0.1845	-0.1874		
	$(0.0121)^{***}$	$(0.0137)^{***}$	$(0.0126)^{***}$	$(0.0118)^{***}$	$(0.0157)^{***}$	$(0.0167)^{***}$		
	$[0.0166]^{***}$	$[0.0217]^{***}$	$[0.0251]^{***}$	$[0.0232]^{***}$	$[0.0215]^{***}$	$[0.0210]^{***}$		
Nonwhite \times Women	-0.0034	0.0078	0.0186	0.0229	0.0099	0.0219		
	(0.0224)	(0.0235)	(0.0194)	(0.0188)	(0.0269)	(0.0249)		
	[0.0183]	[0.0200]	[0.0192]	[0.0224]	[0.0214]	[0.0163]		
$\mathrm{TP}_{ au}$	0.0181	0.0120	-0.0104	-0.0245	-0.0387	0.0022		
	$(0.0064)^{***}$	(0.0101)	(0.0209)	(0.0202)	(0.0241)	(0.0227)		
	$[0.0054]^{***}$	[0.0104]	$[0.0047]^{**}$	$[0.0128]^*$	$[0.0164]^{**}$	[0.0102]		
$TP_{\tau} \times Women$	-0.0056	-0.0092	-0.0106	-0.0016	-0.0010	-0.0004		
	$(0.0024)^{**}$	$(0.0033)^{***}$	$(0.0023)^{***}$	(0.0019)	(0.0060)	(0.0038)		
	$[0.0020]^{**}$	$[0.0018]^{***}$	$[0.0018]^{***}$	[0.0016]	[0.0057]	[0.0040]		
$TP_{\tau} \times Nonwhite$	-0.0018	-0.0021	-0.0034	0.0006	0.0078	0.0060		
	(0.0037)	(0.0032)	(0.0033)	(0.0024)	$(0.0041)^*$	$(0.0030)^*$		
	[0.0027]	[0.0039]	[0.0043]	[0.0036]	$[0.0042]^*$	$[0.0030]^*$		
$TP_{\tau} \times Nonwhite \times Women$	-0.0075	-0.0109	-0.0115	-0.0079	-0.0002	-0.0031		
	(0.0053)	$(0.0063)^*$	$(0.0041)^{***}$	$(0.0038)^{**}$	(0.0067)	(0.0047)		
	$[0.0042]^*$	$[0.0046]^{**}$	$[0.0036]^{***}$	$[0.0045]^*$	[0.0056]	[0.0038]		
N	549716	487338	424252	359430	292574	292574		
Year FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		
State-Urban FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		
State trends	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		
Sector FE	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		
Individual controls	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark	\checkmark		

Table (A17) Trade protection and hourly wage. Changing standard errors

Notes: * p < 0.10, *** p < 0.05, *** p < 0.01. Ordinary least squares (OLS) estimates reported with robust standard errors clustered at the state level in parentheses and state-urban-rural sector level in brackets. Tariff protection measured at time t - 5, t - 4, t - 3, t - 2, t - 1, and t. The sample includes individuals aged 25 to 64. Period is 1987–1996. Survey weights from PNAD are used. Control variables include age, squared age, educational attainment and number of children.



Figure (A5) Contribution of trade liberalization to unexplained term over selected quantiles of the wage distribution: with and without employment correction

Notes: Common Y-axis for all subfigures.



Figure (A6) Contribution of trade liberalization to unexplained term over selected quantiles of the wage distribution and across different lag specifications

Notes: Common Y-axis for all subfigures.



(a) Racial gap among men (white vs. nonwhite)

(b) Racial gap among women (white vs. nonwhite)



(c) Gender gap among whites (men vs. women)

(d) Gender gap among nonwhites (men vs. women)

Figure (A7) Contribution of trade liberalization to unexplained term over selected quantiles of the wage distribution and across different lag specifications: with and without controlling for sector fixed effects.

Notes: Common Y-axis for all subfigures.



Figure (A8) Heterogeneity TP coefficients across sectors

Notes: The plots show the heterogeneous effects of TP across sectors and gender-race groups. In the x-axis the numbers denote different sectors, as follows: 1. Agricultural products, 2. Apparel, 3. Automobiles, trucks and buses; parts, components and other vehicles, 4. Cellulose, paper and printing, 5. Chemical elements and products, 6. Electrical and electronic equipment, 7. Footwear, 8. Machinery and Tractors, 9. Meat packing, dairy industry, vegetable and other food products, 10. Mining Products, 11. Non-metallic minerals, 12. Nontradables, 13. Oil and Coal extraction, 14. Oil refining and petrochemicals, 15. Pharmaceutical and perfumery products, 16. Plastic products, 17. Rubber products, 18. Steel, non-ferrous and other metal products, 19. Textile products, 20. Unclassified manufacturing, 21. Wood products and furniture.