Competition and the Racial Wage Gap: Testing Becker's Model of Employer Discrimination[*]

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Abstract

According to Becker's (1957) theory of taste based discrimination, pure economic rents are necessary for discrimination to be observed in the labor market. Increased competition and reduced rents in the market for final goods should therefore lead to reduced labor market discrimination. We look at the episode of trade liberalization in Brazil in the beginning of the 1990s to study the effect of increased competition in the final goods market on racial discrimination in the labor market. Changes in tariffs and initial employment structures are used to show that, in locations where there was a larger increase in exposure to foreign competition between 1990 and 1995, there were also larger declines in the conditional racial wage gap between 1991 and 2000. As predicted by theory, the initial wage gap and its decline seemed to be more pronounced in regions with more employment in concentrated sectors and with stronger preferences for discrimination. The change in the racial wage gap was not associated with changes in returns to productive attributes, in the structure of employment, or in other observed labor market outcomes.

Keywords: discrimination, racial wage gap, competition, labor market, trade reform, Brazil
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1. Introduction

Becker's (1957) theory of taste based discrimination has a clear yet surprising implication. According to the model, higher competition in the market for final goods should lead to lower discrimination against minorities in the labor market. This clear cut conclusion comes from the fact that discrimination in the labor market requires the existence of pure economic rents. Only in the presence of rents can discriminating employers choose to pay certain workers below the value of their marginal productivities. This behavior creates opportunities for non-discriminating employers to enter the market or to hire more workers, paying higher wages and earning higher profits. Under perfect competition, when firms earn zero profit and pay each input the value of its respective marginal productivity, there is simply no scope for discrimination. Therefore, some sort of market imperfection – such as the absence of free entry in the presence of decreasing returns to scale, or oligopolistic or monopolistic market structures – is required for labor market discrimination to be observed as an equilibrium outcome. In these contexts, increased competition in the market for final goods, by reducing pure economic rents, should lead to reduced discrimination in the labor market. Despite having been established almost 60 years ago, this implication of the theory of taste based discrimination has been subject to surprisingly little scrutiny.

This paper uses the episode of trade liberalization that took place in Brazil during the early 1990s to test whether increased competition in the final goods market is associated with reduced discrimination against blacks in the labor market. Specifically, we analyze whether local labor markets that experienced a larger increase in exposure to international trade also experienced a larger reduction in the conditional wage gap between white and black workers. Local labor markets are defined as sets of geographically contiguous municipalities, representing roughly self-contained labor markets, classified as "micro-regions" by the Brazilian Census Bureau (Instituto Brasileiro de Geografia e Estatística, IBGE). We use changes in tariffs by sector and initial employment structures to calculate the relevant change in average tariffs from the perspective of each local labor market. This local tariff is obtained as a weighted average of the sector specific tariffs, where the weights are functions of employment shares and elasticities of demand for labor (Kovak, 2013). We concentrate on the reductions in tariffs that took place between 1990 and 1995, the main period of trade liberalization in Brazil, and look at microdata from the 1991 and 2000 censuses.

Our empirical strategy is implemented in two stages. First, by running Mincerian regressions, we estimate the conditional wage gap between white and black workers for each local labor market (micro-region) in 1991 and 2000. Following, in the second stage, we estimate the impact of increased openness on labor market discrimination by running, at the level of local labor markets, a regression of the estimated change in the conditional racial wage gap on the change in tariffs.

The results show that the conditional wage gap between whites and blacks fell more in regions associated with larger reductions in tariffs, or, in other words, in regions that experienced larger increases in exposure to international competition, as predicted by the theory of taste based discrimination. According to our preferred specification, a reduction in tariffs equivalent to the average observed in the sample (9.7 percentage points) would lead to a reduction in the racial wage gap of 18%. In fact, during the period of analysis, the conditional racial wage gap in Brazil remained roughly stable, so the liberalization process seems to have helped to offset a trend towards increased racial inequality in the labor market. Our main result is robust to the composition of the sample and is not correlated with changes in returns to productive attributes, in the structure of employment, or in other observable labor market outcomes. In particular, the results are not associated with a Stolper–Samuelson effect, which might lead to relative gains to low skill workers and, possibly, to a reduction in the racial wage gap (if blacks are relatively less qualified in terms of unobserved skills).

In order to provide additional supporting evidence that the estimated effect is indeed driven by the mechanism highlighted in the theory of taste based discrimination, we also analyze its heterogeneity along certain margins. We show that the initial conditional racial wage gap and the impact of trade liberalization tended to be stronger in locations with more employment in concentrated sectors and with stronger tastes for discrimination. According to theory, these are locations that should display initially higher levels of labor market discrimination and, therefore, that should have responded more to increased competition.

There is a small but already established literature on the effect of increased competition on discrimination against minorities, focusing mostly on gender. Ashenfelter and Hannan (1986) look at the banking sector in the US and conclude that women have lower employment rates in more concentrated markets. Other papers use the deregulation of the banking and transportation sectors in the US as natural experiments on increased competition, finding results that support the theory (Black and Strahan, 2001, Peoples and Talley, 2001, and Levine et al, 2008). Zweimüller et al (2008) analyze cross-country data and report a negative correlation between market friendly institutions – such as openness to trade and protection of property rights – and the gender wage gap, meaning that higher economic freedom is associated with lower discrimination against women in the labor market.

A recent literature, more closely related to this paper, explores the impact of international trade on wage inequality across genders. Black and Brainerd (2004) and Jacob (2006) analyze the impact of increased competition from international trade on the gender wage gap in the US and India, respectively. Black and Brainerd (2004) do not have a period of institutional change corresponding to a trade reform, so their empirical setting cannot be seen as a natural experiment. Jacob (2006), on the other hand, does analyze a period of tariff reductions and institutional reforms in India. Both papers find a negative correlation between exposure to international trade and the gender wage gap.[1] Various other papers apply similar methodologies to analyze the gender wage gap in other contexts – sometimes involving an explicit process of trade liberalization and others not – including Mexico, South Korea, Taiwan, and groups of developed and developing countries (Artecona and Cunningham, 2002, Berik et al, 2004, Oostendorp, 2009, Wolszczak-Derlacz, 2013).[2] As a whole, this body of research finds conflicting evidence on the impact of trade liberalization on the gender wage gap (see, for example, the review in Anderson, 2005).

A limitation of these papers is the use of industry (or, in one case, occupation) as the unit of analysis, which can be problematic given the relatively small number of observations and the fact that often they cannot be seen as unified and independent labor markets. In the presence of geographic labor market segmentation and spillovers across industries in the same area, this modelling choice is difficult to justify. This is particularly worrisome in the case of developing countries, where migration, and therefore reallocation of labor across locations within the same industry, is limited (see, for example, Topalova, 2010 for the case of India, or Dix-Carneiro and Kovak, 2014 for Brazil; for a review of the literature on reallocation costs and the impacts of trade reforms, see Dix-Carneiro, 2012). In addition, the focus on wage differentials across genders presents considerable drawbacks. Women's labor supply decisions on the extensive margin can be very important. This affects the predictions of Becker's (1957) theory in non-trivial ways, potentially weakening the link between competition in the market for final goods and labor market discrimination (as observed in wage differentials). In this case, the pattern and evolution of selection into the labor market, mostly ignored in the literature, become of first order relevance in determining the behavior of wage differentials across genders and the impact of increased competition.

Our paper combines different strategies adopted before in the economics literature, but that have not yet been jointly applied to analyze the relationship between competition and discrimination. As Topalova (2010) and Kovak (2013), we focus on local labor markets as the unit of analysis and use the initial structure of employment to calculate the relevant tariff reduction from the perspective of each market. Following, we use the Brazilian trade reform from the 1990s as a natural experiment generating an exogenous increase in competition in the final goods market. We then look at the impact of this exogenous change on the conditional racial wage gap using an approach inspired by Charles and Guryan (2008). By looking at markets that are relatively self-contained and exploring an exogenous shock, we are able to arguably identify the change in equilibrium outcomes of specific labor markets, therefore improving upon

[1] Jacob (2006) also analyzes the impact of trade on discrimination against lower castes, but finds no robust effect.

[2] Juhn et al (2013) analyze the impact of NAFTA on the gender wage gap in Mexico, but highlight a mechanism that is different from discrimination. They show that the reduction in tariffs associated with NAFTA led industries to adopt new technologies that reduced the demand for physical labor, favoring women employed in blue collar occupations and reducing the gender inequality in the labor market. Though this mechanism is different from that discussed in the paper, it will be one of the alternative hypotheses considered when conducting our robustness exercises.

the existing literature on trade liberalization and discrimination. In addition, differently from this literature, we concentrate on racial discrimination among prime aged men, rather than on gender discrimination, making participation decisions a second order issue and bringing the empirical exercise closer to the theory.

This paper also speaks to a broader literature on the impacts of globalization on inequality in developing countries, based on the Heckscher-Ohlin model and the Stolper-Samuelson theorem. Goldberg and Pavnik (2007) review this literature and do not find robust evidence supporting the predictions of the theory. In fact, most of this literature documents increased inequality as a result of increased openness to international trade. In the case of Brazil, controversy still persists, with some studies finding a reduction in inequality due to the 1990s trade reform, and others pointing to null or even opposite effects (see, for example, Arbache and Menezes-Filho, 2000, Arbache and Corseuil, 2004, Gonzaga et al, 2006, Ferreira et al, 2010, Kovak, 2013, Dix-Carneiro and Kovak, 2015). We present evidence on one specific impact of trade reforms that, in the case of Brazil, has unequivocally led to a reduction in labor market inequality.

The remainder of the paper is organized as follows. Section 2 sketches the simplest version of the model of taste based employer discrimination proposed by Becker (1957) and discusses its empirical implications. Section 3 describes the process of trade reforms implemented in Brazil between 1988 and 1994. Section 4 discusses our identification strategy, explains the implementation of our empirical exercise, and presents the methodology for constructing tariffs at the level of local labor markets. Section 5 describes the data and the variables used, while Section 6 presents the results. Finally, Section 7 concludes the paper.

2. The Model of Taste Based Employer Discrimination

This model follows closely the original framework of Becker (1957). We outline the basics of the classic employer discrimination model to help guide our empirical discussion. Consider a population that is heterogeneous in terms of race, a non-productive attribute. There are individuals who belong to the racial minority – blacks (b) – and individuals who belong to the racial majority – whites (w) –, and both groups possess the same set of productive skills. In other words, blacks and whites are assumed to be perfect substitutes in production.

Employers potentially discriminate against members of the minority (blacks), in the sense that they attach a negative value to interacting with them or to having them as employees (assume, for example, that employers belong to the majority and that there is prejudice against the minority in this society). Following Becker (1957), we assume that this prejudice can be summarized by a coefficient of discrimination $\delta \geq 0$, which measures in relative monetary units the disutility that a given employer has when interacting with a member of the minority. In this setting, employers do not maximize profits, but instead a combination of profits and the disutility from interacting with members of the minority. Under these assumptions, the problem of an employer with coefficient of discrimination δ is

$$\max_{\{L_b,L_w\}}\{F(L_b+L_w)-(1+\delta).W_b.L_b-W_w.L_w\}, \tag{1}$$

where the price of the final good is normalized to 1, $F(\cdot)$ is the production function, L_i indicates the number of workers of race i, and W_i is the market wage for race i, with $i \in \{b,w\}$. δ can be interpreted as an additional subjective cost, above that represented by the wage, that the employer perceives when hiring someone from the minority group. The fact that δ represents tastes for discrimination as a proportion of the real wage is a simplifying assumption and has no consequence in terms of the qualitative implications of the model.

The problem of the employer is to choose L_b and L_w to maximize the function above. In a competitive labor market, where wages are taken as given, the first order conditions to this problem are

$$F_{L_w^*} \leq W_w, \text{ with equality if } L_w^* > 0, \text{ and} \tag{2}$$
$$F_{L_b^*} \leq (1+\delta).W_b, \text{ with equality if } L_b^* > 0. \tag{3}$$

Since L_b and L_w are assumed to be perfect substitutes in production, $F_{Lb} = F_{Lw}$. This implies that, typically, an employer hires only black or white workers, but not both simultaneously. If the coefficient of discrimination is such that $W_w < W_b.(1+\delta)$, the employer hires only white workers and, otherwise, he hires only black workers. In this setting, market forces induce employers who do not discriminate or who discriminate less to hire only black workers, and those with higher coefficients of discrimination to hire only white workers.

In order to discuss some features of the equilibrium characterizing this economy, assume that there is a continuum of measure N_e of employers, over which δ is distributed according to some distribution function $H(\delta)$. Given the equilibrium wages W_b and W_w, there must be some employer with $\delta = \delta_m$ for which the following condition holds

$$W_w = W_b(1+\delta_m). \tag{4}$$

We call this employer with $\delta = \delta_m$ the marginal employer. He is the employer who is indifferent between hiring workers from the minority or the majority. Alternatively, he is the employer with the highest coefficient of discrimination who is still willing to hire workers from the minority.

The coefficient of discrimination of the marginal employer, δ_m, corresponds to the equilibrium wage gap between whites and blacks: $(1 + \delta_m) = W_w/W_b$. If $\delta_m = 0$, there is no racial wage gap in equilibrium, despite the fact that there may be some employers with $\delta > 0$ in the population. So the determination of the equilibrium racial wage gap in this economy is equivalent to the determination of the identity of this marginal employer. A simple example helps clarify the relevant forces at work here.

Following Becker (1957), consider a fixed supply S_b of workers from the minority group and focus on a partial equilibrium analysis looking at the demand for minority workers as a function of the racial wage gap. For an employer who hires only minority workers, define the optimal demand for labor as a function of δ as $L_b^*(\delta)$, determined implicitly from $F_L(L_b^*) = (1 + \delta).W_b$. Employers hiring minority workers are those with $\delta \leq \delta_m = (W_w/W_b) - 1$. So the demand for minority workers as a function of W_b/W_w can be written as

$$D_b\left(\frac{W_b}{W_w}\right) = N_e \int_0^{(W_w/W_b)-1} L_b^*(\delta).\,dH(\delta), \qquad (5)$$

where N_e is the number of employers.

The equilibrium racial wage gap in this economy, as well as the identity of the marginal employer of minority workers, is determined from the equality between the supply and demand of minority workers: $D_b(W_b/W_w) = S_b$. This simple model highlights the forces intervening in the determination of W_b/W_w and informs our empirical analysis.

First, there can only be labor market discrimination in equilibrium if the number of employers N_e is given and if there are decreasing returns to scale. With free entry, and a pool of potential employers with $\delta = 0$, non-discriminating employers would enter the market until discrimination were eliminated. This would have to be the case, since $F_{Lb} > W_b$ implies an allocative inefficiency that opens up opportunities of increased profits for non-discriminating firms. Similarly, without decreasing returns to scale, employers with $\delta = 0$ would grow and eventually take over the market, also eliminating any observed wage gap between minority and majority. So taste based discrimination in the labor market requires some degree of inefficiency and the existence of pure economic rents. Increased competition in the form of new entrants and reduced rents should reduce the equilibrium level of discrimination.

In addition, the model also reveals other characteristics of markets where we should expect to see higher levels of labor market discrimination. For a given supply of minority workers, the distribution of preferences for discrimination (δ) is the key determinant of the observed wage differential between races. Not surprisingly, a homogeneous rightward shift in the coefficient of discrimination should increase observed labor market discrimination. More surprising maybe is the fact that increases in the size of the minority, for a given distribution of δ and number of employers, should also increase labor market discrimination. This implication comes from the fact that an increase in the supply of minority workers, in equilibrium, would have to induce an increase in δ_m (the marginal employer of minority workers would have to be someone with a higher δ), which could only happen through an increase in the racial wage gap.

Figure 1 illustrates these two points in a graph where the equilibrium demand for minority workers (in the horizontal axis) is plotted as a function of the racial wage gap (in the vertical axis). The relative demand curve D_b^2 represents a rightward shift in the discrimination coefficient in comparison to D_b^1, while the supply curve S_b^2 represents an increased supply of minority workers in comparison to S_b^1. The movement from point I to point II summarizes the effect of an increase in prejudice among employers on observed labor market discrimination. The movement from point I to point III portrays the effect of a rightward shift in the supply of minority workers. As illustrated in the figure, both changes lead to an increase in the equilibrium racial wage gap.

Two implications of this model should be kept in mind for our later empirical discussion. First, increased competition should lead to reduced labor market discrimination. Second, one should expect to see higher levels of discrimination where production in the final goods market is more concentrated, where there is a higher level of prejudice among the overall population, and where there is a higher share of minority workers.

3. The 1990s Trade Reform in Brazil

From 1957 to 1988, there was little change in trade legislation in Brazil.[3] During this period, there was widespread use of non-tariff barriers, including quotas and lists restricting the variety and quantity of goods that could be imported. The redundancy of tariffs and the existence of various additional taxes – such as, for example, the additional freight fee for renewal of the Merchant Navy –, besides 42 special regimes allowing for tariff exemptions or reductions, generated a heavily bureaucratic structure, distorting relative prices. As a result, Brazil had very little exposure to competition from foreign goods.

[3] Our description of the trade reforms is based to a great extent on Kume et al (2003). For further details on the process of trade liberalization in Brazil, refer to these authors.

Planning of the trade reform started in 1987, during the Sarney government. But implementation was halted due to pressure from interests groups who wanted to maintain the trade barriers in certain sectors. Between 1988 and 1989, the government managed to eliminate only the redundant part of the tariff structure. The process of liberalization was reinitiated under the Collor and Franco governments. From 1991 to 1993, there was virtual elimination of non-tariff barriers and special regimes. Additionally, a timeline for the gradual reduction of tariffs was approved and implemented. Initially planned to be executed until 1994, the timing was anticipated and by the end of 1993 the major part of tariff reductions had already taken place. In a further movement towards openness, the Cardoso government reduced some additional tariffs in 1994, as part of a broader effort focused on economic stabilization (Real Plan).

Figure 2 portrays the evolution of nominal tariffs in Brazil between 1987 and 1998 for the 10 sectors with the highest shares of employment (data from Kume et al, 2003). There is a clear pattern of generalized reduction and homogenization of tariffs up until 1994, when the minimum levels are attained in most sectors. During this period, nominal tariffs fell, on average, by 43 percentage points (75%). As a result, the share of trade in the Brazilian GDP increased from around 15% in the second half of the 1980s, to 22% in 2000 (data from the World Development Indicators).

As the figure also makes clear, there was a mild reversion in the trend towards increased openness after 1995. This was mostly a response of the Brazilian government to domestic pressures derived from the international financial crises of the late 1990s. In this context, the government raised tariffs and reintroduced some red tape in the imports of certain manufactures, but these changes were minor in comparison to the extent of the reductions in tariffs and non-tariff barriers from the first half of the 1990s.

Figure 3 presents the evolution of Brazilian imports by sector from 1985 to 1999. For each sector, the real value of imports (in 1999 Brazilian Reais) is normalized to 1 in 1985 (data from Gonzaga et al, 2006). Apart from apparel and textiles, imports are stable until 1991, without any clear trend. But, starting in the first half of the 1990s, there is a sharp change in trend towards increased imports in most of the sectors. Even for apparel and textiles, for which imports started increasing already before 1990, there is a strong acceleration in the growth rate after 1992.

Two characteristics of the trade reform in Brazil are particularly important for our empirical strategy. First, it was very sharp and concentrated in time: in a period of roughly 5 years, trade barriers were aggressively reduced and large increases in imports were observed. Second, liberalization was driven by a centralized decision at the federal level, unrelated to economic conditions in local labor markets. We come back to these two points in our discussion on the empirical strategy in the next section.

4. Methodology

4.1 Identification

We use the episode of trade liberalization in Brazil as a natural experiment to assess the impact of increased competition on discrimination in the labor market. We combine the reduction in tariffs triggered by the reforms with the regional variation in the structure of employment to explore the heterogeneous effects of the reforms on local labor markets.

As mentioned in the previous section, the reforms were concentrated in time almost in a discrete fashion. Though some tariffs were eliminated and others reduced between 1987 and 1990, these first changes represented mostly a rationalization of the tariff structure and had little impact on the Brazilian economy (Kume et al, 2003). It was really only in the 1990s that the effects of liberalization started being felt. Following Kovak (2013), our analysis focuses on the reductions in tariffs observed between 1990 and 1995, and uses data from the 1991 and 2000 censuses as representing, respectively, the starting point before the reforms and the new equilibrium in the Brazilian labor market following liberalization.

A potential concern with this identification strategy is that reductions in tariffs might have been determined by the political influence of interest groups, which in turn might have been affected by labor market conditions. In this hypothetical setting, tariff reductions would be endogenous to labor market conditions and the identification strategy would be compromised. Figure 4 shows that this does not seem to be the case. The figure plots, by sector of activity, the 1990-1995 tariff reduction in the vertical axis (percentage points) against the initial tariff level in the horizontal axis. The pattern shows that the reforms led to a homogenization and rationalization of tariffs: sectors with initially higher tariffs experienced larger subsequent reductions in tariffs. The average tariff reduction of 60% during this period was accompanied by a reduction of 53% in the sectorial dispersion of tariffs (standard deviation).

The unit of analysis in our empirical exercise is a local labor market, defined as a micro-region, not a sector of economic activity as in Figure 4. So Figure 5 reproduces the same diagram from Figure 4, but for average tariffs at the level of local labor markets. We discuss how these average tariffs are constructed at the end of this section, but mention the data before to inform our discussion on identification. The pattern is even more extreme than that observed in Figure 4: micro-regions with initially higher tariffs experienced larger subsequent reductions in tariffs. In Figure 5, this relationship is linear and close to deterministic. Again, average reductions in tariffs in micro-regions did not seem to be correlated with specific labor market conditions. It is worth mentioning that the average tariff reduction by micro-regions is lower than that observed across sectors. This comes from the fact that employment shares are used to construct average tariffs by micro-region, and some sectors with large employment shares had very small reductions in tariffs after 1990 (this is the case, for example, for the agricultural sector, which accounted for half of the employment outside of the services sector; see Appendix Table A.1).

4.2 Empirical Strategy

Combining the strategies developed by Charles and Guryan (2008) and Kovak (2013), we estimate the impact of the reduction in tariffs on the racial wage gap in two stages. First, we run individual level Mincerian regressions to estimate the conditional wage gap between whites and blacks in each local labor market for 1991 and 2000. Following, the estimated conditional wage gaps are used to construct the dependent variable for the second stage: the change in the wage gap between 1991 and 2000. The change in the racial wage gap is then regressed on the change in tariffs between 1990 and 1995. In the second stage, the unit of analysis is a local labor market, which we define as a micro-region (as mentioned before, a set of contiguous municipalities with similar geographic and socioeconomic conditions, defined by the Brazilian Census Bureau, IBGE). We use a micro-region as a local labor market, instead of a municipality, due to the reduced number of observations for smaller municipalities in the census microdata, which makes it difficult to estimate the racial wage gap with precision in these cases.

4.2.1 First Stage

In the first stage, we estimate the conditional racial wage gap for 1991 and 2000, controlling for correlates of individual productivity. For each year *t*, we estimate individual level Mincerian regressions by OLS. Our basic specification is the following:

$$\ln wages_{ijt} = \alpha_t + \sum_j \delta_{jt} white_{ijt} \times micro\_region_{jt} + \gamma_t{}' \boldsymbol{X}_{ijt} + \varepsilon_{ijt}, \quad (6)$$

where *i* indicates individual, *j* indicates micro-region, *wages* denote hourly earnings, *white* is a dummy for race (equal to 1 for whites and Asians, and 0 for blacks and mixed), *micro_region* is a dummy equal to 1 for region *j*, $\boldsymbol{X}$ is a vector of demographic controls, and ε is a random term. In the benchmark specification, the vector $\boldsymbol{X}$ includes age, age squared, an entirely flexible function of years of schooling (one dummy for each completed year of schooling), a dummy indicating urban residence, and a dummy for each micro-region. This same specification is estimated separately for 1991 and 2000.

Our focus in this first stage is on the coefficient δ_{jt}, which we call the conditional racial wage gap for local labor market *j* in year *t*. Specifically, δ_{jt} indicates the wage advantage (in approximate percentage terms) of a white worker in comparison to a black worker with similar observable characteristics. The fact that we estimate the equation separately for each year means that parameters can change from one year to the other, reflecting potential changes in returns to productive attributes due to labor market conditions. In some robustness exercises, we run equation 6 separately for each micro-region, allowing for the parameters in γ also to vary with j. Though this specification is more flexible, allowing the model to better capture the conditions of each local labor market, it also demands much more from the data, leading to estimates of the conditional racial wage gap δ_{jt} that can be less precise in smaller samples. Therefore, we only use this specification to assess the robustness of our benchmark results.

4.2.2 Second Stage

The estimated racial wage gaps, $\hat{\delta}_{jt}$'s, are used to construct the change in the racial wage gap over time for each local labor market j: $\Delta(\hat{\delta}_j) = \hat{\delta}_{jt} - \hat{\delta}_{jt-1}$. This variable becomes the dependent variable in our second stage regression, estimated by WLS:

$$\Delta(\hat{\delta}_j) = \mu + \beta\Delta(tariff_j) + \lambda'\boldsymbol{W}_j + \omega_j, \tag{7}$$

where $\Delta(tariff)$ represents the change in average tariffs between 1990 and 1995, $\boldsymbol{W}$ is a vector of controls, and ω is a random term. The controls included in the vector $\boldsymbol{W}$ capture changes in aggregate market conditions in the micro-regions, which might affect the determination of wages and, indirectly, the racial wage gap. We discuss the specific variables included later on in this section. Following Charles and Guryan (2008), the second stage regression is weighted by the precision of the first stage estimates (inverse of the standard-error of $\Delta(\hat{\delta}_j)$).

Our parameter of interest in the second stage is β, which captures the impact of the change in average tariffs on the conditional racial wage gap. The theory of taste based discrimination presented in Section 2 predicts that increased competition in the market for final goods should lead to reductions in the conditional racial wage gap, so that we should expect $\beta > 0$. In other words, reductions in tariffs should be associated with reductions in the wage advantage that whites have in relation to blacks.

Our discussion on identification makes it clear that changes in tariffs were not driven by local labor market conditions and, therefore, were not endogenous to the issue that we want to analyze here. Still, there remains the possibility that changes in tariffs might have affected the racial wage gap through channels other than that predicted by the theory of taste based discrimination. This is the main concern in the estimation of our second stage and guides our choice of control variables to be included in $\boldsymbol{W}$.

One possibility in this direction comes from other labor market effects of trade liberalization. The Heckscher-Ohlin model predicts that, after liberalization, a country should shift its production towards goods intensive in its relatively abundant factor, leading to an increase in the relative return to this factor. From this perspective, in the case of Brazil, one should expect to see a shift in production towards sectors intensive in low skill labor, accompanied by a reduction in the wage differential between high and low skill workers. Our first stage specification already controls for schooling, partially accounting for the effects associated with changes in returns to productive attributes. Still, if one thinks that this same Heckscher-Ohlin effect should operate in relation to unobserved skills, it might be the case that it would interfere in the relationship between changes in tariffs and changes in racial wage gaps. This would be the case, for example, if blacks had access to education of lower (unobserved) quality, and liberalization also reduced wage differentials across different (unobserved) qualities of education. For this to be the case, the change in returns to unobserved productive attributes (quality of schooling) would have to accompany the change in returns to observed productive attributes (years of schooling), and blacks would have to be less skilled than whites in terms of unobserved attributes.

To minimize this potential problem, we include as controls in our vector $\boldsymbol{W}$ the changes in average wages between 1991 and 2000 by level of schooling: up to 7 years (less than elementary), from 8 to 10 years (complete elementary and high-school drop-outs), from 11 to 14 years (complete high-school and college drop-outs), and 15 years or more (college graduates). By controlling for changes in returns to productive attributes, we are accounting for Heckscher-Ohlin effects in local labor markets. Even if we cannot measure returns to unobserved attributes, this strategy should go a long way towards shedding light on whether the issue discussed in the previous paragraph is a threat to identification. If changes in returns to unobserved attributes are driving the results, it must be the case that they are similar to changes in returns to observed productive attributes. So, by controlling for the latter, we are capturing labor market equilibrium conditions associated with returns to different skill levels, and indeed partially controlling for the former. If the inclusion of these controls does not affect the coefficient of interest, it is because the correlation between reduction in tariffs and changes in the racial wage gap is not driven by changes in returns to productive attributes (observed or unobserved).

There was continuous improvement in schooling levels in Brazil during this period. So we also control directly for the change in the supply of workers by skill level and race, which might be associated with similar changes in the distribution of unobserved skills. We include in $\boldsymbol{W}$ the share of workers by years of schooling and race, using the same educational classification discussed before. To control for other potential changes in educational policies, which might as well affect unobserved skills, we control for the change in the supply of public education. Since blacks are relatively poorer, increases in the quality of the public educational system may affect relatively more blacks than whites. We do not have a direct measure of the quality of the educational system – such as results of standardized exams – at the micro-region level for the period of analysis, so we control for the change in the number of public schools normalized by the number of children.

Another potential effect of trade liberalization is through investments in technology. Juhn et al (2013) find that NAFTA reduced the gender wage gap in Mexico not because of reduced discrimination, but because of investments in technology that reduced the demand for physical labor. Though this possibility seems less plausible in the context of races, we still account for it by including in our vector of controls $\boldsymbol{W}$ the share of blue collar workers in each micro-region. If there is some technological change in response to liberalization, one should expect it to be partly reflected on changes in the relative share of blue collar workers. We also control for unemployment and informality rates (among salaried workers), to account for other margins of labor market adjustment, and for migration, which might affect the response of a local labor market to exogenous shocks (as suggested by Cadena and Kovak, 2013). These labor market changes could have been affected by the trade reforms and could have heterogeneous effects across races. Additionally, they help control for broader patterns in the Brazilian labor market.

Our benchmark specification also includes dummies for the 5 main geographic regions in Brazil. In some specifications, we replace the 5 geographic region dummies by 27 state dummies.

4.3 Calculating Average Tariffs for Local Labor Markets

Trade policy in Brazil is determined at the federal level, so tariffs are the same for each sector irrespectively of location. But the structure of employment varies across locations, so the impact of a given reduction in tariff is not homogeneous across the territory. To take advantage of this fact and explore the differential impact of the trade reform across local labor markets with different initial structures, we follow Kovak (2013). Kovak (2013) proposes a methodology for calculating average tariffs for local labor markets based on a model specifically developed to analyze the regional impacts of trade liberalization. His model treats each region (local labor market) as a specific-factors economy with two inputs: labor and an immobile factor. Labor is supplied inelastically in each region and can move across sectors, but cannot migrate across regions. The immobile factor, which we call capital here, cannot move across sectors or across regions, and represents location specific factors that augment the productivity of labor in a given industry.[4] Technology is assumed to have constant returns to scale and to vary across sectors, but not within sectors across regions. Finally, there is a single national market for the goods produced in the different regions.

This model justifies the use of a measure of tariffs at the subnational level that is similar to a formulation that was already present in the empirical literature, but had no theoretical basis. Consider an economy with sectors $r = 1,\ldots, R$, where R represents the non-tradable sector. From the perspective of local labor market j, the relevant variation in tariffs between period t-1 and t is

[4] These could include natural resources, land, agglomeration effects, and specific fixed capital, as suggested by Kovak (2013).

$$\Delta\left(tariff_j\right) = \sum_{r \neq R} \psi_{jr} \{ln(1 + tariff_{rt}) - ln(1 + tariff_{rt-1})\}, \tag{8}$$

where $\psi_{jr} = \frac{L_{jr}\epsilon_{jr}}{\sum_{r \neq R} L_{jr}\epsilon_{jr}}$, L_{jr} indicates employment in sector r in local labor market (micro-region) j, $\epsilon_{jr} = \frac{\sigma_{jr}}{\theta_{jr}}$ is the elasticity of the demand for labor, σ_{jr} is the elasticity of substitution between inputs, and θ_{jr} is the share of capital in total cost.

The relevant change in tariff faced by a local labor market is a weighted average of the changes in tariffs experienced by the different sectors, where the weights are functions of the elasticities of labor demand and employment levels observed in each sector. Notice that the non-tradable sector is not explicitly considered in the weighted average, a result that comes directly from the theoretical model (in fact, according to the model, the relevant variation for the non-tradable sector is equivalent to the average variation across sectors). Since changes in employment and elasticities may be endogenous, only values from the initial period (1991) are considered in the calculation.

In practical terms, given the limited information available, some simplifying assumptions are needed. Following Kovak (2013), we assume that the technology is Cobb-Douglas, which implies a constant elasticity of substitution for every j and r: $\sigma_{jr} = 1$. Second, we assume that the share of capital in total cost (θ_{jr}) varies across sectors, but not across regions, so that $\theta_{jr} = \theta_j$. The value of each θ_j is calculated from the National Accounts as the fraction of value added not associated with labor earnings: $\theta_r = \frac{VA_r - LE_r}{VA_r}$, where VA_r is the value added in sector r and LE_r indicates labor earnings in sector r. In fact, under these additional assumptions, the incorporation of the elasticity of demand in the calculation of average tariffs is of little consequence.[5]

When conducting our empirical exercises, we also test the robustness of our results to other commonly used measures of trade openness: the ratio of imports to production (M/P) and the import penetration coefficient (MPC, defined as MPC ≡ Imports/(Production + Imports – Exports)). These data are only available by sectors at the national level. We use equation 8 and apply the same strategy used for tariffs to calculate M/P and MPC by micro-regions.

5. Data

5.1 Sources of Data and Sample

We use data from the Brazilian 1991 and 2000 censuses to estimate the conditional wage gaps in the first stage of our empirical strategy (equation 6). These data are also used to calculate the aggregate

[5] Kovak (2013) reports a correlation of 0.996 between the results of calculations with and without the inclusion of ϵ_{jr}.

variables introduced as controls in the second stage (changes in aggregate labor market conditions, including average wages and employment by educational categories, shares of blue collar occupations, share of informal employees, unemployment, and migration). We define blue collar occupations as those that typically do not require formal (technical or college) training or education, as opposed to professional occupations. In the 1991 census, blue collar occupations are those associated with codes 301-928. These include, among various others, fishermen, miners, mechanics, shoemakers, bricklayers, merchandise packers, sellers, cashiers, drivers, cleaners, and dustmen. Informal employees are defined as those who do not have a registered labor contract (or, in terms of the Brazilian legislation, do not have their "labor card" signed by the employer). Regarding migration, census data allow us to calculate the percentage of the population that immigrated to a given micro-region within the previous five years.

We also use the change in the number of schools per capita as control in some specifications of our second stage. We construct this variable as the number of public schools (preschools, elementary schools, and high schools) per 1,000 individuals aged between 0 and 17 in each micro-region (data from the Brazilian School Census). Unfortunately, there are no data for number of schools in 1991, so we use information from 1995 and 2000 to construct the change in this variable.

Tariff reductions are calculated from information provided by Kume et al (2003). Kume et al (2003) compute average tariffs for 32 sectors directly from international trade legislation. These 32 sectors are not entirely consistent with the sectorial classification used by the Brazilian census, so we merge some of them in order to make the two datasets compatible (Appendix Table A.2 describes how the two sectorial classifications were merged). This gives us 20 sectors, plus services. The tariff of the "new" merged sectors is calculated as a weighted average of its subsectors, where the weights are given by the relative value added of each subsector.

Value added and total labor earnings by sector, used to calculate the change in average tariffs by micro-region, are provided by the National Accounts from the Brazilian Census Bureau (IBGE). The National Accounts also provide the value of production needed to compute our alternative measures of exposure to trade (import-production ratio and import penetration coefficient). Import and export data are from Gonzaga et al (2006), while data on market concentration in Brazil, used in some heterogeneity analyses, are from Ferreira and Fachini (2005).

As mentioned before, our benchmark specification uses the changes in tariffs between 1990 and 1995, as Kovak (2013), because this period concentrates the main and most aggressive part of the reforms. Since we look at changes in wages between 1991 and 2000, we implicitly assume that: (i) the change in policy was perceived as permanent; (ii) the main labor market adjustments due to the trade reform were already completed by 2000; and (iii) the minor additional changes to trade legislation introduced after 1995 were not critical for labor market outcomes in 2000. Still, in some robustness exercises, we also consider changes in tariffs from 1990 to 1998, and from 1987 to 1995.

Our main results use a sample of male employees (excluding public servants, self-employed, employers, and domestic workers), with positive earnings, aged between 20 and 60. We choose to focus on prime aged male employees to come closer to the concept of employer discrimination from Becker (1957), and also to emulate the hypothesis of inelastic labor supply present in both Becker (1957) and Kovak (2013). Under these restrictions, there are 1.8 million observations in the 1991 census and 2.3 million observations in the 2000 census. To assess the robustness of the results to potential market imperfections associated with labor market attachment, insertion, and mobility, some alternative samples are also considered. For example, we present results including self-employed men, women, and restricting the sample to full time workers.

Finally, in the second stage of our analysis, the unit of observation is a micro-region, taken to represent a local labor market. We use compatible definitions of micro-regions in the 1991 and 2000 censuses, resulting in a total of 480 observations.[6]

5.2 Descriptive Analysis

Table 1 presents descriptive statistics for the 1991 and 2000 censuses, based on our main sample (male employees between ages 20 and 60, with positive earnings). In addition to providing a broad overview of the labor market conditions in Brazil during the period of our analysis, the table also helps guide our later discussion of the results.

The typical individual in the sample in both years has around 34 years of age, works full time (more than 90%), did not complete elementary school, works in the services sector, and in a blue collar occupation. It is worth noticing the reduction of 7 percentage points in the fraction of workers with less than complete elementary education between 1991 and 2000, the increase in the fraction of workers attending school, and the reduction in the share of workers in manufacture. Real wages are approximately stable during the period, driven mostly by the change in the composition of the labor force, since wages fell for most educational levels (with the exception of college, which comprises a small fraction of the population; real wages in 2000 values, deflated by the National Consumer Price Index, following Corseuil and Foguel, 2002).

[6] Appendix B describes the procedure adopted to make the definition of micro-regions compatible across censuses. This procedure leads to an initial sample of 488 micro-regions. We lose 8 micro-regions due to missing observations on some of the key variables (for example, the smallest micro-regions do not have observations for employees with positive wages for certain combinations of race and educational group). We use the 480 observations with a complete set of variables throughout to keep a consistent sample in all empirical exercises. Results are very similar if, where possible, we use the complete sample of 488 observations.

The last rows in the table present numbers on the trade variables used and on the conditional racial wage gap.[7] The average tariff reduction across micro-regions was 9.7 percentage points (corresponding to 48% of the initial level), and was accompanied by increases of 85% in the ratio of imports to production and 81% in the import penetration coefficient. The seemingly small initial level and subsequent reduction in tariffs when looking at micro-region averages comes from the role of the agricultural sector.[8] Tariffs were already low in agriculture by 1990 and, when excluding the services sector, agriculture employed a substantial fraction of the labor force. The combination of these two facts dampens the sectorial tariff reductions discussed before in Figure 2. Still, in relative terms, the reduction in tariffs represented a substantial change in exposure to foreign competition, which ended up reflected on the measures of import penetration.

It is also important to notice that the conditional racial wage gap remained roughly stable in Brazil during this period (in fact, it was reduced by 0.4 percentage point). So, if trade liberalization did work towards reducing labor market discrimination, other factors must have worked in the opposite direction. One possibility is that the expansion in basic schooling observed in Brazil, which benefited relatively more the black population, may have reduced the quality of schooling. This would lead to an increase in the unexplained portion of the wage differential across races, therefore augmenting the racial wage gap. A definitive answer to this question is beyond the scope of this paper. Still, this highlights the fact that we are exploring the effect of increased competition in reducing the racial wage gap in a context where there is no widespread trend in this direction.

6. Results

We concentrate on the results from the second stage, since our first stage reproduces commonly used estimation procedures for Mincerian regressions. Still, when useful, we briefly mention the specification used to estimate the racial wage gap in the first stage.

Table 2 presents the main result from our empirical exercise. Column 1 shows the coefficient of a univariate regression of the change in the conditional racial wage gap on the change in tariffs, without additional controls. Column 2 introduces dummies for geographic regions, corresponding to the five great geographic regions in Brazil: North, Northeast, Center-West, Southeast, and South. Column 3 adds controls for changes in earnings by level of schooling (primary, elementary, high school, and college), and column 4 replaces the geographic region dummies with state dummies (26 states plus the Federal District).

[7] The conditional racial wage gap in the table is the average of the gaps estimated for each micro-region in our first stage.

[8] Appendix Table A.1 presents the sectorial distribution of employment in the Brazilian economy for 1991. As mentioned before, the 1991 sectorial shares are used as weights in the calculation of average tariff changes by micro-region.

Panel A corresponds to our benchmark specification, where a single Mincerian regression is used to estimate the racial wage gap for all micro-regions in a given year, while Panel B corresponds to an alternative specification where a different Mincerian regression is estimated separately for each micro-region in each year.

Column 1 in Panel A shows that there is a positive correlation between changes in tariffs and changes in the conditional racial wage gap, as predicted by theory. This means that local labor markets that experienced larger reductions in average tariffs also experienced larger reductions in the conditional wage differential across races (remember that our race dummy indicates white workers). The introduction of regional dummies in column 2 increases slightly the magnitude and the precision of the estimated coefficient.

Maybe the most important result from Table 2 is the change in the coefficient of interest once we move from column 2 to column 3. In column 3, we control for changes in wages by levels of schooling, which should account for the effects of trade on local labor markets that would emerge from a Heckscher-Ohlin model. By controlling for these variables, we are accounting for the relative change in demand for skills driven by shifts in production towards sectors intensive in the abundant factor (in the case of Brazil, low skill labor). If the positive and significant coefficients on *Δ(tariff)* in columns 1 and 2 were only capturing a relative increase in the demand for unskilled labor – with blacks being less skilled than whites – one should expect to see a reduction in the estimated coefficient as we move from column 2 to 3. Notice that this should be the case even if the coefficients estimated in columns 1 and 2 were related to unobserved skills, as long as the change in returns to unobserved skills followed the same pattern of the change in returns to observed skills (meaning that locations with reduced wage differentials across observed skills also displayed reduced wage differentials across unobserved skills). But once we control for changes in earnings by level of schooling, the coefficient of interest increases in magnitude and remains strongly significant. So the correlation between changes in tariffs and changes in racial wage gaps captured by our empirical strategy does not seem to be driven by changes in returns to productive attributes (remember, in addition, that our first stage already controls for schooling and allows for changes in the return to schooling between 1991 and 2000).[9] This conclusion is further reinforced by the fact that there is no consensus in the literature on the effect of trade liberalization on returns to skill in Brazil (see, for

[9] Notice that the coefficient on the change in primary schooling wages is negative and statistically significant. This is consistent with the argument made in the text that increases in the relative gains of low skill workers in terms of observed attributes (in the case of Brazil, primary educated workers, in which blacks are over represented) should be correlated with relative gains to low skill workers also in terms of unobserved attributes, which should in turn be correlated with reductions in the conditional racial wage gap. This effect can be seen in the significant and negative coefficient on the change in primary wages in column 3. Still, this is not driving the estimated impact of the change in tariffs and, therefore, does not interfere with our identification strategy.

example, Arbache and Menezes-Filho, 2000, Arbache and Courseuil, 2004, Gonzaga et al, 2006, and Ferreira et al, 2010, Kovak, 2013, Dix-Carneiro and Kovak, 2015).[10]

In column 4, we replace the 5 geographic region dummies by 27 state dummies. The estimated impact of the change in tariffs increases slightly in magnitude, but remains very similar to that from column 3. Specific characteristics of regions or states do not seem to be correlated with our estimated coefficient. Since we have an average of only 18 micro-regions per state, the specification with state dummies becomes too heavy and compromises the precision of the estimates in some of our following empirical exercises, where we look at more restrictive samples. So we proceed with the controls from column 3, including geographic region dummies, as our benchmark specification.

According to the Hecksher-Ohlin model, relative returns to inputs should summarize all the relevant impacts of trade reforms on labor markets. So the benchmark specification from Table 2 is the closest in spirit to theory, and it is the one we adopt when conducting additional exercises in the remainder of the paper. But we check the robustness of our results to other labor market changes, which could be relevant in the presence of market imperfections. Some of these may be endogenous to increased competition in the market for final goods and, therefore, could be seen as what Angrist and Pischke (2009) call "bad controls." Still, one may wonder whether the significance of the estimated impact of the reduction in tariffs survives their inclusion. So we incorporate these additional variables in Table 3, but do not carry them over to other specifications in the paper.[11]

Column 1 in Table 3 includes controls for the shares of employment by level of schooling and race (where the excluded category, due to perfect multi-colinearity, is primary schooling). Column 2, instead, controls for the number of public schools per 1,000 children. Column 3 controls for labor market changes associated with occupational structure (share of blue collar occupations), informality, and unemployment. Finally, column 4 adds the control for immigration and column 5 includes all previous controls simultaneously.

In columns 1, 3, and 4 of Table 3, results are very close to those from Table 2. In column 2, the coefficient is slightly smaller in magnitude, but remains strongly significant. Most important, when all

[10] For the interested reader, Appendix Table A.3 repeats the same specification from column 3 of Table 2 by educational levels. It shows that, even when estimated separately for each level of schooling, we still find positive and statistically significant coefficients for the cases of primary and elementary schooling. Coefficients are estimated much less precisely for high school and college education, since the number of first stage observations for these cases is very small in various micro-regions. In this exercise, we re-estimate the conditional racial wage gap in the first stage by interacting the racial dummy with the dummies for years of schooling. The second stage remains the same as before. If we include self-employed workers in our first stage, increasing the size of the sample and the precision of the estimates, results for all schooling levels become positive and statistically significant (results available upon request). So the results in the last two columns of Table A.3 seem indeed to be driven by the low number of first stage observations and imprecision in the estimation of the conditional racial wage gap.

[11] In any case, the vast majority of qualitative results reported in other tables in the paper remains unchanged under any of the other specifications from Table 2. These results are available from the authors upon request.

controls are included simultaneously in column 5, the coefficient is again very similar to that estimated in column 3 of Table 2 and remains statistically significant.

Overall, as we compare column 3 in Table 2 to column 5 in Table 3, there is little change in the estimated coefficient once all controls are incorporated into the analysis. Therefore, concerns related to differential changes in schooling or access to education across races, other labor market changes driven by the trade reform, migration, or heterogeneous effects of broader macroeconomic trends do not seem to be a first order issue. Some of these dimensions did affect the racial wage gap during the 1991-2000 period, but in ways that are, on average, orthogonal to the relationship between competition in the market for final goods and labor market discrimination captured here.

Our benchmark specification (column 3 in Table 2) implies that a reduction in tariffs of 9.7 percentage points (equivalent to the average observed in the sample) leads to a reduction of 2.2 percentage points in the conditional racial wage gap (or 18% of its 1991 value, which was 12.3). Alternatively, a reduction in tariffs corresponding to one standard-deviation in the initial period (7 percentage points in 1991) would lead to a reduction of 13% in the racial wage gap. Though this effect may not seem particularly large, one should bear in mind that, over the 1991-2000 period, there was hardly any reduction at all in the conditional racial wage gap (0.4 percentage point). In other words, according to our estimates, had the liberalization process not taken place, the conditional racial wage gap would have increased by 1.8 percentage points. Increased competition may have reversed the increase in the racial wage gap that otherwise would have been observed.

Panel B in Table 2 reproduces the same sequence of results from Panel A, but using a separate Mincerian regression in the first stage to estimate the conditional racial wage gap in each micro-region and year.[12] This strategy is more flexible in that it allows for returns to productive attributes – such as education or experience – to vary across micro-regions in the same year, therefore better capturing the specific characteristics of the equilibrium in each local labor market. On the other hand, it demands much more from the data. Results in Panel B are very similar in magnitude and significance to those in Panel A, particularly so in columns 3 and 4. In addition, the qualitative pattern of change in coefficients as we move from columns 1 to 4 remains the same, so the discussion related to Panel A also applies here.[13]

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[12] The equation estimated for each micro-region j and year t is: $\ln(wage)_{ijt} = \alpha_{jt} + \delta_{jt} white_{ijt} + \gamma_{jt}' \boldsymbol{Z}_{ijt} + \varepsilon_{ijt}$, where the vector $\boldsymbol{Z}$ includes all variables included in $\boldsymbol{X}$, with the exception of the micro-region dummies.

[13] Appendix Table A.4 presents yet an additional alternative specification, where we estimate the impact of the reduction in tariffs directly, in one single step, together with the Mincerian equation. In this strategy, we estimate a single Mincerian regression including both years and adding year and micro-region dummies. Akin to a difference-in-difference strategy, the effect of the tariff reduction on the gender wage gap is identified, in this case, from the interaction of the micro-region specific tariff (which changes between years) with the race dummy (indicating white). To come as close as possible to our two-stage strategy, we include the same individual and aggregate (micro-region) level controls used in our previous first and second stages, and let the coefficients on the individual level variables vary across years. Table A.4 reproduces, in this setting,

6.1 Alternative Timing and Measures of Trade Liberalization

Our first robustness exercise considers alternative timings and measures of trade liberalization. The benchmark specification uses changes in tariffs between 1990 and 1995, which corresponds to the period containing the most aggressive part of the reforms. One might think that this would exaggerate the extent of liberalization, possibly biasing our estimates. To address this concern, we consider alternative specifications that use the change in tariffs between 1990 and 1998 (the last year for which consolidated data on tariffs by sector are computed by Kume et al, 2003), and between 1987 and 1995.[14] The results are presented in columns 1 and 2 of Table 4. In column 1, the estimated coefficient remains almost identical to that in Table 2 (column 3). In column 2, it increases in magnitude and remains strongly significant. So the specific timing of measurement of the change in tariffs does not seem to interfere with the results.

Other concern related to the measurement of the reforms refers to the use of tariffs as sufficient statistics for trade liberalization. Various other dimensions of economic policy and regulation affect the effective degree of protection in a given economy, including non-trade barriers, exchange rate regimes, and red tape. For these reasons, some consider that variables related to the flow of international trade are more adequate measures of the actual degree of openness in an economy. In fact, Brazil experienced changes in exchange rate regimes during the 1990s as well as successive elimination and reintroduction of non-tariff barriers. Therefore, this concern may indeed be relevant. In addition, as argued by Gonzaga et al (2006), the pass through of tariff changes to prices may vary across sectors, so that changes in tariffs alone might not be an adequate measure of changes in exposure to foreign competition. Still, trade flows are endogenous to economic and labor market conditions and possibly to the very issue analyzed here. In any case, we also consider alternative measures of exposure to foreign competition based on trade flows: the ratio of imports to production (M/P) and the import penetration coefficient (MPC). The results obtained when these are used as independent variables are presented in columns 3 and 4 in Table 4. It is worth remembering that, contrary to tariffs, increases in these variables indicate increased exposure to international trade. So the prediction of the theory is that these two variables should be negatively associated with changes in the racial wage gap (increased exposure to international trade associated with reduced advantages of whites in the labor market).

As predicted by theory, and consistently with the results for tariffs, both measures of trade flows indicate that increased imports during the period of trade reforms were associated with reductions in the conditional racial wage gap. The coefficients in columns 3 and 4 are negative and statistically significant.

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specifications analogous to those from Tables 2 and 3. Estimated coefficients are again positive and statistically significant, being typically very similar to the corresponding coefficients presented in Tables 2 and 3.

[14] To calculate average tariffs for the merged sectors (see section 5.1) in 1998, we use value added from the 1995 National Accounts, and, for 1987, we use value added from the 1985 National Accounts.

Though the scales of the three independent variables are different and, therefore, estimated coefficients cannot be directly compared to one another, their quantitative implications in normalized units are very similar. An increase in M/P corresponding to one standard-deviation in the initial period (0.011 in 1991) would be associated with a reduction of 12% in the conditional racial wage gap, a result identical to that obtained with the MPC variable and very close to the 13% mentioned before for tariffs. In other words, the specific variable used to represent the process of trade liberalization does not affect the results either qualitative or quantitatively.

6.2 Falsification Exercises

The timing of the reforms and the measures of exposure to trade based on flows also provide us with an opportunity to falsify our identification strategy. Though we do not have tariffs by sector for 1980 (Kume et al, 2003 do not compute these numbers and, besides, there were few changes between 1980 and 1990), we do have data on imports, exports, and production. If our identification strategy is indeed capturing the effect of the trade reform from the 1990s, we should find no significant results once we repeat an analogous exercise using data from 1980 and 1990, since there was no major change in trade policies during this period. Otherwise, if we are just capturing some spurious correlation between changes in imports and changes in labor market outcomes, we should also find a significant effect when looking at 1980 and 1990.

Columns 2 and 3 in Table 5 reproduce the same exercise from Table 2, but comparing data from the 1980 and 1991 censuses and using the two measures of exposure to trade that we have for 1980 and 1990 (M/P and MPC). Since there was a different political organization in Brazil in 1980, we have 284 micro-regions in this analysis (micro-regions were aggregated to be made compatible across 1980 and 1991, and there were fewer municipalities and micro-regions in 1980 than in 1991; see Appendix B for details). For purposes of comparison, we first estimate again the specification from column 3 in Table 2, using data from 1991 and 2000, with this new geographic division including 284 micro-regions. As shown in column 1, results remain positive and statistically significant, increasing slightly in magnitude. So the different number of observations does not affect the results obtained before.

In columns 2 and 3, we present the results of regressions using the measure of exposure to competition based on trade flows and using the 1980-1991 period. Both estimated coefficients are small in magnitude and far from statistically significant. This evidence suggests that the reduction in the racial wage gap in response to tariff reductions documented in Table 2 is indeed associated with the process of trade reforms from the 1990s. There is no indication of a spurious correlation between level of trade and labor market outcomes before the reforms were implemented.

Yet another possible falsification exercise is to assess whether the change in exposure to foreign competition between 1990 and 1995 was correlated with changes in the racial wage gap between 1980 and 1991, before the changes were actually implemented. If the trade reforms were truly exogenous to local labor market conditions, we should expect such a regression to deliver non-significant results. Otherwise, if the change in exposure to trade in the 1990s was associated with specific characteristics of the labor market before that date, which might have continued to be observed between 1991 and 2000, this regression might lead to significant results.

Columns 4, 5, and 6 in Table 5 show the results from these regressions, where the change in the wage gap between 1980 and 1991 is regressed on our three measures of change in exposure to foreign competition between 1990 and 1995. Again, all estimated coefficients are very small in magnitude and far from statistically significant. Our empirical strategy seems to be indeed isolating the effect of the exogenous shock represented by the 1990s trade reform.

6.3 Alternative Samples

The results presented up to now use a sample of male employees, aged between 20 and 60, with positive earnings. This brings the empirical exercise closer to theory of taste based employer discrimination (Becker, 1957), but may call into question the representativeness of the results for the Brazilian labor market as a whole. Our results would still be representative of the overall impact of the trade reform under the assumption of perfect labor mobility across labor market statuses within regions. With imperfect labor mobility, and differential entry and exit, the impact of increased competition on wages may be heterogeneous across different groups of workers. To assess this possibility, we re-estimate our first stage with different samples. The results from these exercises are presented in Table 6.

The following sample variations are considered in the first stage, always restricting to individuals between 20 and 60 years of age: in column 1, all male workers; in column 2, all male and female workers; in column 3, all male and female employees and self-employed; in column 4, all male and female employees; and, in column 5, all male employees not attending school, and working full time (at least 35 hours per week).

Results in column 1, where we look at all men, are larger in magnitude than those from Table 2 and still significant. This is consistent with part of the labor market adjustments after the reform taking place through differential transition of workers across occupational categories (most importantly, between employees and self-employed). As we move to columns 2 and 3, considering all men and women and then employees and self-employed, the coefficient is reduced somewhat in magnitude in comparison to column 1, but remains strongly significant. When we consider only male and female employees, in column 4, the coefficient drops substantially in magnitude, but remains statistically significant. Finally, in column 5, we consider only men, who do not attend school, and are employed full time, coming closer to the hypothesis of inelastic labor supply from both Becker (1957) and Kovak (2013). Results rise again in magnitude and remain statistically significant.

So differences in sample across genders, labor market insertion, and labor market attachment do not seem to affect the qualitative results. This pattern suggests a reasonable degree of flexibility in local labor markets in Brazil, which is supported by evidence of a high degree of mobility across formal and informal sectors (see, for example, review in Ulyssea, 2006). For all samples considered, we detect a significant impact – mostly of similar magnitude – of increased competition on labor market discrimination.

6.4 Heterogeneity

An additional strategy to assess whether we are indeed capturing the effect of increased competition on labor market discrimination is to explore some dimensions of heterogeneity predicted by Becker's (1957) theory. As discussed in Section 2, the model of taste based employer discrimination predicts that some characteristics of local labor markets should be associated with higher initial levels of discrimination and, therefore, with larger responses of discrimination to reductions in tariffs. First, the model demonstrates that discrimination requires the existence of pure economic rents. So labor markets dominated by firms that face low competition in the market for final goods should respond more to liberalization than markets dominated by firms that face more competition. Second, for given market structure and distribution of tastes for discrimination, the theory predicts that markets with a larger share of minority workers should display initially higher levels of discrimination. Finally, markets with stronger prejudice against the minority should also be associated with more discrimination in the labor market.

We explore these three dimensions to assess whether there is heterogeneity in the initial level of the racial wage gap and in the impact of increased liberalization, and whether this heterogeneity supports the predictions of the theory. These are patterns that would be hard to rationalize outside the framework of taste based discrimination, as, for example, when one considers the possibility that the estimated coefficients are just capturing returns to unobserved skills. By looking at these dimensions, we provide further evidence that our empirical strategy is indeed capturing the effect of increased competition on labor market discrimination.

In order to measure the degree of monopoly power in the market for final goods, we follow Ferreira and Fachini (2005) and use the four-firm concentration ratio (CR4), which considers a sector to be concentrated if the share of the four largest firms in total revenue is above 40%.[15] We use the classification of concentrated sectors from Ferreira and Fachini (2005) and, to translate it to the level of local labor markets, calculate the share of workers occupied in concentrated sectors by micro-region. Following, we use the fraction of black workers in 1991 (black and mixed) as indicating the share of the minority in the local labor market. To measure the strength of tastes for discrimination, we use the index based on inter-racial marriages proposed by Levine et al (2008), also constructed from the 1991 census. The index captures unobserved factors associated with the frequency of inter-racial marriages in a given micro-region (netted out of effects of education, age, and racial composition of the population),[16] with higher values indicating higher discrimination. The basic idea is that, conditional on observables, locations with a higher incidence of inter-racial marriages should have a lower underlying level of prejudice against the minority.[17]

Panel A in Table 7 starts by showing the results of simple micro-region OLS regressions for 1991, where the conditional racial wage gap is regressed on these three local labor market characteristics – market concentration, percentage of black workers, and the prejudice index – plus geographic region dummies. As predicted by theory, the three variables are indeed positively related to the initial level of labor market discrimination in our data, either considered separately (columns 1 to 3) or together (column 4). All coefficients are positive and statistically significant, indicating that higher market concentration, higher percentage of the minority, and higher average prejudice are all positively correlated with observed labor market discrimination. The distribution of the racial wage gap in the cross-section is consistent with the theory of taste based discrimination. Still, this is simply a descriptive pattern and lacks a clear source of identification.

In Panel B of Table 7, we incorporate heterogeneity along these three dimensions in our main empirical exercise. First, in separate second stage regressions, we include interactions of the change in tariffs with these three variables, one at a time. Columns 1 to 3 present these results. For the case of market concentration (column 1), the coefficient is positive and statistically significant. Micro-regions with higher

[15] Ferreira e Fachini (2005) classify as concentrated the following sectors in 1985: transportation, rubber, chemicals, perfumery, and tobacco. The authors do not analyze extractive sectors. Since petroleum, natural gas, and charcoal are monopolies or concessions, they are also considered concentrated.
[16] Specifically, we estimate an OLS regression for couples, where the dependent variable is a dummy indicating an inter-racial marriage, and the independent variables are age, age squared, dummies for years of schooling, share of blacks in the micro-region, share of blacks squared, a dummy for urban areas, and state dummies. The average residual for a micro-region, q_j, indicates the component of the probability of an inter-racial marriage in micro-region j that is not explained by socio-demographic factors. To renormalize it as an index where higher values are associated with higher discrimination, we define $q_j' = \max_i(q_i) - q_j$. This is the measure of tastes for discrimination that we use in our analysis.
[17] Becker's (1957) theory predicts that the coefficient of discrimination of the marginal employer, not the average coefficient of discrimination, should determine the observed racial wage gap. We do not have a measure of the distribution of preferences for discrimination by micro-region, simply a proxy for its average. We proceed under the assumption that shifts in this average would be typically associated with shifts in the entire distribution of preferences for discrimination.

market concentration in 1991 experienced larger declines in the racial wage gap following the trade reform from the 1990s. For the prejudice index (column 3), the coefficient is positive, but estimated imprecisely. Still, the point estimate suggests that higher levels of measured prejudice are associated with a stronger effect of the reduction in tariffs on the racial wage gap. The coefficient on the fraction of black workers (column 2), in turn, is very small quantitatively (negative) and far from statistically significant. This may be due to the fact that the coefficient on the fraction of blacks in Panel A was already relatively small and estimated less precisely than the other coefficients. There seems to be no clear pattern and no support for the theory in the case of the fraction of the minority in the labor force.

In column 4 of Panel B, we consider a specification including the three dimensions of heterogeneity together. In line with the pattern from columns 1 to 3, we estimate a positive and significant coefficient on the interaction of the change in tariff with market concentration, and a positive but imprecisely estimated coefficient on the interaction with the prejudice index. As before, the coefficient on the fraction of blacks is very small (but now positive) and not statistically significant.

Despite the lack of precision in some of the estimates, we read the results from Table 7 as providing indicative evidence that some dimensions of heterogeneity predicted by the employer discrimination model seem to be present in the data. The cross-sectional distribution of the conditional racial wage gap is consistent with the theory. In addition, though the evidence is clearly not overwhelming, the point estimates seem to indicate that higher market concentration and higher prejudice are both associated with a stronger impact of increased exposure to competition on labor market discrimination. These dimensions of the results agree with the predictions of the theory along dimensions that are otherwise difficult to rationalize.

7. Concluding Remarks

We use the episode of trade liberalization in Brazil during the 1990s to test the effect of increased competition in the final goods market on discrimination in the labor market. We show that local labor markets that experienced more exposure to international competition due to the trade liberalization process also observed larger reductions in the conditional wage differential between white and black workers. As predicted by the theory of taste based discrimination, the initial racial wage gap and the impact of increased competition seemed to be larger in local labor markets dominated by firms in more concentrated sectors and with stronger preferences for discrimination. Our empirical setting provides a clean identification of the effect of trade liberalization on labor market discrimination, therefore improving upon the results available from the current literature. By exploring the issue of wage differentials across races, our paper also speaks to the broader literature on trade liberalization and inequality, and identifies a specific dimension over which increased trade openness contributed to reduce earnings inequality.

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Appendix A: Additional Tables

Appendix B: Constructing Compatible Micro-regions across Censuses

The division of the Brazilian territory into municipalities and micro-regions changes substantially between 1991 and 2000. In particular, due to political secession, 1015 new municipalities are created during this period. Some of them incorporate fractions of territory that previously belonged to different municipalities and, sometimes, to different micro-regions.

To deal with this problem, we use Minimum Comparable Areas (MCAs), which define geographic units that are comparable across censuses (Reis et al, 2011). We start with the 4492 MCAs that are comparable across 1991 and 2000 and use the geographic definition of micro-regions from 2000. Whenever a given AMC crosses a micro-region boundary in 2000, we aggregate the micro-regions to which this AMC belongs into a new consolidated micro-region. Proceeding like this, we are able to generate a consistent set of micro-regions that correspond exactly to the same geographic areas in 1991 and 2000. By aggregating some of the micro-regions in this process, we reshape the 558 units that existed in 2000 into 494 units.

For our benchmark sample of men, aged between 20 and 60, with positive wages, and information on schooling, occupation, and urban residence, we further aggregate micro-regions that have too few observations for the conditional racial wage gap to be estimated precisely. So micro-regions with fewer than 500 observations in at least one of the census years are merged with the neighboring micro-region with lowest population. With this additional change, our set of micro-regions is reduced further by 6 units, leaving us with a final set of 488 micro-regions.

For the 285 compatible micro-regions from the 1980 and 1991 censuses used in the robustness exercises, we follow steps analogous to those adopted in the construction of the 1991-2000 areas, but using the micro-regions from 1991 as initial reference points (also following Reis et al, 2011). This procedure also keeps track of the changes in municipality borders between 1980 and 1991.

Figure 1: Comparative Statics in the Taste Based Model of Employer Discrimination

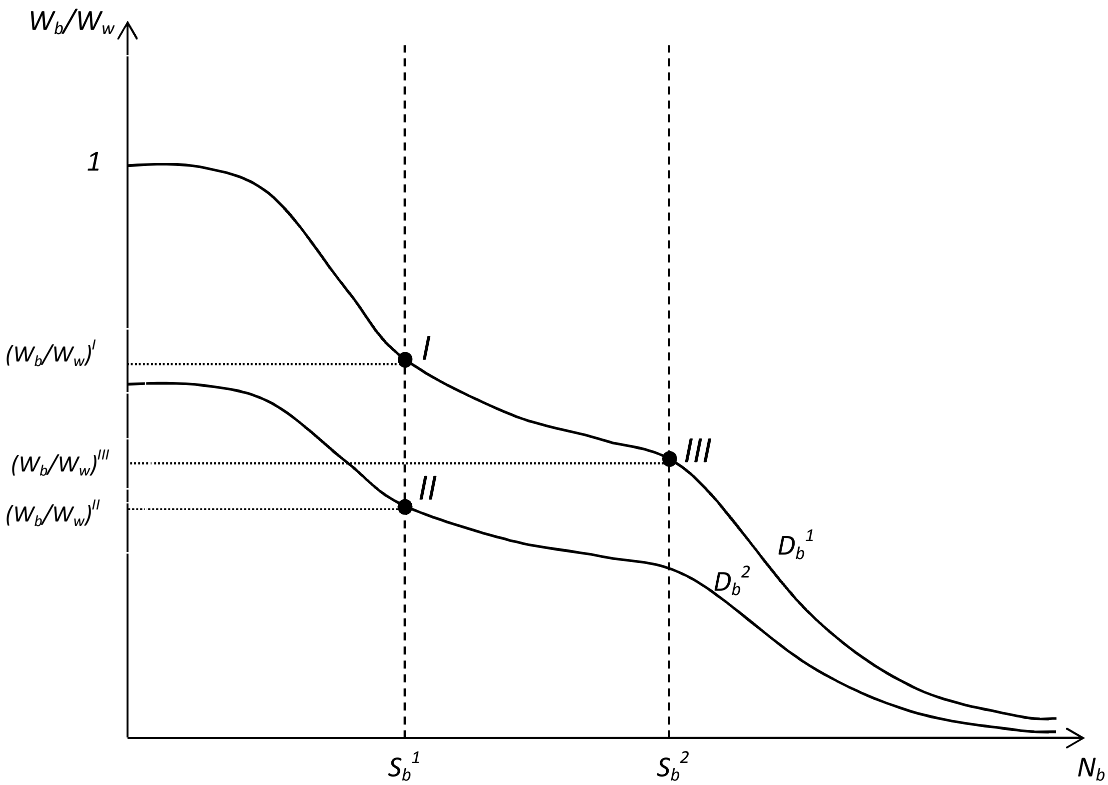
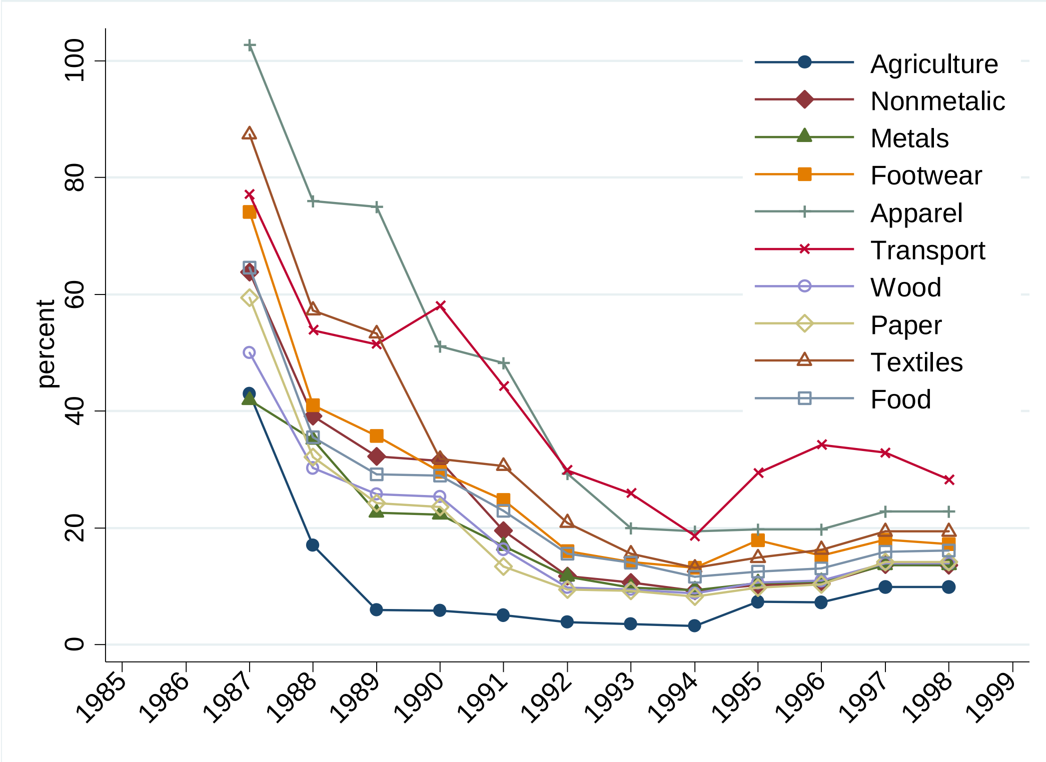
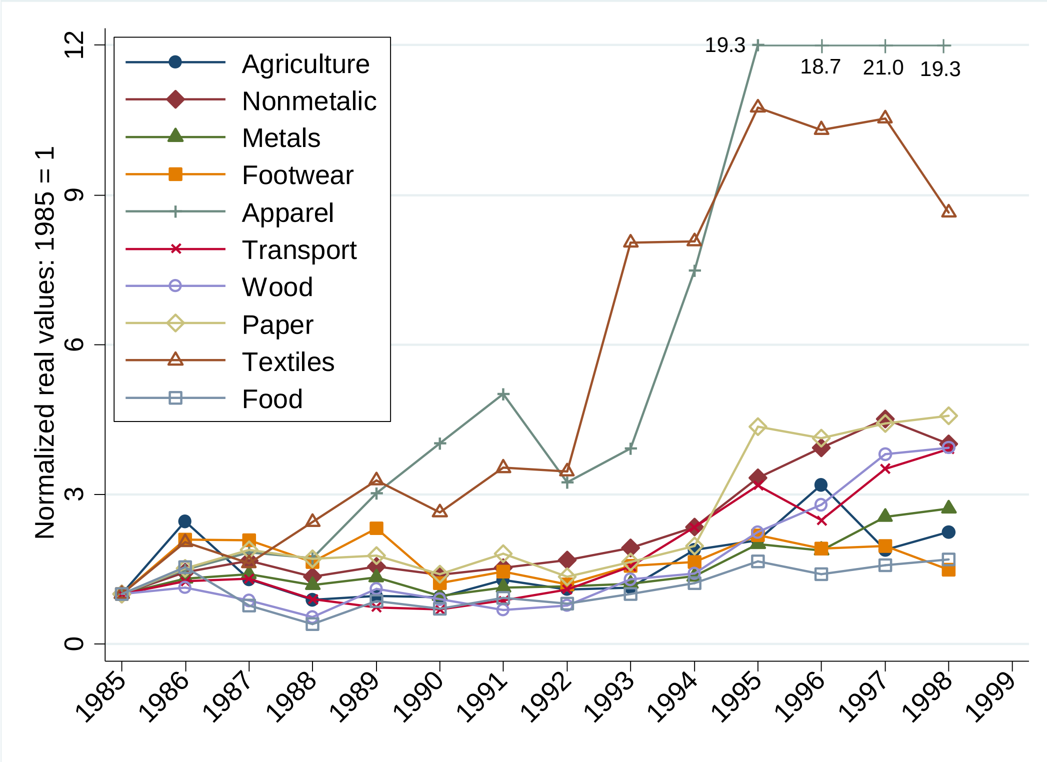


Figure 2: Nominal Import Tariffs during the Late 1980s and 1990s, Brazil



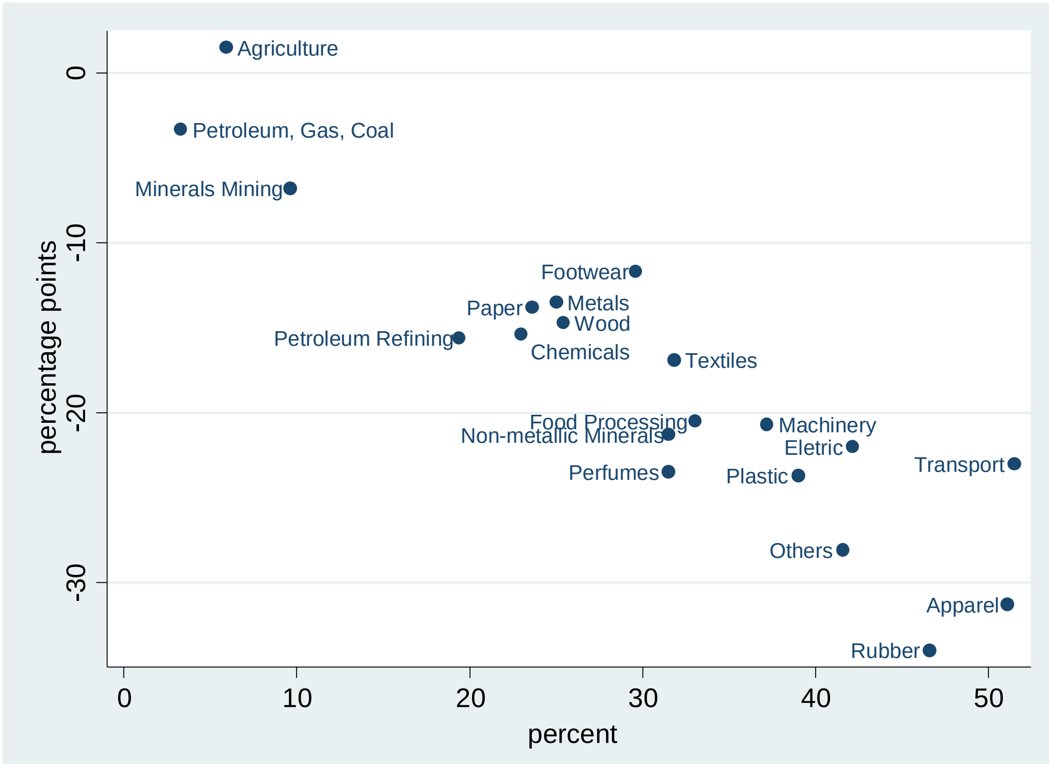
Source: Data from Kume et al (2003).

Figure 3: Imports during the Late 1980s and 1990s, Brazil (1985 value normalized to 1)



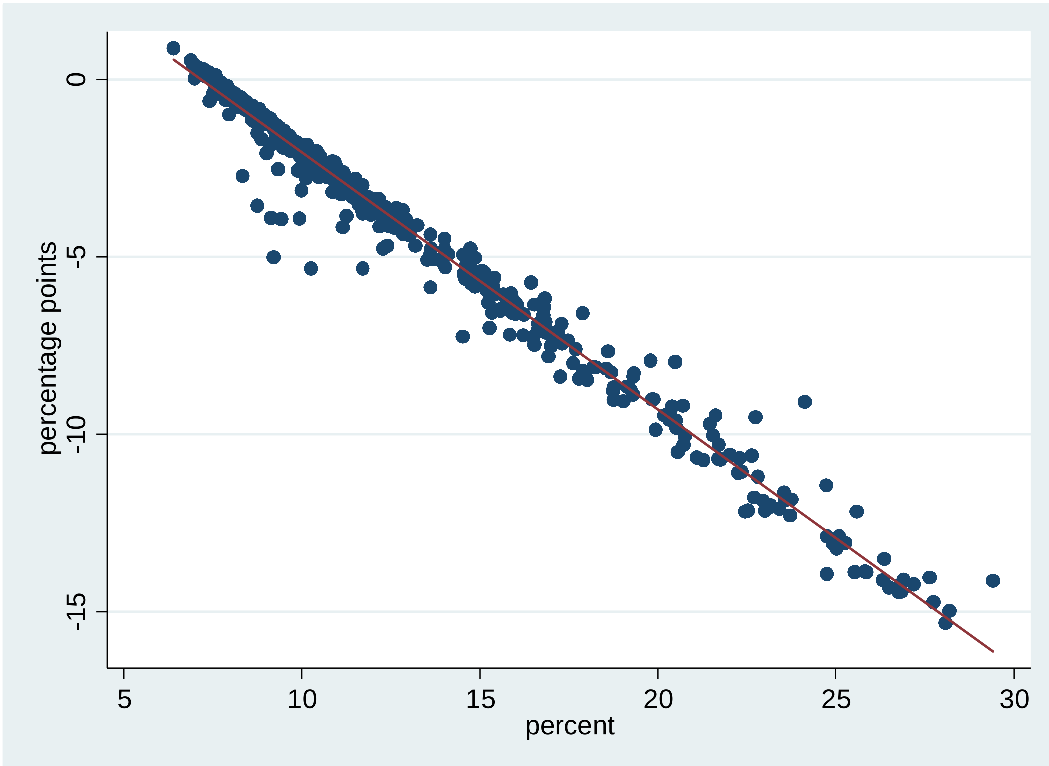
Source: Data from Gonzaga et al (2006).

Figure 4: Tariff Reductions between 1990 and 1995 and Initial Tariff Levels in 1990 by Sector of Economic Activity, Brazil



Source: Data from Kume et al (2003).

Figure 5: Tariff Reductions between 1990 and 1995 and Initial Tariff Levels in 1990 by Micro-Region, Brazil



Source: Authors' calculations based on data from the 1991 census and from Kume et al (2003).

Table 1: Descriptive Statistics, Brazil, 1991 and 2000, Male Employees, ages 20-60

	1991		2000	
	Mean	s.d.	Mean	s.d.
Black	0.445	0.497	0.451	0.498
Age	33.7	10.0	34.1	10.2
Full-time Work (35 hours/week or more)	0.945	0.227	0.928	0.258
Migrant	0.105	0.306	0.087	0.282
Attend school	0.045	0.207	0.091	0.288
Primary School (up to 7 years of schooling)	0.621	0.485	0.551	0.497
Elementary School (8 to 10 years of schooling)	0.154	0.361	0.175	0.38
High School (11 to 14 years of schooling)	0.163	0.369	0.216	0.412
College/University (above 15 years of schooling)	0.062	0.242	0.058	0.234
Blue collar	0.786	0.41	0.816	0.388
Informal	0.196	0.397	0.332	0.471
Agriculture	0.116	0.32	0.148	0.355
Mineral Mining	0.016	0.127	0.008	0.089
Manufacture	0.278	0.448	0.208	0.406
Services	0.59	0.492	0.636	0.481
Wage per hour (R$)	3.37	6.41	3.29	7.29
Wage per hour - Primary School	1.92	3.18	1.80	3.14
Wage per hour - Elementary School	3.09	4.39	2.73	4.15
Wage per hour - High School	5.29	6.81	4.62	7.56
Wage per hour – College/University	13.52	16.04	14.34	20.23
Observations (millions)	1.8		2.3	
Local Market Characteristics*				
Mean tariff **	0.202	0.07	0.105	0.021
Imports/Product (M/P)**	0.026	0.011	0.048	0.023
Import Penetration Coefficient (MPC)**	0.027	0.011	0.049	0.022
White-Black conditional wage gap***	0.123	0.072	0.119	0.052

Notes: * Average across micro-regions (488), weighted by sample size in each micro-region. ** Calculated for 1990 and 1995. *** Average racial wage gap across micro-regions estimated following our first stage strategy. Numbers based on census data from 1991 and 2000. Real wages in 2000 values (deflated by the National Consumer Price Index, following Corseuil and Foguel, 2002).

Table 2: Impact of Tariff Change on Racial Wage Gap, Brazilian Micro-regions, 1991-2000 – Dependent Variable: Change in Conditional Racial Wage Gap

	(1)	(2)	(3)	(4)
Panel A: One Regression in the First Stage				
Δ(tariff)	0.133*	0.157**	0.222***	0.242***
	(0.069)	(0.072)	(0.066)	(0.067)
Δ(primary wage)			-0.166***	-0.147***
			(0.030)	(0.035)
Δ(elementary wage)			0.041	0.033
			(0.035)	(0.036)
Δ(high school wage)			0.035	0.057**
			(0.027)	(0.029)
Δ(college wage)			0.008	0.011
			(0.012)	(0.012)
Region Dummies		X	X	
State Dummies				X
Observations	480	480	480	480
R-Squared	0.008	0.044	0.110	0.165
Panel B: Regressions by Micro-region in the First Stage				
Δ(tariff)	0.114**	0.190***	0.218***	0.228***
	(0.048)	(0.051)	(0.051)	(0.047)
Observations	480	480	480	480
R-Squared	0.017	0.101	0.122	0.178

Notes: Robust standard errors in parentheses. ***p < 0.01; **p < 0.05, *p < 0.10. Panel A first stage is a regression for 20-60 year-old male employees. First stage independent variables in Panel A: age, age squared, dummies for years of schooling, urban area, and micro-region, and interactions between a dummy for white and micro-region dummies. Panel B first stage is a set of independent regressions, estimated separately for each micro-region. First stage independent variables in Panel B: age, age squared, dummies for years of schooling and urban area, and dummy for white. Second stage independent variables: region and state dummies (not shown), and changes in micro-region average wages by level of schooling (primary, elementary, high school, and college). Unit of observation is a micro-region. Census data from 1991 and 2000. Regressions weighted by the precision of first-stage estimates of the dependent variable (inverse of the standard errors).

Table 3: Additional Controls, Impact of Tariff Change on Racial Wage Gap, Brazilian Micro-regions, 1991-2000 – Dependent Variable: Change in Conditional Racial Wage Gap

	(1)	(2)	(3)	(4)	(5)
Δ(tariff)	0.244***	0.178***	0.223***	0.245***	0.215**
	(0.083)	(0.067)	(0.071)	(0.066)	(0.086)
Δ(% elementary)$_{blacks}$	0.288**				0.262*
	(0.141)				(0.142)
Δ(% high school)$_{blacks}$	0.071				0.053
	(0.151)				(0.146)
Δ(% college)$_{blacks}$	0.656				0.649
	(0.538)				(0.551)
Δ(% elementary)$_{whites}$	-0.013				-0.058
	(0.179)				(0.181)
Δ(% high school)$_{whites}$	-0.119				-0.159
	(0.134)				(0.135)
Δ(% college)$_{whites}$	0.207				0.088
	(0.285)				(0.287)
Δ(# schools/1,000 children)		-0.010**			-0.008*
		(0.004)			(0.004)
Δ(% blue collar)			0.120		0.123
			(0.208)		(0.212)
Δ(% informal)			-0.172***		-0.135**
			(0.055)		(0.058)
Δ(% unemployed)			-0.118		-0.025
			(0.135)		(0.147)
Δ(% migrant)				-0.185	-0.069
				(0.157)	(0.150)
Δ Avg W by Schooling	X	X	X	X	X
Region Dummies	X	X	X	X	X
Observations	480	480	480	480	480
R-Squared	0.126	0.120	0.134	0.114	0.151

Notes: Robust standard errors in parentheses. ***p < 0.01; **p < 0.05, *p < 0.10. Independent variables: region dummies and changes in average wages by level of schooling (not shown), and changes in the composition of the labor force by level of schooling and race, in the number of public schools per 1,000 children, in the share of informal employees, in the % of unemployed, and in the % of migrants. Unit of observation is a micro-region. Census data from 1991 and 2000. Regressions weighted by the precision of first-stage estimates of the dependent variable (inverse of the standard errors). First stage is a regression for 20-60 year-old male employees. First stage independent variables: age, age squared, dummies for years of schooling, urban area, and micro-region, and interactions between a dummy for white and micro-region dummies.

Table 4: Impact of Alternative Measures of Trade Liberalization on Racial Wage Gap, Brazilian Micro-regions, 1991-2000 – Dependent Variable: Change in Conditional Racial Wage Gap

Indep. Var.:	$\Delta(\text{tariff})_{1990\text{-}1998}$ (1)	$\Delta(\text{tariff})_{1987\text{-}1995}$ (2)	$\Delta(\text{M/P})_{1990\text{-}1995}$ (3)	$\Delta(\text{MPC})_{1990\text{-}1995}$ (4)
Coefficient	0.234***	0.318***	-1.331***	-1.369***
	(0.066)	(0.119)	(0.265)	(0.278)
Observations	480	480	480	480
R-Squared	0.112	0.104	0.127	0.126

Notes: Robust standard errors in parentheses. ***p < 0.01; **p < 0.05, *p < 0.10. Independent variables: region dummies and changes in average wages by level of schooling (not shown). Columns 3 and 4 use, respectively, the share of imports and the import penetration coefficient as measures of the trade reform, instead of tariffs. Unit of observation is a micro-region. Census data from 1991 and 2000. Regressions weighted by the precision of first-stage estimates of the dependent variable (inverse of the standard errors). First stage is a regression for 20-60 year-old male employees. First stage independent variables: age, age squared, dummies for years of schooling, urban area, and micro-region, and interactions between a dummy for white and micro-region dummies.

Table 5: Falsification Exercises, Impact of Tariff Change on Racial Wage Gap, Brazilian 1980 Micro-regions, 1980-1991 – Dependent Variable: Change in Conditional Racial Wage Gap

Dep. Var:	$\Delta(\text{gap})_{1991\text{-}2000}$	$\Delta(\text{gap})_{1980\text{-}1991}$				
Indep. Var.:	1990-1995	1980-1990		1990-1995		
	Δ(tariff)	Δ(M/P)	Δ(MPC)	Δ(tariff)	Δ(M/P)	Δ(MPC)
	(1)	(2)	(3)	(4)	(5)	(6)
Coefficient	0.297***	-0.021	-0.207	-0.032	0.128	0.159
	(0.066)	(0.106)	(0.553)	(0.062)	(0.221)	(0.248)
Obs.	284	284	284	284	284	284
R-Squared	0.222	0.130	0.130	0.130	0.130	0.131

Notes: Robust standard errors in parentheses. ***p < 0.01; **p < 0.05, *p < 0.10. Independent variables: region dummies and changes in average wages by level of schooling (not shown). Unit of observation is a micro-region, according to the 1980 definition. Census data from 1980, 1991, and 2000. Regressions weighted by the precision of first-stage estimates of the dependent variable (inverse of the standard errors). First stage is a regression for 20-60 year-old male employees. First stage independent variables: age, age squared, dummies for years of schooling, urban area, and micro-region, and interactions between a dummy for white and micro-region dummies.

Table 6: Alternative Samples in the 1st Stage, Impact of Tariff Change on Racial Wage Gap, Brazilian Micro-regions, 1991-2000 – Dependent Variable: Change in Conditional Racial Wage Gap

	Men (1)	Men & Women (2)	Employees & Self-empl. (3)	Employees (4)	Men, Not in School, Full-time (5)
Δ(tariff)	0.367***	0.295***	0.317***	0.187***	0.334***
	(0.064)	(0.055)	(0.057)	(0.061)	(0.063)
Observations	480	480	480	480	480
R-Squared	0.193	0.269	0.209	0.219	0.173
1991 w gap	0.151	0.151	0.133	0.127	0.149
Δ(w gap)	0.006	-0.001	0.004	-0.008	0.006

Notes: Robust standard errors in parentheses. ***p < 0.01; **p < 0.05, *p < 0.10. Independent variables: region dummies and changes in average wages by level of schooling (not shown). Unit of observation is a micro-region. Census data from 1991 and 2000. Regressions weighted by the precision of first-stage estimates of the dependent variable (inverse of the standard errors). First stage is a regression for 20-60 year-old male employees. First stage independent variables: age, age squared, dummies for years of schooling, urban area, and micro-region, and interactions between a dummy for white and micro-region dummies.

Table 7: Heterogeneity of the Initial Wage Gap and of the Impact of Tariff Changes, Brazilian Micro-regions, 1991-2000 – Dependent Variable: 1991 Conditional Racial Wage Gap in Panel A and Change in Conditional Racial Wage Gap in Panel B

	(1)	(2)	(3)	(4)
Panel A: Initial Wage Gap and Local Labor Market Characteristics (1991)				
Market Concentration	0.306***			0.365***
	(0.117)			(0.110)
% Black		0.075**		0.095***
		(0.031)		(0.031)
Prejudice Index			0.601***	0.631***
			(0.168)	(0.158)
Region Dummies	X	X	X	X
Observations	480	480	480	480
R-squared	0.065	0.050	0.066	0.126
Panel B: Heterogeneity of the Impact of Tariff Change on Racial Wage Gap				
Δ(tariff)	0.029	0.309*	0.119	-0.040
	(0.092)	(0.164)	(0.280)	(0.354)
Δ(tariff) × Market Concentration	4.770***			4.579***
	(1.545)			(1.493)
Δ(tariff) × % Black		-0.131		0.066
		(0.334)		(0.328)
Δ(tariff) × Prejudice Index			1.301	0.722
			(3.645)	(3.431)
Δ Avg W by Schooling	X	X	X	X
Region Dummies	X	X	X	X
Observations	480	480	480	480
R-squared	0.124	0.115	0.113	0.131

Notes: Robust standard errors in parentheses. ***p < 0.01; **p < 0.05, *p < 0.10. Independent variables in Panel A: region dummies (not shown), industrial concentration, % black, and prejudice index in 1991. Independent variables in Panel B: region dummies and changes in average wages by level of schooling (not shown), and changes in industrial concentration, % black, and prejudice index. Unit of observation is a micro-region. Census data from 1991 and 2000. Regressions weighted by the precision of first-stage estimates of the dependent variable (inverse of the standard errors). First stage is a regression for 20-60 year-old male employees. First stage independent variables: age, age squared, dummies for years of schooling, urban area, and micro-region, and interactions between a dummy for white and micro-region dummies.

Table A.1: Employment Share by Sector – Brazil, 1991 Census

	1991	
	Including services	Excluding services
Agriculture	21.63	55.45
Mineral mining	0.80	2.05
Petroleum, gas extraction	0.11	0.28
Nonmetallic metallic	0.94	2.41
Metals	2.75	7.05
Machinery, equipment	0.50	1.28
Electric, electronic equipment	0.54	1.38
Transport	0.59	1.51
Wood, furniture	1.67	4.28
Paper, publishing, printing	0.87	2.23
Rubber	0.16	0.41
Chemicals	0.73	1.87
Petroleum refining	0.17	0.44
Pharma, perfumes	0.22	0.56
Plastic	0.34	0.87
Textiles	1.06	2.72
Apparel	2.52	6.46
Footwear	0.74	1.90
Food processing	2.33	5.97
Other manufacturing	0.34	0.87
Services	61.00	-
Total	100	100

Table A.2: Matching of Sectors between Kume et al (2003) and the 1991 Census

Kume et al (200)		1991 Census	Aggregated	
1	Agriculture	11-37, 41, 42, 581	Agriculture	1
2	Mineral mining	50, 53-59	Mineral mining	2
3	Petroleum, gas extraction	51, 52	Petroleum, gas extraction	3
4	Nonmetallic mineral	100	Nonmetalic mineral	4
5	Metals	110	Metals	5
6	Nonmetallic manufacturing	110	Metals	5
7	Other nonmetalic manufacturing	110	Metals	5
8	Machinery, equipment	120	Machinery, equipment	6
10	Electric materials	130	Electric, electonic equipment	7
11	Electonic equipment	130	Electric, electonic equipment	7
12	Automobile, transportation	140	Transport	8
13	Vehicle parts, other vehicles	140	Transport	8
14	Wood, furniture	150, 151, 160	Wood, furniture	9
15	Paper, publishing, printing	170, 290	Paper, publishing, printing	10
16	Rubber	180	Rubber	11
17	Chemicals	200	Chemicals	12
18	Petroleum refining	201, 202, 352, 477	Petroleum refining	13
19	Other chemicals	200	Chemicals	12
20	Pharma, perfume	210. 220	Pharma, perfume	14
21	Plastic	230	Plastic	15
22	Textile	240. 241	Textile	16
23	Apparel	250, 352	Apparel	17
24	Footwear	190, 251	Footwear	18
25	Coffee	260, 261, 270, 280	Food processing	19
26	Vegetables	260, 261, 270, 280	Food processing	19
27	Animal Slaughter	260, 261, 270, 280	Food processing	19
28	Dairy	260, 261, 270, 280	Food processing	19
29	Sugar	260, 261, 270, 280	Food processing	19
30	Vegetable oils	260, 261, 270, 280	Food processing	19
31	Other food processing	260, 261, 270, 280	Food processing	19
32	Other manufaturing	300	Other manufaturing	20

Table A.3: Impact of Tariff Change on Racial Wage Gap by Level of Schooling, Brazilian Micro-regions, 1991-2000 – Dependent Variable: Change in Conditional Racial Wage Gap – Conditional Racial Wage Gap in the 1st Stage Estimated by Level of Schooling

	Schooling levels:			
	Primary School	Elementary School	High School	College/ University
	(1)	(2)	(3)	(4)
Δ(tariff)	0.116*	0.386***	0.118	-0.358
	(0.064)	(0.102)	(0.103)	(0.331)
Observations	480	479	479	464
R-Squared	0.060	0.353	0.414	0.366
1991wage gap	0.105	0.113	0.112	0.113
Δ(w gap)	0.000	-0.010	0.003	0.023

Notes: Robust standard errors in parentheses. ***p < 0.01; **p < 0.05, *p < 0.10. Independent variables: region dummies and changes in average wages by level of schooling (not shown). Unit of observation is a micro-region. Census data from 1991 and 2000. Regressions weighted by the precision of first-stage estimates of the dependent variable (inverse of the standard errors). First stage is a regression for 20-60 year-old male employees. First stage independent variables: age, age squared, dummies for years of schooling, urban area, and micro-region, and interactions between a dummy for white and micro-region dummies.

Table A.4: Individual Level Estimates of the Impact of Tariff Change on Racial Wage Gap, Brazilian Micro-regions, 1991-2000 – Dependent Variable: Natural Logarithm of Hourly Wage

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
white × tariff	0.355***	0.242***	0.255***	0.241***	0.249***	0.243***	0.268***
	(0.061)	(0.039)	(0.038)	(0.039)	(0.040)	(0.038)	(0.039)
Controls:							
Individual Characts.	X	X	X	X	X	X	X
Agg. Micro-region Vars.:							
Avg. Wage by School.		X	X	X	X	X	X
School. Comp. by Race			X				X
# Schools/1,000 Children				X			X
Labor Market Characts.					X		X
Migrantion						X	X
Year dummies	X	X	X	X	X	X	X
Micro-region dummies	X	X	X	X	X	X	X
Observations (millions)	4.01	4.01	4.01	4.01	4.01	4.01	4.01
R-squared	0.512	0.516	0.516	0.516	0.516	0.516	0.516

Notes: Robust standard errors in parentheses (clustered at micro-region level). ***p < 0.01; **p < 0.05, *p < 0.10. Individual controls: age, age squared, dummies for years of schooling, dummy for urban area, and interactions with the year dummy. Aggregate (micro-region) controls: region dummies, average wages by level of schooling, composition of the labor force by level of schooling and race, number of public schools per 1,000 children, share of informal employees, % of unemployed, and % of migrants. Unit of observation is an individual. Census data from 1991 and 2000.