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Trade and labor market segregation in Colombia

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Abstract

Gary Becker's theory of discrimination argues that increasing competition will reduce discrimination by reallocating market share to less discriminatory firms. We develop a simple model in which increased competition can also affect discrimination by affecting firm-level hiring decisions. We use the 1984–1991 Colombian trade liberalization episode and plant-level data to investigate this claim. We find that plants in industries that faced the greatest reductions in tariff protection increased the female share of their workforce more than plants in industries that saw little or no reduction in tariffs. In addition, we find that exporting plants tended to employ a higher share of female workers than non-exporters did. In contrast, we find little evidence that trade liberalization drove discriminating plants from the market.

KEYWORDS

gender discrimination, international trade, labor market segregation

JEL CLASSIFICATION

F16, F10, J7, F6

1 | INTRODUCTION

There is significant interest surrounding the issue of gender equality in development economics. This partly reflects well-established research showing strong correlations between female earnings and child outcomes, especially with regard to education and health (e.g., see Duflo, 2003;

Qian, 2008; Thomas, 1990), suggesting a link between gender equality and long-run development. Since many developing countries have experienced significant trade liberalization in recent decades, this raises the question of how trade liberalization in developing countries has affected the labor market for female workers. Specifically, we are interested in whether trade liberalization (and the accompanying increased foreign competition) has a differential impact on female labor by reducing employment discrimination as in Gary Becker's (1957) canonical model.

This paper makes three main contributions to the large literature on the link between trade and gender inequality. First, the Becker model of competition and discrimination argues that firms that have a "taste for discrimination" (i.e., place a negative valuation on hiring women) tend to have higher costs and thus lower profits than their less discriminating counterparts. To the extent that capital investment only flows to more profitable, non-discriminating firms, competition will, in the long run, result in the expansion of non-discriminating firms and the eventual exit of discriminating firms. Thus, competition reduces *industry-level* discrimination by reallocating market shares across firms. However, it is noteworthy that the Becker framework provides no mechanism by which a marginal change in competition would affect *firm-level* discriminatory behavior. Specifically, in the Becker model, firm-level hiring decisions are a function of preferences (the disutility of hiring female workers), technology (the marginal rate of substitution between male and female workers) and wages (the wage differential). A marginal change in market competition does not directly affect any of these, and so it would not be predicted to affect firm-level hiring decisions. However, as we show in a brief theoretical section, competition, including increased foreign competition, can potentially reduce firm-level discrimination in the short run by increasing the marginal cost of discriminating behavior. That is, firms that perceive a tradeoff between profits and the female share of their workforce could potentially respond to increased competition by hiring more women. Intuitively, to the extent that increased competition results in an increase in the elasticity of demand for a firm's products, such competition will also result in an increased cost to firm-level discriminatory behavior since the higher costs resulting from discrimination will result in a greater loss in market share. We provide a theoretical mechanism by which competition can not only reduce discrimination through traditional selection effects (e.g., by forcing discriminating firms from the market), but can also reduce *firm-level* discriminatory behavior by influencing firm-level hiring decisions.

Second, previous studies of employment in developing countries have typically found that women are concentrated in the export-oriented industries of the manufacturing sector such as textiles and food products (e.g., Catagay & Berik, 1991; Ozler, 2000).[1] Given that our theoretical framework suggests that competition can influence *firm-level* discriminatory behavior, it seems that since exporting firms face more competition than firms that produce only for the domestic market, exporters should discriminate less than non-exporters. However, it is difficult to determine if female concentration in export industries is due to the exporting nature of those industries, or to some unobserved industry characteristics. In this paper we investigate whether women are more likely to be employed in exporting *plants* within an industry. Using Colombian manufacturing data we find that this is the case: the female share of employment is higher among exporting plants.

Finally, in contrast to previous studies, we use plant-level data to directly examine the impact of changes in foreign competition on firm selection and hiring decisions. Previous research on the gender-specific effects of trade on the labor market have relied, almost exclusively, on household- or individual-level data (e.g., see Benguria & Ederington, 2023; Berik et al., 2004; Black & Brainerd, 2004; Gaddis & Pieters, 2017; Hazarika & Otero, 2004; Keller & Utar, 2022; Mansour et al., 2022; Menon & Rodgers, 2009; Yahmed, 2023). The main focus of these papers is to estimate the effect of increased trade competition on gender-specific wages (at either regional or sectoral levels) or labor market outcomes (e.g., labor force participation).[2] However, the primary implication of the Becker model concerns the effect of competition on industry hiring: increases in competition within a sector lead to the growth of less discriminatory firms, which proportionally hire more women. The impact on relative wages or labor force participation occurs only indirectly through the relative growth in demand for female workers, and need not occur at all if there are a sufficient number of nondiscriminatory employers. Our use of plant-level data allows us to more directly investigate Becker-type impacts of trade, and also differentiate between two mechanisms through which competition may reduce discrimination: the traditional Becker mechanism of reallocating market shares across firms, and our proposed mechanism (see Section 2) of firm-level hiring decisions. Specifically, we can measure the extent to which increased foreign competition induces discriminating plants (i.e., plants with low female shares) to exit relative to the extent to which it induces all plants to increase the female share of their labor force. Our primary conclusion is that increased competition (in the form of exogenous trade liberalization) reduced discrimination primarily by affecting the hiring practices of plants. That is, we find that plants in industries that faced the greatest reductions in tariff protection increased the female share of their workforce more than plants in industries that saw little or no reduction in tariffs. In contrast, we find little evidence that trade liberalization drove plants with low female shares from the market.

In this paper, we exploit a natural experiment: the Colombian trade liberalization episode of 1984. Starting in 1985, and following its entry into the GATT/WTO, Colombia undertook major unilateral trade liberalization of its manufacturing sector. This liberalization entailed both a reduction in the average level of protection and a collapse of the distribution of protection as Colombia moved to a more uniform tariff structure. We exploit this cross-sectional variation in tariff reductions to see if a greater increase in foreign competition (i.e., a larger tariff reduction) resulted in higher shares of female workers across Colombian plants. Both of these differences enable us to more directly examine the implications of the Becker model.

The most related research in the literature is a sequence of papers (Aguayo-Tellez et al., 2013; Juhn et al., 2013; Juhn et al., 2014) that argue that access to export markets might induce firms to acquire new technologies that are more complementary to female labor. As evidence for their channel, they use the North American Free Trade Agreement (NAFTA) to demonstrate how tariff reductions induced both firm investment and improved female labor market outcomes (in terms of both labor share and earnings). Clearly, our mechanism (decreased discrimination) and theirs (technology upgrading) are not mutually exclusive,[3] but there are important differences. First, they utilize a Melitz (2003)/Bustos (2011) export-pull type model in which firm dynamics are driven by the export side (i.e., it is increased access to foreign markets that induces firm selection effects and technology upgrading), so their empirics concentrate on bilateral trade agreements and how reductions in U.S. tariffs affected the Mexican labor market.[4] In contrast, our model and empirics provide a channel through which an episode of unilateral trade liberalization (such as the Colombian experience) can differentially impact the female labor market. Second, our model provides a mechanism by which trade liberalization can reduce gender inequality even in the absence of "technology upgrading." We show in Section 4 that exporting plants hire a larger percentage of women employees, even after controlling for a range of plant-level characteristics, and we show in Section 6 that domestic tariff reductions increase the female labor share, even after controlling for changes in plant inputs. We see our mechanism (decreased discrimination) as working in concert with technology upgrading to explain how trade liberalization might impact female labor market outcomes.

In what follows, Section 2 provides a model of how increased trade competition can influence firm-level hiring decisions by firms with a "taste for discrimination." In Section 3 we discuss the data and the Colombian trade liberalization episode, and then, in Section 4 we look at the relative female shares of exporting and non-exporting plants. In Section 5 we utilize the Colombian trade liberalization to investigate the impacts of increased foreign competition on plant selection and in Section 6 we look at plant-level hiring. We conclude in Section 7.

2 | GENDER DISCRIMINATION: THEORY

In this section we present a partial equilibrium model of competition and gender discrimination. We follow Becker (1957) and Arrow (1973) in assuming that firms are not strictly profit maximizers but rather maximize a utility function that evidences a tradeoff between profits and the number of male and female employees. However, in contrast to Becker (1957) and Arrow (1973), we explicitly assume a monopolistically competitive environment that allows both discriminating and non-discriminating firms to coexist in equilibrium. The question of interest is the effect of an increase in competition on the equilibrium.

We assume an economy with two sectors: one sector consists of a numeraire good, x_0, while the other sector is characterized by differentiated products. The preferences of a representative consumer are defined by the following utility function:

$$U = x_0 + \log\left[\int_0^n y(j)^\rho dj\right]^{1/\rho} \tag{1}$$

where x_0 is consumption of the numeraire good, $y(j)$ represents consumption of brand j of the differentiated product good and n represents the number of available varieties (firms) in the differentiated product sector. It is straightforward to show that with these preferences, the elasticity of substitution between any two products is $\sigma = 1/(1-\rho) > 1$, and aggregate demand for brand i is given by:

$$y(i) = \frac{p(i)^{-\sigma} E}{\int_0^n p(j)^{1-\sigma} dj} \tag{2}$$

where $p(i)$ is the price of good i and E represents the total number of consumers in the country.

We assume that production of the differentiated product good requires a sequence of tasks to be performed (e.g., in the automobile sector, one task might involve installing the brakes and another might be installing the windshield). This treatment of production is similar to that of Becker and Murphy (1992) and Kremer (1993). Rather than assuming a discrete set of tasks, it will be convenient to assume production is defined by the completion of a continuum of tasks along the unit interval. Letting t be the index for tasks and letting the cost of task t be given by $w(t)$, then the marginal cost of producing a variety of the differentiated product good is given by:

$$c = \int_0^1 w(t) dt. \tag{3}$$

We assume that either a male employee can be hired to complete a task at cost w_m or a female employee can be hired at cost w_f where $w_m > w_f$ (thus, we assume that male and female employees are equally productive in producing the differentiated product, but that a wage differential exists in the economy).[5] Defining $z_i \in [0, 1]$ as the female share employed by firm i, the marginal costs of firm i are given by:

$$c_i = w_m - z_i(w_m - w_f). \tag{4}$$

It should be clear that, given the existence of a wage differential, a cost-minimizing firm will choose to hire only women (i.e., set $z_i = 1$). However, as discussed previously, we assume firms maximize a utility function that encompasses both profits and a "taste for discrimination," which we capture by assuming that the firm owner/manager derives extra disutility from hiring female workers, defined by $\phi_i(z_i, y_i)$. Thus, firms choose price, p_i, and the female share of the labor-force, z_i, to maximize:

$$\max_{p_i, z_i} (p_i - c_i)y_i - \phi_i(z_i, y_i) \tag{5}$$

While we are thinking of $\phi_i(z_i, y_i)$ as a disutility function, it could, as easily, reflect alternative (non-wage) costs to hiring women relative to men. For example, in Bøler et al. (2018), the underlying assumption was that male and female labor were imperfect substitutes because female workers were perceived as less flexible in their working hours. This naturally leads to questions about the functional form of $\phi_i(z_i, y_i)$. If we assume the disutility takes the form of increased marginal costs to hiring female workers (i.e., $\phi_i(z_i, y_i) = \phi_i(z_i)y_i)$), then we get a corner solution in which each firm chooses to hire only male or female workers. Since we are interested in exploring how competition influences individual firm-level hiring decisions, we instead follow Arrow (1973) in which firms care only about the fraction of their workforce that is female (i.e., $\phi_i(z_i, y_i) = \phi_i(z_i)$). Note, however, that with the Arrow assumption the disutility costs are taking on the form of fixed costs in the profit maximization condition, leading to firm scale effects (which we discuss below).

Assuming $\phi_(z_i, y_i) = \phi_i(z_i)$ where $\phi_i'(z_i) > 0$ and $\phi_i''(z_i) > 0$, from the first-order condition with respect to p_i, one can derive that firms use a constant mark-up pricing rule where:

$$p_i = \frac{\sigma}{\sigma - 1} c_i \tag{6}$$

From the first-order condition with respect to z_i, one can derive that z_i is implicitly defined by:

$$\phi_i'(z_i) = \frac{\sigma - 1}{\sigma} \frac{(w_m - w_f)E[w_m - (w_m - w_f)z_i]^{-\sigma}}{\int_0^n (c_j)^{1-\sigma} dj}. \tag{7}$$

The left-hand side of (7) represents the marginal cost to the firm of increasing the female share of its employees, while the right-hand side represents the marginal benefit (in lower costs of production). Firms will choose to employ men (i.e., $z_i < 1$) if and only if the marginal disutility of hiring women is sufficiently high (and outweighs the cost of the wage differential).[6]

There are two items to note about the above derivations. First, firms with the greatest "taste for discrimination" (i.e., with the highest values of $\phi_i'(z_i)$ for any z_i) will employ the lowest share of female workers (i.e., choose the lowest z_i). As in Becker (1957) and Arrow (1973), the female share of the workforce will vary across firms, with more discriminating firms employing a lower share of women and less discriminating firms employing a higher share. In addition, the more discriminating firms, since they have higher marginal costs of production (i.e., higher c_i), will be less profitable than the less discriminating firms. As Becker (1957) and Arrow (1973) note, one would expect that, in a competitive environment, capital would flow to the more profitable (less discriminating firms), driving the more discriminating firms out of the market in the long run. Thus, the traditional theory of discrimination argues that competition reduces discrimination in the long run through a selection effect where only the most profitable (least discriminating) firms survive.

However, these derivations suggest a second mechanism through which competition can affect discrimination: by affecting firms' hiring decisions (i.e., by affecting the optimal choice of z_i). For expositional simplicity, assume that firms are symmetric and have identical utility functions (i.e., $z_i = z$ in equilibrium). In this case, the first-order condition (7) reduces to:

$$\phi_i'(z_i) = \frac{\sigma - 1}{\sigma} \frac{(w_m - w_f)E}{[w_m - (w_m - w_f)z]n} \tag{8}$$

It is common in the industrial organization literature (e.g., see Vives, 2008) to decompose increases in production market competition into (1) changes in scale (i.e., changes in n) and (2) changes in the elasticity of substitution between product varieties (i.e., changes in σ). We consider the comparative statics of each below.

First, consider an exogenous increase in competition captured by an increase in the elasticity of substitution (i.e., σ) between firm varieties. This requires some discussion since σ in is ostensibly a taste parameter. The reason that increases in product market competition are modeled as an increase in σ is that an increase in product market competition could equivalently be seen as a reduction in the regulatory segmentation of the market, allowing consumers to more freely substitute between available varieties (e.g., see Aghion et al., 2001). This potential regulatory segmentation of the market also includes geographic segmentation (e.g., customs duties or regulations) where a reduction in trade costs is viewed as increasing σ (see Syverson, 2004).[7]

We take a derivative of the right-hand side of (8) with respect to σ to derive that:[8]

$$\frac{\partial RHS}{\partial \sigma} = \frac{1}{\sigma^2} \frac{(w_m - w_f)E}{[w_m - (w_m - w_f)z]n} > 0 \tag{9}$$

An increase in competition will increase the marginal benefit of employing more women, increasing the female share of a firm's workforce (i.e., increasing z_i). This negative impact of competition on discrimination is due to an elasticity effect: firms that face a more elastic demand for their product will incur a higher cost to discriminating behavior, as the resulting increase in costs and prices causes a correspondingly greater loss of market share. To the extent that greater competition increases the elasticity of demand for a firm's product, such competition will increase the cost of discriminating behavior resulting in the hiring of more female employees.

However, one could also think of an increase in competition as an exogenous increase in the number of firms in the market (perhaps due to a decline in entry costs). Taking a derivative of the right hand side of (8) with respect to n (and holding σ constant) yields:

$$\frac{\partial RHS}{\partial n} = \frac{1 - \sigma}{\sigma n^2} \frac{(w_m - w_f)E}{[w_m - (w_m - w_f)z]} < 0 \tag{10}$$

In this case, an increase in competition (modeled as an exogenous increase in the number of firms), decreases the marginal benefit of hiring women and so decreases the share of female employees (increasing gender discrimination). This positive effect of competition on discrimination is due to a scale effect: larger firms suffer a greater cost to discriminating behavior since the increased marginal costs of production resulting from such behavior affect a larger volume of production. To the extent that greater competition reduces firm size, it might reduce the costs of discrimination and result in the hiring of more male employees. Differentiating between these different types of increases in competition is, to our knowledge, a unique feature of our model.

In the empirics that follow, we consider an episode of significant unilateral trade liberalization as a proxy for increased competition (i.e., the Colombian trade liberalization episode of the 1980s). In this case, we can think of increased import competition as both (1) increasing the number of competitors in the market (i.e., an increase in n) and (2) increasing the elasticity of demand for a firm's products (i.e., an increase in σ). Indeed, Krugman (1979) shows that, under reasonable parameter assumptions, one of the pro-competitive aspects of international trade is that it increases the elasticity of demand faced by firms.[9] A contribution of our model is that it suggests that competition does not simply affect industry-level discrimination by reallocating market shares across firms, but can also affect firm-level discriminatory behavior. To tie our empirics to the model, we use plant-level data to investigate whether increased foreign competition does in fact affect plant-level hiring decisions, which we assume are made at the plant level following Fernandes (2007) and Roberts and Tybout (1997).

However, we can also consider the impact of exporting on firm-level hiring decisions. In this case, while we would argue that exporting firms face more competitors (i.e., an increase in n) and less geographical/regulatory segmentation from other competing varieties (i.e., an increase in σ), and they also enjoy an increase in market size (i.e., an increase in E). Taking the derivative of the right hand side of (8) with respect to E (and holding σ and n constant), it is direct to derive that $\frac{\partial RHS}{\partial E} > 0$ (i.e., an increase in market size increases the female share), suggesting that exporting firms employ a higher share of female workers due to both: (1) greater scale and (2) higher elasticity of demand. In Section 4 of the paper, we investigate the hiring decisions of exporting and non-exporting plants.

To summarize, the overall effect of competition on discrimination reflects a combination of elasticity and scale effects. Note that the force of these effects depends partially on the degree of existing competition in the market. For example, it is direct to derive from the above calculation that the elasticity effect is decreasing in σ (see Equation 9) and the scale effect is increasing in σ (see Equation 10). Thus, in industries that are already sufficiently non-competitive (i.e., have a sufficiently small σ), the elasticity effect will tend to dominate, and competition should reduce the degree of gender discrimination. This might not be the case in industries that are sufficiently competitive (i.e., have a sufficiently large σ), where the scale effect will tend to dominate. In other words, the impact of competition on hiring decisions and the degree of gender discrimination could be a function of the existing degree of competition in the industry.[10] Some empirical evidence for this conditional result can be found in Black and Brainerd (2004), in which the impact of trade volume on gender wage differentials depends on the degree of concentration of the industry. While we do not have industry-level data on elasticity, we do control for changes in plant size in our empirics.

3 | DATA: COLOMBIAN TRADE LIBERALIZATION

In this paper we use plant-level panel data from Colombia to examine the prediction of the Becker model that changes in the level of competition reduce discrimination in the labor market.

Like many developing countries, Colombia followed a policy of import substitution in the 1950s and 1960s. In the 1970s, this policy was reconsidered as Colombia sought entry into the GATT. Starting in 1985 and culminating in 1991, Colombia systematically lowered its trade barriers with the aim of creating a relatively uniform structure of protection comparable to those in developed countries. Since high tariffs shield domestic producers from competitive pressures from producers in other countries, we treat changes in tariffs as changes in the level of competition faced by firms in Colombia.

This episode of trade liberalization has been extensively studied in the international trade literature since several of its features make it attractive from an empirical standpoint.[11] First, prior to liberalization, Colombia had relied primarily on tariffs as a means of trade protection and so the decline in tariffs was significant (the average tariff reduction was 31.4 percentage points between 1984 and 1991), and also provides an accurate measure of the overall change in trade policy.[12] Second, the Colombian trade liberalization had a significant impact on the structure of protection, with some industries receiving extensive tariff cuts while other industries were not significantly affected. It is this variation in tariff reductions that we exploit in our empirics. Finally, the main policy objective of the Colombian government was to achieve uniformity in tariff levels across industries, and to make their tariff levels comparable to those of other WTO members as part of Colombia's entry into the GATT/WTO.[13] Thus, the Colombian tariff reductions were less susceptible to industry pressure or governmental preferences, and can be plausibly treated as exogenous (see Attanasio et al., 2004; Goldberg & Pavcnik, 2003; Goldberg & Pavcnik, 2007 on this point as well as for a more detailed description of the reforms). Evidence on both the decline in overall tariffs and the narrowing of tariff differences across sectors is provided in Figure 1,

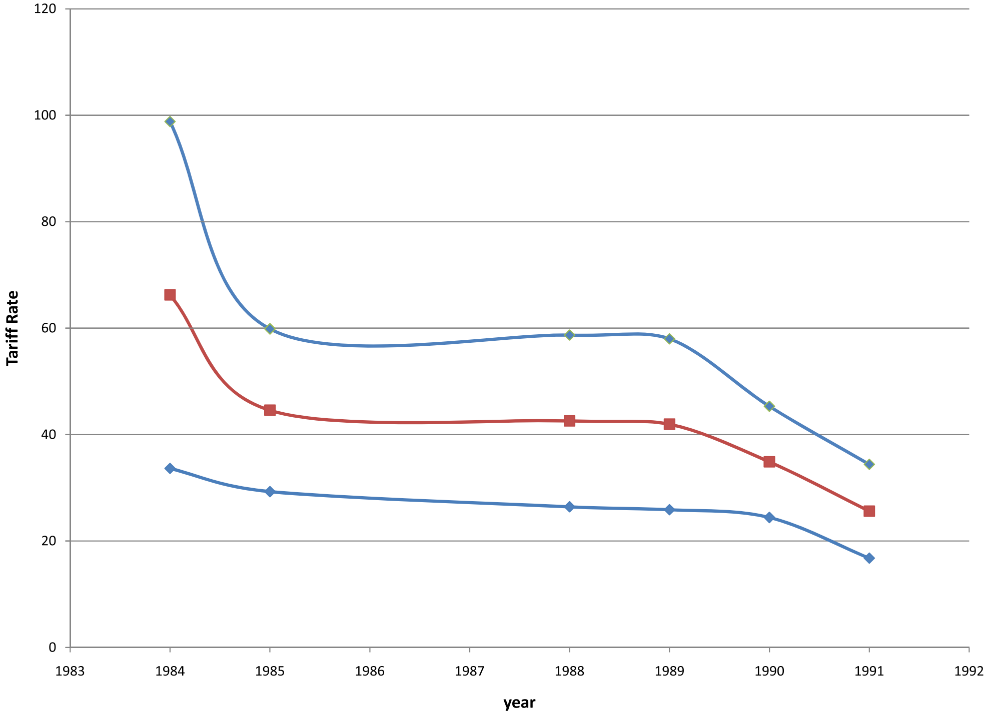

FIGURE 1 Mean tariff ± 1 standard deviation by year. The red line plots the mean ad-valorem tariff by year from 1984 to 1991 for the plants in our sample. The blue lines plot the mean tariff plus and minus one standard deviation. [Colour figure can be viewed at wileyonlinelibrary.com]

which shows the mean tariff, and the mean tariff plus and minus one standard deviation, by year across the plants in our sample. Note both the significant average decline, and the narrowing of the dispersion of tariffs.

All data come from a plant-level dataset produced from the Colombian Manufacturing census by DANE (National Statistical Institute) for the years 1977 through 1991.[14] This is an unbalanced panel covering industrial production in all three-digit ISIC industries in Colombia. For each year, the survey has collected data on production, sales, inputs, inventories, exports and the industry code (four-digit ISIC) of each Colombian plant. It should be noted that plants with fewer than 10 employees are included in the census prior to 1982, but excluded after 1983 (although a small proportion are included following 1985). Our sample includes a total of 6035 plants in 1984; 3760 remain in the sample in 1991, which totals 6972 plants.[15] The Colombian manufacturing industry is characterized by relatively small-scale plants (70% of production is by plants employing fewer than 50 workers), and high levels of entry and exit (with average entry rates of 11% and exit rates of 10%).[16] However, the distribution of plants across industries is relatively constant over time, with the major industries being food, apparel, textiles, printing, and metal products. For measures of trade liberalization, we employ ad-valorem tariff levels at the four-digit ISIC level from the Departamento Nacional de Planeación.[17]

We concentrate on the female share of workers in Colombian plants. We should note that our data are plant level, and unfortunately do not include information about the firm, although most firms in Colombia during this period were single-establishment firms.[18] Our analysis follows the conventional Becker treatment of discrimination in which male and female wages are set by the aggregate labor market, and individual firms take these wages (and the presumed wage differential) as given in making employment decisions. In this framework, discriminatory behavior is revealed by the more discriminatory firms being less willing to hire the (cheaper) female workers relative to an otherwise similar firm. Therefore the focus of our empirics is on the reallocation of workers across plants (and industries) as opposed to the impact on aggregate labor market outcomes such as the gender wage gap. Angel-Urdinola and Wodon (2003) provide estimates of the gender wage gap in Columbia for both the whole economy and by sector over the time period and find that the gender wage gap is essentially zero in 1985 (the starting year of their data) but rises to around 4% by 1991, while the wage gap in the manufacturing sector is relatively flat over this period.

Table 1 presents summary statistics for our data. Most of our analysis concentrates on the change between 1984 and 1991, so Table 1 includes means for both years where applicable, as well as the mean change in the relevant variables. Some trends stand out: between 1984 and 1991, average tariffs dropped dramatically; the percentage of plants exporting increased and exports (as a share of sales) increased; and plants became more productive, used more energy, and paid higher wages.[19] There was virtually no change, however, in the average female share of total workers over this period, although the female share of skilled workers increased slightly.

Since we are examining the female share of workers in Colombian plants, it is worth briefly discussing female labor force participation in Columbia more generally. Colombia's female labor force participation was higher than that of most Latin American countries: in the early 1990s, it was 43%, approximately ten percentage points higher than the Latin American average. This is generally attributed to more widely available contraception in Colombia (between 1964 and 1994, fertility rates fell from 7.4 to 2.7 children per woman) and very high violent crime rates for men, leading to a high probability of widowhood.[20] In addition, between 1984 and 1991 female labor force participation rates rose fairly dramatically, particularly for younger women (see Arango & Posada, 2005, 2007). Female workers are, not surprisingly, concentrated in the apparel and textile

TABLE 1 Summary statistics.

Variable	1984 Mean (SD)	1991 Mean (SD)	Change Mean (SD)
Female share, total	0.365 (0.286)	0.355 (0.265)	0.007 (0.145)
Female share, unskilled	0.341 (0.354)	0.314 (0.333)	−.005 (0.164)
Female share, skilled	0.497 (0.301)	0.504 (0.269)	0.024 (0.261)
Exports (0/1)	0.106 (0.308)	0.182 (0386)	
Exports/Total sales	0.018 (0.101)	0.042 (0.151)	
Plant age	14.55 (12.23)		
Employment (log)	3.55 (1.02)	3.34 (1.15)	
Salary (log)	5.25 (0.43)	6.83 (0.46)	
Skill ratio	0.213 (0.168)	0.269 (0.213)	
Energy use (log)	2.67 (1.15)	4.65 (1.30)	
Productivity (log)	6.31 (0.80)	8.19 (0.93)	
Capital/labor (log)	4.94 (1.41)	6.58 (1.58)	
Office equipment	0.083 (0.111)	0.094 (0.128)	
Female management	0.202 (0.283)	0.237 (0.295)	
Corporation	0.170 (0.376)	0.174 (0.379)	
Proprietorship	0.139 (0.346)	0.123 (0.329)	
Partnership	0.691 (0.462)	0.703 (0.457)	
Tariff	0.662 (0.328)	0.257 (0.088)	−0.403 (0.269)
Four-plant concentration	0.400 (0.192)		
Observations	6035	6972	3760

Note: See Appendix A for variable definitions. Changes are the average of the changes for plants for which we have data in both 1984 and 1991.

industries; in 1984, the female share of workers in these (four-digit) industries was as high as 80%, relative to the overall average of 35%.

4 | EXPORTS AND THE PLANT-LEVEL FEMALE LABOR SHARE

As mentioned in the Introduction, one interpretation of exporting plants is that they face higher degrees of competition than non-exporting plants, because they compete in international markets. According to our model, one might expect exporters to discriminate less than non-exporters. Studies of employment in developing countries have typically found that women are concentrated in the export-oriented industries of the manufacturing sector (e.g., Catagay & Berik, 1991; Ozler, 2000), which typically includes the textile, garment, electronics, leather and agricultural processing industries. However, it is difficult to determine if this female concentration is due to the exporting nature of those industries, or simply to unobserved industry characteristics. Indeed,

TABLE 2 OLS: Which plants are more female?

Variable	(1) Total $\hat{\beta}$ (SE)	(2) Unskilled $\hat{\beta}$ (SE)	(3) Skilled $\hat{\beta}$ (SE)
Exports (0/1)	0.044 (0.005)***	0.053 (0.007)***	0.019 (0.006)***
Plant age	−0.000 (0.000)	−0.000 (0.000)**	0.000 (0.000)
Employment (log)	0.010 (0.002)***	0.030 (0.003)***	−0.072 (0.003)***
Salary (log)	−0.042 (0.005)***	−0.008 (0.007)	−0.071 (0.008)***
Skill ratio	0.061 (0.010)***	−0.106 (0.015)***	−0.295 (0.015)***
Energy use (log)	−0.022 (0.002)***	−0.030 (0.002)***	−0.013 (0.002)***
Capital/labor (log)	−0.010 (0.001)***	−0.010 (0.002)***	−0.004 (0.002)
Productivity (log)	−0.011 (0.003)***	−0.015 (0.003)***	0.009 (0.004)***
Office equip	0.034 (0.015)**	0.032 (0.020)	0.077 (0.021)***
Female mgnt	—	0.084 (0.007)***	0.023 (0.009)***
Corporation	−0.032 (0.006)***	−0.035 (0.008)***	0.003 (0.007)
Proprietorship	−0.015 (0.005)***	−0.032 (0.006)***	0.036 (0.010)***
Observations	53,521	52,244	49,082
Plants	10,935	10,801	10,144
Industries (4-digit)	96	96	95
R^2	0.562	0.593	0.205

Note: The dependent variable is the plant's female share of workers; annual data are pooled 1984–1991. Time and three-digit industry fixed effects are included, and standard errors are clustered by plant. See Appendix A for data definitions and notes. "Total" includes owners and managers, in addition to skilled and unskilled workers, in the calculation of the female share variable. *** indicates statistical significance at the 99% level or better; ** at 95%; and * at 90%.

Catagay and Berik (1991) and Ozler (2000) also find that women in developing countries are typically employed in low-paying, low capital-intensity, small-scale firms, hardly the standard characteristics of exporting firms (see Roberts & Tybout, 1997 for empirical evidence on the types of plants that export). While Ozler (2000) has plant-level data from Turkey on male and female employment, as well as some other plant-level characteristics, her measure of export status is the export intensity of the industry (her data do not include export status at the plant level). Our data allow us to investigate whether, in a given industry, women are more likely to be employed by exporting plants.

In Table 2 we provide the results of a set of regressions where we regress the plant's female share of labor on a variety of plant-level characteristics including an indicator of whether the plant exports:

$$FS_{it} = \beta_1 \cdot E_{it} + \beta_2 \cdot \mathbf{X_{it}} + \delta_j + \delta_t + \epsilon_{it} \tag{11}$$

Each observation represents a Colombian plant i in year t where we use the full panel of plants from 1984 to 1991 (including both exporting and non-exporting plants).[21] The dependent variable in column 1 of Table 2 is FS_{it}, the female share of total employment in the plant.[22] The coefficient of interest is β_1 where E_{it} is the export status of a plant (E_{it} is an indicator variable which takes the value one if the plant exported that year). Our main hypothesis is that $\beta_1 > 0$ since exporting plants are likely to face greater competition in world markets (i.e., higher elasticity of demand) and so will hire proportionally more female employees. However, it is well established that exporters differ from non-exporters in multiple ways: for example, they tend to be larger, more productive, and more capital intensive. To address these issues, our estimation also includes $\mathbf{X_{it}}$, a vector of plant-level characteristics including the age of the plant, total employment in the plant (to proxy for size), and productivity/profitability of the plant (value added per employee).[23] We also control for the skilled-labor share of the plant (i.e., skilled workers out of total employment) which is a more aggregate measure of the occupational composition of the plant (see Appendix A). In addition, as additional explanatory variables, we include other measures of the production process including the energy share, capital share and office equipment share of the plant. Given that existing literature suggests that women are concentrated in low-paying jobs, we also include a measure of average wages of workers in the plant (the log of the total salary and wages of all workers in the plant divided by total employment in the plant). Finally, δ_j are 3-digit ISIC industry fixed effects to control for industry-level characteristics and δ_t are year fixed effects to control for aggregate time effects.[24]

Column 1 of Table 2 suggests that exporting plants do employ a higher share of female employees than non-exporting plants, controlling for other plant characteristics. The point estimate suggests that this difference is fairly large: controlling for the other variables included in the model, exporting plants have, on average, a female share 4.4% points higher than non-exporters (the mean female share for the plants in the sample is 36.3%). This is consistent with our hypothesis that plants facing more competition are less likely to discriminate: since exporting plants face more competition, we expect them to hire more women.[25]

The negative coefficient estimate on salary is consistent with our theoretical model, as well as standard models of discrimination in Becker (1957) and Arrow (1973), where the benefit of hiring female workers is precisely that a firm can pay them lower wages. These results are also consistent with work examining inter-firm segregation in the U.S. (e.g., Carrington & Troske, 1998; Hellerstein et al., 2002). The positive coefficient on plant size is also consistent with our model, since large plants produce more output and therefore suffer larger losses by hiring men.

Since women tend to be concentrated in certain occupations, in regressions 2 and 3 of Table 2 we present results where the dependent variable, FS_{it}, is the female share of unskilled (column 2) and skilled labor (column 3), respectively. In regressions 2 and 3, we also add the female share of managers and owners as additional explanatory variables, since women managers and owners may have less taste for discrimination of female workers; this seems to be the case, since in both regressions, the coefficient estimate on female management share is positive and statistically significant (although the effect seems to be stronger for unskilled workers than for skilled workers).[26]

The results for the unskilled female labor share (regression 2) and skilled female labor share (regression 3) are quite similar to the overall sample: the coefficient estimate on exports is positive and remains highly statistically significant, controlling for the other variables included. The main difference is that the coefficient estimate on the skilled labor ratio is smaller, suggesting most of the differences between exporting and non-exporting plants is due to exporting plants hiring a higher percentage of female workers in unskilled-labor intensive occupations.

While Table 2 employs a dummy variable for export status, we also have data on the volume a plant exports. When we repeat the analysis of Table 2 with the plant's export intensity (plant exports as a percentage of total sales) replacing the export indicator variable, the coefficient estimate on export intensity remains positive and highly statistically significant for the overall female share and the female share of unskilled labor. However, an intriguing difference is that the coefficient estimate on export volume for the skilled female labor share becomes negative (and is statistically significant).[27] This suggests that the larger exporting plants, although they hire proportionally more women overall, and proportionally more female unskilled workers, tend to have *lower* female shares of skilled workers. One possible reason for this effect is that large exporting firms tend to have multiple establishments with separate headquarters and production facilities, and clerical workers are located in the headquarters, rather than in the plants that appear in our data.[28]

5 | COMPETITION AND PLANT EXIT

As mentioned previously, both Becker (1957) and Arrow (1973) argue that competition should reduce discrimination by inducing the exit of discriminating firms (which would have higher costs and lower profits than non-discriminating firms). While this is primarily a long-run argument, our interest is whether we can observe this effect in the short run. That is, does an exogenous increase in the degree of foreign competition (as we observe with the Colombian trade liberalization episode) induce discriminating plants (i.e., plants with lower female shares) to exit the market? Indeed, the recent literature on firm heterogeneity and trade stresses the ability of trade to improve aggregate industry productivity through exactly such selection effects. Specifically, trade can induce productivity improvements by causing more productive firms to expand while less productive firms shrink or exit the market (e.g., see Melitz, 2003). The obvious parallel is the potential for trade competition to reduce discrimination by inducing discriminating (i.e., less efficient) firms, or plants, to shrink or exit the market.[29] For example, Yahmed (2023) provides a model in which trade impacts the gender wage gap partly through selection effects.

That competitive forces will drive discriminating employers from the market is one of the strongest predictions of the Becker model. The typical approach to testing this prediction is by comparing employment or wage discrimination in in highly concentrated markets to discrimination in markets with a less concentrated market structure (e.g., see Ashenfelter & Hannan, 1986; Black & Brainerd, 2004; Haessel & Palmer, 1973; Hellerstein et al., 2002; Jones & Walsh, 1991; Kawaguchi, 2007; Oster, 1975). However, this approach has been critiqued by Ederington and Sandford (2016) who formalize the Becker model in a dynamic context and show that there is no consistent theoretical relationship between industry market concentration and discrimination. We use a natural experiment, the Colombian trade liberalization episode of the 1980s, to observe the effects of changing levels of competition on plant-level hiring.[30]

In Table 3, we regress industry exit rates on the tariff change induced by the Colombian trade liberalization episode of 1984–1991, as well as past exit rates to control for industry-specific effects. As can be seen, industries that experienced the most significant tariff reductions (largest *negative* tariff change) also saw higher exit rates, so the increase in foreign competition does appear to have induced exit by Colombian plants.

The question of interest in this section is whether plants with low female shares are disproportionately represented among these exiting plants. To investigate this question, in regression 1 of Table 4, we first run a logit regression predicting the probability of plant exit (by 1991) for plants in the sample in 1984, including the plant's female share of its labor force as an explanatory variable. As can be seen, an establishment with a higher female share was, on average, *more* likely to exit the industry between 1984 and 1991.

However, recall from Table 2 that women also tend to be concentrated in less capital-intensive, low-wage plants, so it is possible that women are overrepresented in exiting plants simply due to

TABLE 3 Industry exit rates 1984–1991.

Variable	(1) $\hat{\beta}$ (SE)
Exit rate 1977–1981	0.076 (0.108)
Tariff change 1984–1991	−0.141 (0.057)**
constant	0.259 (0.038)***
Observations	92
R^2	0.074

Note: The dependent variable is the industry exit rate between 1984 and 1991. *** indicates statistical significance at the 99% level or better; ** at 95%; and * at 90%. Mean probability of exit: 0.340 (0.163). 2 industries (3842, 3902) have all plants present in 1984 exit by 1991.

TABLE 4 Logit: Probability of plant exit.

Variable	(1) $\hat{\beta}$ (SE)	(2) $\hat{\beta}$ (SE)
Female share	0.324 (0.143)**	−0.234 (0.476)
Tariff		−0.007 (0.003)**
Tariff × female share		0.006 (0.005)
Exports		0.707 (0.252)***
Years of existence		−0.002 (0.003)
Total employment (log)		−0.625 (0.058)***
Salary (log)		0.015 (0.107)
Productivity (log)		−0.474 (0.067)***
Skill ratio		−0.123 (0.219)
Energy use (log)		−0.049 (0.031)*
Capital-labor ratio (log)		0.026 (0.039)
Office equipment		0.611 (0.403)
Corporation		0.488 (0.129)***
Proprietorship		−0.028 (0.071)
Observations	6044	6032
Of which, exited	2281	2273
Log likelihood	−3921.2	−3604.0

Note: This is a cross-section of plants in the sample in 1984. The dependent variable is the probability that the plant exits between 1984 and 1991 (note: change in SIC not treated as exit). All explanatory variables are measured in 1984. See Appendix A for data definitions. Three-digit industry dummy variables are also included in both regressions, and standard errors in Column 2 are clustered by four-digit industry (tariffs are measured at the four-digit level). *** indicates statistical significance at the 99% level or better; ** at 95%; and * at 90%.

the fact that they are concentrated in the types of plants that would be most likely to exit. We are interested in whether the trade liberalization episode disproportionately drove out plants with lower female share *controlling for other plant characteristics*. In regression 2 of Table 4 we run our base specification:

$$\Pr(exit = 1) = F(\phi_1 \cdot FS_i + \phi_2 \cdot \Delta\tau_j + \phi_3 \cdot (FS_i \cdot \Delta\tau_j) + \phi\mathbf{X_i} + \delta_j) \tag{12}$$

where $F(\cdot)$ is the cumulative logistic distribution and $exit_i$ is an indicator variable taking the value of one if the plant exits the market at some point between 1984 and 1991 (note that a simple change in SIC code is not treated as exit, and that all plants are present in 1984). As before, FS_i is the female labor share of the plant (measured in 1984) and X_i is a vector of plant characteristics that control for the probability of plant exit (e.g., capital and skill-intensity).[31] Finally, $\Delta\tau_j$ is the change in ad-valorem tariffs (between 1984 and 1991) for industry j at the four-digit SIC level and δ_j are (more aggregated) 3-digit industry fixed effects.[32]

First, note from column 2 of Table 4 that $\phi_2 < 0$. Consistent with Table 3, plants in industries that faced larger decreases in tariffs were more likely to exit, controlling for other plant characteristics. In addition, other coefficient estimates are also in line with previous literature: exporters, larger plants, and more productive (profitable) plants are all less likely to exit. However, in this estimation, our primary interest is whether $\phi_1 < 0$ (plants with higher female have a higher probability of exit) or $\phi_3 > 0$ (the positive impact of trade liberalization on exit is lower for plants with higher female shares). As can be seen in Table 4, while the signs of the coefficients are consistent with these hypotheses, neither coefficient estimate is statistically significant. We fail to find evidence for the proposition that an exogenous increase in competition drives discriminating plants from the market, at least in the short run.

6 | COMPETITION AND PLANT HIRING DECISIONS

While the previous section fails to find evidence that increased competition due to trade liberalization forced discriminating plants from the market, a second possibility suggested by our model is that the increase in foreign competition influenced plant hiring decisions. In this section, we look for evidence that an exogenous increase in foreign competition (i.e., the Colombian trade liberalization episode) induced plants to increase the female share of their workforce.

Specifically, we investigate whether surviving plants in industries that experienced larger decreases in tariff protection responded by increasing the female share of their labor force (relative to industries where tariffs remained relatively constant). Our baseline (long-difference) regression is given by:

$$\Delta FS_i = \alpha_1 \cdot \Delta\tau_j + \alpha_2\mathbf{X_i} + \delta_j + \epsilon_i \tag{13}$$

where the dependent variable is the change in plant i's female labor share between 1984 and 1991. 1984 corresponds to the high point of average protection in Colombia, and 1991 is the final year of data available to us.[33] The main coefficient of interest is α_1 where $\Delta\tau_j$ is the change in ad-valorem tariffs (between 1984 and 1991) for industry j. Thus, we exploit cross-sectional variation in the degree of trade liberalization across industries to examine whether plants in industries that lost more tariff protection responded by increasing the female share of their labor force (i.e., $\alpha_1 < 0$).

TABLE 5 Effect of tariff change on plants' female share of workers.

	(1)	(2)	(3)	(4)
Tariff chg 1984–1991	−0.162 (0.065)**	−0.163 (0.055)***	−0.161 (0.054)***	−0.265 (0.069)***
Female share 1984	−0.327 (0.051)***	−0.319 (0.042)***	−0.318 (0.041)***	−0.322 (0.043)***
Plant age	−0.000 (0.000)**	−0.000 (0.000)	−0.000 (0.000)**	−0.000 (0.000)**
Corporation	−0.015 (0.006)**	−0.013 (0.006)**	−0.013 (0.006)**	−0.014 (0.006)**
Proprietorship	0.011 (0.012)	0.010 (0.011)	0.006 (0.008)	0.005 (0.008)
Chg log employment		0.049 (0.014)***	0.021 (0.010)**	0.019 (0.009)**
Chg exports			0.003 (0.028)	0.008 (0.027)
Chg log salary			−0.030 (0.009)***	−0.029 (0.008)***
Chg log prody			−0.006 (0.004)	−0.006 (0.004)
Chg skill ratio			−0.004 (0.028)	−0.003 (0.028)
Chg log energy			−0.007 (0.003)**	−0.007 (0.003)**
Chg log K/L			0.000 (0.002)	0.001 (0.002)
Chg office equip			0.002 (0.018)	0.006 (0.017)
Four-plant concentration				0.122 (0.044)***
Conc4×chg tariff				0.249 (0.082)***
N	3726	3726	3677	3677
R^2	0.193	0.241	0.232	0.238

Note: The dependent variable is the change in the plant's share of (total) female workers between 1984 and 1991. Three-digit industry fixed effects are also included. Standard errors in parentheses are clustered at the four-digit industry level. Tariffs and concentration ratios are measured at the four-digit industry level. *** indicates statistical significance at the 99% level or better, ** at 95%, and * at 90%.

The regressions also include a vector of plant-level ($\mathbf{X_i}$) characteristics that might influence hiring decisions, including the female share of the plant in 1984.[34] Finally δ_j are three-digit industry fixed effects to control for industry-level characteristics.[35]

Results of estimating Equation (13) are provided in column 1 of Table 5, where an increase in foreign competition (as captured by a decline in tariff protection) is associated with plants increasing their share of female employees.[36] Recall that over this period, industries experienced (on average) a significant decrease in tariffs. Indeed, from Table 1, the average industry experienced a tariff change of −0.403, which corresponds to an increase in the female labor share of their workforce of 0.065 relative to an industry which received no change in tariff protection.[37] As a point of comparison, the average female share in 1984 of plants in the dataset is 0.365, so an increase of 0.065 is quite large. The Colombian trade liberalization episode appears to have had a significant impact on plants' hiring decisions and women's employment.

However, as discussed in Section 2, import competition could affect firm hiring decisions through two channels: (1) by affecting the elasticity of demand and (2) by affecting firm size. In column 2 of Table 5 we undertake a mediation exercise where we include the change in plant size (measured by employment) to control for the scale effect discussed in Section 2. Including the change in plant size as an explanatory variable allows us to better isolate the "elasticity effect," since we cannot measure the change in elasticity directly. Specifically, in column 1 we estimate the total average casual effect (ACE) of the trade liberalization episode on plant-level female labor shares. In contrast, by controlling for the change in employment in column 2, we can estimate the controlled direct effect (CDE) of the trade liberalization episode on the plant's female share (see Cinelli et al., 2022).[38]

As can be seen in column 2 of Table 5, the coefficient estimate on the change in plant employment (the "scale effect" discussed in Section 2) has the predicted sign and is highly statistically significant, implying that an increase in employment is correlated with an increase in a plant's female share of workers. The intuition from our model is that in a larger firm, the cost of discriminating is higher, since the increased marginal costs of discriminating affect a higher volume of production. However, importantly, controlling for the change in plant size does not impact the coefficient on the change in tariffs (α_1), which is virtually unchanged from regression 1. This implies that the decrease in tariffs did not only affect plant hiring decisions through scale effects.

Of course, it is also possible that trade liberalization induces some change in the production process (e.g., a change in the capital-labor ratio) which, in turn, affects the female share of the labor force (e.g., similar to the mechanism of Juhn et al., 2014). In the third column of Table 5, we expand the mediation analysis to include changes in plant-level measures of the production process (e.g., capital intensity, energy intensity, skill levels). Intuitively, if trade liberalization primarily affected plant hiring decisions through changes in the production process, then controlling for changes in technology (e.g, changes in capital intensity or energy intensity) should attenuate the estimate of α_1. As can be seen, including additional variables to control for potential changes in the plant's production process has little impact on the main coefficient of interest.[39]

Finally, as discussed in the theoretical section, the effect of competition on hiring could also be a function of the existing degree of competition in the industry. To examine this hypothesis, in the fourth column, we also include an interaction term between the four-plant concentration of the industry and the tariff change. Presumably, plants in more concentrated industries face less competition and should be more affected by any given change in tariffs. The sign on the interaction term suggests that the marginal effect of a tariff change on the female share is, somewhat surprisingly, lower in absolute value for firms in more concentrated industries, controlling for other plant characteristics.

In Table 6, we present regression results examining the impact of tariff changes on plant-level female shares of skilled and unskilled workers separately. These results largely mirror the results seen in Table 5, in that trade liberalization leads plants to increase the female share of both skilled and unskilled workers. Finally, to further examine the hypothesis that the effect of tariffs could vary by the degree of competition in the industry prior to the tariff change, in Table 6 we estimate the regressions separately by the export status of plants in 1984. Since exporting plants already compete in international markets, it seems plausible that a decline in domestic tariffs would have a more significant impact on the competition facing non-exporting plants. Indeed, as can be seen in Table 6, our results are driven primarily by non-exporting plants.

6.1 | Plant fixed effects

In the previous section, we employed a long-difference regression approach and found that plants in industries that faced an increase in foreign competition (i.e., greater decline in tariffs) responded by increasing the female share of their workforce. A drawback to that cross-sectional approach is that we cannot use plant fixed effects to control for unobserved plant characteristics. As a robustness check, in this section we exploit year-to-year changes in tariffs and plant-level

TABLE 6 Effect of tariff change on plants' female share.

Variable	(1) All workers $\hat{\beta}$ (SE)	(2) Unskilled $\hat{\beta}$ (SE)	(3) Skilled $\hat{\beta}$ (SE)
All plants			
Tariff chg 1984–1991	−0.162 (0.065)**	−0.197 (0.083)**	−0.210 (0.047)***
Female share 1984	−0.327 (0.051)***	−0.281 (0.051)***	−0.576 (0.030)***
Plant age	−0.000 (0.000)**	−0.000 (0.000)**	−0.000 (0.000)
Corporation	−0.015 (0.006)**	−0.017 (0.007)***	−0.061 (0.009)***
Proprietorship	0.011 (0.012)	−0.000 (0.013)	0.062 (0.020)***
Observations	3726	3632	3390
R^2	0.193	0.169	0.340
Non-exporting plants only			
Tariff chg 1984–1991	−0.166 (0.056)***	−0.189 (0.068)***	−0.230(0.050)***
Female share 1984	−0.337 (0.043)***	−0.284 (0.038)***	−0.585 (0.031)***
Plant age	−0.000 (0.000)	−0.000 (0.000)	−0.000 (0.000)
Corporation	−0.013 (0.007)*	−0.018 (0.010)*	−0.058 (0.012)***
Proprietorship	0.008 (0.013)	−0.004 (0.014)	0.057 (0.020)***
Observations	3209	3116	2880
R^2	0.196	0.168	0.347
Exporting plants only			
Tariff chg 1984–1991	−0.136 (0.109)	−0.229 (0.145)	−0.054 (0.062)
Female share 1984	−0.274 (0.086)***	−0.249 (0.093)***	−0.527 (0.059)***
Plant age	−0.001 (0.000)**	−0.001 (0.001)**	−0.001 (0.001)
Corporation	−0.011 (0.018)	−0.006 (0.026)	−0.031 (0.018)*
Proprietorship	0.104 (0.054)*	0.071 (0.062)	0.125 (0.093)
Observations	517	516	510
R^2	0.324	0.304	0.337

Note: The dependent variable is the change in the plant's share of female workers 1984–1991. The dependent variable and the 1984 female share explanatory variable are the overall female share in Column 1, the unskilled female share in Column 2, and the skilled labor share in Column 3. Three-digit industry fixed effects are also included. Tariff change is at the four-digit industry level. Standard errors are clustered at the four-digit industry level. *** indicates statistical significance at the 99% level or better; ** at 95%; and * at 90%.

hiring decisions to undertake a panel regression approach that allows us to include plant fixed effects. Specifically, the regression is given by:

$$FS_{it} = \alpha_1 \cdot (\tau_{jt} \cdot qFS_{1984}) + \delta_i + \epsilon_{it} \tag{14}$$

where the dependent variable, FS_{it}, is again the female share of total employment in plant i at time t and τ_{jt} is the ad-valorem tariff of industry j in time t. The panel nature of the regression allows us

TABLE 7 Plant-level fixed effects: Female share.

	(1)	(2)	(3)	(4)
Q1×tariff	−0.148 (0.010)***	−0.154 (0.010)***	−0.177 (0.011)***	−0.273 (0.014)***
Q2×tariff	−0.106 (0.009)***	−0.098 (0.009)***	−0.120 (0.009)***	−0.176 (0.013) ***
Q3×tariff	−0.029 (0.008)***	−0.022 (0.008)***	−0.047 (0.008)***	−0.102 (0.012)***
Q4×tariff	0.070 (0.005)***	0.059 (0.005)***	0.032 (0.005)***	0.002 (0.007)
Log employment		0.057 (0.002)***	0.032 (0.002)***	0.032 (0.002)***
Log salary			−0.003 (0.002)	−0.002 (0.002)
Log prody			−0.002 (0.002)	−0.002 (0.002)
Skill ratio			−0.000 (0.001)	−0.000 (0.001)
Log energy			−0.001 (0.001)	−0.001 (0.001)
Log K/L			−0.000 (0.001)	−0.000 (0.001)
Office equip			−0.008 (0.009)	−0.007 (0.009)
conc4×tariff				−0.144 (0.023)***
Observations	25,973	25,973	25,459	25,459
Plants	5937	5937	5895	5724

Note: Data are panel data 1984–1991 (excluding 1986 and 1987 due to missing tariff data). The dependent variable is the plant's share of female workers. Quartiles are based on 1984 female shares: that is, the coefficient estimate for Q1×tariff is the coefficient estimate on tariff for the lowest quartile of female share. *** indicates statistical significance at the 99% level or better; ** at 95%; and * at 90%.

to include plant-level fixed effects (δ_i). As before, the coefficient of interest is α_1 where we expect that $\alpha_1 < 0$ (a larger decrease in tariff is associated with an increase in the plant's female share). Note that we can no longer include the 1984 plant female share to control for censoring (female share is bounded by zero and one), so we interact the tariff variable with indicator variables based on the quartiles of the cross-plant female share in 1984 (i.e., $Q1$ indicates that a plant fell in the lowest quartile (below 11%)). We drop the firm structure variables since they do not vary over time (and we drop plant age). Finally, since industry concentration is largely fixed, we interact the four-plant concentration ratio in 1984 with the year's tariff to capture how plants in concentrated industries change their female share in response to changing tariffs.

The estimates we see in Table 7 from estimating (14) largely mirror our results from Table 5. For example, the estimates in column 1 show that decreases in tariffs over time are associated with the largest increase in female share among plants with the lowest female share in 1984, which in our model are the most discriminating plants, and that the impact monotonically declines (in absolute value) as the initial female share increases. The estimates in column 2 also show that plants that increase in size increase the share of women in the plant, but controlling for changes in scale has very limited impact on the estimated effects of changes in tariffs on the female share. Finally, these results are robust to the inclusion of controls for the plant's production process and the concentration of the four-digit industry in 1984.

7 | CONCLUSION

Using plant-level data, we show that Colombia's 1984–1991 trade liberalization episode resulted in proportionally more women being hired by Colombian plants. This evidence is consistent with

Becker's theory of discrimination, as well as with our slightly modified version of Becker's theory. We find that this change occurs primarily because increasing competition leads existing plants to hire more women, not because they exit the market. The effect of tariff liberalization on the female share of workers is quantitatively large: the average decrease in tariffs over this period corresponds to a 6.5% point increase in a plant's female share of employment, relative to a plant in an industry with no change in tariffs. Consistent with the predictions of our model, we find that large plants employ a higher share of women. We are also able to elaborate on earlier research that found that women are concentrated in exporting industries, by showing that women are also concentrated in exporting plants. Finally, we show that the effect of tariff liberalization is larger for non-exporting plants than for plants that were exporting prior to liberalization—a finding consistent with our model's prediction that plants that initially faced little competition would be the most affected by increasing competition.

Importantly, our findings provide some evidence on how changes in competition can lead to less labor market discrimination and help resolve a puzzle in the existing literature. In Becker's original model, competition leads to less discrimination by driving discriminating firms from the market. However, despite evidence that discrimination leads to lower wages and higher levels of employment for women and that discriminating firms earn higher profits (see Black & Brainerd, 2004; Black & Strahan, 2001; Hellerstein et al., 1999; Hellerstein et al., 2002), there is little evidence that competition drives discriminating firms from the market (see Hellerstein et al., 2002). In our modified version of the Becker model, we show that increases in competition can also reduce discrimination by raising the cost of discrimination, pushing discriminating employers to hire more women. Thus, our results suggest the potential importance of enhancing competition as a way of reducing the extent of discrimination in the labor market.

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CONFLICT OF INTEREST STATEMENT

The authors have no declared conflict of interest.

DATA AVAILABILITY STATEMENT

The data that support the findings of this study are available from the corresponding author upon reasonable request. We can check with Mark Roberts about placing the data in a public repository.

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ENDNOTES

1 Although Ozler (2000) has plant-level data from Turkey, the data do not include the export status of the plant. She shows that the female share of labor tends to be higher in plants that are in industries that export more of their output. Our data allow us to directly link the female share of labor in a plant with the export status of that plant.

2 An exception of a paper that uses a more firm-level approach is Bøler et al. (2018) which utilizes matched employer-employee data from the Norwegian manufacturing sector to exhibit a higher gender wage gap among exporting firms. They argue this is due to exporting firms requiring more employee flexibility.

3 Nor are they the only possible mechanisms; for example, Keller and Utar (2022) show, using individual-level data from Denmark, that women were more likely than men to leave the formal labor market for family reasons in response to increased imports from China, and Bøler et al. (2018) argue that exporting firms might be biased against female workers since they require more commitment and flexibility from employees in order to work with firms in other time zones.

4 Juhn et al. (2013, 2014) do control for changes in Mexican tariffs but do not get statistically significant results.

5 This wage differential is simply taken as exogenous in the differentiated product sector. It can be generated either by discrimination or by productivity differences in the numeraire product sector. For example, assuming each male employee can produce w_m units of the numeraire good and each female employee can produce w_f units of the numeraire good, production of the numeraire good in positive amounts would fix wages in the economy at w_m and w_f respectively.

6 There also exists a second-order condition on $\phi_i''(z_i)$, which we assume is satisfied, that ensures an interior solution.

7 In addition, σ corresponds to standard measures of competition used in empirical research (e.g., it is monotonically increasing transformation of the elasticity of demand faced by the firm and it is a monotonically decreasing transformation of firm profitability). Note that as $\sigma \to \infty$, the CES preferences take the linear form: the goods become perfect substitutes and the market mimics perfect competition.

8 Note that these comparative statics hold the gender wage gap constant. In a world where everyone is working and equally productive, a universal increase in product market competition should lead to a rise in women's wages and a fall in the male-female wage gap. However, in a world where there is heterogeneity in productivity and where there are differences in the selection of men and women into the labor market and into the traded sectors, it is not clear how changes in the demand for labor and changes in discrimination will impact the average wages of women or the observed gender wage gap. Thus, the focus of theory and empirics in this paper is on the reallocation of workers across firms/industries as opposed to aggregate changes in the labor market.

9 Krugman (1979) models a monopolistically competitive industry as we do in this paper. Similarly, Markusen (1981) shows that trade can increase the "perceived" elasticity of demand in oligopolistic industries.

10 Both the scale and elasticity effects are decreasing as one increases the number of firms, n, in the market.

11 A partial list of papers that have studied Colombia's trade liberalization includes Roberts and Tybout (1997), Fernandes (2007), Goldberg and Pavcnik (2005), Brooks (2006) and Goldberg and Pavcnik (2007).

12 Unfortunately, while non-tariff barriers were also reduced as part of the reform, complete data on the reductions in such barriers is not available. However, as mentioned, tariffs were the primary policy instrument and the available data suggest that tariff levels (and changes) are positively correlated with NTB levels (and changes) across industries in our dataset. See Goldberg and Pavcnik (2005) for more details.

13 Colombia entered the GATT in 1981, but used developing-country exemptions to avoid tariff reductions until the trade liberalization episode studied in this paper. For an overview of the evolution of trade policy in Colombia see Goldberg and Pavcnik (2005) and Fernandes (2007).

14 We would like to thank Mark Roberts for making the Colombian manufacturing census data available to us. For a complete description of all the variables used in our estimation, see Appendix A, and Roberts (1996) for a description of this dataset.

15 In some of the analysis that follows, sample sizes are reduced due to missing data. In addition, we drop approximately 2% of the sample classified as "other" types of enterprise (cooperatives, collectives, etc.) in the analyses that follow.

16 See Fernandes (2007).

17 We would like to thank Jorge Garcia-Gracia at the World Bank for making these data available.

18 See Eslava et al. (2022); fn. 6.

19 All wages are measured in nominal dollars, so it is not surprising that wages rose over this period.

20 "Women in Colombia Move to Job Forefront," *The New York Times*, July 15, 1994.

21 We do not include plant-level fixed effects, since our interest is in the within-industry variation in female share as a function of exports.

22 All signs and significance levels are identical when we use the fractional logit model of Papke and Wooldridge (1996).

23 Controlling for plant productivity also helps control for the the possibility that more profitable (i.e., exporting) plants might feel more freedom to discriminate since they are less threatened by the prospect of market exit.

24 Results are robust to the use of less aggregate two-digit SIC industry fixed effects as well as two-digit industry-year fixed effects to control for differential trends across industries.

25 Another potential explanation is that plants that hire more women are more efficient (in our model, non-discriminating firms can produce the same output at lower cost), and more efficient plants are more likely to select into exporting. See, for example, Bernard and Jensen (1995), Roberts and Tybout (1997), and Bernard et al. (2003), who provide empirical evidence (from both developing and developed countries) that exporting is an activity primarily undertaken by successful firms. However, it is intriguing that the characteristics of plants that hire more women (in terms of wage, capital intensity and energy intensity) differ substantially from the standard characteristics of export-oriented firms (e.g., see Roberts & Tybout, 1997).

26 Note that regression 1 includes these owners and managers, a classification in addition to skilled and unskilled workers, in the dependent variable so regression 1 is not simply an average of regressions 2 and 3.

27 The coefficient estimates on export intensity are 0.090 (standard error 0.014) for the overall female share, 0.121 (standard error 0.018) for the unskilled female share, and −0.033 (standard error 0.016) for the skilled female share. The other coefficient estimates are very similar to those in Table 2; results are available on request from the authors.

28 Proprietorships may be less likely than partnerships or corporations to have multiple establishments. When we restrict the sample to only proprietorships (approximately 10% of our sample), the coefficient estimate on exports in this regression becomes positive, although it is not statistically significant.

29 It should be noted that these models of trade competition and productivity typically work through the export side of the market, so the parallel to the Colombian trade liberalization episode is not direct, as unilateral trade liberalization in Melitz (2003) has no impact on the productivity distribution of firms.

30 Other papers using natural experiments include Black and Strahan (2001) and Heyman et al. (2008), although neither is in the context of international competition. Black and Strahan (2001) examine deregulation in the banking industry and find that women's relative wages increased after deregulation, as did their share of managerial positions. Heyman et al. (2008) use firm data on takeovers in Sweden to examine the change in firm-level female employment, and find that, particularly in less competitive industries, firm takeovers result in a significant increase in female shares of employment.

31 Results are very similar for both subsamples when we estimate regression 2 on exporting and non-exporting plants separately.

32 Results are robust to two-digit industry fixed effects.

33 For this analysis, our sample consists of plants that are in operation both in 1984 and 1991. One possibility is that the change in female share is related to plant exit which, if true, would lead to a bias in our estimated coefficients. To investigate this possibility, we estimated a standard two-stage selection model where the first stage regression estimates the probability of plant exit. Since the results from the two-stage model are largely identical to the results reported in Table 5, and since a standard Hausman test fails to reject the hypothesis of no selection, we have chosen to report the results from the single equation model. Results from the two-stage model are available from the authors on request.

34 Note that the inclusion of the plant's 1984 female share also partially corrects for censoring, in that a plant with an already high female share (i.e., close to 1) cannot increase its female share further.

35 We can only include three-digit industry fixed effects since our tariff-change variable is at the four-digit level. Results were similar when we employed two-digit SIC industry fixed effects, although the magnitude of the estimates on tariff change was smaller.

36 Some related papers instrument for the change in tariffs; for example, Goldberg and Pavcnik (2005) use initial tariffs and several time-varying measures to instrument for tariff changes. This is more common in studies using annual panel data due to concerns about the endogeneity of year-to-year tariff changes. Non-panel approaches like ours, such as Dix-Carneiro and Kovak (2017) on the effects of trade liberalization on Brazilian labor markets, are more likely to use tariff changes. In our data, the correlation between initial tariff and the change in tariffs is −0.97, and our results throughout are robust to using the initial tariff.

37 Of 92 four-digit industries in the sample, five actually received increases in tariff protection over this period.

38 As noted in Fagereng et al. (2021) and Heckman and Pinto (2015), our estimation of the CDE relies on the assumption that any unobserved mediator variables are uncorrelated with the observed mediator variables and FS_i. To assess the validity of this assumption, we repeated the analysis with the control variables entered in initial levels, which yielded similar results.

39 Part of the reason that controlling for changes in the production process has little impact on our estimate of α_1 is that we find little measurable impact of trade liberalization on a plant's production process. Specifically, if we repeat the analysis and estimate Equation (13) with the 1984–1991 change in other measures of the plant production process (e.g., capital or energy intensity) as the dependent variable, we find very little evidence of any effect (the coefficient estimates are not statistically significant).

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APPENDIX A. DATA

All plant-level data are taken from a plant-level dataset produced from the Colombian Manufacturing census by DANE (National Statistical Institute) for the years 1977 through 1991. From 1983 on, the census covers industrial production for plants with greater than 10 employees. Our empirics concentrate on plants that were operating in both 1984 and 1991. For a thorough description of this dataset see Roberts (1996). All variables are measured at the plant level unless otherwise noted.

Female share: female share of workers (total, unskilled, and skilled as described in text).
Skilled Workers: number of individuals classified as professionals, local or foreign technicians, or specialists (e.g., mechanical, chemical, industrial, electrical, mining, and petroleum engineers).
Unskilled Workers: number of workers classified as unskilled workers or apprentices (performing activities such as manufacturing, processing, assembly, installation, maintenance, inspection, storage, packing, loading, and unloading).
Productivity: value added for the plant divided by total employment.
Employment: number of individuals working in the plant in the reference year.
Plant Age: years since the plant's establishment until 1984.
Exports: plant exports scaled by total sales.
Salary: total payroll divided by total employment.
Skill Ratio: share of skilled employment in skilled and unskilled employment.

Capital/Labor Ratio: ratio of fixed capital to total employment. A small number of plants with fixed capital reported as zero are dropped. (In Table 4, e.g., this reduces the sample size by 23.)

Energy Use: one plus the ratio of energy consumed to total employment.

Office Equipment: office equipment's share of total capital equipment.

Female Management: percentage of management and owners that is female.

Type of Enterprise: The data set classifies plants by 10 different enterprise types. We omit plants classified as collectives, cooperatives, official entities, and religious communities (overall, these comprise less than 2% of the sample). We construct dummy variables for *Corporations* (this includes plants classified as corporations, de facto corporations, and joint stock companies), *Proprietorships*, and *Partnerships* (including limited partnerships and joint partnerships).

Industry Tariff: ad-valorem tariff at the four-digit ISIC level. Provided by Jorge Garcia at the World Bank. The tariff change for a four-digit industry is simply the difference between 1991 and 1984 ad-valorem tariffs.

Industry Concentration: Four-plant concentration ratio for the four-digit industry (authors' construction).

Female Share quartiles Q1–Q4: Dummy variables based on the 1984 cross-plant distribution of female share (0–0.11, 0.11–0.28, 0.28–0.60, 0.60–1).