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Summary: This paper examines whether there exists a glass ceiling to women's ascension in the largest firms of the Brazilian manufacturing industry. In particular, we check whether gender matters in the time it takes to get a promotion to a managerial position. Although we employ parametric accelerated failure time models in our preliminary analysis, our main results rely on partial rank estimates of a semiparametric duration model that allows for covariate-dependent censoring. We find that foreign-owned firms feature less gender differences in promotions than domestic firms. The same applies in other dimensions of career progress, namely, wage growth and promotion likelihood. It turns out that wage gains after promotion contribute to generating wage differential between males and females only within domestic firms. This is consistent with statistical discrimination and with the self-selection that results from employees optimally choosing which jobs to apply for. Jobs in domestic firms offer more flexibility in terms of hours per week, whereas multinationals compete for the most career-concerned workers.

JEL classification numbers: C24, C41, J16, J71, M51
Keywords: career mobility, duration, foreign capital, gender differences, promotion, multinational

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1 Introduction

There is a widespread consensus that occupational segmentation is the chief determinant for gender differences in wages in the Brazilian labor market (see, e.g., Oliveira, 2001). Although women correspond to 39% of the legal workers in Brazil, they amount to only 23% of the employees in the sectors paying relative higher wages in 2004. As well, segmentation increases if one restricts attention to managerial positions in that women hold only 14% of the latter. This suggests that there may exist a glass-ceiling effect in the Brazilian job market to the extent that women may experience implicit barriers to ascension in their jobs.

This paper contributes to the literature by examining data from the largest firms of the Brazilian manufacturing industry. In particular, our goal is to verify whether there indeed exists a glass ceiling to women's ascension by checking whether gender matters in how much time it takes to get a promotion to a managerial position. The aim is to complement the literature on gender differences in promotions, whose papers mainly focus on computing the promotion rates. Time to promotion offers a different angle in that samples with a large time span could well present similar likelihoods of promotion for both genders even if the promotion durations are very different.

As far as we know, this is the first study to examine gender differences in promotions using micro-data from a developing country. Apart from the obvious interest in verifying whether the main stylized facts hold in a Latin American country, the motivation also lies on the quality and availability of the Brazilian data. In contrast to the many studies that employ data from individual firms (see, e.g., Cabral, Ferber and Green, 1981; Gerhart and Milkovich, 1989; Baker, Gibbs and Holmstrom, 1994), we rely on a homogeneous sample of recently hired workers from the Brazilian manufacturing industry in the period running from 1996 to 2005. Our data set is particularly convenient. First, the data include a wide array of controls for worker and firm characteristics. Second, as opposed to Blau and deVaro (2007), we observe multiple workers per establishment at different occupations and hierarchical levels as well as their career paths in terms of occupation and wage. A potential limitation is that our data set does not include any direct measure of on-the-job productivity and hence we must come up with indirect controls.

Interestingly, we find that there are significantly less gender differences in the time to promotion within foreign-owned firms than within domestic firms in the Brazilian manufacturing industry. Our findings complements to some extent the literature on the differences between multinationals and domestic firms (Doms and Jensen, 2006; Greene, Hornstein, White and Yeung, 2009) as well as the evidence that gender differences may depend on the nature of the firm.[1] See deVaro and Brookshire (2007) for an interesting comparison between nonprofit and for-profit organizations, for instance.

We argue that such gender differences are consistent with statistical discrimination and self-selection. We reason as follows. Suppose that there are relatively more women than men who prefer to dedicate more time to their family than to their careers. Berk (2001) shows that gender differences in average career concern do not ensure statistical discrimination against the group with lower average career concern if employees optimally choose which jobs to apply for. The idea is that a career-concerned woman who applies for a job in a firm with a preference for dedicated employees only does so because she rationally believes that her odds of being hired compensate the effort to go through the application process. This means that her qualifications for the job have to be sufficiently good to stand a chance despite the discrimination. To complete the argument, we rely on two anecdotal evidence concerning the impact of multinationals in the Brazilian labor market (OECD, 2008). The first is that domestic firms in Brazil offer a more flexible package in terms of working hours and business trips than multinationals. This makes them more appealing to less career-concerned individuals (regardless of their gender). The second is that multinationals compete more fiercely for highly skilled workers. Under these circumstances, career-concerned women prefer jobs in multinationals and so statistical discrimination will become more prominent within domestic firms.

We find some indirect evidence supporting the above explanation. On the one hand, male workers tend to officially work similar hours in foreign-owned and domestic firms regardless of whether they have been promoted or not. Figure 1 reveals that the main difference is due to the concentration of the distribution: at least 75% of the male employees of domestic firms work exactly 44 hours per week, whereas there is a bit more of variation in multinationals, with 75% of the male employees working from 40 to 44 hours per week. It also documents some minor differences in the lower support of the distribution. In particular, the minimum number of hours worked by promoted male workers are slightly higher than by nonpromoted male workers. The same applies to female workers. This is consistent with the the fact that promoted workers are more likely to be of the career-concerned type. On the other hand, non-promoted female employees work relatively much less only in domestic firms (10 to 44 hours). This is in line with domestic firms offering more flexible packages. In addition, women in multinationals work much more than females in domestic firms

[1] In this paper, we employ multinational and foreign-owned firm interchangeably even if, in recent years, we have been witnessing the rise of many Brazilian multinationals such as CVRD (mining and metals), Petrobras (oil and gas), Gerdau (steel), and EMBRAER (aviation). See Amann (2009) for a historical perspective and a number of case studies.

and than male workers in general. This is well in line with self-selection. As the latter alleviates the impact of statistical discrimination, we fail to observe as much gender difference in foreign-owned companies as in domestic firms.

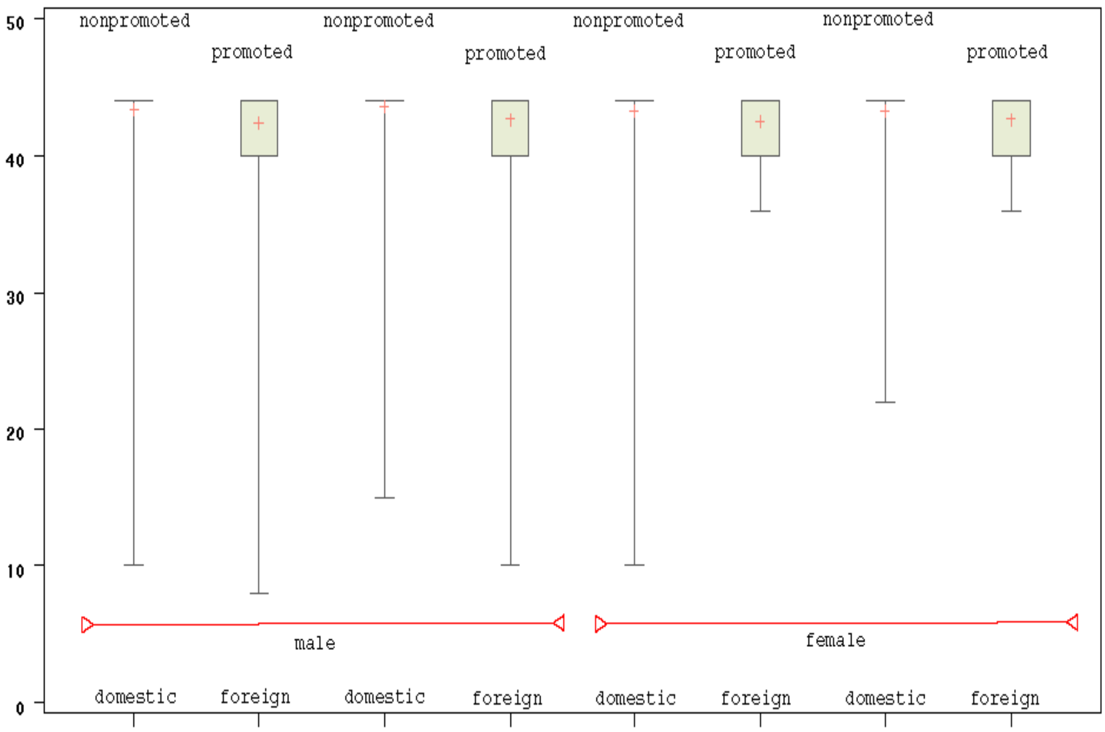

Figure 1: Box plot for working hours per week

This paper mainly relates to the literature on gender differences in career mobility. The theoretical literature on career mobility mainly focus on schooling (Sicherman and Galor, 1990), abstracting away from gender differences. There are a few exceptions, though. Booth, Francesconi and Frank (2003) derive a model that distinguishes between the initial pay increase upon promotion and subsequent pay increases. Under the assumption that women have worse market alternatives, the model implications are consistent with their empirical findings that gender does not affect promotion rates in UK, though women receive lower wage gains. Baldwin, Butler and Johnson (2001) identify the effects of occupational segregation on gender wage gaps using a hierarchical discrimination model in which men dislike supervision by female managers. They predict an exponential decline in the relative proportion of female workers in the top tiers of the job ladder, which is in line with the evidence from a 1988 CPS sample of workers in the insurance industry.

In contrast, it is common practice in the empirical literature to also include gender among the determinants of job mobility and promotion likelihoods (see, among others, Groot and van den Brink, 1996; Booth and Francesconi, 2000; Blau and deVaro, 2007) in view that men and women may differ in alternative opportunities (Mincer and Ofek, 1982; Lazear and Rosen, 1990; Royalty, 1998) as well as in job search costs (Meitzen, 1986). For instance, women may face more constraints either to work longer hours or even to remain in the labor market. If women are more likely to quit, firms will have less incentives to train and promote them. Moreover, if women view promotion as unlikely due to discriminatory promotion practices, they may refrain to put themselves forward for training programmes at the firm (Arrow, 1972). See, among others, Cabral et al. (1981), Spurr (1990), Cannings and Montmarquette (1991), McCue (1996), Barnett, Baron and Stuart (2000), Ransom and Oaxaca (2005), Blau and deVaro (2007), and Acosta (2010) for supporting evidence as well as Lewis (1986), Powell and Butterfield (1994), Paulin and Mellor (1996), Petersen and Saporta (2004), and Giuliano, Levine and Leonard (2011) for evidence against gender differences.

The papers closest to ours are McCue (1996) and Pekkarinen and Vartiainen (2006). The former documents that it takes on average more time for women and black men to get a promotion than for white males using data from PSID (1976-1988). Pekkarinen and Vartiainen (2006) analyze gender differences in time to promotion for workers within the metallurgical industry in Finland. They evince that women usually take more time to get a promotion than men with similar jobs, even if women are consistently more productive than men. Our findings are interestingly different in that the Brazilian manufacturing industry appears to feature more gender differential for domestic firms relative to foreign-owned.

The remainder of this paper ensues as follows. Section 2 describes the main features of our database, whereas Section 3 discusses the econometric methodology that we employ to assess whether there is gender differences in promotions. Section 4 then reports the findings of our empirical analysis. Section 5 offers some concluding remarks.

2 Data description

The data set we employ gathers information from several databases. In particular, it combines data from the Annual Report of Social Information (RAIS), covering the period running from 1991 to 2005, as well as data from the 1996 Foreign Trade Census of the Foreign Trade Secretary (SECEX) and from the 2000 Census of Foreign Capital (CEB).

RAIS is the administrative registry of the Ministry of Labor that provides socioeconomic information regarding the employees of every firm in the Brazilian formal sector. It reports the employees' identifying security number, age, gender, schooling level, job tenure, average monthly salary (including performance bonus and commissions), occupation (as reported by the employer), number of hours at work, type of labor contract, and month of admission. In addition, it also documents the firm's identifying fiscal number, sector of activity, and location. We make use of the SECEX data to gather information on how much each firm exports as a proxy for productivity.

The Central Bank of Brazil publishes the CEB every five years, collecting information on the origin of the shareholders' capital for every firm in Brazil. We employ the CEB to classify firms either as domestic or multinational. We define as multinational a firm in which more than 50% of the shareholders' capital is foreign.[2] Matching data from RAIS and CEB reveals that women amount to 21% of the employees in multinationals of the Brazilian manufacturing industry, occupying 13% of their managerial positions. The figures are similar for domestic firms (i.e., with less than 50% of foreign capital): 25% and 15%, respectively.

To form a homogeneous RAIS sample, we focus on individuals meeting the following criteria. First, the individual must work in a profit-seeking private firm with 500 or more employees in the Brazilian manufacturing industry. We focus exclusively on large firms because smaller firms have not enough internal turnover at the managerial level. Although these large firms respond only for 5.29% of the firms in the Brazilian manufacturing industry, they employ 32.9% of the workers in the sector, of which a quarter are female. Second, the individual must hold a university degree. The proportion of workers with a university degree has increased from 9.58% in 1996 to 13.93& in 2005 within our sample of large Brazilian manufacturing firms. Third, the individual must have joined the firm between January 1991 and December 1995. Fourth, the individual must work as an accountant, administrator, director, economist, engineer, intermediate manager, lawyer, manager, or purchases/sales supervisor. Fifth, the individual must also have a labor contract with no expiration date.

The resulting sample includes 1,422 firms, of which 297 (20.9%) are multinationals, that altogether employ 23,737 male and 3,552 female workers. The average individual in our sample is about 34 years old, working around 43 hours per week. As for occupations, engineers are the mode with 32.5% of the observations due to the predominance of male workers in the sample—the proportion of engineers among female workers is only 19.4%. There is a fraction of 29.0% of men occupying managerial positions, whereas this figure drops to 18.7% for women. Table 1 stratifies the sample according to occupation and gender, whereas Table 2 reports the sample averages of the individuals' main characteristics according to censoring and gender.

[2] The threshold at 50% is arbitrary, but also very likely to be harmless. The fraction of foreign capital concentrates either at zero (relative frequency of 65.83%) or at one (18.35%) and hence variations in the threshold are extremely unlikely to change any of the findings we report in the subsequent sections.

3 Duration model for time to promotion

In this section, we describe the duration model that we estimate to address gender differences in time to promotion. Although we also consider a more general semiparametric duration model later, we start with a simple linear regression specification for the log of the time to promotion:

$$\ln T_i^* = \boldsymbol{X}_i \boldsymbol{\beta} + \epsilon_i, \tag{1}$$

where T_i^* gauges how much time it takes to the individual i to obtain a promotion, $\boldsymbol{X}_i$ is a vector of control variates, and ϵ_i is an error term with scale and shape parameters σ and ς, respectively. In the context of duration models, (1) corresponds to an accelerated failure time (AFT) specification.

We construct the duration variable by following from January 1996 to December 2005 every individual that enters the RAIS database between January 1991 and December 1995. As some individuals do not obtain a promotion to managerial positions within the sample, we do not observe the time they take to get a promotion. In contrast, we classify as zero durations individuals whose first position within the firm is at the managerial level (i.e., either as a manager or as a director). We initially exclude these individuals from the AFT regression in view that (1) involves a log transformation of the duration. Altogether, this means that our sample criteria ensure that promotion durations may exhibit only right censoring.

Under right censoring, instead of observing the time to promotion T_i^* for each individual in the sample, we have information only on the following promotion duration variable:

$$T_i = \begin{cases} T_i^* & \text{in the absence of censorship} \\ R_i & \text{under right censoring} \end{cases} \tag{2}$$

where R_i corresponds to how much time the individual i has on the job up to December 2005. If the individual i exits the firm before December 2005, then the right-censoring variable R_i denotes tenure on the job, without receiving a promotion, up to the exit date.

The control variables at the individual level come from the RAIS database and refer to the month at which the individual starts at the firm. Using first-month data for individual-specific controls avoids further endogeneity issues, but it has the disadvantage of ruling out hours worked as a control given that these data are available only from 1995. In contrast, information at the firm level stems from the RAIS database of January 1996, from the SECEX data of the year 1996, and from the 2000 CEB.

We construct the binary variable MALE to control for the individual's gender. We additionally include the dummy variables MULTINATIONAL and MODERN that respectively take value one for firms with more than 50% of foreign capital and for firms from a technology-intensive sector.[3] Also, we consider the interaction dummy MALE×MULTINATIONAL as well as several continuous variables to control for firm-specific factors such as productivity, exposure to international markets, and average quality of jobs. More specifically, SIZE and EXPORTS correspond to the natural logarithm of the number of employees and of the total exports (in USD billions) of the firm that the individual works for, respectively. UNIVERSITY RATIO is the proportion of employees in the firm with a university degree, whereas MEAN WAGE is the average monthly stipend within the firm. We also consider a measure of TURNOVER that gauges the job flow intensity of the firm by means of the ratio of job flow (admissions plus dismissals) to the number of employees in the firm.

Let F denote the cumulative distribution function of the error term in (1), with density function f and survival function $S = 1 - F$, and let $\boldsymbol{\theta} = (\boldsymbol{\beta}, \sigma, \varsigma)$ denote the parameter vector. The log-likelihood function then reads

$$\mathcal{L}_N(\boldsymbol{\theta}) = \sum_{i=1}^{N} (1 - I_i^{(RC)}) \ln \Big[f(\epsilon_i)/\sigma \Big] + \sum_{i=1}^{N} I_i^{(RC)} \ln S(\epsilon_i) \tag{3}$$

where $I_i^{(RC)}$ is the indicator function that takes value one if there is right censoring, zero otherwise. To account for heterogeneity, we specify duration models with frailty, treating the unobserved individual effects as random draws from a gamma distribution with variance λ.

We initially suppose that censoring is independent of the regressors. This is a very strong assumption in that it rules out the situation in which women are more likely to quit their jobs than men, as in Lazear and Rosen's (1990) model.[4] Table 2 documents that 6.2% of the durations relating to male workers exhibit no censoring, whereas 65.1% display right censoring and 28.7% zero durations. These figures are respectively 4.5%, 76.8%, and 18.7% for female workers.

These differences suggest to some extent that censoring depends on gender, violating the independence assumption. We thus control for covariate-dependent censoring by employing Khan and Tamer's (2007) partial rank estimator. The latter entails distribution-free estimates of the regression coefficients without imposing any parametric specification on the link function (see next subsection for more details). In the empirical analysis, we show that accommodating for this more general form of censoring is paramount to examine promotion durations within the Brazilian manufacturing industry.

[3] In particular, we consider the following sectors as modern: electrical and communication materials, mechanical, chemical and pharmaceutical, and automobile and other transport materials. Their codes at the Brazilian Institute of Geography and Statistics (IBGE) are 4506, 4510, 4512 and 4515, respectively.

[4] The empirical evidence is conflicting at best. Pekkarinen and Vartiainen's (2006) results confirm that women quit more often than men, whereas Blau and Kahn (1981) and Ransom and Oaxaca (2005) find no evidence supporting such gender differences.

3.1 Partial rank estimation of duration models

Khan and Tamer (2007) propose a partial rank estimator for duration models that imposes no parametric specification on the baseline hazard function and allows for general forms of censoring. For instance, in the right-censored version of Ridder's (1990) generalized accelerated failure time (GAFT) model, one observes $\boldsymbol{Y}_i = (T_i, I_i^{(RC)})'$, where the duration $T_i = \min\{\ell^{-1}(\boldsymbol{X}_i\boldsymbol{\beta} + \epsilon_i), \ell^{-1}(R_i)\}$ with R_i denoting the censoring variable and $\ell(\cdot)$ some unknown monotone link function, and $I_i^{(RC)} = \boldsymbol{I}\left\{\ell^{-1}(\boldsymbol{X}_i\boldsymbol{\beta} + \epsilon_i) \leq \ell^{-1}(R_i)\right\} = \boldsymbol{I}\{\boldsymbol{X}_i\boldsymbol{\beta} + \epsilon_i \leq R_i\}$ indicates whether there is right censoring. As before, $T_i = T_i^*$ for uncensored observations, and $T_i = R_i$ otherwise.

Similarly to Han's (1987) maximum rank correlation estimator, the idea is to find a transformation $F_{ij} = f(\boldsymbol{Y}_i, \boldsymbol{Y}_j)$ such that

$$\mathbb{E}\left[\boldsymbol{I}\{F_{ij} \geq 0\} \,\middle|\, X_i, X_j\right] = \mathbb{E}\left[\boldsymbol{I}\{F_{ji} \geq 0\} \,\middle|\, X_i, X_j\right] \tag{4}$$

if and only if $\boldsymbol{X}_i\boldsymbol{\beta} \geq \boldsymbol{X}_j\boldsymbol{\beta}$. Han (1987) considers $F_{ij} = T_i^* - T_j^*$ in the context of uncensored transformation models, which turns out to produce inconsistent estimates if the censoring variable R_i depends somehow on the covariates $\boldsymbol{X}_i$. Instead, Khan and Tamer set $F_{ij} = \bar{T}_i - T_j$, where $\bar{T}_i = I_i^{(RC)}\, T_i + (1 - I_i^{(RC)}) \times (+\infty)$ with $0 \times (+\infty) = 0$. As such,

$$\boldsymbol{I}\{F_{ij} \geq 0\} = \boldsymbol{I}\{\bar{T}_i - T_j \geq 0\} = 1 - I_i^{(RC)} + I_i^{(RC)}\, \boldsymbol{I}\{T_i \geq T_j\}. \tag{5}$$

It then follows that $\ell(T_i) \leq \boldsymbol{X}_i\boldsymbol{\beta} + \epsilon_i \leq \ell(\bar{T}_i)$ and hence

$$\boldsymbol{X}_i\boldsymbol{\beta} \geq \boldsymbol{X}_j\boldsymbol{\beta} \quad \Rightarrow \quad \Pr(\bar{T}_i \geq T_j) \geq 1/2 \tag{6}$$

by monotonicity of $\ell(\cdot)$.

Identification is possible only up to scale given that the function $\ell(\cdot)$ is unknown and hence it is more convenient to reparameterize the model as follows: $\boldsymbol{\beta} = (1, \boldsymbol{\theta}')'$. The partial rank estimator of $\boldsymbol{\theta}$ then is

$$\widehat{\theta} = \operatorname{argmax}_{\boldsymbol{\theta}\in\Theta} \frac{1}{n(n-1)} \sum_{i\neq j} I_i^{(RC)}\, \boldsymbol{I}\{T_i < T_j\}\, \boldsymbol{I}\{\boldsymbol{X}_i\boldsymbol{\beta} < \boldsymbol{X}_j\boldsymbol{\beta}\}, \tag{7}$$

where Θ is the parameter space. Note that the rank correlation function depends only on uncensored observations, though their ranks consider all observations with longer or equal durations. The partial rank estimator thus combines the information on both censored and uncensored observations, just as in the partial maximum likelihood method put forth by Cox (1972, 1975). In addition, it is straightforward to observe that Khan and Tamer's (2007) partial rank estimator is numerically equivalent to Han's (1987) maximum rank correlation estimator in the absence of censoring as well as in the case of fixed censoring (e.g., $R_i = R$).

Finally, it is also possible to extend the partial rank framework to consider double censoring. Let $T_i^* = \ell^{-1}(\boldsymbol{X}_i\boldsymbol{\beta} + \epsilon_i)$ as before. In the doubly censored regression model, we observe the pair $(T_i, I_i^{(DC)})$, where

$$\begin{aligned} T_i &= \boldsymbol{I}\{I_i^{(DC)} = 1\}(\boldsymbol{X}_i\boldsymbol{\beta} + \epsilon_i) + \boldsymbol{I}\{I_i^{(DC)} = 2\}\, L_i + \boldsymbol{I}\{I_i^{(DC)} = 3\}\, R_i && (8) \\ I_i^{(DC)} &= \boldsymbol{I}\{L_i < \boldsymbol{X}_i\boldsymbol{\beta} + \epsilon_i \le R_i\} + 2\,\boldsymbol{I}\{\boldsymbol{X}_i\boldsymbol{\beta} + \epsilon_i \le L_i\} + 3\,\boldsymbol{I}\{\boldsymbol{X}_i\boldsymbol{\beta} + \epsilon_i > R_i\} && (9) \\ & && (10) \end{aligned}$$

with L_i and R_i denoting left- and right-censoring variables such that $\Pr(L_i < R_i) = 1$. Applying the same trick as before, we let

$$\begin{aligned} \bar{T}_i &= \boldsymbol{I}\{I_i^{(DC)} < 3\}\, T_i + \boldsymbol{I}\{I_i^{(DC)} = 3\} \times (+\infty) && (11) \\ \underline{\mathrm{T}}_i &= \boldsymbol{I}\{I_i^{(DC)} \neq 2\}\, T_i + \boldsymbol{I}\{I_i^{(DC)} = 2\} \times (-\infty), && (12) \end{aligned}$$

and accordingly define $F_{ij} = \bar{T}_i - \underline{\mathrm{T}}_j$. This yields

$$\widehat{\theta} = \mathrm{argmax}_{\boldsymbol{\theta}\in\Theta}\, \frac{1}{n(n-1)} \sum_{i\neq j} \boldsymbol{I}\{\bar{T}_i \ge \underline{\mathrm{T}}_j\}\boldsymbol{I}\{\boldsymbol{X}_i\boldsymbol{\beta} \ge \boldsymbol{X}_j\boldsymbol{\beta}\} \qquad (13)$$

as the corresponding partial rank estimator for the doubly censored transformation model.

Khan and Tamer (2007) characterize the consistency and asymptotic normality of the partial rank estimators in (7) and (13) under the usual regularity conditions for rank-based semiparametric estimators. See, for instance, Sherman (1993) for similar regularity conditions in the context of maximum rank correlation.

4 Promotions in the Brazilian manufacturing industry

4.1 Preliminary analysis

The main parametric duration model we estimate assumes a generalized gamma distribution for promotion durations. There are two distributional parameters, namely, the scale and shape parameters σ and ς. We employ the generalized gamma distribution because it encompasses most of the distributions that appear in the duration literature. In particular, the generalized gamma coincides with the lognormal distribution if $\varsigma = 0$ and with the Weibull distribution if $\varsigma = 1$. We thus also examine using log-likelihood ratio tests whether the distributional parameter estimates are consistent with the lognormal and Weibull distributions.

Table 3 reports the estimation results of the accelerated failure time models with Gamma frailty for the different error distributions.[5] The estimates of the regression coefficients are very similar regardless of the distribution assumption. They indicate that, within firms of domestic capital, there are very significant gender differences in promotions. In particular, male employees wait significantly less than female employees to get a promotion. The time to promotion for men is on average between 32.71% and 37.39% shorter than that for women depending on the specification. Gender differences are less pronounced in multinationals, though still significant. Male workers now take on average between 27.42% and 31.13% less time to get a promotion than women. In addition, time to promotion is relatively shorter, though not significantly, within multinationals than within domestic firms regardless of gender.[6] Table 4 validates these claims through formal hypothesis testing. Panel A in Table 4 documents that, even though multinationals display less gender differences than domestic firms, they are still significant at the 1% level of significance regardless of the distribution one employs. Panel B also shows that males tend to get a promotion faster within multinationals than within domestic firms.

As for the other controls, it is interesting to observe that the size effect is pretty insignificant as in Blau and deVaro (2007), though time to promotion is increasing with productivity and exposure to international markets given the sign of the coefficient estimates for EXPORTS, MEAN WAGE, and MODERN. This is consistent with a more competitive environment as what concerns internal turnover within more productive firms (including Brazilian multinationals; see Footnote 1). In contrast, it is decreasing with UNIVERSITY RATIO and TURNOVER, reflecting competition effects within the firm and the sector, respectively. The remaining regression coefficients are all as expected. For instance, there is a significant negative relationship between time to promotion and age, which is not surprising given that the latter acts as a proxy for experience. In addition, intermediate managers and supervisors wait substantially less to obtain a promotion to a managerial position.

Finally, the frailty parameter that regulates the variance of the individual random effects does not differ from zero as long as one considers either a generalized gamma or lognormal distribution. The absence of a frailty factor perhaps results from the homogeneity of our sample.[7] In contrast, the frailty variance is quite close to one for the Weibull distribution. This is consistent with the

[5] Although we report only nonrobust standard errors, clustering by firm changes only very marginally the confidence intervals of the coefficient estimates. All qualitative results thus remain valid for cluster-robust standard errors.

[6] We compute these effects by exponentiating the (sum of) regression coefficient(s) and then subtracting one. For instance, specification (2) in Table 3 gives way to a change of $\exp(-0.4524) - 1 = -36.39\%$ in the average time to promotion if the individual is male. This effect reduces to $\exp(-0.4524 + 0.0935) - 1 = -30.16\%$ if the individual is male and works at a multinational.

[7] Not surprisingly, the coefficient estimates for the duration models without frailty are almost identical to the ones in Table 3 regardless of the distribution. These results are available from the authors upon request.

fact that a frailty-implied gamma mixture of Weibull variates results in a Burr distribution, which is very similar to the generalized gamma distribution (Rodriguez, 1977). Table 5 confirms that, despite the similarity of the regression coefficients, the statistical evidence favors the extra flexibility of the generalized gamma and Burr distributions (i.e., either generalized gamma without frailty or Weibull with frailty) as opposed to the more parsimonious lognormal distribution.[8]

4.2 Semiparametric analysis

We now investigate the extent to which the log-linear specification and the assumption of covariate-independence censoring affect the results. As frailty does not seem to matter much, we estimate the duration model using Khan and Tamer's (2007) semiparametric estimator. As the latter relies on rank-based methods, it consistently estimate the relative magnitude of the regression coefficients for any strictly monotonic link function (in particular, we fix the coefficient of EXPORTS to unit). This results in a semiparametric variant of the AFT model in (1) under which the link function is strictly monotonic, but otherwise unknown. This is in stark contrast to the simple log transformation that (1) imposes. Moreover, Khan and Tamer's (2007) partial-rank estimator does not require specifying a parametric family for the error distribution.

We employ a SAS/IML genetic algorithm to compute the partial rank estimator, whereas we obtain standard errors by means of bootstrap methods as in Subbotin (2008). The latter hinges on 100 artificial samples with the same number of observations as the original sample. We consider two specifications. The first accounts for the position at which the individual starts at the firm and hence we must exclude every individual that already begins at a managerial position, that is to say, any individual with time to promotion equal to zero. The second specification does not control for the starting occupation and so we consider samples both with and without individuals with zero duration. Adding individuals that start at a managerial position to the sample ameliorates the precision of the estimates due to the increase in the sample size. However, it may bring about a sample selection bias. It is beyond the scope of this paper to propose an estimation procedure that also accounts for sample selection bias and hence we proceed by simply checking whether there is any qualitative change in the estimation results across these two samples.

Table 6 reveals that the qualitative findings are quite robust to the specification as well as to the sample. We indeed observe no variation in the sign of the coefficient estimates across the different specifications and samples. Accounting for previous occupation does not affect much the coefficient

[8] We report the p-value coming from usual chi-squared distribution for the LR test of the generalized gamma versus the lognormal even if the shape parameter is in the boundary.

estimates of interest, but it improves considerably their precision. The partial rank estimates unveil that females are in disadvantage in domestic firms relative to male workers, but significantly less so within foreign-owned companies. Including individuals with zero duration in the sample reinforces the moderating effect of multinationals to the extent of seemingly eliminating gender differences in promotion.

Table 7 confirms these results by means of formal Wald tests for linear restrictions. In fact, we cannot even reject the null of no gender differences within foreign-owned companies for the specifications that do not control for previous occupation. The partial rank estimates also indicate that it takes less time for a male worker to obtain a promotion within multinationals. As before, including individuals with zero durations reverses the result in that we cannot anymore reject at the 5% level of significance that time to promotion for men does not depend on whether the firm is domestic or foreign-owned.

The above findings contribute to the literature that compares multinationals and domestic firms (Doms and Jensen, 2006; Greene et al., 2009) as well as to the literature showing that gender differences may depend on the nature of the firm (see, e.g., deVaro and Brookshire, 2007). The impact of individuals that start at managerial positions is also interesting inasmuch as it reflects self-selection effects. Before discussing self-selection issues and the reasons why gender differences are less pronounced in multinationals, we first brief look into two other dimensions in which gender differences in promotions might arise, namely, pay rise due to promotion and likelihood of a promotion.

4.3 Wage growth and promotion likelihood

In this section, we investigate gender differences in promotion likelihoods and in wage growths. As per the latter, we employ wage information of workers over time to estimate longitudinal models of the effect of promotions on wages. The dummy variable PROMOTION takes value one for every period after a promotion and zero otherwise. We estimate both random- and fixed-effects panel regressions for males and females and then compare the coefficient estimates of interest to assess whether the impact of promotions on wages differs according to the gender. Apart from the previous firm-specific covariates, we control for job tenure and starting occupation as well as for interaction effects due to promotions within multinationals. The latter is to capture, if any, differential wage effects of promotions in foreign-owned firms.

Table 8 reports the coefficient estimates with their robust standard errors as well as Wald tests for the equality of the coefficients of interest as estimated from the samples of males and females. As expected, the results show that promotion increases wages of both males and females.

The specification with random effects suggest that women receive on average higher increases in wages after a promotion within domestic firms. In contrast, we cannot reject similar increases for men and women in domestic firms if restricting attention to fixed individual effects as well as in foreign-owned companies regardless of whether the individual effects are random or fixed. Finally, the coefficient estimate of the interaction dummy evince that pay increases due to promotions are somewhat smaller in multinationals.

To verify whether the probability of promotion depends on gender, we run logit regressions in which the dependent variable takes value one if the individual obtains a promotion to a managerial position, or else it equals zero. We control once more for the gender, occupation, and job tenure of the individual as well as for the same firm-specific covariates as before. The results in Table 9 evince that domestic firms seems to display gender differences in promotions not only in terms of how much time it takes to get a promotion, but also in terms of promotion likelihood. They also suggest that, apart from featuring less gender differences, it is more likely to obtain a promotion in a multinational for both genders. This is similar to the findings of the preliminary parametric analysis concerning time to promotion, reflecting perhaps the weaknesses they share, namely, the restrictiveness of the parametric specification and of the regressors' exogeneity assumption.

In sum, we document that wage gains after promotion does not contribute to creating gender differential at least within foreign-owned firms. The evidence is weaker for domestic firms. The logit regressions for the probability of a promotion indicate that women are in disadvantage within domestic firms but less so in multinationals, thereof confirming the time-to-promotion results. Finally, multinationals are also characterized by higher likelihoods of promotion regardless of gender.

4.4 Foreign ownership and gender differences

Our findings about gender differences are interesting, but it remains to explain how they arise and why they vary according to ownership. The first point to notice is that a taste for discrimination as in Becker (1957) does not survive long if there is enough competition amongst firms/employers (Arrow, 1972). Of course, one could always claim that, in addition to strong cultural differences, labor market institutions in Brazil are such that they actually curb the sort of competition that would drive away taste for gender discrimination. As we have no means to falsify such a conjecture, we turn our attention to the next suspect, namely, statistical discrimination (Arrow, 1972; Phelps, 1972; Spence, 1973).

Assessing the quality of a worker involves costs and hence some employers might consider costless inference methods such as observing the worker's gender. These employers would then apply their prior beliefs about the expected qualification of the worker conditional on gender as a hiring/promotion criterion. This would lead to statistical discrimination against a particular gender. Berk (2001) extends the statistical discrimination model to a world in which workers compete for jobs/promotions. It turns out that the self-selection that results from individuals rationally selecting which jobs/promotions to apply for helps mitigating (and sometimes even overcompensating) the effects of statistical discrimination. The self-selection mechanism is pretty simple. An individual from the gender with lower average qualifications would only apply for a job/promotion if the probability of getting the job/promotion compensates the application costs. This is more likely to occur if the individual has above-average qualifications.

We next argue that our results are consistent with Berk's (2001) model implications under the assumption that, on average, women face more constraints than men to work long hours or to do business trips. This is enough to generate statistical discrimination against women. A career-minded woman who applies for a job/promotion in a firm with a preference for dedicated employees only does so because she rationally believes that her odds of being hired/promoted compensate the effort to go through the application process. This means that her qualifications have to be sufficiently high to stand a chance despite discrimination. The latter entails the self-selection bit. To establish the differences between multinationals and domestic firms, we rely on two anecdotal evidence concerning the Brazilian labor market (OECD, 2008). The first is that domestic firms offer a more flexible package in terms of working hours and business trips than multinationals. This makes them more appealing to less ambitious/career-concerned individuals (regardless of their gender). The second is that multinationals compete more fiercely for ambitious/career-minded highly-skilled workers, promoting on average better pay than domestic firms (see, e.g., Martins and Esteves, 2008).

We find some indirect evidence supporting the above claims.[9] On the one hand, male workers tend to work similar hours in foreign-owned and domestic firms regardless of whether they have been promoted or not. The only difference is in the lower support of the distribution, with promoted men working much harder than non-promoted ones (22 to 44 hours versus 8 to 44 hours, respectively). This is consistent with the the fact that promoted workers are more likely to be career-minded. On the other hand, non-promoted female employees work relatively much less only in domestic firms (10 to 44 hours). This is in line with domestic firms offering more flexible packages. In addition,

[9] Note that, unfortunately, we only observe the effects of self-selection. We cannot properly model the selection mechanism given that we do not observe many characteristics that determine individual career-mindness, such as marital status and number of child dependents.

the fact that women work more in multinationals (36 to 44 hours regardless of whether promoted or not) is consistent with self-selection. As the latter alleviates the impact of statistical discrimination, we fail to observe as much gender difference in foreign-owned companies as in domestic firms.

The impact of individuals that start at managerial positions in the coefficient estimates reinforces the self-selection story. These are precisely the workers that are most likely to have the highest qualifications and hence their inclusion in the sample potentiates self-selection effects. As the latter reduces the imprint of statistical discrimination, it is not surprising that we cannot reject gender differences in multinationals for the sample that includes individuals with zero durations. This is because the repercussion to self-selection is higher in multinationals given that they compete more intensively for career-minded workers.

Before concluding, it is important to make a caveat. Our sample is very homogeneous as what concerns skill levels in that we consider only individuals with university degrees and pursuing specific career types. Also, we have no reason to believe that the skill distribution is gender specific. Differentiating between low- and high-skilled workers would then bring not much to the glass ceiling discussion even if the impact of career-mindness in the time to promotion may differ across different levels of skills.

5 Conclusion

This paper examines whether there exists a glass ceiling to women's ascension in the largest firms of the Brazilian manufacturing industry. Our main goal is to check whether gender matters in the time it takes to get a promotion to a managerial position. The motivation lies not only on the natural interest in assessing gender differences in a developing country, but also on the general features of the Brazilian data set. In particular, the latter include observations of multiple workers per establishment at different occupations and hierarchical levels as well as their career paths in terms of occupation and wages.

Our main results hinge on the partial rank estimation of a semiparametric AFT model for the time to promotion as in Khan and Tamer (2007). Apart from relaxing the parametric restrictions on the specification of the duration model, their semiparametric approach also allows for covariates that are not necessarily independent from censoring. The latter is pretty convenient for it is very likely that censoring depends on gender.

We find that there are significative gender differences in promotions within domestic firms. The evidence for foreign-owned firms is a bit weaker, however. These findings complement well the recent evidence that the nature of the firm may entail substantial differences in managerial practices and in the role of promotions (Doms and Jensen, 2006; deVaro and Brookshire, 2007; Greene et al., 2009). We argue that the differences in gender differences are a result of a combination between statistical discrimination and job selection (Berk, 2001). If career-concerned women prefer to work in multinational firms, statistical discrimination would become more apparent within domestic firms.

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Table 1
Sample size according to gender and occupation

The label 'absolute' refers to the number of observations in that cell, whereas 'relative' corresponds to the relative sample size as a percentage of the total number of observations in that column. We group under 'AAEL' all individuals that work as accountants, administrators, economists, and lawyers.

occupation	female		male		total	
	absolute	relative	absolute	relative	absolute	relative
ENGINEER	689	19.4%	8,171	34.4%	8,860	32.5%
AAEL GROUP	864	24.3%	2,637	11.1%	3,501	12.8%
INTERMEDIATE MANAGER	773	21.8%	3,334	14.0%	4,107	15.1%
SUPERVISOR	562	15.8%	2,721	11.5%	3,283	12.0%
MANAGER AND DIRECTOR	664	18.7%	6,874	29.0%	7,538	27.6%
total	3,552	100%	23,737	100%	27,289	100%

Table 2

Sample statistics according to censoring and gender

The columns 'mean' and 'st.dev.' correspond to the sample averages and standard deviations. We gauge PROMOTION DURATION in months, AGE in years, and HOURS AT WORK in hours per week. All other individual-specific variables are binary assuming value either one or zero according to the occupational group. As for the variables relating to the firm at which the individual works, SIZE) and EXPORTS denote the number of employees and how much the firm exports in USD billions, respectively. UNIVERSITY RATIO is the fraction of employees with a university degree, whereas TURNOVER gauges the job flow intensity of the firm. MEAN WAGE of the firm is the average monthly stipend across employees. Finally, MODERN is a binary variable that takes value one if the firm is within a technology-intensive sector.

variable	zero duration				no censoring				right censoring			
	female		male		female		male		female		male	
	mean	st.dev.	mean	st.dev.	mean	st.dev.	mean	st.dev.	mean	st.dev.	mean	st.dev.
PROMOTION DURATION	0	0	0	0	41.1	35.1	41.9	34.8	33.7	33.0	35.2	33.9
ENGINEER	0	0	0	0	0.22	0.41	0.44	0.50	0.24	0.43	0.49	0.50
AAEL GROUP	0	0	0	0	0.27	0.44	0.13	0.34	0.30	0.46	0.16	0.37
INTERMEDIATE MANAGER	0	0	0	0	0.35	0.48	0.27	0.44	0.26	0.44	0.19	0.39
SUPERVISOR	0	0	0	0	0.16	0.37	0.16	0.36	0.20	0.40	0.16	0.37
MANAGER AND DIRECTOR	1	0	1	0	0	0	0	0	0	0	0	0
AGE	34.4	7.2	39.6	7.5	30.0	5.8	33.2	7.6	30.2	6.1	33.1	7.7
HOURS AT WORK	43.0	2.5	43.3	1.9	42.9	2.5	43.1	1.8	43.0	2.5	42.9	2.5
MULTINATIONAL	0.34	0.47	0.32	0.47	0.48	0.50	0.49	0.50	0.37	0.48	0.41	0.49
SIZE	2,314	2,983	2,419	3,370	3,033	3,081	3,046	3,604	3,555	6,177	3,568	6,125
EXPORTS	0.038	0.120	0.043	0.131	0.076	0.184	0.073	0.163	0.059	0.151	0.085	0.194
UNIVERSITY RATIO	0.19	0.14	0.16	0.12	0.20	0.13	0.18	0.13	0.17	0.13	0.17	0.13
TURNOVER	0.56	0.24	0.56	0.24	0.52	0.21	0.52	0.25	0.53	0.24	0.52	0.24
MEAN WAGE	2,107.0	1,108	2,017	961.2	2,238	973.9	2,280	1,003	2,180	1,055	2,291	1,030
MODERN	0.37	0.48	0.37	0.48	0.37	0.49	0.41	0.49	0.38	0.49	0.49	0.50
number of observations	664		6,874		161		1,465		2,727		15,398	

Table 3

Maximum likelihood estimates of the AFT models with gamma frailty

The dependent variable is the natural logarithm of the time to promotion. We report maximum likelihood estimates of the regression coefficients as well as of the distributional and frailty parameters. Figures within brackets correspond to standard errors. We exclude the AAEL group from the set of occupational dummies in columns (2), (4) and (6).

covariate	generalized gamma		lognormal		Weibull	
	(1)	(2)	(3)	(4)	(5)	(6)
MALE	−0.3979	−0.4524	−0.4077	−0.4683	−0.3962	−0.4524
	(0.0928)	(0.0930)	(0.0940)	(0.0944)	(0.0925)	(0.0930)
MALE×MULTINATIONAL	0.0736	0.0935	0.0872	0.0954	0.0746	0.0932
	(0.1350)	(0.1347)	(0.1402)	(0.1400)	(0.1347)	(0.1348)
MULTINATIONAL	−0.4142	−0.4195	−0.4547	−0.4436	−0.4164	−0.4195
	(0.1307)	(0.1302)	(0.1351)	(0.1349)	(0.1304)	(0.1303)
AGE	−0.0087	−0.0045	−0.0135	−0.0086	−0.0090	−0.0045
	(0.0030)	(0.0031)	(0.0030)	(0.0031)	(0.0030)	(0.0031)
ln(SIZE)	0.0254	0.0255	0.0348	0.0345	0.0246	0.0256
	(0.0272)	(0.0269)	(0.0292)	(0.0291)	(0.0271)	(0.0270)
EXPORTS	0.3975	0.3649	0.4101	0.4168	0.3974	0.3657
	(0.1440)	(0.1424)	(0.1482)	(0.1485)	(0.1432)	(0.1425)
UNIVERSITY RATIO	−1.2538	−1.0992	−1.3779	−1.2028	−1.2478	−1.1004
	(0.2776)	(0.2778)	(0.3010)	(0.3011)	(0.2773)	(0.2780)
TURNOVER	−0.2599	−0.2274	−0.2484	−0.2066	−0.2650	−0.2272
	(0.0930)	(0.0942)	(0.1020)	(0.1022)	(0.0933)	(0.0942)
MEAN WAGE	0.0001	0.0001	0.0001	0.0001	0.0001	0.0001
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
MODERN	0.4650	0.4515	0.5053	0.4830	0.4686	0.4518
	(0.0529)	(0.0518)	(0.0546)	(0.0546)	(0.0521)	(0.0518)
ENGINEER		0.0339		0.0896		0.0344
		(0.0640)		(0.0663)		(0.0640)
INTERMEDIATE MANAGER		−0.3222		−0.3027		−0.3219
		(0.0676)		(0.0709)		(0.0677)
SUPERVISOR		−0.1572		−0.0979		−0.1564
		(0.0746)		(0.0782)		(0.0746)
CONSTANT	6.0976	5.8156	6.2913	6.2496	5.9026	5.8795
	(0.3465)	(0.3323)	(0.2840)	(0.2879)	(0.2694)	(0.2749)
scale σ	0.9904	0.6521	1.5919	1.5838	0.7385	0.7305
	(0.2983)	(0.2303)	(0.0280)	(0.0278)	(0.0223)	(0.0218)
shape ς	0.6912	1.1263	0	0	1	1
	(0.3146)	(0.4028)				
frailty variance λ	0.2247	1.7164	0.0000	0.0001	1.2212	1.3379
	(1.0610)	(1.2459)	(0.0013)	(0.0024)	(0.4301)	(0.4225)
sample size	19,751		19,751		19,751	

Table 4
Likelihood-ratio tests for linear restrictions in the AFT coefficients

The restricted log-likelihood value refers to constraining the maximum likelihood estimator so that the sum of the regression coefficients is zero. Panel A tests whether there is gender differences in time to promotion within multinational firms, whereas Panel B examines whether it takes less time for men to get a promotion in multinationals. All results are for specifications that include occupational dummies.

distribution	log-likelihood		LR statistic	p-value
	unrestricted	restricted		
Panel A: MALE + MALE×MULTINATIONAL = 0				
generalized gamma	-5,929.35	-5,936.36	14.01	0.0002
lognormal	-5,950.79	-5,957.32	12.97	0.0003
Weibull	-5,929.39	-5,936.40	14.02	0.0004
Panel B: MULTINATIONAL + MALE×MULTINATIONAL = 0				
generalized gamma	-5,929.35	-5,949.19	39.68	0.0000
lognormal	-5,950.79	-5,970.37	39.15	0.0000
Weibull	-5,929.39	-5,949.23	39.68	0.0000

Table 5
Likelihood ratio tests for the shape parameter

The unrestricted log-likelihood value refers to the maximum likelihood estimation under the generalized gamma distribution, whereas the restricted log-likelihood value corresponds to constraining the shape parameter ς either to zero or to one, so that the generalized gamma distribution reduces to the lognormal or Weibull distributions, respectively.

distribution	log-likelihood		LR statistic	p-value
	unrestricted	restricted		
lognormal ($\varsigma = 0$)	-5,929.35	-5,950.79	42.88	0.0000
Weibull ($\varsigma = 1$)	-5,929.35	-5,929.39	0.08	0.7831

Table 6

Partial rank estimates of the semiparametric time-to-promotion model

We estimate a model for time to promotion in which the link function is strictly monotonic, but otherwise unknown. We consider two samples. The first excludes individuals that begin at the firm either as managers or directors (i.e., with zero duration), whereas the second includes these individuals and hence does not allow one to control for the starting occupation. The partial rank estimator identifies the regression coefficients only up to scale and so we fix the coefficient for EXPORTS to one. We obtain the pointwise estimates by means of a genetic algorithm, whereas we employ subsampling to compute the standard errors that we report within parentheses.

covariate	no zero durations		with zero durations
	with occupation	without occupation	
MALE	−1.6872 (0.2367)	−1.5774 (0.4310)	−0.4835 (0.0523)
MALE×MULTINATIONAL	0.4035 (0.1850)	0.4576 (0.7971)	0.4838 (0.0640)
MULTINATIONAL	−1.4063 (0.2079)	−1.6530 (0.8481)	−0.5463 (0.0411)
AGE	−0.0598 (0.0045)	−0.0937 (0.0212)	−0.1486 (0.0297)
ln(SIZE)	0.1305 (0.0686)	0.1241 (0.1102)	0.0911 (0.0316)
EXPORTS	1.0000	1.0000	1.0000
UNIVERSITY RATIO	−5.2572 (0.6845)	−6.6980 (1.0881)	−3.5195 (0.4713)
TURNOVER	−0.3169 (0.2834)	−0.5950 (0.3894)	−0.1406 (0.0488)
MEAN WAGE	0.0005 (0.0000)	0.0007 (0.0001)	0.0005 (0.0000)
MODERN	2.0510 (0.1506)	2.2914 (0.3475)	0.4490 (0.0866)
ENGINEER	0.6561 (0.9823)		
INTERMEDIATE MANAGER	−1.2572 (1.4152)		
SUPERVISOR	−0.1089 (5.1821)		
sample size	19,751	19,751	27,289

Table 7

Wald tests for linear restrictions in the semiparametric model coefficients

We report partial rank estimates for the sum of coefficients with their subsampling-based standard errors within parenthesis and with their p-values within brackets. Panel A tests whether there is gender differences in time to promotion within multinational firms, whereas Panel B examines whether it takes less time for men to get a promotion in multinationals as compared to domestic firms.

	no zero durations		with zero durations
	with occupation	without occupation	
Panel A: MALE + MALE×MULTINATIONAL = 0			
partial rank estimate	-1.2837	-1.1198	0.0003
standard error	(0.2858)	(1.1530)	(0.0831)
p-value	[0.0000]	[0.3315]	[0.9966]
Panel B: MULTINATIONAL + MALE×MULTINATIONAL = 0			
partial rank estimate	-1.0028	-1.1954	-0.0624
standard error	(0.1737)	(0.2682)	(0.0359)
p-value	[0.0000]	[0.0000]	[0.0817]

Table 8

Wage regressions with individual effects for both female and male workers

The dependent variable is the logarithm of real monthly wages. The dummy variable PROMOTION takes value one for every period after a promotion and zero otherwise, whereas TENURE corresponds to the number of years the individual has worked for the firm. We report both random- and fixed-effects estimates with their robust standard errors within parentheses for both female and male wage regressions as well as the p-values of the Wald test for the equality some (linear combinations of) coefficients in both regressions.

covariate	random effects		fixed effects	
	female	male	female	male
PROMOTION	0.2569 (0.0312)	0.1874 (0.0105)	0.1937 (0.0387)	0.1304 (0.0124)
MULTINATIONAL×PROMOTION	−0.0996 (0.0418)	−0.0358 (0.0141)	−0.0984 (0.0520)	−0.0282 (0.0168)
MULTINATIONAL	0.0700 (0.0240)	0.0770 (0.0096)		
lnSIZE	0.0412 (0.0123)	−0.0039 (0.0052)		
EXPORTS	0.3079 (0.0803)	0.3746 (0.0243)		
UNIVERSITY RATIO	−0.6454 (0.1408)	−0.6342 (0.0559)		
TURNOVER	0.0606 (0.0487)	0.1168 (0.0201)		
MEAN WAGE	0.0003 (0.0000)	0.0002 (0.0000)		
MODERN	0.0742 (0.0247)	0.0613 (0.0097)		
ENGINEER	0.0550 (0.0257)	−0.0101 (0.0120)		
INTERMEDIATE MANAGER	0.1084 (0.0280)	0.1477 (0.0140)		
SUPERVISOR	−0.1100 (0.0304)	−0.0640 (0.0152)		
TENURE	0.1324 (0.0041)	0.1217 (0.0016)	0.1309 (0.0047)	0.1210 (0.0017)
TENURE^2	−0.0059 (0.0003)	−0.0051 (0.0001)	−0.0057 (0.0004)	−0.0050 (0.0001)
CONSTANT	7.0700 (0.1039)	7.7659 (0.0454)	8.0128 (0.0111)	8.3169 (0.0043)
Wald test for equality of coefficients				
PROMOTION		0.0348		0.1190
PROMOTION + MULTINATIONAL×PROMOTION		0.8487		0.8585
number of observations	10,343	67,125	10,343	67,125

Table 9
Logit regressions for the probability of promotion

The dependent variable is a dummy that takes value one if the individual obtains a promotion to a managerial position, or else it equals zero. We report both random- and fixed-effects estimates with their robust standard errors within parentheses as well as the p-values of the Wald test for whether the sum of coefficients is equal to zero.

covariate	(1)	(2)	(3)
MALE	0.6029 (0.1184)	0.6238 (0.1191)	0.6297 (0.1195)
MALE×MULTINATIONAL	−0.1409 (0.1727)	−0.1596 (0.1733)	−0.1628 (0.1737)
MULTINATIONAL	0.7145 (0.1657)	0.7119 (0.1662)	0.6912 (0.1669)
lnSIZE	0.0179 (0.0322)	0.0251 (0.0320)	0.0086 (0.0323)
EXPORTS	−0.2507 (0.1801)	−0.2634 (0.1765)	−0.3758 (0.1773)
UNIVERSITY RATIO	1.6929 (0.3106)	1.6144 (0.3098)	1.4631 (0.3173)
TURNOVER	−0.0918 (0.1249)	−0.1408 (0.1264)	−0.0732 (0.1251)
MEAN WAGE	−0.0002 (0.0000)	−0.0002 (0.0000)	−0.0002 (0.0000)
MODERN	−0.5037 (0.0649)	−0.4875 (0.0655)	−0.5049 (0.0656)
ENGINEER		0.1017 (0.0798)	0.0609 (0.0801)
INTERMEDIATE MANAGER		0.4718 (0.0852)	0.4593 (0.0857)
SUPERVISOR		0.1030 (0.0945)	0.1143 (0.0948)
TENURE			0.0203 (0.0022)
TENURE^2			−0.0001 (0.0000)
CONSTANT	−2.9127 (0.2943)	−3.1655 (0.2988)	−3.5113 (0.3045)
Wald test for whether the sum of coefficients is equal to zero			
MALE + MALE×MULTINATIONAL	0.0002	0.0003	0.0002
MULTINATIONAL + MALE×MULTINATIONAL	0.0000	0.0000	0.0000
log likelihood	-5,523.80	-5,502.38	-5,450.79
number of observations	19,751	19,751	19,751