

## OCCUPATIONAL FEMINIZATION AND PAY: THE CASE OF BRAZIL

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The sex wage gap is widely documented in the literature. Although several theories emerged to explain it, the devaluation and queuing views have been dominant due to women's concentration in lower-paying occupations. They differ in the direction of the causal effect: while the devaluation theory predicts lower wages for higher occupation's female share, the queuing theory predicts causal effect in the opposite direction. No studies to support these theories have been found for developing countries and this paper aims to test them for Brazil. Using Censuses data, three results emerge: 1) negative correlation between sex composition and wages; 2) no evidence to support neither theory; 3) some evidence to support another view: the theory of equalizing differences. This last theory predicts no causal relationship between the occupational share of females and wages because wage differentials would emerge from human capital differences and the value attributed to jobs' non-monetary advantages.

KEYWORDS: Gender inequality, Occupations, Earnings, Longitudinal analysis.

### 1. INTRODUCTION

Worldwide, women receive on average less than men, even controlling for education. A sizable literature shows that part of this pay gap comes from women's concentration in lower-paying occupations (Blau and Kahn, 2006; 2017).

Among the explanations for the association between occupations' sex composition and payment, two sociological theories have gained ground: the devaluation theory and the queuing view. The first one assumes that employers attribute a lower social value, hence a lower pay, for the work done by women (England, 1992). Therefore, if women's share in an occupation grows, its pay falls, although there is no degradation of work processes or of skill utilization.

The queuing view also assumes that there is discrimination against women, but it occurs during the hiring. According to this theory, all workers prefer occupations with higher pay relative to their educational attainments, but employers prefer to hire men. Consequently, better-paying jobs are filled mostly by men, leaving scarce places for women, who have to find jobs in occupations with lower pay (Reskin and Roos, 1990; Strober and Catanzarite, 1994).

Longitudinal studies have proven to be more appropriate than cross-sectional analysis to test these theories because they enable the investigation of the direction of the causal relationship between occupations' sex composition and payment. Considering occupations as the units of analysis, if the occupations' female proportion is affected by prior wages, there is evidence for the queuing view. Instead, if earlier sex composition causes changes in occupations' median (or mean) pay, then data speak in favor of the devaluation thesis.

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Since the late 1970s, many studies on panel data have been carried out, bringing support for the devaluation view (see, for example, Snyder and Hudis (1976); Baron and Newman (1989); Semyonov and Lewin-Epstein (1989); Karlin et al. (2002); Catanzarite (2003); England et al. (2007); Levanon et al. (2009)). However, to the best of our knowledge, none of them are related to developing countries.

In Brazil, women receive about 77.7% of men's earnings, despite being, on average, more educated (Instituto Brasileiro de Geografia e Estatística, 2021). As in other countries, a significant part of this pay gap comes from the female concentration in lower-paying occupations (Madalozzo, 2010). Another striking feature of the Brazilian labor market is the high degree of informality, which accounted for 41.6% of workers in 2019 (Instituto Brasileiro de Geografia e Estatística, 2020, p. 22). These two stylized facts—sex segregation and informality—appear combined in many occupations. For instance, more than 90% of paid domestic workers in Brazil are women, and only one-third of them have a formal labor contract. One of the reasons women are very present in the informal labor market is that this type of employment allows for more flexible working hours. The difficulties reconciling reproductive work with that done out of home make them prefer jobs with more flexible schedules (Bruschini, 1994, p. 186).

Many developing countries face the same labor market issues, suggesting that, for these countries, a third theory could explain the selection bias due to women's preference for flexible working hours and family/mother-friendly jobs. Indeed one framework has been developed among economists in contrast to the devaluation and queuing sociological views: the theory of equalizing differences. This theory establishes that workers value jobs' non-monetary (dis)advantages, so jobs associated with some (dis)amenity must pay (above)below-average wages to attract workers (Rosen, 1986). Therefore, besides human capital, wages would also be determined by a compensating differential.

If the theory of equalizing differences is correct, neither the occupation's wages nor the proportion of women affect each other because differences in wages across occupations arise from non-monetary (dis)advantages and human capital. This theory should not be disregarded since non-monetary attributes play a significant role in the decision for women to participate in the labor market.

Considering the literature gap on empirical works that test the devaluation and queuing theories for developing countries, this article intends to fill this gap by testing both views, for men and women, for Brazil, with Census data from 1980, 1991, and 2000.

This article is organized as follows. Section 2 reviews the major sociological views about the relationship between occupations' sex composition and payment and presents previous empirical results on the topic. Section 3 presents past evidence on sex occupational segregation in Brazil and its relation to the wage gap. Section 4 presents the data and models used and the empirical strategy. Results are presented and discussed in Section 5, and Section 6 concludes the study.

## 2. THEORY AND PAST EVIDENCE

According to the devaluation view, employers attribute a lower social status for female jobs. Thus, if an occupation is filled mostly by women, it is seen as less pay-worth and hence is assigned a lower pay (England, 1992). If this explanation stands, supposing occupational skill demands are appropriately controlled, occupations' sex composition has a causal effect on their later wage setting.

The queuing view, also known as "relative attractiveness" theory, states that employer's discrimination operates in a different manner. Although all workers prefer to work in occupations with higher pay relative to their educational attainments, employers generally prefer to

hire men. Such hiring discrimination against women results in better-paying jobs being filled mostly by men. As the queue moves forward, there is usually no place left for women in better-paid occupations, although they would prefer to work on it, and they will find work only in not so well-paid occupations, leading to female concentration in occupations with lower pay (Reskin and Roos, 1990; Strober and Catanzarite, 1994). Therefore, in this case the direction of the causal relationship between occupations' sex composition and payment is reversed, with occupations' wage levels at one time affecting later gender composition.

Another view contrasting the previous ones is the theory of equalizing differences. In such theory, neither the wage nor the sex composition causes each other, and any correlation between these variables is spurious. This theory has its origins in the seminal Adam Smith's 1776 work *The Wealth of Nations*, but it was only rigorously summarized almost two centuries later (Friedman, 1962). Since then, many empirical works have been dedicated to establishing the determinants of the wage structure based on a fundamental assumption: some jobs offer favorable working conditions while others don't. The first ones attract workers at lower wages than average because of non-monetary advantages. On the other hand, jobs that offer unfavorable working conditions must pay wages above-average to attract workers (Rosen, 1986). As women value jobs' non-monetary attributes such as more flexible working hours or more mother-friendly work places, they are attracted to jobs that offer these amenities, even if they pay less than the average. Men, on the other hand, give less value to such non-monetary attributes (Karlin et al., 2002). The wage gap is observed due to the difference between the total monetary and non-monetary (dis)advantages. Thus an "equalizing difference" is necessary to lead the labor market to a (long-run) equilibrium.

Although the devaluation and queuing explanations are not mutually exclusive and causal relationships in both ways could be going on simultaneously (Karlin et al., 2002), empirical researchers have been more successful in finding support for the devaluation view than for the queuing theory. Also, longitudinal studies have proven to be more appropriate than cross-sectional analysis to test for a causal order between gender composition and occupations' wages, and include, in the recent period, Karlin et al. (2002), Catanzarite (2003), England et al. (2007), Levanon et al. (2009), and Busch (2018). In common, these authors consider occupations (instead of workers) as the unit of analysis. They all run models to test the devaluation hypothesis, therefore adopting occupations' sex composition as a predictor of median (or average) wages. Models control for occupational characteristics believed to affect wages, such as average education level and experience (or proxies for it), returns to education and experience, race composition, urban/rural mix, regional concentrations, among others.<sup>1</sup> All of them include some modeling strategy to treat the omitted variable bias associated with non-observable occupations skill requirements or disamenities. Finally, another common feature is that they carry national analysis of US data.

Catanzarite (2003) investigates if wages depreciate in occupations with a large share of minorities (black, Latina, or white women or black or Latino men). She uses data from the Current Population Survey (CPS) of 1972, 1982, 1983, and 1993. Due to changes in occupational classifications, separate analyses are run for the 1972-1982 and 1983-1993 panels. White men's median hourly wages are modelled as a function of percentages of race-gender groups in the previous period — the so-called cross-lagged panel model —, along with occupational characteristics. Controls also include lagged wages in order to minimize omitted variable bias. Results show that the percent of white and black women contributes to pay deterioration within occupations, supporting the devaluation view (the author does not test the queuing view).

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<sup>1</sup>Catanzarite (2003) also controls industrial distribution, proportions of incumbents who are part-time, in public employment, and self-employed.

Karlin et al. (2002) use a more detailed occupational classification system, created by cross-tabulations of three digits occupational categories with one-digit industry codes. According to the authors, this “controls for changes in occupations that would occur if, over time, an occupation expands within one industry while contracting in another” (p. 10). Additionally, cross-lagged panel models are run separately by sex, using sex-specific occupation-level variables. Two basic equations are estimated for each CPS year-pair from 1984 to 1991, using 1991 as an endpoint in each case: one with the proportion of females as the dependent variable and wage in the previous period as a predictor, and other that predicts wages from the earlier proportion of females. In both cases, controls include the lagged dependent variable. Results show that jobs with a higher percentage of females at one point have slower wage growth for both men and women in the ensuing years. No support for the reverse causal order is found.

England et al. (2007) improve on previous studies by using a fixed-effects model with lagged independent variables, therefore giving a more appropriate treatment to the omitted variable bias problem associated with unmeasured characteristics of occupations. They also pool multiple years of data—instead of using just pairs of years—from the 1983-2001 CPS. No evidence is found for the queuing theory, and there is only a slight effect of earlier sex composition on later wages.

Levanon et al. (2009) advance over past research by combining a longer period of analysis—US Census data from 1950 to 2000, therefore expanding Snyder and Hudis (1976) time span. They use a fixed-effects model with lagged independent variables, and detailed units of analysis obtained from cross tabulations of IPUMS harmonized occupational and industrial categories. Results reveal that female proportion has a negative effect on the levels of reward, consistent with the prediction of the devaluation view. Results are robust to changes in the model, to the division of the data into period groups of decades and to the adoption of alternative occupational classification systems. No support for the queuing theory is found, except in the 1950 decade.

Busch (2018) shows that the negative effect of occupational feminization on wages is heterogeneous, varying by period, by occupational wage groups, and by gender. Using US Census data from 1960 to 2010 and a similar methodology to Levanon et al. (2009), he finds that the devaluation effect is mostly observed between the 1960s and the 1980s, is stronger for women than for men, and is located until the fourth quintile of the wage distribution. Furthermore, he also verifies that effect sizes have decreased in time for the second, third and fourth wage quintiles. In the top fifth, however, no devaluation effect is observed in any period or for either gender. The author also argues that the literature has neglected the theory behind the causal relationship proclaimed by the devaluation hypothesis, making it necessary, first, a better understanding of the mechanisms that lead to an occupation being stigmatized and devalued.

One of the mechanisms behind the influence of gender composition on occupational pay is cultural devaluation of work done by women (Catanzarite, 2003). Busch (2018) emphasizes two underlying mechanisms governing the way in which feminization leads to stigmatization and devaluation: first, the formation of occupation gender stereotypes, given the share of women within occupations; second, the interaction of gender stereotypes with a cultural bias against the value of female labor. These processes may operate differently in alternative historical contexts and over time.

Outside the United States evidence on the relationship between the sex composition of a job and its wages also speaks in favor of the devaluation theory, but it is concentrated in developed countries. Longitudinal analysis with a similar methodology than the aforementioned studies are available for Switzerland (Murphy and Oesch, 2016), Germany (Murphy and Oesch, 2016; Leuze and Strauß, 2016), UK (Murphy and Oesch, 2016; Brynin and Perales, 2016; Perales, 2013), Sweden (Grönlund and Magnusson, 2013; Magnusson, 2013), Netherlands (De Ruijter et al., 2003) and Spain (Polavieja, 2008, effect non-significant).

Although the devaluation is a well-established fact for the United States and for some European economies, scarce evidence is available for developing countries. To the best of our knowledge, no study on a longitudinal basis has ever been conducted for Brazil for testing the devaluation and queuing theories, nor cross-sectional analysis. The next section aims to give an overview on gender pay gap and occupational segregation in Brazil.

### 3. GENDER SEGREGATION AND PAY GAP IN BRAZIL

In Brazil, women's labor force participation began rising in the mid-20th century. According to Census data (IBGE), in 1950, 13.6% of women above ten years old were engaged in the labor market. This rate rose gradually during the next 20 years, reaching 18.5% in 1970. From the 1970s to the 2000s it speeded up, reaching 32.9% in 1991 and 44.1% in 2000.

This process resulted from demographic, cultural, and social changes (Bruschini, 2000). The drop in fertility rates, the increase in life expectancy, the postponement of marriage and child-bearing, and changes in cultural norms about women's roles were key-factors (Bruschini et al., 2003). Industrialization, urbanization, and the rapid economic growth of the 1970s created job opportunities for women. The economic crisis of the 1980s was another relevant factor because it decreased family income, pushing married women to the labor market to supplement it. It is also worth mentioning that in 1961 the education of teachers for the beginning grades of primary school—known as “Habilitação Específica para o Magistério”<sup>2</sup> and attended by most women—, was recognized as a secondary course. This legal change enabled them to compete for college places. From this moment on, female presence in higher education has grown, and the gender gap in education has narrowed until it reversed in the 1980s (Beltrão and Alves, 2009). Higher levels of schooling increased both job offers for women and their opportunity cost of not participating in the labor market.

From the very beginning, women entered the workforce in a markedly segregated manner, mainly in occupations in the service sector. In 1981, the activities that most employed women were services (with 31.8% of them), followed by agricultural (19.8%) and social (16.6%), according to the Brazilian National Household Sample Survey (PNAD in Portuguese) microdata tabulated by (Bruschini, 1994).

During the 1980s and 1990s, women advanced in commerce, social activities, and administration sectors. According to Wajzman et al. (1998), the most common occupations for them at that time were: self-employed merchandise traders (for example, street vendors of cosmetics); paid domestic service (registered or not); public-sector employees in social activities (nurses, teachers, cleaning staff and cooks); and public administration employees (such as administrative aides and clerks, cleaning staff and personnel).<sup>3</sup> This diversification was reflected in occupational segregation measures. Between 1989 and 2001, King (2009) found a drop of almost 9% in the dissimilarity index<sup>4</sup>—a traditional measure of occupational segregation—, calculated based on a harmonized code of 356 occupations.

In the 2000s, women's labor force participation rate increased more slowly: from 44.1% in 2000 to 48.9% in 2010, according to Census data. Occupational diversification among women, however, continued. Sex segregation calculated using the Karmel-MacLachlan index<sup>5</sup> based

<sup>2</sup>Specific Teaching Qualification Course.

<sup>3</sup>Wajzman et al. (1998) analysis comprised the population living in urban areas, aged between 25 and 50 years old, working at least 20 hours per week in non-agricultural occupations, not attending school, and with positive labor income.

<sup>4</sup>The dissimilarity index is calculated as  $D = \frac{1}{2} \sum_{j=1}^J |m_j - h_j|$ , where  $J$  is the total number of occupations,  $m_j$  is the  $j$ -th occupation's participation in the total number of women and  $h_j$  is the analogous for men.

<sup>5</sup>The Karmel-MacLachlan index is defined as  $2a(1-a)D$ , where  $a$  is the proportion of men in total labor force and  $D$  is the dissimilarity index.



on the one-digit industries' classification fell 4.8% between 1995 and 2004 (Salas and Leite, 2007). The sector that most contributed to this fall was the manufacturing industry, where female participation rose from 27.3% in 1995 to 37.1% in 2004. Salardi (2016) shows that although segregation increased slightly in the early 1990s, it decreased substantially from 1995 to 2006.

Although women raised their participation rate and entered new career fields over the past decades, overall sex segregation remained high in Brazil. Also, in the most recent period, it reversed the downward trend. Relying on indices of segregation for two-digit occupations, Botassio and Vaz (2020) find that sex segregation increased between 2006 and 2015. The dissimilarity index, for instance, was 0.447 in 2015, compared to 0.417 in 2006. These values represent the proportion of women (or men) who would have to change occupations in order to eliminate segregation.

Along with the integration of women into the labor market, gender wage gaps have been shrinking in Brazil. According to PNAD data, women's average monthly income in the main job was 47% lower than men's in 1995. This difference dropped continuously and reached 28% in 2015. More recent data from a different database—PNAD Contínua—shows that the sex differential, this time considering income from all jobs, declined from 26% in 2012 to 22% in 2019.

The narrowing gender wage gaps are to a great extent associated with the evolution of women's endowments for the labor market and the decrease in segregation. Haussmann and Golgher (2016)<sup>6</sup> find an increase in the female labor supply in the intensive margin, and hence a decrease in the work-hour gap between sexes. Besides, women's educational levels have risen more rapidly than men's. Madalozzo (2010) shows that occupational segregation decreased for careers traditionally dominated by men. Given that this type of occupation tends to pay better wages, this fact positively influenced the convergence of wages between the sexes. However, women continue to dominate occupations previously considered female (Madalozzo, 2010). Thus, gender differences in occupational distribution still play a role in explaining the pay gap. Empirical studies based on cross-sectional data have found evidence of this relationship.

Barros et al. (1997) study the low-educated workers from the São Paulo Metropolitan Area, the richest urban area in Brazil. The occupations of these workers are aggregated in 19 groups according to their similarities. There are substantial differences in occupational distribution by sex. For instance, the two most important groups for women—domestic services and clothing sector—represent about 48% of female employees, but only 3% of the male ones. Furthermore, female occupations are not only different from males but also paid less. There is a negative relationship between the feminization degree of an occupational group and its income level. Finally, the authors show that the differences in occupational distributions between men and women explain a relevant part of the gender wage gap amongst low-educated workers. This difference would decrease by 33% if men and women had the same occupational distribution.

Ometto et al. (1997) compare the evolution of occupational sex segregation in São Paulo and Pernambuco using PNAD data from 1981 to 1990 (excepting 1982 and 1986). These states are

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<sup>6</sup>The authors analyze gender wage gaps in the Brazilian labor market using a hierarchical model based on the Age-Period-Cohort approach. They apply it by building synthetic cohorts using PNAD from the years of 1992, 1997, 2002, 2007 and 2012. Results show that for all cohorts there are significant gender wage gaps in favor of men, after controlling for period, individual, employment, geographical and educational variables. Besides, there is stability in the unexplained part of the gender wage gap across cohorts. Re-estimating models dropping controls shows that women's endowments positively influence wages relative to men. The authors conclude that gender wage gaps are shrinking due to the evolution of women's endowments for the labor market and the decrease in labor market segregation. However, a sizable proportion of the wage gap continues to be unexplained, possibly due to discrimination. Examining the data grouped by cohort and schooling level shows similar wages for individuals with higher education, indicating a gender wage convergence for more qualified workers.

amongst the most populous in Brazil but differ in productive structure complexity and income level. The authors show that, in both cases, women remain concentrated in low-prestige occupations that represent an extension of household tasks, such as domestic workers, seamstress, laundry workers, cooks, and primary school teachers. The dissimilarity index, calculated based on a three-digit classification of occupations, shows a high degree of overall segregation, which has declined continuously in São Paulo between 1981 and 1990 but has increased considerably in Pernambuco.<sup>7</sup> Earning equations are estimated with dummies for sex and the sex composition of occupations as explanatory variables and controlling for schooling, age, activity sector, position in the occupation, and if the worker is legally employed. Results show that belonging to a male occupation raises hourly income by 24% in São Paulo and 21% in Pernambuco compared to belonging to a female or an integrated occupation.<sup>8</sup>

In another study based on the same period and regional comparison, Hoffmann et al. (1999) show that differences in observable endowments and in occupational distribution due to those endowments explain an irrelevant part of the gender pay gap in São Paulo. Instead, if female productive attributes were paid as well as men's, women's earnings would increase by 30% in the first half of the 80s. Their earnings would also increase substantially—between 15.50% in 1988 and 37.05% in 1984—if differences in the occupational distribution not explained by differences in endowments disappeared. In Pernambuco, however, occupational segregation plays a minor role in explaining the pay gap. Most of it is due to differences in the returns to female endowments compared to men's.

Madalozzo et al. (2015) use the dissimilarity index and men's and women's percentages in an occupation as a base to classify it as integrated, predominantly female, or predominantly male. The authors then estimate Mincerian equations separately by sex and occupation type, using Heckman's procedure to cope with sample selection bias. Finally, the Oaxaca-Blinder methodology (1973) is applied to decompose the gender wage gap into two parts. The first one is related to differences in men's and women's observable characteristics and is called the "explained" part of the gap. The other is attributable to different returns to the characteristics between the two groups—the unexplained part, often attributed to discrimination. Results show the importance of occupational segregation for understanding the wage gap. Women earn less than men in predominantly female and integrated occupations but have higher wages in male occupations. In female occupations, both components of the Oaxaca-Blinder decomposition contribute to the wage differential, although the unexplained one accounts for 87% of it. When it comes to integrated occupations, however, the gap is entirely associated with differences in returns. In male occupations, female characteristics are also paid less. However, since women are, on average, much more educated than men in these occupations, the net effect is positive for them. If workers were paid according only to their observable characteristics, women would earn 16.6%, 17.8%, and 22.6% more on female, male and integrated occupations, respectively.

Although these studies present evidence on the relationship between occupational segregation and wages, they do not allow us to conclude anything about the causal relationship between an occupation's wage level and its sex composition. This article aims to bring evidence on this relationship for Brazil.

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<sup>7</sup>Between 1981 and 1990, the dissimilarity index passed from 0.6349 to 0.5902 in São Paulo and from 0.5901 to 0.6506 in Pernambuco.

<sup>8</sup>These values are the averages of the 1981-1990 effects.

#### 4. DATA AND METHODS

##### 4.1. *Data and unit of analysis*

Data for our analysis come from the 1980, 1991 and 2000 editions of the Brazilian Population Census, conducted by the Brazilian Institute of Geography and Statistics (IBGE). 1980 data refers to a 25% sample of the population, whereas 1991 and 2000 data are a 10% fraction of the population in municipalities with estimated populations greater than 15,000 and 20% in the remaining municipalities.

Occupational information provided by respondents must be coded to allow for statistical treatment. Until the 1991 Census, IBGE used its own classification for workers' occupations. At the time of censuses, i.e., every ten years, IBGE reviewed its classification, adding or eliminating occupations, to reflect the changes observed in the labor market.

Between 1980 and 1991, IBGE's classification suffered only minor changes. Before 1980, however, it was substantially different. Neither IBGE nor any other research institute or university has, up to date, proposed a harmonization that would make the occupational categories adopted in the 1970 and 1980 Censuses comparable.

Because IBGE's nomenclature didn't have descriptions and presented low international comparability, it was abandoned in the 2000 Census. From then on, a new classifying system, based on the 1988 International Standard Classification of Occupations (ISCO-88), has been adopted. As this classification represented a major change to the one used previously, in the 2000 Census, occupations have also been coded according to the classification used in the 1991 Census to allow the assessment of existing differences. As a result, 1980, 1991, and 2000 Censuses have comparable occupational categories. For this reason, our analysis comprises this time span.

Our units of analysis are occupations that are followed over 1980, 1991, and 2000 Censuses. Therefore, variables included in the models are characteristics of workers in the different occupations in each year.

IBGE's occupational classification suffered some changes between 1980 and 1991 because new occupations have been included and others have been collapsed. So, by combining occupations and dropping a few, we were able to construct a set of 358 categories that are consistent for the entire period (1980 to 2000). Procedures to make the occupations compatible are described in Appendix A. Furthermore, up to 1991, the Census data concerning employment and income was related to a worker's labor activities during the previous year. In 2000, the reference period changed to the previous week. For our purposes, this is not an issue for two reasons: i) for the vast majority of workers, his/her occupation in the 12 months before the survey is the same as the occupation in the previous week, and ii) for those whose occupation has changed, it could impact only analysis for individuals, but not when considering longitudinal analysis by occupations.

Military occupations<sup>9</sup> were excluded from the sample.<sup>10</sup> We also excluded other occupations not elsewhere classified, undefined and undeclared occupations.<sup>11</sup>

<sup>9</sup>Military occupations in IBGE's classification refers to Officers in the armed-forces (861), Enlisted men in the armed-forces (862) and Fire department officers and personnel (863).

<sup>10</sup>In Brazil, until the 1990s, women couldn't join the armed forces. In 1996 legislation allowed them to join the auxiliary army corps as physicians, pharmaceuticals, dentists, vets, and nurses. Only in 2012, they were permitted to serve in combat posts. The Air Force accepted female appliance to officer training earlier, in 2002. In the Navy, however, they are still allowed only in auxiliary tasks.

<sup>11</sup>Considering the employed population, the proportion of people in undefined/undeclared occupations is 3.16% (1980), 3.48% (1991), and 7.26% (2000). There is little difference in the proportion of women considering only undefined/undeclared occupations compared to the proportion of women in the employed population. For example, in 1980, 27.22% of employed people were women, and they were 28.71% of the people in undefined/undeclared



#### 4.2. Variables

Our universe of analysis comprises the civilian labor force aged between 25 and 64 years old working at least 30 hours/week in the main occupation.

Our variables of interest are occupation's median wage and sex composition. Census data provides month income, along with usual hours worked per week in the main occupation, which is the one performed during most part of the previous year. Using this information, we constructed an hourly wage variable for each individual in each census year.<sup>12</sup> Median hourly wages were then computed, separately for men and women, for each occupation in each year. To make 1980 and 1991 values comparable to 2000 ones, we inflated them using the deflator for Census incomes proposed by [Corseuil and Foguel \(2002, p. 7\)](#).<sup>13</sup> The monetary values are in 2000 R\$ (Real—reais—is the Brazilian currency). In our models, median wages are always logarithmized.

Sex composition within occupations was measured as the natural logarithm of female proportion divided by male proportion—which is commonly referred to as logit of the proportion of females. This specification is preferable over female proportion for two reasons: first, it avoids the effect simple proportion takes when it approaches 0 or 1; second, as median wages were also logarithmized, the effect of wages on sex composition, and vice-versa, is given in terms of elasticities ([Levanon et al., 2009](#)).

Control variables used in the models include sex-specific averages that measure an occupation's male or female workers' characteristics. To measure human capital from schooling, we used the proportion of individuals in each occupation with up to four years of college education and the proportion with more than four years of college education (individuals with less than 12 years of education are the basis). For our period of analysis, educational reforms did not change the duration of compulsory schooling nor its division into primary and secondary stages. However, there have been differences in how data were collected across Censuses.<sup>14</sup> To deal with this, we adopted the procedures proposed by [Soares and Lima \(2002\)](#).

Alongside the set of education variables, we included a measure of returns to experience to control for varying skill demands. This is obtained by extracting the potential experience coefficients from wage equations conducted separately for each occupation in each Census year. Only male workers were considered because, as pointed out by [Levanon et al. \(2009\)](#), women's intermittent labor force participation may lead to inaccurate rates of returns to experience.

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occupations. For 1991 these values are 31.7% and 31.5%, and for 2000, 37.6% and 32.81%. For 2000, it is interesting to note that the proportion of men in undefined/undeclared occupations (67.19%) is higher than that observed in the employed population (62.4%).

<sup>12</sup>For the 1980 Census, hours worked in the main occupation were provided only in ranges (30-39, 40-48, and 49+). To overcome this issue, we considered the 1981 Brazilian National Household Sample Survey (PNAD)—also carried out by IBGE—that provides more detailed information on the labor force. First, we filtered the sample so that it represented our universe of analysis, and then we computed (separately by sex) average hours worked in the main job in each of these three ranges. Finally, we attributed this information to our observations in the 1980 Census.

<sup>13</sup>As Brazil underwent several currency changes in the 1980s and 1990s, first, it is necessary to convert monetary values to the same currency. For that, 1980 nominal values were divided by 2,750,000,000,000 and 1991 values by 2,750,000. Then we divided the 1980 values by the deflator 0.00000000006457 and the 1991 values by 0.000075501186024.

<sup>14</sup>In the 1960 to 1991 Censuses, students were asked the level and year of school they were attending. Non-students were asked the highest level of school they have completed and the highest grade they have attended. From these four questions, it was possible to calculate the years of schooling as a discrete quantitative variable. However, there were cases when individuals answered both sets of questions so that it was not possible to distinguish between those who studied and those who did not. Furthermore, incoherent responses generated ambiguities in the calculation of years of study. This problem was solved in the 2000 Census with the inclusion of the question "Do you attend school?" ([Rigotti, 2004](#)).

Specification to estimate potential return to experience follows the empirical Human Capital literature on the role of education on wages.<sup>15</sup>

To control for racial differences, we followed the most common practice in Brazilian labor market studies, which combines the racial classifications of “pardo” (brown or mixed-race) and black into one, as contrasted with white. Individuals who self-classified as “amarelos” (Asians) were excluded from the analysis due to small sample size. The category “indigenous” did not exist in 1980 — people who described themselves as “indigenous” were considered “brown/pardo”. So, individuals who were categorized as “indigenous” in 1991 and 2000 were reclassified as “brown/pardo”.

We also controlled for the proportion of incumbents residing in Brazil’s five macro-regions: North, Northeast, Center-West, Southeast, and South. In 1988 the new Constitution created the state of Tocantins by separation of the north area of the state of Goiás. Goiás belongs to the Center-West region, but Tocantins was considered part of the Northern region—changing the boundaries of the regional division. To maintain the five macro-regions comparable over time, we placed Tocantins in the Center-West region in the 1991 and 2000 Census.

Finally, a control that would be especially important in our model is the proportion of incumbents in each occupation that are not formally employed, e.g., that don’t have a labor contract and do not contribute to the social security system. In case of disease, pregnancy, retirement, etc. these workers do not have the right to any benefit. The 1980 and 1991 Census do not allow the identification of these workers. Therefore, we assume that the proportion of informal workers within an occupation remained stable across decades so that this characteristic is eliminated in the fixed-effect transformation—to be presented in the next section.

### 4.3. *Models and estimation*

To assess the causal relationship between sex composition and median wage, two models must be considered: one with sex composition as the dependent variable, and wage as the interest regressor, to test for the queuing view, and the other with median wage as the dependent variable and sex composition as a regressor, to check the devaluation theory. Without loss of generality, first, let us consider a specification to test the devaluation theory.

For the  $i$ th occupation in year  $t$ , a simple cross-sectional model is

$$\ln(w_{it}) = \alpha + \beta_1 P_{it} + \mathbf{x}_{it} \beta' + \varepsilon_{it}. \quad (1)$$

The dependent variable is the logged median wage per hour, and  $P_{it}$  is the logit of the proportion of females into the occupation. Both  $\mathbf{x}_{it}$  and  $\beta$  are row vectors. The vector of covariates  $\mathbf{x}_{it}$  includes two variables for education—the proportion of high school level workers and of graduated workers (less than 12 years of schooling is the base). It also includes returns to experience, the proportion of blacks (whites as the base), and proportions of workers located in the North, Northeast, Center-West, and Southeast (South is the base). Equation (1) is estimated by ordinary least squares (OLS). A model like (1), with logit of proportion female as the dependent variable and logged median hourly wage as a predictor, is estimated to test for the

<sup>15</sup>The dependent variable is the logarithm of hourly wages. Regressors are eight binary variables for education level (in years): 1-4, 5-8, 9, 10, 11, 12, 1-3 college years, and at least four years of college education, and two variables for potential experience (age-schooling-6) considering a quadratic specification (potential experience and potential experience to square). After the parameters’ estimation, the return to experience (in a given occupation in a given year) is the experience coefficient plus twice the squared experience coefficient multiplied by the average experience (average age – average schooling – 6).

queuing view. Omitted variables in (1) correlated with  $P_{it}$  or  $\mathbf{x}_{it}$  lead to biased estimations when the model is estimated by OLS regression. Using longitudinal data can overcome this issue.

The first empirical studies that have used longitudinal data on a range of occupations to test for the queuing and devaluation theories considered pooled OLS estimations (Snyder and Hudis, 1976; Baron and Newman, 1989; Semyonov and Lewin-Epstein, 1989). The dependent variable in  $t$  was modeled as a function of the interest regressor in  $t - k$ , along with control variables. Using lagged independent variables was important to minimize time-variant bias.

As changes in non-observable characteristics of occupations may lead to changes in both wage and sex composition, another important way to minimize omitted variable bias is controlling for the lagged dependent (endogenous) variable (Karlin et al., 2002; England et al., 2007; Levanon et al., 2009). Due to the inclusion of the lagged dependent variable as a predictor, these models are commonly referred to as cross-lagged panel models or dynamic models.

Using longitudinal data, one important resource for overcoming bias due to omitted time-invariant variables (including those that are not observable) is to use a fixed-effect model. In a fixed-effect model, occupations' heterogeneous characteristics are controlled, i.e., the bias due to time-invariant measurable or non-measurable characteristics within an occupation is eliminated from the analysis. According to Levanon et al. (2009), for this reason, the fixed effect estimator is more efficient in removing omitted variable bias. For notation purposes, to incorporate the occupations' fixed effects,  $\alpha$  should be replaced by  $\alpha_i$  in (1). This allows correlation among fixed effects and the predictors.

Thus, considering fixed effects and the reciprocal lagged dependent variable, (1) turns into

$$\ln(w_{it}) = \alpha_i + \beta_1 P_{i,t-k} + \beta_2 \ln(w_{i,t-k}) + \mathbf{x}_{i,t-k} \beta' + \varepsilon_{it}, \quad (2)$$

where  $k$  is the number of years that the variable is lagged. Equation (2) is used to test the devaluation theory. From the queuing theory point of view, occupations' wage level at one time affects later gender composition. Thus, the specification for testing such theory, considering the fixed and lagged effects discussed previously, is

$$P_{it} = \mu_i + \delta_1 \ln(w_{i,t-k}) + \delta_2 P_{i,t-k} + \mathbf{x}_{i,t-k} \delta' + v_{it}. \quad (3)$$

From (2) and (3),  $\varepsilon_{it}$  and  $v_{it}$  are idiosyncratic errors, assumed to be independent of each other and uncorrelated with predictors, which is the key identification assumption of our model called strict exogeneity.  $\alpha_i$  and  $\mu_i$  are the occupation's fixed effects.

Dynamic fixed-effect models, as represented by (2) and (3), pose difficulties for conventional estimation methods. For instance, when the lagged dependent variable is included as a predictor, the usual fixed effect OLS estimator leads to biased estimations, because  $\varepsilon_{it}$  and  $v_{it}$  are correlated with  $P_{it}$  and  $\ln(w_{it})$  in later periods, violating our identification assumption (Levanon et al., 2009; Wooldridge, 2010). One way to overcome these issues is estimating each model separately for each year, in a structural equation modeling approach, imposing the coefficients for each variable to be the same (Allison, 2005). As we have three cross-sections (1980, 1991, and 2000), we estimate models (2) and (3) separately considering  $t$  for 1991 and 2000 and constraining the coefficients in both equations to be equal.<sup>16</sup> Of course, we cannot estimate the equations for 1980, because we do not have data available for one decade before.

Although (1) and its respective equation for the queuing view do not allow us to establish a causal relationship between the sex composition and wages, they are useful to have an idea of

<sup>16</sup>This same procedure is used in England et al. (2007) and Levanon et al. (2009) and can be estimated in SAS using PROC CALIS. For details see Allison (2005).

the correlation (either direction and magnitude) between these variables after controlling other effects (predictors). Equations (2) and (3) allow us to test for causal effects in both directions.

An estimation procedure should be considered to not overestimate the results due to a gap in pay within-occupations (Levanon et al., 2009). For instance, consider (2) to test the devaluation theory, where median wage per hour is the median wage per hour in the occupation considering both men and women, and the predictors are also calculated considering both sexes. Suppose that all men receive the same wage, and all women receive another value (equal among them), and different occupations have different proportions of women. The sex wage gap comes from sex differences in pay within-occupation and, thus, the coefficient for the sex composition is affected by the within-occupation gender gap in pay. Pooling data for both sexes is not a good idea because it would superestimate the coefficients (Levanon et al., 2009). According to the devaluation theory, men and women are negatively affected by working in a predominantly female occupation. A similar argument holds for the queuing theory. Then, when estimating our models, the dependent variable and predictors are calculated separately for women and men and by occupation,<sup>17</sup> and the models are run separately by sex to lead to the desirable hypothesis testing of each theory.

## 5. RESULTS

To provide context, we start with descriptive statistics on the main variables of interest. Table I provides a detailed comparison of the profiles of men and women in the three years of analysis. Women are, on average, more educated than men, although their advantage in total schooling years decreased from 1.82 years in 1991 to 1.53 in 2000. There are between six (in 1980) and eight (in 1991) percentage points more women than men with at least 12 years of education, which corresponds to having attended higher education.

Women are on average younger than men, although the difference in age means declined from 1.64 years in 1980 to 0.96 in 2000. They are also more often part-time employed and much more frequently unmarried. This profile is in line with previous studies that show that many women leave the labor market after starting a family (see Kleven et al. (2019b) for Denmark, Kleven et al. (2019a) for a cross country comparison, and Hecksher et al. (2020) for Brazil).

Concerning the spatial distribution of workers, women are more concentrated than men in urban areas, although this difference declined. In 1980 87% of female workers were in urban areas, against 69% for men, a difference of 18 p.p. In 2000 both groups raised their participation in urban areas, and the difference dropped to about 9 p.p. (Bruschini, 2007) shows that Brazilian women are more frequent than men in precarious job positions, such as in unpaid work and work for production for self-consumption. These jobs are predominant in the agricultural sector, in small farms, ranches, or dwellings in the outskirts of towns, explaining why paid female workers—the focus of this study—are more concentrated in urban areas.

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<sup>17</sup>Of course, within an occupation, the values of the variables sex composition and return to experience (as explained in Subsection 4.2) are the same, either in the female or the male equation.

TABLE I  
DESCRIPTIVE ANALYSIS: WORKERS' PROFILE BY GENDER AND YEAR

Variable	1980			1991			2000		
	Female	Male	Diff	Female	Male	Diff	Female	Male	Diff
Prop by Sex	0.25	0.75	-0.50***	0.28	0.72	-0.45***	0.29	0.71	-0.42***
Age (mean)	37.43	39.07	-1.64***	37.47	38.94	-1.47***	38.39	39.35	-0.96***
Wage/hour (mean)	3.07	4.23	-1.17***	3.06	3.78	-0.72**	3.68	4.50	-0.82*
Wage/hour (median)	2.40	3.04	-0.64**	2.13	2.50	-0.36 <sup>†</sup>	2.57	2.93	-0.36 <sup>†</sup>
Prop. Full-Time	0.83	0.98	-0.15***	0.82	0.96	-0.15***	0.80	0.95	-0.14***
Prop. in North	0.03	0.05	-0.01**	0.04	0.06	-0.02***	0.05	0.06	-0.01***
Prop. in Northeast	0.24	0.25	-0.00ns	0.23	0.25	-0.02ns	0.22	0.23	-0.01ns
Prop. in Southeast	0.52	0.47	0.05**	0.50	0.46	0.04***	0.49	0.46	0.03**
Prop. in South	0.15	0.17	-0.02***	0.16	0.17	-0.00ns	0.17	0.17	-0.00ns
Prop. in Centerwest	0.05	0.07	-0.01***	0.07	0.08	-0.01**	0.08	0.08	-0.00 <sup>†</sup>
Schooling (mean)	5.60	3.98	1.63***	7.07	5.25	1.82***	8.07	6.54	1.53***
Prop. Scho. 0	0.21	0.28	-0.07***	0.12	0.19	-0.07***	0.06	0.09	-0.04***
Prop. Scho. 1 to 4	0.39	0.48	-0.09***	0.33	0.42	-0.09***	0.26	0.34	-0.09***
Prop. Scho. 5 to 8	0.12	0.11	0.01ns	0.17	0.18	-0.01ns	0.23	0.25	-0.03**
Prop. Scho. 9	0.01	0.01	0.00***	0.01	0.01	0.00*	0.02	0.02	0.00ns
Prop. Scho. 10	0.01	0.01	0.00**	0.02	0.02	0.00 <sup>†</sup>	0.03	0.03	0.00ns
Prop. Scho. 11	0.13	0.05	0.08***	0.19	0.11	0.09***	0.23	0.16	0.08***
Prop. Scho. 12	0.01	0.00	0.00***	0.01	0.01	0.00***	0.01	0.01	0.01***
Prop. Scho.13 to 15	0.09	0.04	0.06***	0.12	0.05	0.07***	0.12	0.06	0.06***
Prop. Scho. 16	0.01	0.02	-0.00ns	0.03	0.02	0.00ns	0.03	0.03	0.01ns
Prop. Scho. 17 or more	0.01	0.01	0.00ns	0.01	0.01	0.00ns	0.02	0.01	0.00ns
Prop. White	0.59	0.58	0.02ns	0.58	0.54	0.04***	0.61	0.57	0.04***
Prop. Black	0.41	0.42	-0.02ns	0.42	0.46	-0.04***	0.39	0.43	-0.04***
Prop. Married	0.45	0.74	-0.29***	0.57	0.83	-0.26***	0.49	0.61	-0.11***
Prop Urban	0.87	0.69	0.18***	0.90	0.75	0.15***	0.92	0.83	0.09***
Return Experience		0.01			0.01			0.01	
Observations (occup.)	355	358		354	358		353	357	

Note: \*\*\* p<0.001; \*\* p<0.01; \* p<0.05; <sup>†</sup> p<0.1; ns not significant; wage in 2000 R\$; age and schooling in years. Elaborated by the authors based on Brazil Census Data from 1980 to 2000, with individual sample weights.

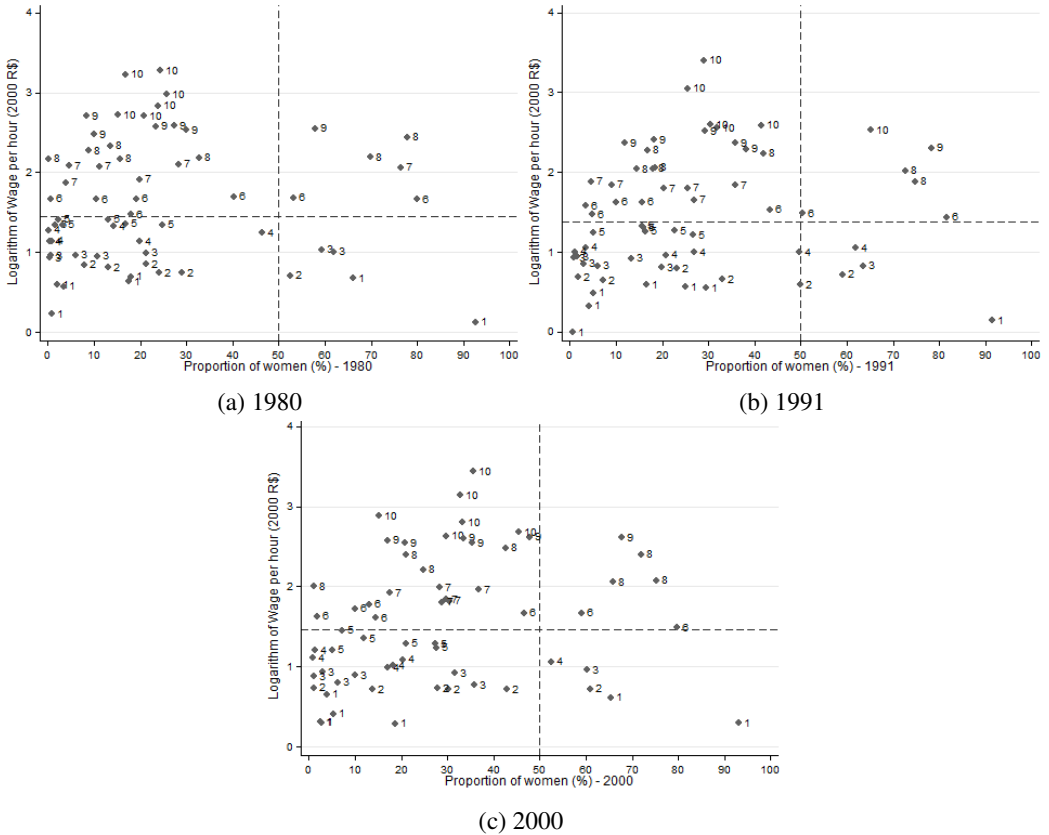


FIGURE 1.—Scatter plot of proportion of women against logarithm of wage per hour. Brazil: (a) 1980, (b) 1991, and (c) 2000. Source: Elaborated by the authors based on Brazil Census Data from 1980 to 2000.

Regarding pay, women have a lower mean hourly wage compared to men, although the gap between them decreased. In 1980 they received on average 72% of men's income, and in 2000 it reached 82%. For both groups, median wages are lower than means due to right-skewed wage distributions, although the asymmetry is stronger in male distribution.

Figure 1 presents the distribution of groups of occupations<sup>18</sup> according to the logarithm of wage per hour and the proportion of women for 1980 (panel a), 1991 (panel b), and 2000 (panel c). The number in each dot indicates the tenth of the wage distribution to which the occupational group belongs. The horizontal dashed line indicates the median wage.

For all years the pattern is the same. For example, in 2000, from the 64 occupational groups, only six were in the first quadrant (which combines a high proportion of women with above-median wages). Indeed most of the occupational groups with higher pay (first and second quadrants) were male predominant. The bottom-right observation (for the three years) is the group associated with domestic services (wage per hour below the first decile and proportion of women above 90%).

We start our regression analysis by the cross-sectional models estimated by OLS (1). Table II presents these results in two panels: Panel A shows results for the devaluation view model

<sup>18</sup>The 358 occupations are classified in 64 groups, according to IBGE's own classification.



and Panel B for the queuing. The models are estimated separately by sex and year (1980, 1991, and 2000). Further, pooled models by sex are also estimated.

From the devaluation theory panel in Table II, for men, the coefficient associated with sex composition increased from  $-0.064$  to  $-0.045$  in the period. On the other hand, for women, it decreased from  $-0.023$  to  $-0.039$ . As the dependent variable is the logarithm of wage per hour, these values are the elasticities of wages in relation to the ratio of the proportion of females by proportion of males (odds).<sup>19</sup>

As the devaluation view predicts that a higher proportion of females negatively affects the wages, either for women or men, i.e., the coefficient associated with the sex composition should be negative, we may have found some evidence for this theory. In methodological terms and considering the same period, the closest work for comparison is [Levanon et al. \(2009\)](#). For the United States, from 1980 to 2000, the authors found similar results to ours. However, in our analysis, the absolute effect of the sex composition on wages increased for women and decreased for men. [Levanon et al. \(2009\)](#) found that this effect remained stable for women and increased for men. Also, the coefficients' trends for Brazil from 1980 to 2000 are more similar to the results found for the US from 1950 to 1970.

Panel B of Table II presents the estimated coefficients for the queuing view. For women, the coefficient associated with the logarithm of wages per hour varied from  $-1.04$  (1980) to  $-1.09$  (2000), with a maximum value in 1991 ( $-0.57$ ). For men, it increased from  $-3.63$  to  $-1.32$ . Similar to the previous results, the significant and negative coefficients, as predicted by the queuing view, require further analysis to identify causality.

The cross-sectional models presented in Table II do not allow us to establish a causal relationship between the proportion of females and median wages but give some idea on the correlation between these two variables. For both devaluation and queuing theories, either for women or men, for all years of analysis and the pooled model, the correlations between sex composition and the logarithm of wage per hour are negative and statistically significant at different levels (0.1%, 1% or 5%). Additional analysis must be carried out to lead to conclusions supporting either the devaluation or the queuing theory for Brazil.

Table III presents results from the fixed-effect models, estimated separately for women and men. In Panel A, median wages are a function of the lagged logit proportion of females, as predicted by the devaluation theory (2). Panel B model, instead, considers that previous wages affect the logit proportion of females, in line with the queuing view (3). In both cases, models are estimated without controls (Model 1) and with lagged controls for region, education, race, and experience (Model 2) plus the lagged dependent variable (Model 3).

Model 3 is used to test the causal relationship for both theories because it corresponds to the complete model with controls and the lagged dependent variable. The devaluation thesis predicts negative coefficients associated with the lagged logit proportion of females. From Panel A of Table III, however, we obtained positive coefficients—although non-significant—for both men and women regressions. Thus, there is no evidence for devaluation. The queuing view predicts the coefficient associated with the lagged median wages to be negative. As can be noticed in Panel B, results also do not support this assumption, either for women or men.

In general, contradicting the expectations found by the cross-sectional models, there is no support neither for the devaluation theory nor for the queuing view, either for women or men.

<sup>19</sup>A more detailed interpretation of the parameters of the regressions does not contribute to the discussion proposed in this article. But for those interested in quantitative analysis, considering men in 2000, an increase of 10% in the odds decreases the wage per hour by 0.45%. For an occupation with 50% of females, an increase of 10% means 2.4 p.p.

TABLE II

COEFFICIENTS FOR LOGIT OF PROPORTION FEMALE PREDICTING MALE AND FEMALE LOG OF MEDIAN HOURLY WAGE (PANEL A) AND FOR LOG MEDIAN FEMALE AND MALE HOURLY WAGE PREDICTING LOGIT OF PROPORTION FEMALE (PANEL B) FROM CROSS-SECTIONAL OLS MODELS. BRAZIL, 1900, 1991, AND 2000

	Women				Men			
	1980	1991	2000	Pooling	1980	1991	2000	Pooling
PANEL A. Devaluation Theory								
Logit prop. Female	-0.023*	-0.023*	-0.039***	-0.030***	-0.064***	-0.057***	-0.045**	-0.057***
	(0.010)	(0.012)	(0.010)	(0.008)	(0.007)	(0.011)	(0.014)	(0.010)
R-squared	0.844	0.821	0.832	0.815	0.871	0.839	0.820	0.828
Observations	355	354	353	1062	355	354	353	1062
PANEL B. Queuing Theory								
Median wage per hour (logarithm)	-1.039*	-0.574*	-1.087***	-0.904***	-3.631***	-1.687***	-1.323**	-2.037***
	(0.412)	(0.269)	(0.238)	(0.225)	(0.463)	(0.342)	(0.448)	(0.416)
R-squared	0.133	0.163	0.174	0.134	0.377	0.274	0.258	0.269
Observations	355	354	353	1062	355	354	353	1062

Note: \*  $p < 0.05$ ; \*\*  $p < 0.01$ ; \*\*\*  $p < 0.001$  (two-tailed test). Standard errors in parentheses. Controls for region, education, race and experience. Elaborated by the authors based on Brazil Census Data from 1980 to 2000.

TABLE III

COEFFICIENTS FOR LAGGED LOGIT OF OCCUPATIONS' PROPORTION FEMALE PREDICTING LOG OF LATER (MALE OR FEMALE) MEDIAN HOURLY WAGE (PANEL A) AND FOR LAGGED LOG MEDIAN (FEMALE OR MALE) HOURLY WAGE PREDICTING LOGIT OF PROPORTION FEMALE (PANEL B), USING POOLED LONGITUDINAL DATA, 1980-2000.

	Women			Men			
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3	
PANEL A. Devaluation Theory							
Lagged logit prop. female		-0.00	0.03	0.02	0.01	0.01	0.00
		(0.02)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
Lagged median wage per hour (logarithm)				-0.20 <sup>†</sup>			-0.37 <sup>†</sup>
				(0.11)			(0.20)
Lagged controls for region, education, race, and experience	0	1	1	0	1	1	1
PANEL B. Queuing Theory							
Lagged median wage per hour (logarithm)		-0.01	0.18	0.15	0.24	0.27	0.19
		(0.09)	(0.11)	(0.11)	(0.17)	(0.19)	(0.19)
Lagged logit prop. female				-0.12**			-0.14**
				(0.05)			(0.05)
Lagged controls for region, education, race, and experience	0	1	1	0	1	1	1

Note: <sup>†</sup>  $p < 0.1$ ; \*  $p < 0.05$ ; \*\*  $p < 0.01$ ; \*\*\*  $p < 0.001$  (two-tailed test). Standard errors in parentheses. 0 NO, 1 YES. Elaborated by the authors based on Brazil Census Data from 1980 to 2000.

There is also no evidence to support these theories when models without controls (Model 1) and with controls but without the lagged dependent variable (Model 2) are considered.<sup>20</sup>

The results suggest that we may have evidence to support a third theory, the equalizing differences. This theory refutes either the devaluation and the queuing view, suggesting that the correlation between sex composition and wages may be spurious. In the last 20 years of the 20th century, Brazil has undergone several structural changes, such as re-democratization

<sup>20</sup>We note that Model 3 for the devaluation theory (either for women or men) at 10% and the queuing theory (either for women or men) at 1% exhibits state dependence, although the expected sign would be positive. State dependence means that the current state (logit of proportion of females or logarithm of median wage per hour) depends on the last period's state, even after controlling the fixed effects and other variables.

in 1984/85, trade opening in the 1990s, and the monetary stabilization in 1994. During this period, women's and men's labor supply has still been strongly unbalanced (less than 30% of workers were female in 2000). The support found to the theory of equalizing differences may suggest that significant non-monetary occupations' amenities have played a considerable role in these differences. Due to gender norms, men are less susceptible to value non-monetary advantages.

## 6. CONCLUSIONS

Although a lot has been said on the relationship between occupational feminization and pay in developed countries, no such effort has been devoted to developing countries. Taking Brazil as an example, in this paper we proposed to test the two main sociological theories on this issue: the devaluation and the queuing theories.

Using Census data from 1980, 1991, and 2000 and dynamic fixed-effect models, there is no evidence to support that occupational feminization affects wages (devaluation theory) or vice-versa (queuing theory). Notwithstanding that a negative correlation between those variables is significant, a third theory can be evoked. The equalizing differences view stands that neither variable causes each other and that gender differences in pay appear due to non-monetary advantages valued differently by male and female workers. As in many other developing countries, in Brazil, female over-representation and informality appear combined in many occupations. Compared to developed countries, women tend to choose more frequently flexible working hours jobs that enable them to reconcile productive work with taking care of children and the house. These arguments endorse the theory of equalizing differences for Brazil in the 1980s and 1990s.

Finally, a remark should be made: the Brazilian labor market and the female labor supply have changed substantially in the first decades of the 21st century. Thus, recent data may not support the theory of equalizing differences anymore. Further studies with more recent data are necessary to a better understanding of this topic and to help design public policies that promote equal pay.

## APPENDIX A: ADDITIONAL TABLES

TABLE A.I

OCCUPATIONAL CLASSIFICATION HARMONIZATION: BACKWARD COLLAPSING OF CATEGORIES.

1980 Census category	Categories merged in the 1991/2000 Census
Others owners (013)	Street vendors (013) + Street market merchants (014) + Other owners (015)
Religious workers (251)	Priests (251) + Self-employed religious workers (252)
Laundresses (805) + Other workers (801)	Chambermaids (801) + Nannies (802) + Cooks (female) (803) + Cleaning personnel (female) (804) + Laundresses (805) + Governesses and butlers (806) + Non-specialized domestic help (807) + Other occupations in domestic service (808)

TABLE A.II

OCCUPATIONAL CLASSIFICATION HARMONIZATION: FORWARD COLLAPSING OF CATEGORIES.

Categories merged in the 1980 Census	1991/2000 Census category
Administrative aides (064) + Office helpers (065)	Administrative aides (or clerks) (064)

TABLE A.III

OCCUPATIONS DROPPED DUE TO NON-CORRESPONDENCE.

1980 Census	1991/2000 Census
Telephone lineman (776)	Census agents (193) Agricultural workers (305) Governesses and butlers (except for domestic help) (816) Hotel master (817) Master of food services (818) Owners in self-owned farming and cattle services not classified above (851) Owners in self-employed services not previously classified (852) Nannies (except for domestic help) (926)

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