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Abstract

In France, in 2014, women's hourly wages were on average 14.4 % lower than men's. Beyond differentials in observed characteristics, is this gap explained by segregation of women in low-wage firms, or by gender inequality within a given firm? To answer that question, we apply the approach of Card, Cardoso, and Kline (2016) on French data to disentangle the role of between-firm (*sorting*) and within-firm heterogeneity (*bargaining*) on the gender wage gap. We use a two-way fixed effect wage model, in which firm fixed effects differ between male and female employees to account for within-firm gender differences in bargaining power and wage policy. We estimate this model with linked employer-employee data covering French private sector from 1995 to 2014. The *sorting* effect accounts for almost 11 % of the gender wage gap, whereas the *bargaining* effect is close to zero. This last result could be related to the protective role of the high French minimum wage level. We have access to very rich administrative data that allow us to recover information on family events. Hence, we can analyze sorting and bargaining effects all along the family life cycle. Our analysis shows that firm effect gap appears clearly around the first childbirth and deepens over the life cycle: in addition to the direct effects of childbirth on wages, mothers also experience wage losses associated to sorting into low-paying firms.

JEL Codes: J31, J71, J16
Keywords: gender wage gap, gender inequalities, linked employer-employee data, two-way fixed effect models, discrimination

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1 Introduction

In spite of the increase in female education and professional experience over the last decades, women still earn lower wages compared to men. In France, in 2014, women's hourly wages are on average 14.4 % lower than men's in the private sector. Part of this gender gap can be attributed to differences in observed characteristics between the two groups. However, large gender wage discrepancies remain once productivity differentials are accounted for: after taking into account seniority, professional experience, age, level of education, occupation, part time work, industry, firm size, and region there is still a 8.4 % unexplained gap between men and women.

Several hypotheses can be put forward to understand this gender wage gap. First, females could obtain a lower wage than comparable male coworkers while doing the same job in the same firm. This situation can occur because of discrimination (Blau and Kahn, 2016, reckon this cannot be completely dismissed) or, more subtly, because women do not bargain their wages as well as men do. As a matter of fact, different contributions have shown that women tend to initiate less negotiation, or to perform poorer when bargaining their own wages (Bertrand, 2011 and the references therein, especially Bowles, Babcock, and Lai, 2007; Bowles, Babcock, and McGinn, 2005; Small, Gelfand, Babcock, and Gettman, 2007 for laboratory studies and Babcock, Gelfand, Small, and Stayn, 2006; Babcock and Laschever, 2003; Greig, 2008; Manning and Saidi, 698 for some (mitigated) field results, see also Rigdon 2012).

A second explanation of the gender wage gap could lie in segregation across industries, jobs, and firms. This is the point of Groshen (1991) who evidences that occupation sorting accounts for a large part of the gender wage gap. Her contribution motivates a large literature on the role of segregation across occupations and establishments in gender inequalities (Barth, Bryson, Davis, and Freeman, 2016; Bayard, Hellerstein, Neumark, and Troske, 2003; Nekby, 2003; Trond Petersen, 1995). Regarding job opportunity discrepancies between men and women in France, Gobillon, Meurs, and Roux (2015) documented the lower access to high-paid jobs for full-time executive women aged 40-45 in French private and public sectors relative to men. They find that the probability for women to get a job at the bottom of the wage distribution (5^{th} percentile) is 9 % lower than for males but 50 % lower at the top of the wage distribution (95^{th} percentile).

Card, Cardoso, and Kline (2016) provide a first assessment of the joint effect of firms sorting and bargaining differences on the gender wage gap using linked employer-employee models. Using Abowd, Kramarz, and Margolis (1999) and Lentz and Mortensen (2010) methods, they disentangle a segregation component of the gender wage gap from unobserved individual productivity, and from other within-firm differences between men and women. More precisely, they build a model to decompose the role of firms in gender earning gap into a bargaining/within-firm effect - the gender wage difference of a given employer after controlling for individual heterogeneity - and a sorting/between firm effect - the wage impact of different distributions of women and men in firms.

The first contribution of our paper is to replicate the decomposition proposed by Card, Cardoso, and Kline (2016) on the French labor market using matched employer-employee data from 1995 to 2014. To our knowledge, this method has only been applied to Portuguese data. We discuss the exogenous mobility condition validity, showing that it is likely to hold in the general population but not for some job occupation subgroups such as executive workers. The comparison between our findings and those from Card, Cardoso, and Kline (2016) is of interest given the existing differences between the French and the Portuguese labor markets. We find the sorting effect accounts for almost 11 % of the gender wage gap, whereas the bargaining effect is very small, and if anything negative. These results differ from Card, Cardoso, and Kline (2016) who find that firms account for 21 % of the average gender log-wage gap, with a 15-20 % sorting channel, and a (positive) 1-6 % bargaining channel; see also Cardoso, Guimaraes, and Portugal (2012). Our second contribution is to relate our results to specific aspects of the French institutional settings. In particular, we argue that the discrepancies between our results and the ones obtained for Portugal can be related to the high level of the minimum wage in France, which is more than twice higher than the one in Portugal in 2016. Indeed, women are more often paid at the minimum wage than men, and a high minimum wage is likely to attenuate the importance of bargaining in the decomposition. We explore this channel by replicating our analysis on subsamples before and after the fast increase in the French minimum wage that occurred around 2005. Our findings are consistent with this mechanism. Furthermore, we take advantage of the availability of detailed firm level variables to better describe the within firm gap. It appears that a non negligible share of the variance of the bargaining effect is actually explained by collective agreements in the context of a three-level (firm, contractual industry and national level) hierarchical wage bargaining system. It appears also that a higher share of workers at the minimum wage decreases the firm bargaining effect.

Card, Cardoso, and Kline (2016) find that the firm and the sorting effects grow dramatically with age, that sorting effects are larger among less educated workers while bargaining effects are more important among highly-educated ones. Our study confirms these findings. Our third contribution is then, to relate explicitly sorting and bargaining sources of inequality to family circumstances of workers. These inequalities are indeed likely to be linked to gender-specific family duties, and their importance may vary over the life cycle. Women take care of children more often than men, which may lead them to choose to work in less-paying and probably more family-friendly firms when their children are young, and to bargain differently with their employers. Related to this, Wilner (2016) identifies child wage penalties in French firms for mothers relative to both non-mothers and fathers, even after controlling for human capital depreciation and unobserved heterogeneity. Childbirths generate sex-specific behavior regarding job and firm mobility. This is a point developed in Albrecht, Bronson, Thoursie, and Vroman (2017) in examining the career dynamics of high-skill women and men in Sweden. They document how firms contribute to the gap that widens between men and women around the age of the first childbirth. They find that nor firm heterogeneity nor mobility rate differences can explain the gender wage gap totally. Instead, they show than men experience faster wage growth than women do, be it when they leave a firm for another or when they stay in a firm. Barth, Kerr, and Olivetti (2017) also focus on the life-cycle wage profiles of men and women, assessing the share of the increase in the gender earnings gap over the life-cycle that comes from between or within establishment inequalities. Interestingly, they point out that the between-establishment gender wage gap component is almost entirely due to married workers.

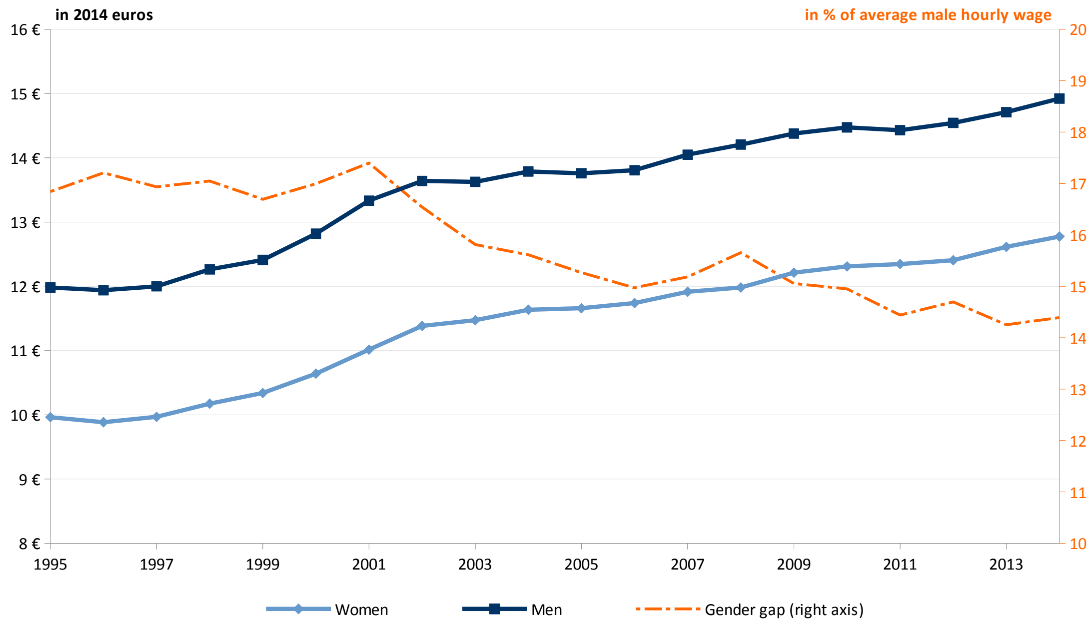
Taking advantage of the richness of our data upon family events, we relate those to both within and between firm gaps along the life cycle. First, we document a sorting effect almost three times bigger for parents than for non-parents. Then, we describe sorting and bargaining effects at different moments in time before and after the birth of the first child of parents. We find that the sorting effect between women and men starts rising just before workers have their first child, clearly widens in the first years after the birth, and strikingly never decreases even thirty years after the child first birth. Our findings indicate that the dynamic approach is important to understand the generation of gender inequalities.

The remainder of the paper is as follows. Section 2 describes the French context and the data. Section 3 presents the model and discusses conditions for identification. Section 4 presents the results, which are further analyzed in the light of family events along worker life cycles in Section 5. Section 6 relates our findings to institutional settings in France, and the last section concludes.

2 French context and Data

2.1 Trends in the Gender Wage Gap

Figure 1: Average hourly wage for women and men since 1995



Source: DADS, *Panel Tous Salariés*. Scope: Metropolitan France. Workers aged less than 16 or more than 65 are excluded. Self-employed farmers, craftsmen, shopkeepers, trainees, apprentices and private household workers are excluded. Note: in 2014, the average hourly wage for private sector employees is 12.8 € for females and 14.9 € for males. The gender gap corresponds to 14.4 % of the average hourly male wage.

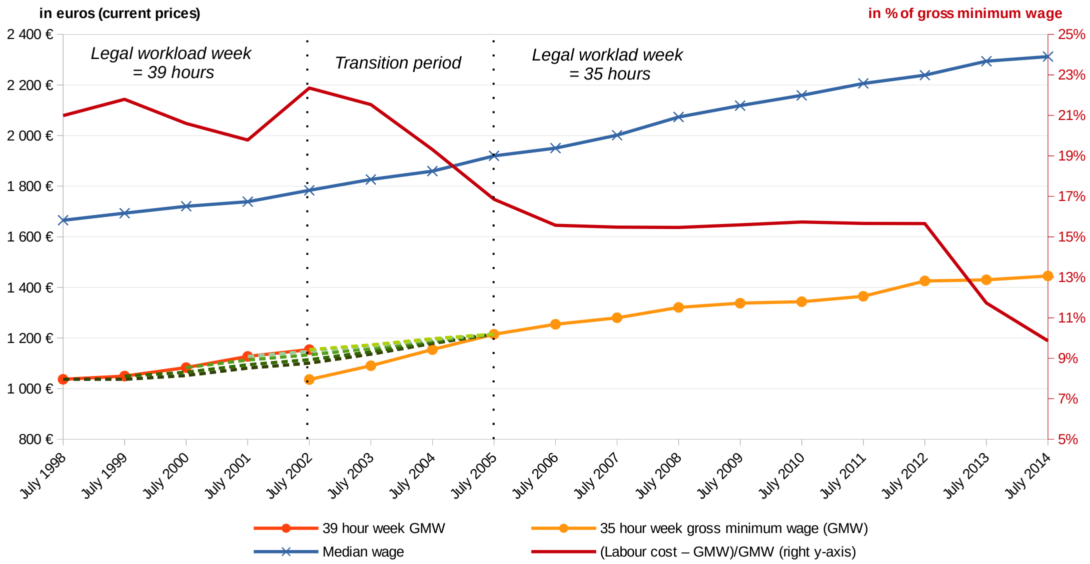
From 1995 to 2014, the gender hourly wage gap in the private sector has remained stable (about 2 €) over the period, but the relative gender wage gap has decreased (16.8 % in 1995 vs 14.4 % in 2014) because of hourly wage gains (see Figure 1). Male and female wages have increased regularly -in particular around year 2000 when workers have benefited from the working time reduction (35 hour week) while monthly wages were held constant. Over the period, participation of French women to the labor force increased from 40.6 to 43.3 %.[1] As in many OECD countries, the gender wage gap stagnated in the recent years despite the high educational achievement of women.

[1]Though the rate of full time workers among the participating women decreased from 70.2 to 66.2 %.

2.2 High minimum wage but contained labor cost

Although the situation of women in France is comparable to the one of other developed countries, the specificity of the French labor market has to be considered when analyzing the gender wage gap. In particular, a key feature is the relatively high level of the minimum wage in France.

Figure 2: Gross minimum and median wages, and relative labor cost at the minimum wage since 1998



Note: Labor cost at the minimum wage includes gross minimum wage and legal social contributions paid by the employer. It is net of social exemptions (Fillon exemptions for instance), and of the tax credit for competitiveness and employment (that applies for workers paid less than 1.6 minimum wage since 2013). The gross median wage is computed by applying worker social contribution rates to net median wage of full-time workers in the private sector.

The national level of minimum wage applies for all employees in all industries, except for apprentices and some workers under 18, who may be paid down to 80 % of the minimum wage. Since 1998, the national minimum wage has grown as rapidly as the median wage in France (figure 2), representing about 62 % of the full-time private sector median wage in 1998 and in 2014.[2] Between 1998 and 2005, the *35 hour working week* laws were gradually implemented.[3] To maintain monthly earnings of workers at the bottom of the wage distribution, monthly guaranteed salaries (GMR) were enforced. As shown by figure 2, these GMR then converged in 2005 to the 35 hour working week minimum wage.

[2]The minimum wage or *Smic* (salaire minimum interprofessionnel de croissance) is adjusted such as to grow at least by half the rate of the purchasing power increase of the hourly gross wage of blue collar workers (SHBO). In addition to the automatic reevaluation rule, the government can pass *coups de pouce* (boosts), often after presidential elections (2007, 2012).

[3]*Loi n° 98-461 du 13 juin 1998 d'orientation et d'incitation relative à la réduction du temps de travail* also called *loi Aubry* and *Loi n° 2000-37 du 19 janvier 2000 relative à la réduction négociée du temps de travail pour les 35h.*

However, the increase in the gross minimum wage has not resulted in a growing labor cost of low-paid workers since firms have benefited from substantial social security exemptions and tax credits for workers paid at the minimum wage or immediately above. Hence, even though the minimum wage remained stable relative to the median, and increased in absolute terms, the cost of hiring a worker paid at the minimum wage dropped from 22 % to 10 % of the gross minimum wage (figure 2, red line, right y-axis). The first decline in minimum wage labor cost in years 2003-2005 is due to the Fillon exemptions that aimed to compensate the rise in hourly wages implied by the transition to the *35 hour working week*.[4] The exemption rate decreases with the wage level and cancels out at 1.6 minimum wage. Another minimum wage labor cost fall can be attributed in 2013 to the tax credit for competitiveness and employment (CICE), which applies for all workers paid up to 1.6 minimum wage.[5]

This combination of a high minimum wage level with a relatively low labor cost is likely to provide an implicit bargaining power for workers at the bottom of the wage distribution, which may drive part of our results since women are over-represented at the minimum wage level.

2.3 A hierarchical wage bargaining system

Another aspect of the French labor market that matters regarding the gender wage gap, and that differs in some aspects from other countries is the French collective wage bargaining system. In France, wages are bargained at different levels: on top of the binding national minimum wage set by the government, wages can be negotiated among industries and firms.

On the one hand, at the industry level, employers' organizations and unions bargain on wage floors for each occupation and for each level of a productivity grid. Firms cannot opt out from an industry-level agreement, which applies then to all wage-earners (unionized or not). More precisely, each firm belongs to a *contractual industry* (*branche conventionnelle*) depending mainly on its activity, sometimes also on its localization. A *contractual industry* boundary is established jointly by employers and unions. There are more than 700 different contractual industries in France but only around 300 cover more than 5,000 workers. For each *contractual industry*, a collective agreement (*convention collective*) sets general rules regarding

[4] *Loi n° 2003-47 du 17 janvier 2003 relative aux salaires, au temps de travail et au développement de l'emploi.*

[5] *Loi n° 2012-1510 du 29 décembre 2012 de finances rectificative pour 2012.*

relations between workers and their employers such as wages, working conditions, hours worked, lay-off conditions, etc. Importantly, it establishes an industry-specific classification of occupations based on worker skills, job requirements, experience, age or qualifications. A wage grid is set for every position of the classification, defining wage floors for all productivity levels in each cell of the classification. *Contractual industries* are legally assigned to open a bargaining process to discuss wage floors at least once a year. However, there is no obligation to reach an agreement. On the other hand, at the firm level, employers and unions bargain on wage increases provided that wages are set above the industry wage floors. More details on the impact on how the national minimum wage interacts with the industry level wage bargaining can be found in Fougère, Gautier, and Roux (2016).

2.4 Data Sources

Our main data come from the merger of two administrative datasets, the *Déclarations annuelles de données sociales* (DADS, Annual Declarations of Social Data) database and the Census database. We use the panel subsample extracted from the exhaustive DADS administrative database constructed by the Insee (the French National Institute of Statistics) for scientific use. This panel is updated every year since 1967 using the wage bill information firms have to fill in annually for each of their employees for payroll and fiscal tax purposes. The filling is mandatory. This panel has had a linked employer-employee structure since 1975: it contains both the firm unique identifier (SIREN), which comes from the French firm register (SIRENE) and the person unique identifier (social security number, NIR). Agriculture sector and self-employment are excluded from the panel, and the public sector has been phased in during the 1980s (public hospital employees in 1984; local and state public service employees in 1988). The DADS panel sample gathers information on individuals born in October in even years, which gives a representative sample of roughly $1/24^{\text{th}}$ of the French employee population. The panel statistical unit is the worker $\times$ firm $\times$ year level: for each worker, we know for which firm she worked at any time, its occupation, and how much she earned.

We use the same wage concept as Abowd, Kramarz, and Margolis (1999) or Postel-Vinay and Robin (2002) and more recently Wilner (2016). The wage is net of social contributions (but before income tax) and precisely corresponds to the wage information sent by the firm to fiscal services for income tax issues. It is therefore of great quality and contains wages and salaries, paid overtime, advantages in kind, all bonuses and indemnities (including shift work) even those paid once a year, and those paid after contract termination when they exceed the conventional level. So this variable is particularly accurate for our analysis as it covers well parts of the wage that can be negotiated (bonuses) by the employee. The sole limitation is that it does not completely cover profit-sharing schemes, but only the part of the remuneration that is directly paid to the employee and not saved. However, we believe that this caveat is limited as profit-sharing schemes stand on average on the whole economy for 3 % of gross earnings (resp. 4 % for executives), whereas bonuses stand for around 17 % of them (resp. 17 %) according to the 2010 Structure of earnings survey.[6] The data also indicates how many hours the employee worked but only since 1995 and only in the private sector. As we take as dependent variable the hourly wage, we focus on private sector employees during the 1995-2014 period.[7]

We supplement the matched employer-employee data with individual information on workers, especially education, number and dates of birth of children using the Permanent Demographic Sample (EDP). The EDP is a large-scale socio-demographic panel gathering information taken from the official publications of the registry office for births, marriages and deaths since 1968, along with census information from 1968, 1975, 1982, 1990 and 1999. Information derived from the annual census surveys (which have replaced exhaustive census since 2000) is also integrated. The sample corresponds broadly to a 4 % survey of the population living in France. Like DADS, Selection to EDP is based on date of birth, and the merger of the two datasets corresponds to around 13 % of the DADS panel.

We also gather additional administrative information using the firm unique identifier available in the DADS files. We use the *File of Approximated Financial Results* (*Fichier approché des résultats d'Esane*, FARE), which provides us with financial information on the firm, such as valued-added, surplus, income statement, and balance sheet items (sales, exports, investments). Especially, we track back firm value-added for years 2014 to 2012. In addition, the firm level data also allows us to identify the main collective agreement in force at the firm level. Some information from these agreements can be extracted from the *Contractual Industries Floor Wages* database (*Base des Minima de Branches*, BMB). In particular, if there is one, we know in each firm for each type of occupation the bottom and top wages negotiated by unions and

[6] Card, Cardoso, and Kline (2016) use a hourly wage calculated as the worker's base salary plus any regular earnings supplements divided by the worker's normal hours of work. Card, Heining, and Kline (2013) resort to total earnings corrected for top-coding at the Social Security maximum. They divided earnings by days worked at full-time jobs (they have no information on hours worked).

[7] The same analysis on the public sector will be the object of a companion paper.

firms.[8] These wages are only available for firms belonging to a *contractual industry* with 5,000 workers or more from 2003 to 2014 so we do not integrate them in our baseline estimation, but we use them to document the within gender firm gap.

Finally, after taking into account sample restrictions due to identification that we detail below, our sample allows to identify bargaining and sorting effects for 912,784 observations, that is 102,048 workers, working in 89,855 connected firms. Descriptive statistics are provided in Tables 1 and 2 for workers, and 3 for firms. Separate description is given for all individuals from the survey (Whole sample) and individuals actually included in the estimation sample (Estimation sample). The structures of both samples are close, with few remarkable points: executives are slightly overrepresented in the estimation sample, such as short-term contract holders, since they may move more often between firms, and such as workers in firms with less than 10 employees because in the estimation procedure those small firms are put together. High value-added firms are overrepresented in the estimation sample, likely due to their longer life.

[8] Firms are coerced to pay a worker at least its *branche* × occupation bottom wage. What we call top wage specifies the minimum wage for the highest productivity workers in the occupation.

Table 1: Descriptive statistics on before and after-estimation worker samples

	Whole sample				**Estimation sample**			
	Male		**Female**		**Male**		**Female**	
Variable	*N*	*Mean*	*N*	*Mean*	*N*	*Mean*	*N*	*Mean*
Net wage (in 2014 €)	945,589	23,399 (26,853)	686,596	16,588 (13,528)	512,325	23,602 (26,588)	400,459	16,979 (14,176)
Net hourly log-wage	945,589	2.52 (0.48)	686,596	2.35 (0.40)	512,325	2.54 (0.49)	400,459	2.37 (0.40)
Age	945,589	39.4	686,596	39.3	512,325	39.2	400,459	39.1
Professional experience	945,589	15.4	686,596	13.6	512,325	14.9	400,459	13.5
Seniority	945,589	5.1	686,596	5.0	512,325	5.0	400,459	5.0
Education #1 (no degree)	945,589	17.1 %	686,596	12.5 %	512,325	15.7 %	400,459	12.0 %
Education #2	945,589	5.9 %	686,596	6.8 %	512,325	5.4 %	400,459	6.7 %
Education #3	945,589	6.5 %	686,596	8.0 %	512,325	7.0 %	400,459	8.5 %
Education #4	945,589	34.2 %	686,596	25.3 %	512,325	32.7 %	400,459	24.7 %
Education #5	945,589	9.9 %	686,596	11.7 %	512,325	10.6 %	400,459	11.8 %
Education #6	945,589	4.8 %	686,596	7.7 %	512,325	5.5 %	400,459	8.3 %
Education #7	945,589	11.3 %	686,596	16.8 %	512,325	11.7 %	400,459	16.4 %
Educ. #8 (master/PhD)	945,589	10.3 %	686,596	11.2 %	512,325	11.5 %	400,459	11.7 %
No child	945,589	47.7 %	686,596	43.4 %	512,325	48.7 %	400,459	43.7 %
1 or 2 children	945,589	42.2 %	686,596	47.7 %	512,325	41.6 %	400,459	47.7 %
3 or more children	945,589	10.1 %	686,596	8.9 %	512,325	9.6 %	400,459	8.6 %

Note: In the entire sample, workers with no degree account for 17.1 % of male observations and 12.5 % of female ones. In the after estimation sample they represent 15.7 % of male observations and 12.0 % of female ones.

Table 2: Descriptive statistics on before and after-estimation worker samples (end)

	Whole sample				Estimation sample			
	Male		Female		Male		Female	
Variable	*N*	*Mean*	*N*	*Mean*	*N*	*Mean*	*N*	*Mean*
Executives	945,589	16.9 %	686,596	10.3 %	512,325	18.3 %	400,459	11.2 %
Clerks	945,578	20.9 %	686,582	22.2 %	512,316	21.8 %	400,448	22.6 %
White collar workers	945,578	12.4 %	686,582	50.6 %	512,316	14.5 %	400,448	49.2 %
Blue collar workers	945,578	49.8 %	686,582	16.8 %	512,316	45.4 %	400,448	17.0 %
Paid hours	945,589	1,585	686,596	1,396	512,325	1,548	400,459	1,393
Part-time job	945,589	11.1 %	686,596	32.9 %	512,325	13.1 %	400,459	33.1 %
Open-ended contracts	443,416	83.4 %	335,948	83.1 %	235,643	78.3 %	193,667	81.8 %
Fixed-term contracts	443,416	7.4 %	335,948	11.4 %	235,643	7.3 %	193,667	10.4 %
Other short term jobs	443,416	9.2 %	335,948	5.5 %	235,643	14.4 %	193,667	7.7 %
Agriculture	945,589	0.9 %	686,596	0.6 %	512,325	1.0 %	400,459	0.6 %
Manufacturing	945,589	27.8 %	686,596	15.7 %	512,325	24.9 %	400,459	14.9 %
Construction	945,589	11.7 %	686,596	1.8 %	512,325	8.4 %	400,459	1.6 %
Trade	945,589	15.5 %	686,596	19.9 %	512,325	14.8 %	400,459	20.1 %
Services	945,589	44.1 %	686,596	61.9 %	512,325	50.9 %	400,459	62.8 %
10 or less worker firms	945,589	14.6 %	686,596	17.6 %	512,325	26.3 %	400,459	29.4 %

Note: In the entire sample, the average number of paid hours in the year is 1,585 for men and 1,396 for women. In the after estimation sample the annual average paid hours amount to 1,548 for men and 1,393 for women.

Table 3: Descriptive statistics on firms, before and after-estimation

	Whole Sample		**Estimation Sample**	
Variable	*N*	*Mean*	*N*	*Mean*
Number of employees	205,267	78.7 (993)	89,855	101.7 (1,494)
Value added before tax (2014)	107,019	5,863 (96,518)	43,590	9,140 (149,607)
Gross operating surplus	107,019	1,236 (38,244)	43,590	2,019 (58,675)
Operating income before tax	107,019	1,063 (36,804)	43,590	1,596 (51,198)
Net overall sales	107,019	21,553 (288,949)	43,590	32,126 (427,112)
Share of exporting firms	107,019	25.5 % (0.44)	43,590	16.9 % (0.375)
Investments	107,019	31,094 (1,000,794)	43,590	51,819 (1,495,052)

Note: Firms in the entire sample have on average 78.7 workers. Firms in the after estimation sample have on average 101.7 workers.

3 Disentangling Bargaining from Sorting effects

3.1 A rent-sharing model

Following Card, Cardoso, and Kline (2016) and Card, Cardoso, Heining, and Kline (2017), at each period t, wages result from a Nash-bargaining between individual i with outside option a_{it} and firm $J(i,t)$. The surplus associated to the job is $S_{i,J(i,t)}$, and the wage resulting from the process is a sum of a_{it} and $S_{i,J(i,t)}$ weighted by a parameter γ reflecting the bargaining power of the worker. γ differs by gender ($G(i) \in \{F, M\}$):

$$w_{it} = a_{it} + \gamma^{G(i)} S_{i,J(i,t)}, \tag{1}$$

which leads to the reduced form equation:

$$w_{it} = \alpha_i + X_{it}' \beta^{G(i)} + \psi_{J(i,t)}^{G(i)} + r_{it}. \tag{2}$$

Card, Cardoso, and Kline (2016) obtain this reduced form equation by specifying the individual outside option as $a_{it} = \alpha_i + X_{it}' \beta^{G(i)} + \varepsilon_{it}$, and the surplus as the sum of three component, one that is fixed over time, $\bar{S}_{J(i,t)}$, one time varying firm component, $\phi_{J(i,t)t}$, and a firm-worker specific component, $m_{iJ(i,t)}$. As a consequence, α_i reflects the individual fixed component, $\beta^{G(i)}$ are sex-specific returns to productive characteristics X_{it}, and $\psi^{G(i)}_{J(i,t)}$ are gender-specific firm effects, that account for firm-specific pay premia, and are directly linked to the gender specific bargaining power $\gamma^{G(i)}$. The residual term, r_{it}, is thus an unobserved heterogeneity term accounting for both the worker's and firm's period specific unobserved heterogeneity as well as worker-firm shocks.

3.2 Sorting and bargaining effects

The gender specific bargaining effects, $\gamma^{G(i)}$, can be recovered from equation (2) applying a Blinder-Oaxaca decomposition to the average gap due to firm specific components:

$$\begin{aligned}\mathbb{E}\left[\psi^M_{J(i,t)} \mid g=M\right] - \mathbb{E}\left[\psi^F_{J(i,t)} \mid g=F\right] = & \underbrace{\mathbb{E}\left[\psi^M_{J(i,t)} - \psi^F_{J(i,t)} \mid g=M\right]}_{\text{(i) Bargaining effect}} \\ & + \underbrace{\mathbb{E}\left[\psi^F_{J(i,t)} \mid g=M\right] - \mathbb{E}\left[\psi^F_{J(i,t)} \mid g=F\right]}_{\text{(ii) Sorting effect}}\end{aligned} \tag{3}$$

The first term (i) reflects the average difference between men and women firm's component if they were working in equal proportions in the same firms (bargaining effect). The second term (ii) describes differences between the average firm effect for women if they were employed in the same firms as men and their actual average firm effect (sorting effect). Note that, as in the case of any Blinder-Oaxaca decomposition, the decomposition is not unique, and the choice of the reference group may not be innocuous.

3.3 Identification and Estimation

The empirical counterpart of equation (2) is a two-way fixed effect model corresponding to an AKM model. Models are estimated separately for men and women, and only for workers employed in companies hiring both genders. As it is usual in this type of model, we group firms with ten workers or less so as to compute the regression on a maximum number of workers. The comparison of gender specific firm-effects requires additional data restriction, so that our estimation is obtained from the set of workers employed in firms belonging to both male and female largest connected sets. In addition to sample restrictions, the OLS estimates of equation (2) are unbiased provided that:

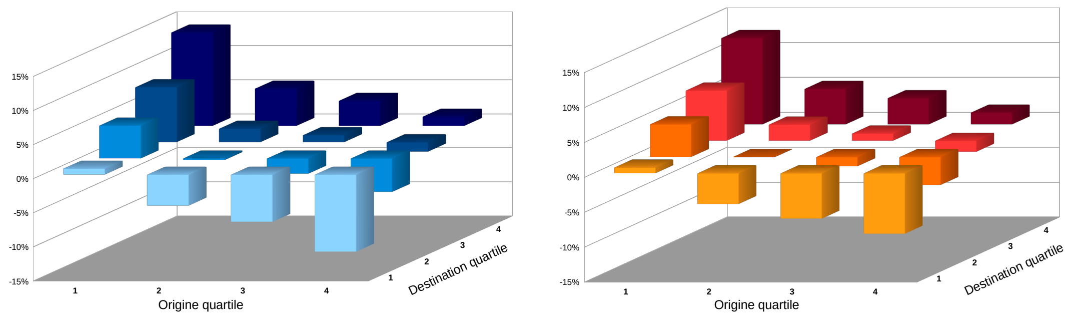
$$\mathbb{E}\left[(r_{it} - \bar{r}_i)\left(\mathbf{1}_{J(i,t)=j} - \frac{1}{T}\sum_{t=1}^{T}\mathbf{1}_{J(i,t)=j}\right)\right] = 0, \quad \forall j \in \{1...J\}. \tag{4}$$

This condition must hold for each firm, j, and states that on average unobserved shocks r_{it} should not depend on the mobility of individuals (one should notice that the condition is only active for firm movers otherwise $\mathbf{1}_{J(i,t)=j} - \frac{1}{T}\sum_{t=1}^{T}\mathbf{1}_{J(i,t)=j} = 0$). In other words, conditional on mobility, the expected effect of individual wage unobserved factors (r_{it}) should not deviate from their average value ($\bar{r}_i$). Given that r_{it} encompasses shocks on worker, firm or match between workers and firms productivity, the exogeneity condition holds if mobility between firms is not correlated with shocks on firm profits, on match surplus, and on individual productivity.

The exogenous mobility assumption is not directly testable from the data. However, following Card, Cardoso, and Kline (2016) we gather elements in line with some of its main predictions. First, wage gains and losses associated with entering or leaving high/low paying firms look symmetric. This is the main message in Figures 3a (men) and 3b (women). The two figures show the average wage evolution for movers according to the average coworkers' wage before and after mobility. Thus, a man moving from a low paying firm (first quartile) to a high paying one (fourth quartile) experiences a 13.7 % wage increase on average. Symmetrically a man going from a high paying firm (Q4) to a low paying one (Q1) can expect a 11.3 % wage drop.

Figure 3: Average wage change for movers conditional on origin and destination firm average wage

(a) Male Workers (b) Female Workers

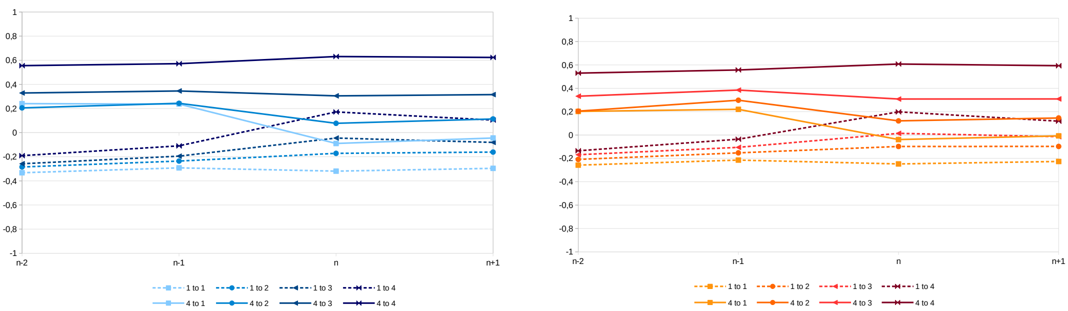


Note: A female worker moving from a firm paying average wages below the bottom quartile (quartile 1) of the wage distribution to a firm above top quartile (quartile 4) gets a average 12.4 % increase in wage. Symmetrically a female going from an above top quartile firm to a below bottom quartile firm can expect a 8.7 % drop in wage.

Besides symmetry, the exogenous mobility condition implies the absence of transitory wage shock driving firm-to-firm mobility of workers. We thus consider the evolution of the residual of the regression of the hourly log wage on individual characteristics (age, seniority, experience, year dummies) for movers. Figures 4a and 4b show the average value of these residuals from two-year before to two-year after mobility. The evolution is broken down according to the average coworkers' wage before and after mobility. The no-transitory wage shock assumption requires that the after-mobility coworkers' wage cannot be predicted by before-mobility wage residual shocks, and reversely the before-mobility coworkers' wage should not be correlated with the after-mobility residual wage trend. From our analysis, this hypothesis is not likely to be rejected for either male or female workers that were working in high paying firms before mobility. The result is perhaps not as clear for workers going from lowest to highest paying firms but we find no evidence of systematic transitory wage shocks correlated with mobility. Based on these elements, exogenous mobility seems to be a reasonable assumption.

Figure 4: Mean wage trends two years before and two years after a mobility conditional on origin and destination firm average wage

(a) Male Workers (b) Female Workers

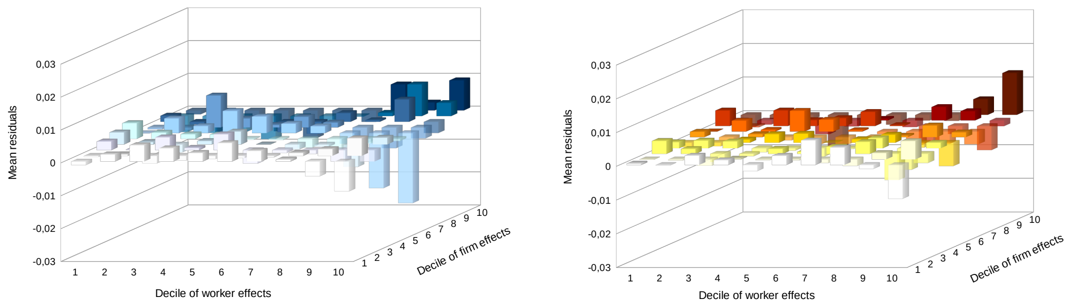


Note: A female worker moving from a firm paying average wages below the bottom quartile (quartile 1) of the wage distribution to a firm above top quartile (quartile 4) has an average residual wage of -0.13 two years before moving, -0.04 the year before her mobility, 0.20 the year she moves and 0.12 the following year. Symmetrically a female going from an above top quartile firm to a below bottom quartile firm can expect a residual wage of 0.20 two years before moving, 0.22 the year before his mobility, -0.04 the year he moves and -0.01 the following year.

Finally, we also provide elements regarding the additive separability of worker and firm fixed effects, which is often viewed as a strong assumption (Eeckhout and Kircher, 2011). We plot the mean wage residuals for either males and females conditional on worker fixed effect and firm fixed effect deciles (figures 5a and 5b). If wages depended not only on worker and firm productivity, but also on the *interaction* of the two factors, the residual should observe specific patterns. For instance, in the case of a supermodular production function, high productivity workers and firms should extract a higher surplus, and we would find larger positive wage residuals for matches between high productivity workers and firms. Figures 5a and 5b show no particular pattern that would suggest the need of an additional interaction in the wage specification. The mean of residuals does not seem to vary as a function of individual workers and firms effects, and it is contained between -0.02 and 0.02 which corresponds to -/+1 % of the average hourly log wage for either men (2.54) and women (2.37). The order of magnitude of the residuals is comparable to the one obtained by Card, Cardoso, and Kline (2016) and Card, Heining, and Kline (2013).

Figure 5: Mean of wage residuals conditional on deciles of worker and firm fixed effects

(a) Male Workers (b) Female Workers



Note: The average wage residual for top-productivity male workers (decile of worker effect= 10) employed in top-paying firms (decile of firm effect= 10) is 0.009.

4 Results

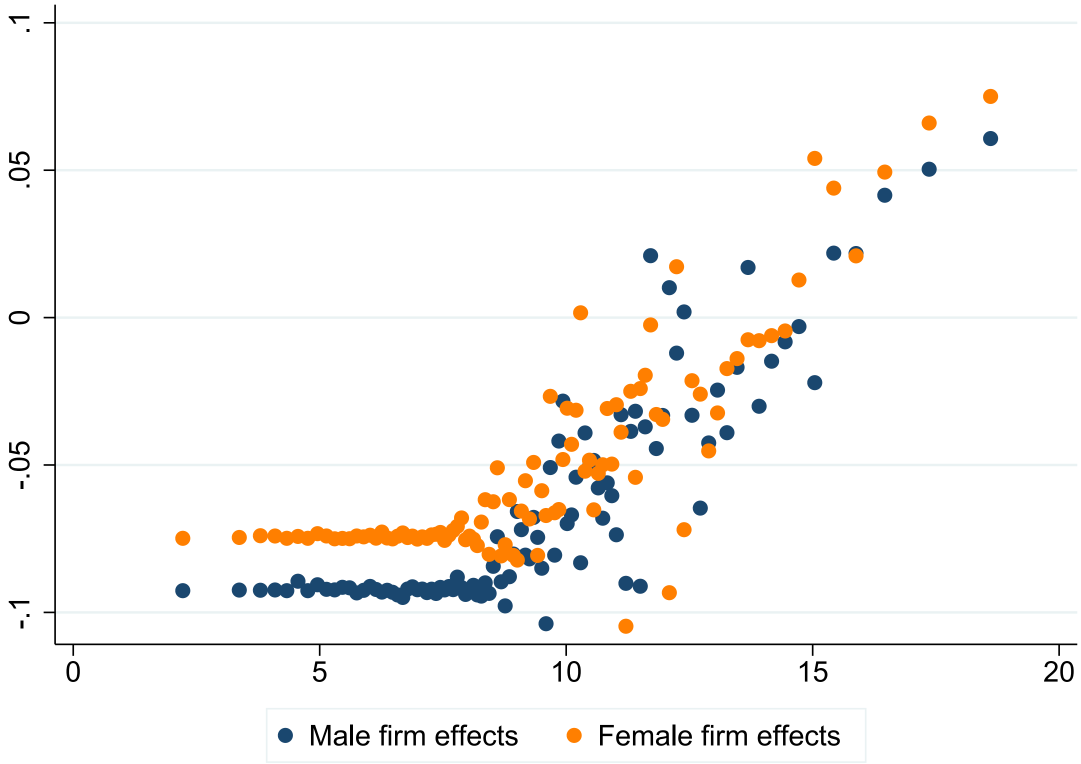
We now present the results obtained from the whole sample of private sector employees. As already mentioned, the estimation sample is composed of 102,048 workers (45,124 women) after identification restrictions.

4.1 Firm-effect normalization

As detailed in Abowd, Creecy, and Kramarz (2002), the model estimation requires a normalization of the firm-fixed effects. In order to be able to compare firm fixed effect estimated from two different groups - men and women, this normalization has to be carefully chosen as it is likely to influence the level of all firm fixed effects estimated on the subpopulation. We thus need to normalize to zero the average fixed effect for firms that are likely to be identical for men and women.

To do so, we follow Card, Cardoso, and Kline (2016) and fix to zero both male and female firm-fixed effects in the firms where there is structurally little rent to share

Figure 6: Firm effects according to log per capita value-added



Lecture: Firms in the dual connected set are grouped into 100 bins according to their log value-added per worker. For each VA per capita bin we plot the corresponding average female and male firm effects obtained with an arbitrary normalization rule (one firm effect set to zero). Note: For firms in the VA per capita top percentile (average VA = 18) female premia before normalization are equal to 0.078 and male premia are equal to 0.061.

between workers, and therefore little risk of sharing differentials between female and male workers. First, we choose to fix to zero the average firm fixed effects in accommodation and food services firms, as this industry is the one for which the value-added per worker is the lowest. As an alternative normalization, and to check for the robustness of our choice, we fix to zero the average firm effects of firms with the lowest log-VA per worker. Figure 6 motivates this normalization: there is a positive relation between the productivity of the firm and the premia female and male workers get. The optimal level of log value-added per worker under which a firm is put in the zero fixed effect group is 8.3. 36,218 firms are below this threshold (8,763 firms with 10 workers or more) representing 17.6 % of firms and 9.4 % of worker observations. Note that in both cases, the normalization is made after the estimation, and does not affect the estimation of the marginal impact of time varying covariates. It only impacts the comparison between the two groups of population.

4.2 Firm fixed effects and firm characteristics

To check whether our results are consistent with the model, we analyze the firm fixed effects as a function of several firm characteristics.[9] For each firm we average the estimated male and female firm fixed effects, and we regress this variable on a set of firm covariates. Findings for different specifications are presented in table 4. Model (1) includes a range of workforce composition variables in addition to firm characteristics such as the value-added per worker, a dummy for exporting firms, the assets per worker, and the number of workers in the firm. In model (2) we add industry dummies. In addition to these industry dummies, model (3) controls for collective agreement dummies. As robustness checks, models (4) and (5) use the same specification as models (2) and (3), but for the sample of firms for which we have information on wage grids (in *contractual industries* with 5,000 or more workers). These wage grids are included in the regression in model (6).

As assumed by the rent-sharing theory, firm fixed effects are higher in firms that generate higher value-added per worker. This result holds in all specifications, whether we control or not for industries, collective agreements, and wage grids. On average, firms belonging to the fourth quartile of value-added per worker pay about 4 to 5 % more than firms of the first quartile. Furthermore, firm fixed effects are higher in firms employing a large share of executives and clerks (relative to blue collars), and of open-ended contracts- though this result is not robust in the restricted sample of firms in large *contractual industries*. In contrast, fixed effects are dramatically lower in firms with a high share of workers paid at the minimum wage level, and of white collars. A higher share of women among white collars is related to somewhat lower firm fixed effects, and a higher share of women among executives to higher fixed effects. The effect cancels out when controlling for the collective agreements. These last results may indicate that firms where occupations are segregated by gender pay lower wages than others. The positive correlation between the share of part time workers and the average firm premia can be interpreted in the same way. Part-time work is indeed considerably more common among women so this variable could capture a positive effect of workforce parity on wages. Finally, we do not find significant differences between exporting and non-exporting firms, nor relation between the level of assets per worker and wages.

[9] Here we use the fixed effects obtained from the accommodation and food services industry normalization. Note that the normalization only affects the intercept of the model since it is a mere translation of firm fixed effects, and the results using the alternative normalization would be identical.

Table 4: Average firm premia (after accommodation and food services normalization)

	(1)	(2)	(3)	(4)	(5)	(6)
Constant	0.057***	0.049**	0.089**	0.051**	0.051	-0.722
	(0.017)	(0.020)	(0.037)	(0.025)	(0.044)	(1.569)
Average age	-0.002***	-0.001**	-0.002***	-0.001	-0.001	-0.001*
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Average prof. experience	0.001*	0.001	0.001	-0.000	-0.000	0.000
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
% part time workers	0.026***	0.031***	0.033***	0.033***	0.029**	0.031**
	(0.010)	(0.010)	(0.011)	(0.012)	(0.013)	(0.013)
% minimum wage earners	-0.243***	-0.239***	-0.262***	-0.244***	-0.256***	-0.252***
	(0.013)	(0.014)	(0.015)	(0.016)	(0.017)	(0.016)
(%F-%M) paid at the min. wage	0.015	0.017	0.014	0.019	0.012	0.013
	(0.015)	(0.015)	(0.016)	(0.016)	(0.017)	(0.016)
% open end contracts	0.049***	0.043***	0.021**	0.017*	0.013	0.013
	(0.007)	(0.008)	(0.009)	(0.009)	(0.010)	(0.010)
% executives	0.185***	0.199***	0.174***	0.194***	0.183***	0.195***
	(0.012)	(0.015)	(0.016)	(0.019)	(0.019)	(0.019)
% clerks	0.047***	0.057***	0.042***	0.041***	0.040**	0.043***
	(0.011)	(0.011)	(0.012)	(0.015)	(0.016)	(0.015)
% white collar workers	-0.043***	-0.030***	-0.031***	-0.042***	-0.030**	-0.028**
	(0.008)	(0.009)	(0.011)	(0.011)	(0.013)	(0.011)
% female among executives	0.022***	0.019**	0.014	0.010	0.008	0.009
	(0.008)	(0.008)	(0.009)	(0.010)	(0.010)	(0.010)
% female among clerks	0.001	0.001	0.003	0.002	0.004	0.001
	(0.006)	(0.006)	(0.006)	(0.007)	(0.008)	(0.008)
% female among white collars	-0.011**	-0.014***	-0.010*	-0.019***	-0.013*	-0.016**
	(0.005)	(0.005)	(0.005)	(0.007)	(0.007)	(0.007)
%female among blue collars	0.004	0.005	0.000	-0.010	-0.012	-0.016*
	(0.008)	(0.008)	(0.008)	(0.010)	(0.010)	(0.010)
Value-added per worker (*ref=1*)						
Quartile 2	-0.014**	-0.015***	-0.008	-0.012*	-0.007	-0.010
	(0.006)	(0.006)	(0.006)	(0.007)	(0.007)	(0.007)
Quartile 3	0.018***	0.014**	0.020***	0.019**	0.019**	0.018**
	(0.006)	(0.006)	(0.007)	(0.008)	(0.008)	(0.008)
Quartile 4	0.036***	0.033***	0.038***	0.055***	0.049***	0.052***
	(0.007)	(0.007)	(0.009)	(0.011)	(0.011)	(0.011)
Exporting firm	0.001	0.001	0.001	-0.004	-0.005	-0.005
	(0.005)	(0.005)	(0.006)	(0.006)	(0.007)	(0.007)
Assets per worker (*ref=1*)						
Quartile 2	-0.008	-0.005	0.003	0.006	0.013	0.012
	(0.006)	(0.006)	(0.007)	(0.009)	(0.009)	(0.009)
Quartile 3	-0.001	0.003	0.005	0.007	0.005	0.009
	(0.006)	(0.007)	(0.007)	(0.009)	(0.010)	(0.009)
Quartile 4	0.015**	0.016**	0.006	0.016	0.010	0.015
	(0.006)	(0.007)	(0.008)	(0.010)	(0.011)	(0.010)
Number of workers /1000	0.000*	0.000	0.001	0.001	0.000	0.001
	(0.000)	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
Industry dummies	-	Yes	Yes	Yes	Yes	Yes
Collective agreement dummies	-	-	Yes	-	Yes	-
Top and bottom wages by position	-	-	-	-	-	Yes
Adjusted R^2	0.257	0.267	0.297	0.293	0.311	0.302
Number of observations	6518	6491	6491	4066	4066	4066

Standard errors in parentheses - * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

4.3 Returns to time varying characteristics

Table 5 shows the estimates associated to time varying characteristics. In addition to firm and individual fixed effects, the first model controls for quadratic functions of age, experience on the labor market, and seniority in the firm. In a second specification, we also control for the number and the age of children.

Table 5: Two-way fixed effects model estimates

	Model (1)		**Model (2)**	
	Female	*Male*	*Female*	*Male*
Age	0.034***	0.043***	0.041***	0.041***
	(0.001)	(0.001)	(0.001)	(0.001)
Age^2/100	-0.032***	-0.042***	-0.038***	-0.039***
	(0.001)	(0.001)	(0.001)	(0.001)
Experience	0.012***	0.013***	0.011***	0.013***
	(0.000)	(0.000)	(0.000)	(0.000)
Experience^2/100	-0.004***	-0.006***	-0.004***	-0.006***
	(0.001)	(0.001)	(0.001)	(0.001)
Seniority	0.010***	0.012***	0.010***	0.012***
	(0.000)	(0.000)	(0.000)	(0.000)
Seniority^2/100	-0.030***	-0.041***	-0.032***	-0.040***
	(0.002)	(0.001)	(0.002)	(0.001)
1 child	-	-	-0.049***	-0.017***
			(0.004)	(0.004)
2 children	-	-	-0.064***	-0.004
			(0.005)	(0.004)
3 children	-	-	-0.099***	-0.005
			(0.006)	(0.005)
4 children or more	-	-	-0.113***	-0.006
			(0.010)	(0.007)
18 - age of the youngest child	-	-	0.001**	0.002***
			(0.000)	(0.000)
Number of observations	403 728	515 967	403 728	515 967
Adjusted R^2	0.80	0.85	0.80	0.85

Standard errors in parentheses - * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Year dummies, individual fixed effects and firm fixed effects are included in the regressions but are not reported here. In addition to the number of children and the difference between 18 and the age of the last born child we include interaction between these two variables and being born on October 2 or 3- demographic data collection was not complete for people born those days.

Some differences between model (1) and model (2) are worth noting. Returns to age are higher and more concave for men relative to women in model (1). However, once we account for family characteristics age parameter estimates are almost identical for men and women. This discrepancy has to do with much larger wage penalties associated to childbirth for women (mothers with one child suffer a 4.9 % wage penalty relative to women without child; the father average penalty is 1.7 %). Wilner (2016) measures parenthood pay gap with a two-way fixed effect model on the same dataset as ours from 1995 to 2011. He finds a 4.7 % wage penalty for women after their first childbirth but no significant effect for men.[10] Interestingly, accounting for children do not change the returns to professional experience nor to seniority of male and female workers whereas one could expect the reverse if during a maternity leave there is no human capital accumulation, if no loss.[11]

4.4 Sorting and bargaining

We now turn to the results of the decomposition of firm contribution to the gender wage gap as detailed in equation (3). The sorting and bargaining effects are presented in table 6. We display the results of model (1) here (results for model (2) are identical, see table 11 in appendix). The gender wage gap amounts roughly to 17 % of average male earnings. 8.2 % of this gap is due to firms. The sorting effect accounts for 10.6 % of the gender wage gap.[12] This result indicates that women are more likely than men to be employed in low-paying firms, even once worker fixed-effects and characteristics are accounted for.

The bargaining effect is very small, and if anything negative, around -2 % of the gender wage gap: once productivity differentials are accounted for, on average women receive the same, and if anything higher firm pay premia than men. Note that estimates of the bargaining effect and of the total contribution of firms on gender wage gap can vary with the normalization choice, but results are exactly similar with either the accommodation and food services, or the lowest value added per worker firms. Overall, the role of firms on the gender wage gap is mainly due to the sorting of women in low-paying firms with respect to men. Card, Cardoso, and Kline (2016) find a higher firm contribution, about 20 % of the Portuguese gender wage gap, due to a higher sorting effect, 75 % of the firm contribution to the gender wage gap comes from sorting, and a small but positive bargaining effect.

[10] Wilner (2016) computes professional experience differently as we do. He uses an experience variable corrected for maternity leaves. Here, maternity leave is not excluded from experience and seniority, and interestingly we find the same child penalty as his.

[11] In an alternative specification we introduce collective agreement dummies as controls in the two-way fixed effect estimation. Results are not affected (see table 10).

[12] We comment sorting effect computed from female premia, and the corresponding bargaining effect computed with male assignment in firms as reference. For completeness of the analysis we

Table 6: Sorting and bargaining contributions to the gender wage gap

	Normalization based on...			
	Accommodation and food services firms		*Lowest log VA per worker firms*	
Total gender gap	0.170	*100 %*	0.170	*100 %*
Gender log wage gap due to firms	0.014	*8.2 %*	0.014	*8.2 %*
including sorting effect				
Male assignment, female premia (a)	0.018	*10.6 %*	0.018	*10.6 %*
Female assignment, male premia (b)	0.018	*10.6 %*	0.018	*10.6 %*
including bargaining effect				
Male assignment, female premia (c)	-0.004	*-2.4 %*	-0.004	*-2.4 %*
Female assignment, male premia (d)	-0.004	*-2.4 %*	-0.004	*-2.4 %*
Number of observations	912,784			
Number of firms (10+ workers)	11,062			
Number of workers	102,048			

Calculation of sorting and bargaining are based on model (1) estimates. Line (a) reports sorting effect calculated using female premia: $\mathbb{E}[\psi^F_{J(i,t)} \mid g = M] - \mathbb{E}[\psi^F_{J(i,t)} \mid g = F]$. The corresponding bargaining effect is in line (c) and computed with male assignment in firms as reference: $\mathbb{E}[\psi^M_{J(i,t)} - \psi^F_{J(i,t)} \mid g = M]$. Oppositely (b) gives the estimates for the sorting effect measured with male premia: $\mathbb{E}[\psi^M_{J(i,t)} \mid g = M] - \mathbb{E}[\psi^M_{J(i,t)} \mid g = F]$ and (d) for the bargaining effect based on female assignment in firms: $\mathbb{E}[\psi^M_{J(i,t)} - \psi^F_{J(i,t)} \mid g = F]$. In both cases sorting and bargaining effects add up to the gender log wage due to firms: (a)+(c)=(b)+(d)= Gender log wage gap due to firms.

4.5 Heterogeneity analysis

Using the firm fixed effect estimates presented above we compute the sorting and bargaining effects for different subgroups depending on job occupation, education level, age, birth cohort and parenthood (table 7). This entails for a group of population P:

$$\begin{aligned}&\mathbb{E}\left[\psi^M_{J(i,t)} \mid g = M, P\right] - \mathbb{E}\left[\psi^F_{J(i,t)} \mid g = F,\, P\right] \\ &\quad = \underbrace{\mathbb{E}\left[\psi^M_{J(i,t)} - \psi^F_{J(i,t)} \mid g = M, P\right]}_{\text{(i) Bargaining effect}} + \underbrace{\mathbb{E}\left[\psi^F_{J(i,t)} \mid g = M, P\right] - \mathbb{E}\left[\psi^F_{J(i,t)} \mid g = F, P\right]}_{\text{(ii) Sorting effect}}.\end{aligned} \tag{5}$$

For each subgroup we report also the total gender wage gap.

Occupations: the total gender wage gap ranges from .067 among white collars to .186 among executives. Similarly the total firm contribution is the highest for executives (.021) but almost null for blue or white collars. The sorting and bargaining effects are rather small among blue and white collars, with a positive sorting

also report results for the opposite reference in Table 6 but the results show no differences.

(.006 and .010), and a negative bargaining (-.003 and -.005), which means that blue and white collar women tend to be employed in less paying firms but do have some negotiation power relatively to their male coworkers (probably due to the minimum wage action). Reversely, sorting effect is small but negative among executives (-.006): women are not segregated in less-paying firms. And above all, the bargaining effect among executives is stronger and positive (0.027). The clerks situation lies somewhere between the blue and white collars and the executives ones.

The bargaining result for executives relates to the fact that the latter are likely to have more opportunity to actively bargain their wages than the rest of workers and that women tend to be less efficient in negotiating their wages than men (Bowles, Babcock, and Lai, 2007; Bowles, Babcock, and McGinn, 2005; Small, Gelfand, Babcock, and Gettman, 2007). The high bargaining effect, as measured here, may also reflect some vertical job segregation within firms: female executives may generally not access to the top executive and highest paid jobs (Gobillon, Meurs, and Roux, 2015), part of which can be reflected by high within firm gender gaps. Besides, both mechanisms may be intricated as suggested by Greig (2008) who documents a correlation between one's propensity to negotiate and rate of advancement.

For these reasons, together with the fact that gender gap is particularly acute among executives (Bertrand, Goldin, and Katz, 2010), we considered estimating separately the model for executives. However, figures 10 to 12 evidence that the exogenous mobility condition and the rent-sharing model are not credible within this group of workers: executives tend to get better wages when they leave a firm for another, whatever the productivity level of their new employers.[13]

Lastly, for every group the sorting effect is smaller than among the entire population (0.018): although we control for individual heterogeneity, this result means that sorting across firms and sorting across job occupations are partly overlapping phenomena.

Education: in contrast to what we find for job position groups, sorting effects are roughly stable across education level. The sorting effect is somewhat smaller among more educated workers but the difference between highly educated and less educated workers is far smaller than the one found by Card, Cardoso, and Kline (2016) on Portuguese data. Moreover, we can note that the total gender wage gap is twice larger among college and university graduates than among less educated workers,

[13]Models with endogenous mobility (Lamadon, Manresa, and Bonhomme, 2015) would probably suit better for executives behavior. This issue should be the object of future research.

Table 7: Sorting and bargaining effects conditional on job positions, education level, age, birth cohort and parenthood

	N	Sorting	Bargaining	Total firm contribution	Total gender wage gap
Blue collars	300,986	0.006	-0.003	0.003	0.141
White collars	271,596	0.010	-0.005	0.005	0.067
Clerks	201,820	0.007	0.007	0.014	0.105
Executives	135,031	-0.006	0.027	0.021	0.186
< High school	518,600	0.018	-0.007	0.011	0.151
High school	162,820	0.019	0.010	0.029	0.186
> High school	231,364	0.014	0.004	0.018	0.297
< 30 year old	211,070	0.010	-0.004	0.006	0.063
30-39	261,637	0.011	-0.002	0.009	0.137
40-49	249,832	0.021	-0.002	0.019	0.213
> 49	190,245	0.026	-0.006	0.020	0.270
Born before 54	140,593	0.021	-0.002	0.019	0.291
1954-1963	242,007	0.021	-0.002	0.019	0.219
1964-1973	275,132	0.010	0.000	0.010	0.145
1974-1983	205,754	0.008	-0.001	0.007	0.082
Parents	627,930	0.020	-0.002	0.018	0.213
Non parents	284,854	0.007	-0.005	0.002	0.078
Parents 45+	216,804	0.027	-0.002	0.025	0.302
Non parents 45+	94,376	0.009	-0.003	0.006	0.145

Note: Calculation of sorting and bargaining are based on model (1) estimates and calculated using female premia and male assignment in firms as reference. We use the accommodation and food services normalization. When we compare mother workers and father workers (no matter when the child is born, all observations are kept) the sorting effect amounts to 0.020 and the bargaining effect is -0.002.

so the relative share of the sorting effect actually decreases with the education level. The bargaining effect for less educated workers is negative and close to the general sample estimate whereas the ones for high-school and college and university graduates are small but positive.

Age and cohort: when we compare age groups with each other we find a clear increase in the sorting effect with age: the estimate is twice larger for workers aged 40 or more than for younger employees. By contrast, the bargaining effect does not seem to vary with age, and remain stable around our baseline estimate. At the end, the total firm contribution follows the same pattern as the sorting effect and increases strongly with age. The cohort analysis conveys the same messages. For cohorts born after 1964 we find a contained sorting effect of the same magnitude as among workers aged 40 or less. Among cohorts born before 1964 the sorting effect doubles whereas the bargaining effect remains close to zero whatever the birth cohort considered.

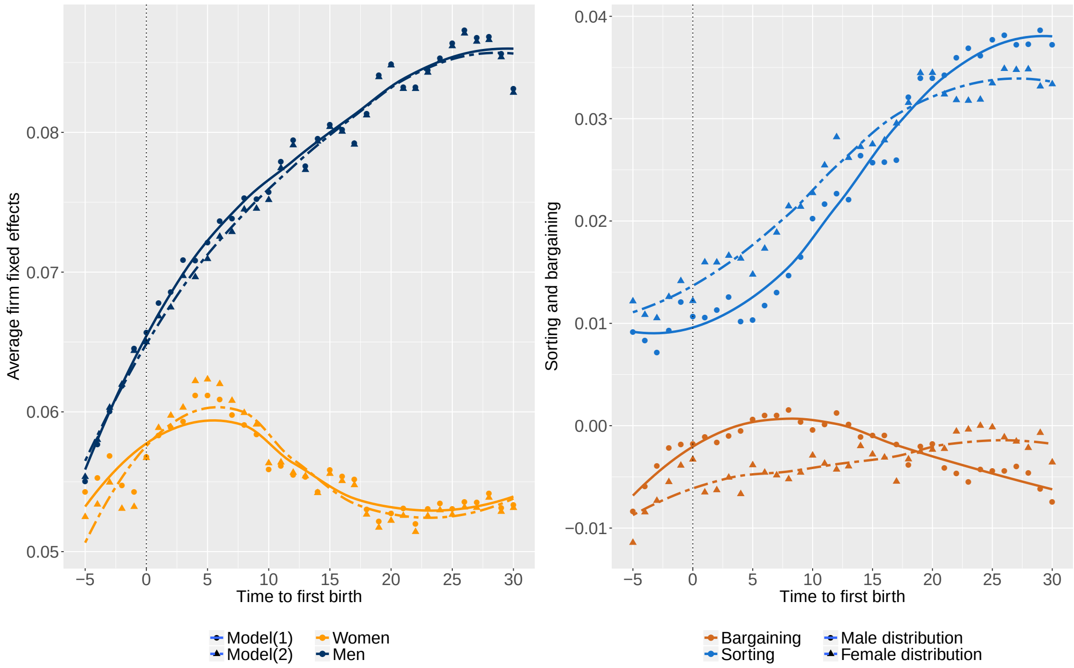
Parents vs non parents: the sorting effect is almost three times bigger for parents than for non-parents. Moreover, the presence of children exacerbates the sorting of women/mothers in less-paying but probably more family-friendly firms with respect to fathers. We also display sorting and bargaining effects for parents and non parents aged 45 or more, because the gap between parents and non parents could also be driven by age effect (non parents being younger). The comparison delivers the same message.

5 Sorting and bargaining over the family life cycle

The findings highlighted in the previous section show that the firm contribution to the gender wage gap and the sorting effect grow dramatically with age, and with the fact of having children. Results from Card, Cardoso, and Kline (2016) show a similar pattern for Portugal, and suggest that sorting may appear because of family duties. From our data, we are able to track childbirth. Thus, we propose a finer picture of the formation of inequality over the family life cycle. Using our firm fixed effects estimates, we build an event study analyzing the link between the birth of the first child, and sorting and bargaining effects. The results indicate that in addition to the direct effects of childbirth on wages (see Table 5), women also experience wage losses associated to sorting after birth.

We restrict our attention to workers who eventually have children, and we center our event study on the time of birth of the first child. Using firm fixed effect estimates, we average sorting and bargaining effect for workers according to the time to this birth, from five years before to thirty years after. Figure 7 plots the results of this event study. The left panel displays the average firm fixed effects for both men (blue) and women (yellow) as a function of time to the event. As shown by this graph, accounting for family events in the wage equation (Model (2)) or not (Model (1)) does not change the results. In the very first years of their careers women and men experience comparable firm premia. However, female firm effects fall rapidly behind male ones especially after the first childbirth. Until five years after the first childbirth, women experience a slower growth rate in firm premia compared to men. Then, we observe a permanent drop in female firm premia in subsequent years.

Figure 7: Firm premia, sorting, and bargaining effects over time to first birth



Note: The average firm fixed effect (normalized based on accommodation and food services) from model (1) for females 10 years after their first childbirth is 0.056; for males, it is 0.076. For the same time to first birth, using male distribution into firms as reference gives a sorting effect of 0.02 and a bargaining effect of 0. The lines were obtained by smoothing the averages.

The right-hand side of the figure shows the corresponding sorting and bargaining effects. The bargaining effect is relatively flat, between -0.01 and 0. It increases slightly in early careers driving part of the increase in gender firm effect gap that occurs before the first childbirth, but it levels off around ten years after it. On the contrary, the sorting effect dramatically increases after the first childbirth from approximately 0.01 in the early career to around 0.04 twenty years after the first childbirth, persisting at this high level until the end of careers. Compared to men, women tend to work in low-paying (and probably family-friendly) firms, or to move less to high-paying firms after the birth of their first child. These differences persist for a long time after it, maybe due in a first place to the births of siblings. But strikingly the sorting effect never decreases even thirty years after the first childbirth: women do not experience upward mobility between firm anymore. This observation is in line with Albrecht, Bronson, Thoursie, and Vroman (2017) who note that men tend to switch more between firms relative to females in the early stages of their careers, when mobility is the most profitable. As mentioned above, there is a double penalty for women having children: childbirths directly affect wages (table 5), and also accentuate a long-term gender divergence in firm effects. Figures 13 and 14 in appendix detail this event study by cohort, and show that younger cohorts are somewhat less concerned by this increasing sorting effect pattern. However, cohort differences are of second-order compared to age differentials.

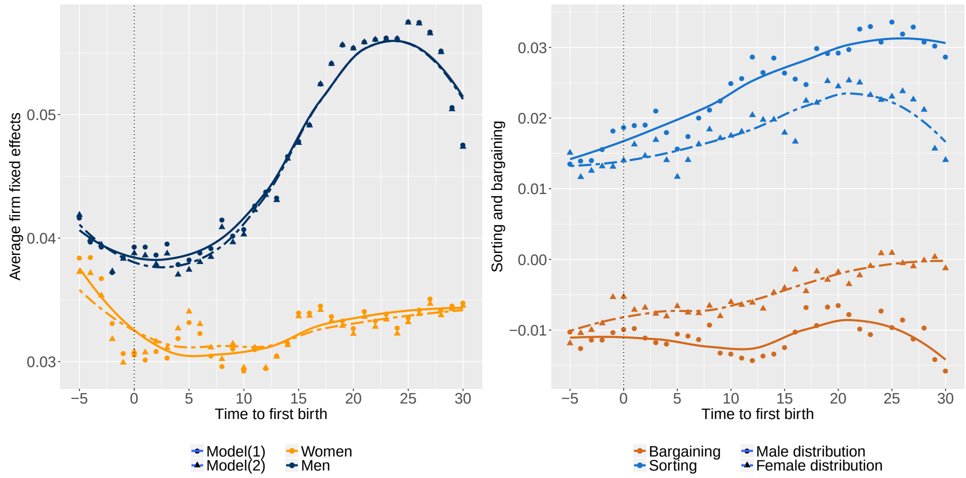
Figure 8 shows the same analysis for the three main groups of occupations: white and blue collars, clerks and executives. For white and blue collars (8a), we can see from the left panel that firm premia trends are close to the ones observed in the entire sample, except that the firm premia magnitudes are smaller (the highest value for average male premia is 0.056 versus 0.085 in the general population). In addition, firm premia decrease earlier in the career than among the entire population. The right panel shows a growing sorting effect over the family life cycle slightly compensated by a stable and negative bargaining effect. Among clerks, female firm premia are higher than male ones in the early careers, but the differences cancel out around five years after first childbirths. Then, the gap turns in favor of men and increases along the family life cycle. It is almost entirely due to the sorting effect whereas the bargaining effect remains around zero all over the career. The picture is less clear for executive workers, maybe because this group is relatively small. Firm fixed effects are high (up to an average 0.145, versus 0.085 in the general sample) and relatively close for men and women, except at the end of careers when firm effects decline more rapidly among women. As in the general population there seems to be an upward trend in sorting effect and a downward bargaining effect evolution from five years after childbirth, but things are less clear before.

Moreover, the sorting/bargaining decompositions of the firm effect gap are very different according to which distribution of workers across firms we use as reference. The sorting effect computed with the male distribution is smaller than the one based on female distribution in firms and reversely the bargaining effect is higher, which implies that male executives tend to work in firms where the within gender gap is more pronounced.

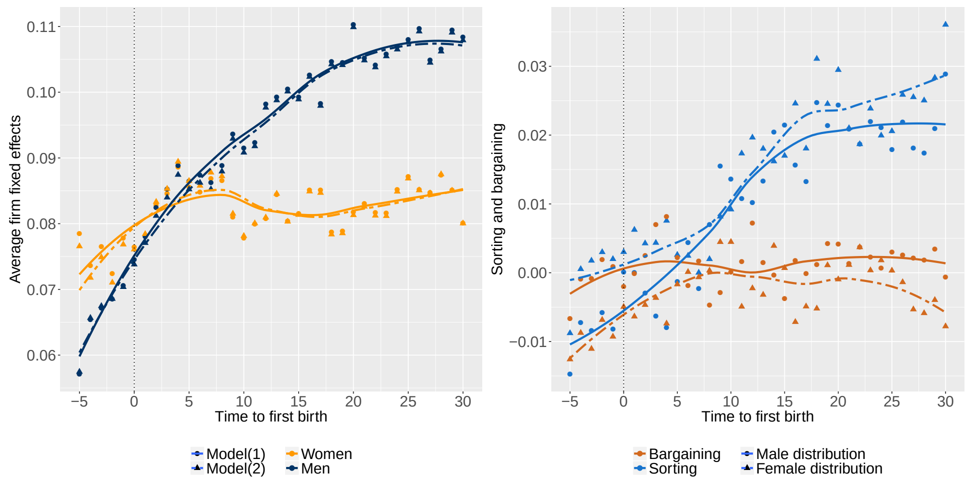
Figure 9 repeats this analysis by level of education. The event studies for workers with low and high-school education (9a and 9b) are close to the whole workers ones. However, the high-school educated group shows a slight decrease in the sorting effect and an increase in the bargaining one that becomes positive about ten years after the first childbirth. College educated workers show a different pattern with inverted u-shaped firm effects. For both men and women, the firm specific wage components climax some years after childbirth and then decrease. This result can be compared to the one of Albrecht, Bronson, Thoursie, and Vroman (2017) who exclusively analyze educated workers, and find that firm contributions on wages are the strongest several years after the childbirth for men, and slightly earlier for women. Finally, the sorting effect seems to increase with time after childbirth whereas the bargaining one is more steady or slightly decreasing.

Figure 8: Firm premia, sorting, and bargaining effects over time to first birth by job position

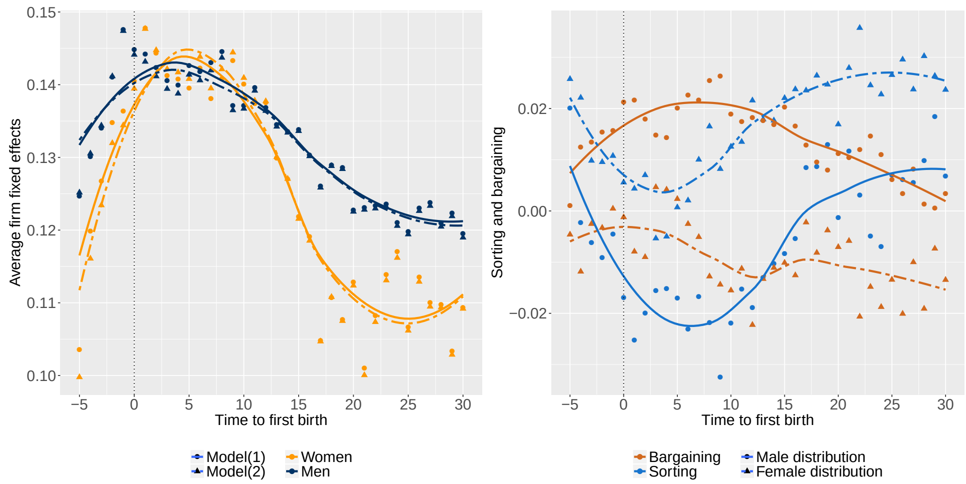
(a) Blue and white collar workers



(b) Clerks



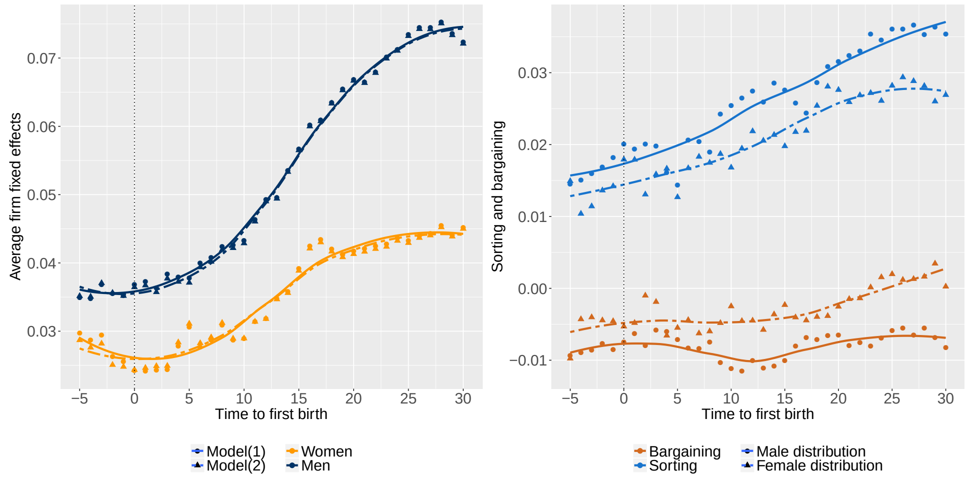
(c) Executive workers



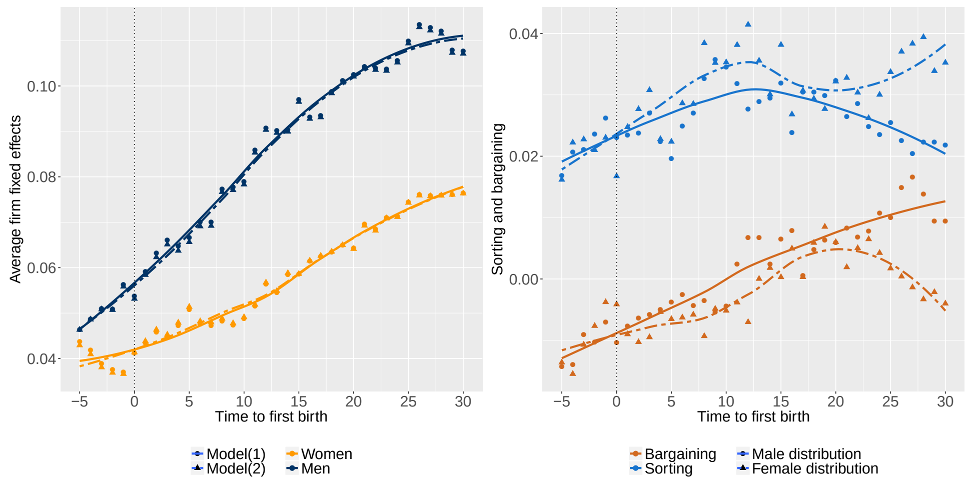
Note: The average firm fixed effect from model (1) for blue or white collar females 10 years after their first childbirth is 0.029; for males, it is 0.041. For the same time to first birth, using male distribution into firms as reference gives a sorting effect of 0.025 and a bargaining effect of -0.013.

Figure 9: Firm premia, sorting, and bargaining effects over time to first birth by education level

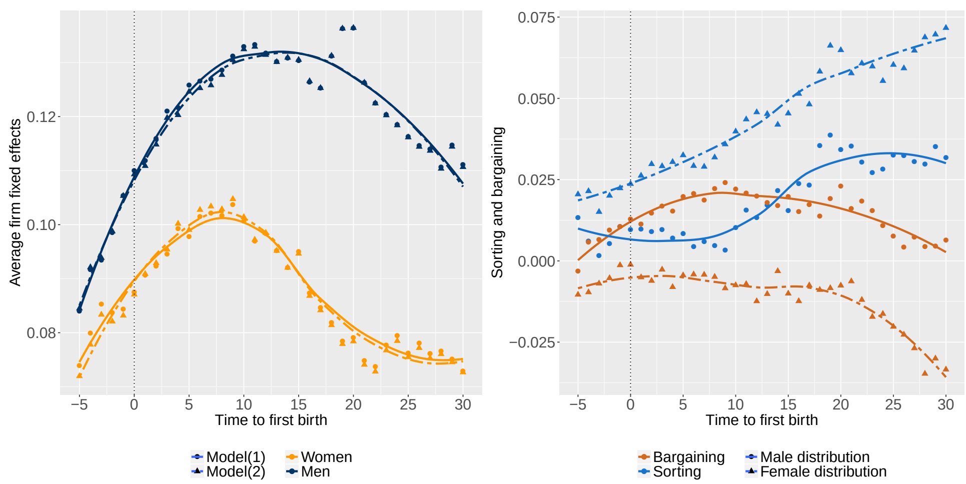
(a) Low-educated workers



(b) High-school educated workers



(c) College educated workers



Note: The average firm fixed effect from model (1) for low-educated females 10 years after their first childbirth is 0.029; for males, it is 0.043. For the same time to first birth, using male distribution into firms as reference gives a sorting effect of 0.025 and a bargaining effect of -0.011.

6 Institutional setting in firm effects

6.1 Collective agreements, firm performance and gender bargaining gap

We now investigate whether the gender gap in fixed effects within a given firm can be related to observed firm characteristics. Note that this exercise does not aim at assessing causal relationships: there is no reason to believe that firms' wage policy derives from their financial reports rather than the other way around. We regress the average within firm fixed effect gap between male and female workers on firm workers' characteristics and firm financial results. Results are presented in table 8. Models (1) to (6) refer to the same specifications as those in table 4.

Compared to the average firm effect analysis shown in table 4, the different specifications only explain little variation of the within firm gap in firm premia: the rent that firms share with their employees (without regard to gender differences) is much more predictable than the gender discrepancies of the rent-sharing process. Only 3.7 % of the variance is explained by our baseline model. Controlling by industry raises the explanatory power to 4.1 %. The model including both industry and collective agreement dummies almost doubles the baseline explained share of the variance. Most of the variance in firm specific gender gap remains unexplained by the main firm characteristics, although some institutional features explain a little share of it.

Only few covariates have significant effects on the firm effect gender gap: the gender gap is increased by 0.125 and 0.145 percentage point when the share of executives is increased by 1 percentage point. However, the higher the share of women among executives is, the lower the gender gap. In addition to that, the firm effect gender gap is negatively correlated with the share of women among white and blue collars in the firm. This finding is consistent with a prominent sorting mechanism across jobs within firms.

Comparing the different models, we can see that a negative effect of the share of minimum wage earners in the firm arises when we control for collective agreements (using dummies or wage grids). A one percentage point in the difference between the share of women and men paid at the minimum wage ((%F-%M) at the min. wage) is associated with a .264 percentage point increase in the firm effect gender gap. This effect reinforces the idea that the minimum wage policy may protect low productivity workers, and thus increases their relative bargaining power.

Table 8: Within firm gender gap in firm premia (after accommodation and food services normalization)

	(1)	(2)	(3)	(4)	(5)	(6)
Constant	0.011	-0.009	0.032	-0.016	0.091	0.850
	(0.031)	(0.035)	(0.064)	(0.043)	(0.074)	(2.973)
Average age	-0.002**	-0.002	-0.002*	-0.001	-0.002	-0.001
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Average prof. experience	0.002	0.001	0.001	0.000	0.000	0.000
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
% part time workers	-0.028*	-0.014	-0.016	-0.003	-0.014	0.000
	(0.017)	(0.017)	(0.019)	(0.021)	(0.022)	(0.022)
% minimum wage earners	-0.035	-0.048*	-0.070**	-0.057**	-0.071**	-0.061**
	(0.024)	(0.025)	(0.028)	(0.029)	(0.031)	(0.030)
(%F-%M) at the min. wage	0.264***	0.257***	0.269***	0.245***	0.251***	0.242***
	(0.028)	(0.028)	(0.030)	(0.030)	(0.031)	(0.031)
% open end contracts	0.010	0.001	-0.009	-0.007	-0.006	-0.011
	(0.013)	(0.014)	(0.017)	(0.016)	(0.018)	(0.017)
% executives	0.125***	0.142***	0.135***	0.145***	0.130***	0.138***
	(0.021)	(0.026)	(0.029)	(0.034)	(0.035)	(0.034)
% clerks	0.036*	0.049**	0.035	0.048*	0.033	0.040
	(0.019)	(0.020)	(0.023)	(0.027)	(0.029)	(0.027)
% white collar workers	0.020	0.029*	0.033	0.021	0.029	0.026
	(0.014)	(0.016)	(0.021)	(0.020)	(0.025)	(0.022)
% female among executives	-0.118***	-0.115***	-0.121***	-0.127***	-0.128***	-0.128***
	(0.015)	(0.015)	(0.015)	(0.019)	(0.019)	(0.019)
% female among clerks	-0.012	-0.013	-0.013	-0.012	-0.016	-0.013
	(0.011)	(0.011)	(0.011)	(0.014)	(0.014)	(0.014)
% female among white collars	0.043***	0.045***	0.042***	0.051***	0.048***	0.051***
	(0.009)	(0.010)	(0.010)	(0.013)	(0.013)	(0.013)
%female among blue collars	0.075***	0.073***	0.073***	0.063***	0.058***	0.060***
	(0.014)	(0.015)	(0.016)	(0.018)	(0.019)	(0.019)
Value-added per worker (*ref=1*)						
Quartile 2	0.002	0.001	0.004	-0.004	0.002	-0.002
	(0.010)	(0.010)	(0.011)	(0.011)	(0.012)	(0.012)
Quartile 3	-0.004	-0.004	-0.004	-0.007	-0.001	-0.007
	(0.011)	(0.011)	(0.012)	(0.014)	(0.015)	(0.014)
Quartile 4	0.010	0.003	0.008	0.004	0.011	0.007
	(0.012)	(0.013)	(0.016)	(0.019)	(0.021)	(0.020)
Exporting firm	0.003	-0.008	-0.008	-0.005	-0.004	-0.010
	(0.008)	(0.009)	(0.010)	(0.011)	(0.012)	(0.012)
Assets per worker (*ref=1*)						
Quartile 2	0.003	0.001	0.008	0.018	0.027*	0.020
	(0.010)	(0.011)	(0.012)	(0.015)	(0.016)	(0.016)
Quartile 3	-0.010	-0.012	-0.005	0.002	0.011	0.004
	(0.010)	(0.012)	(0.013)	(0.016)	(0.018)	(0.017)
Quartile 4	-0.011	-0.007	-0.001	0.007	0.018	0.010
	(0.011)	(0.013)	(0.015)	(0.018)	(0.020)	(0.019)
Number of workers /1000	0.000	0.000	-0.001	0.000	-0.001	0.000
	(0.000)	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
Industry dummies	-	Yes	Yes	Yes	Yes	Yes
Collective agreement dummies	-	-	Yes	-	Yes	-
Top bottom wages by position	-	-	-	-	-	Yes
Adjusted R^2	0.037	0.041	0.062	0.050	0.058	0.049
Number of observations	6518	6491	6491	4066	4066	4066

Standard errors in parentheses - * $p < 0.01$, ** $p < 0.05$, *** $p < 0.01$.

6.2 More on the role of the minimum wage

Given the previous results, and the negative bargaining effect that we find in table 6, we further investigate the role of the minimum wage. A high minimum wage combined with a contained labor cost at the minimum wage imply that low-paid workers are protected in two ways: they are not discarded from employment, and they are paid at a relatively higher wage than their individual productivity would predict. As already highlighted, these two protections are all the more important for our purpose that women are over-represented around the minimum wage: 13.6 % of female workers in the estimation sample have a wage equal to or below 1.1 minimum wage, compared to 7.4 % of male workers. Hence, the minimum wage may reduce the bargaining effect at the bottom of the wage distribution.

Table 9: Sorting and bargaining contributions to the gender wage gap in 1995-2004 and 2005-2014

	1995-2004		**2005-2014**	
Total gender gap	0.180	*100 %*	0.154	*100 %*
Gender log wage gap due to firms	0.015	*8.3 %*	0.010	*6.5 %*
including sorting effect				
Male assignment, female premia (a)	0.015	*8.3 %*	0.018	*11.7 %*
Female assignment, male premia (b)	0.012	*6.7 %*	0.013	*8.4 %*
including bargaining effect				
Male assignment, female premia (c)	0.000	*0 %*	-0.008	*-5.2 %*
Female assignment, male premia (d)	0.003	*1.7 %*	-0.004	*-2.6 %*
Number of observations	383,368		375,121	
Number of firms (10+ workers)	4,907		3,862	
Number of workers	68,903		61,371	

Calculation of sorting and bargaining are based on model (1) estimates, using accommodation and food services normalization. Line (a) reports sorting effect calculated using female premia: $\mathbb{E}[\psi^F_{J(i,t)} \mid g = M] - \mathbb{E}[\psi^F_{J(i,t)} \mid g = F]$. The corresponding bargaining effect is in line (c) and computed with male assignment in firms as reference: $\mathbb{E}[\psi^M_{J(i,t)} - \psi^F_{J(i,t)} \mid g = M]$. Oppositely (b) gives the estimates for the sorting effect measured with male premia: $\mathbb{E}[\psi^M_{J(i,t)} \mid g = M] - \mathbb{E}[\psi^M_{J(i,t)} \mid g = F]$ and (d) for the bargaining effect based on female assignment in firms: $\mathbb{E}[\psi^M_{J(i,t)} - \psi^F_{J(i,t)} \mid g = F]$. In both cases sorting and bargaining effects add up to the gender log wage due to firms: (a)+(c)=(b)+(d)= Gender log wage gap due to firms.

This explanation is in line with the bargaining results obtained separately for periods before (1995-2004) and after (2005-2014) the rapid growth of the minimum wage presented in table 9 (detailed results of the estimation on the two periods can be found in tables 12 to 15). The bargaining effect is indeed only negative for the 2005-2014 period when the minimum wage is higher.[14] In theory, a higher minimum

[14]The financial crisis could also have contributed to a stagnation or a drop in bargaining effect,

wage may also lead to a lesser sorting effect, as low-paid workers are more protected also *between* firms. Hence, the fact that the French minimum wage is higher than the Portuguese one could account for Card, Cardoso, and Kline (2016) finding a greater contribution of firms to the gender wage gap in Portugal, with both higher sorting and bargaining effects. However, there is still a puzzle here: in our data, the sorting effect slightly grows between the two sub-periods 1995-2004 and 2005-2014 in spite of the minimum wage increase. But this could also depend on other underlying trends, and how low-wage jobs are distributed between firms. Lastly, if the minimum wage reduces the bargaining effect, this may have consequences on the interpretation of our reference group. We choose as our reference group the accommodation and food services, an industry where the share of minimum wage workers is high, and where women are overrepresented among minimum wage earners (24.0 % versus 15.7 % for males). Hence our estimate of the bargaining effect may overestimate the bargaining effect we would have estimated in the absence of minimum wage- the underlying bargaining effect may be more in favor of women.

with firms having a smaller rent to distribute, either to men and women.

7 Conclusion

In this paper, we investigate the within and between firm contributions to the gender wage gap. Using matched employer-employee data, we apply Card, Cardoso, and Kline (2016) decomposition of the residual wage gap remaining after controlling for individual unobserved heterogeneity on French private sector data. We display pieces of evidence that our sample fulfills with the requirements for identification of the two way fixed effect models. When we consider the whole population, we find a bargaining effect close to zero, and if anything negative: women tend to be paid as well as their male coworkers after controlling for individual characteristics and unobserved productivity. However, we identify a positive sorting effect (around 11 % of the total gap in hourly wage between men and women), which means that firms contribute to the gender wage gap in the sense that women tend to be at work in firms paying smaller wages than men with comparable productivity.

Sorting is smaller within job positions than in the entire sample – hence, sorting into firms is often combined with sorting into jobs. Bargaining is clearly positive adn stronger among executives than among other groups, which is consistent with the fact that this group shows the highest gender gap, is likely to have more negotiation margins, which benefit more to men who perform better in this exercise than women. This result relates also to some vertical job segregation. We also provide evidence that we cannot conduct separate estimation of two-way fixed effect model on executives only as exogeneity assumptions are not fulfilled for this group.

Our very rich administrative data allow us to recover information on both work and family lifes. Hence, we provide a finer picture of sorting and bargaining effects all along the family life cycle, according to the time after the workers' first childbirth. We document a double child penalty: in addition to the direct effects of childbirth on wages, women also experience wage losses associated to sorting between firms after the birth of their first children. Indeed, gender firm effect gap appears around the first childbirth and deepens over the life cycle. In particular, sorting effect increases after childbirth: when they become mothers women tend to favor less-paying firms and to move less to high-paying firms with respect to men. Then, the sorting effect stays at a high level no matter how their family duties evolve: women do not move upward anymore.

Our results confirm some of Card, Cardoso, and Kline (2016): higher bargaining effect among executives, sorting effect that increases in age. But Card, Cardoso, and Kline (2016) find much larger bargaining effects in Portugal than we do in France.

Possible explanations for these differences lie in the institutional specificity of the French labor market: minimum wage and collective agreements. In particular, there is in France a high minimum wage combined with a contained labor cost at the minimum wage. It implies that low-paid workers are likely to be paid at a relatively higher wage than their individual productivity would predict. Showing that the within firm gender wage gap is attenuated in firms where there are more individuals paid at the minimum wage, our firm level analysis confirms that the minimum wage plays an important role in the bargaining effect. Hence, we believe the bargaining effect we are able to measure would be larger without the minimum wage, which acts as a device reducing gender inequalities at the bottom of the wage distribution.

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A Appendix

A.1 Two-way fixed-effect estimation with collective agreement indicators

Table 10: Two-way fixed effects model estimates on the entire sample with collective agreement indicators

	Model (1)		**Model (2)**	
	Female	*Male*	*Female*	*Male*
Age	0.035***	0.044***	0.042***	0.042***
	(0.001)	(0.001)	(0.001)	(0.001)
$Age^2/100$	-0.031***	-0.042***	-0.038***	-0.040***
	(0.001)	(0.001)	(0.001)	(0.001)
Experience	0.012***	0.012***	0.011***	0.012***
	(0.000)	(0.000)	(0.000)	(0.000)
$Experience^2/100$	-0.005***	-0.006***	-0.005***	-0.006***
	(0.001)	(0.001)	(0.001)	(0.001)
Seniority	0.010***	0.012***	0.010***	0.012***
	(0.000)	(0.000)	(0.000)	(0.000)
$Seniority^2/100$	-0.030***	-0.044***	-0.032***	-0.042***
	(0.002)	(0.001)	(0.002)	(0.001)
1 child			-0.050***	-0.019***
			(0.004)	(0.004)
2 children			-0.066***	-0.007
			(0.005)	(0.004)
3 children			-0.102***	-0.008
			(0.006)	(0.005)
4 children or more			-0.116***	-0.009
			(0.010)	(0.007)
18 - age of the youngest child			0.001**	0.002***
			(0.000)	(0.000)
Collective agreements	Yes	Yes	Yes	Yes
Number of observations	403,728	515,967	403,728	515,967
Adjusted R^2	0.80	0.85	0.80	0.85

Standard errors in parentheses - * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Collective agreement dummies (for the 84 main agreements, covering 90 % of employees), year dummies, individual fixed effects and firm fixed effects are included in the regressions but are not reported here. In addition to the number of children and the difference between 18 and the age of the last born child we include interaction between these two variables and being born on October 2 or 3- demographic data collection was not complete for people born those days.

A.2 Sorting and bargaining effects computed with model (2) estimates

Table 11: Sorting and bargaining contributions to the gender wage gap (Model (2))

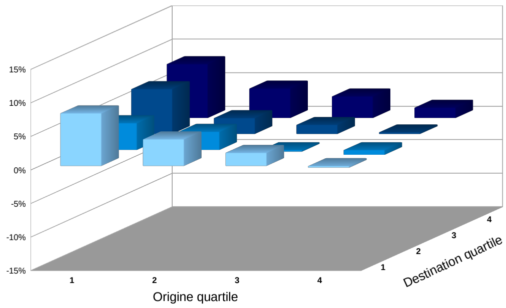
	Normalization based on...			
	Accommodation and food services firms		*Lowest log VA per worker firms*	
Total gender gap	0.170	*100 %*	0.170	*100 %*
Gender log wage gap due to firms	0.014	*8.2 %*	0.014	*8.2 %*
including sorting effect				
Male assignment, female premia (a)	0.018	*10.6 %*	0.018	*10.6 %*
Female assignment, male premia (b)	0.018	*10.6 %*	0.018	*10.6 %*
including bargaining effect				
Male assignment, female premia (c)	-0.004	*-2.4 %*	-0.004	*-2.4 %*
Female assignment, male premia (d)	-0.004	*-2.4 %*	-0.004	*-2.4 %*
Number of observations	912,784			
Number of firms (10+ workers)	11,062			
Number of workers	102,048			

Calculation of sorting and bargaining are based on model (1) estimates. Line (a) reports sorting effect calculated using female premia: $\mathbb{E}[\psi^F_{J(i,t)} \mid g = M] - \mathbb{E}[\psi^F_{J(i,t)} \mid g = F]$. The corresponding bargaining effect is in line (c) and computed with male assignment in firms as reference: $\mathbb{E}[\psi^M_{J(i,t)} - \psi^F_{J(i,t)} \mid g = M]$. Oppositely (b) gives the estimates for the sorting effect measured with male premia: $\mathbb{E}[\psi^M_{J(i,t)} \mid g = M] - \mathbb{E}[\psi^M_{J(i,t)} \mid g = F]$ and (d) for the bargaining effect based on female assignment in firms: $\mathbb{E}[\psi^M_{J(i,t)} - \psi^F_{J(i,t)} \mid g = F]$. In both cases sorting and bargaining effects add up to the gender log wage due to firms: (a)+(c)=(b)+(d)= Gender log wage gap due to firms.

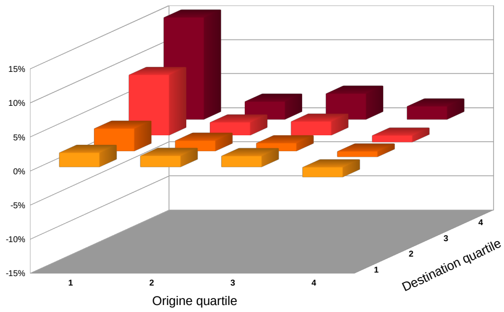
A.3 Firm/Wage Mobility and Additive Separability for Executives

Figure 10: Average wage change for executive movers conditional on origin and destination firm average wage

(a) Male Workers



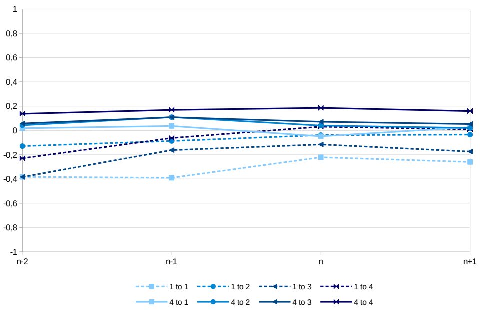
(b) Female Workers



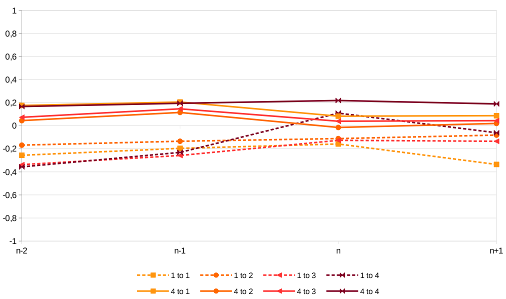
Note: Leaving a Q1 firm for a Q4 firm yields an average wage gain of 8.1 % to male executives.

Figure 11: Mean wage trends for executives two years before and two years after a mobility conditional on origin and destination firm average wage

(a) Male Workers



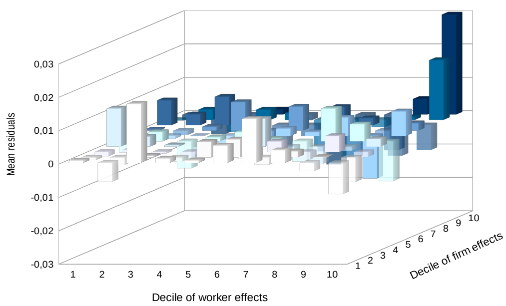
(b) Female Workers



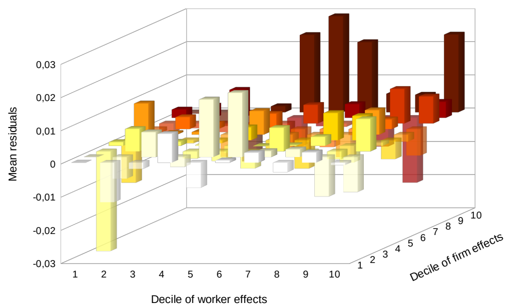
Note: The average wage residual trend for a female executive going from a Q1 firm to a Q4 firm is: -0.36, -0.23, 0.11 and -0.06.

Figure 12: Mean wage residuals for executives conditional on deciles of worker and firm fixed effects

(a) Male Workers



(b) Female Workers



Note: Mean wage residuals for female executives in D10 working in a D10 firm is 0.023.

A.4 Firm effects and time to conception by cohorts

Figure 13: Average gender gap in firm premia by cohort over time to first birth

Note: In cohorts born between 1960 and 1964, the average firm fixed effects (normalized based on accommodation and food services) from model (1) for females 20 years after their first childbirth is 0.045; for males, it is 0.074.

Figure 14: Sorting and bargaining effects by cohort over time to first birth

Note: In cohorts born between 1960 and 1964, we find a sorting effect of 0.034 and a bargaining effect of -0.002 20 years after workers' first childbirth, using model (1) estimates and using male distribution across firms as reference.

A.5 Separate estimations for period 1995-2004 and 2005-2014

Table 12: Summary of two-way fixed effects models and model estimates for the 1995-2004 sample

	Model (1)		**Model (2)**	
	Female	*Male*	*Female*	*Male*
Age	0.055***	0.068***	0.066***	0.067***
	(0.001)	(0.001)	(0.001)	(0.001)
$Age^2/100$	-0.044***	-0.059***	-0.056***	-0.057***
	(0.002)	(0.002)	(0.002)	(0.002)
Experience	0.003*	0.002	0.002	0.002*
	(0.001)	(0.001)	(0.001)	(0.001)
$Experience^2/100$	-0.002	-0.002	-0.003	-0.002
	(0.002)	(0.001)	(0.002)	(0.001)
Seniority	0.018***	0.025***	0.019***	0.024***
	(0.001)	(0.001)	(0.001)	(0.001)
$Seniority^2/100$	-0.095***	-0.126***	-0.098***	-0.123***
	(0.008)	(0.007)	(0.008)	(0.007)
1 child	-	-	-0.058***	-0.046***
			(0.011)	(0.009)
2 children	-	-	-0.095***	-0.059***
			(0.012)	(0.010)
3 children	-	-	-0.190***	-0.057***
			(0.016)	(0.012)
4 children or more	-	-	-0.268***	-0.074***
			(0.029)	(0.017)
18 - age of the youngest child	-	-	-0.000	0.003***
			(0.001)	(0.000)
Number of observations	167,055	220,399	167,055	220,399
Adjusted R^2	0.78	0.85	0.78	0.85

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 13: Sorting and bargaining contributions to the gender wage gap, 1995-2004

	Normalization based on...			
	Accommodation and food services firms		*Lowest log VA per worker firms*	
Total gender gap	0.180	*100 %*	0.180	*100 %*
Gender log wage gap due to firms	0.015	*8.3 %*	0.017	*9.4 %*
including sorting effect				
Male assignment, female premia (a)	0.015	*8.3 %*	0.015	*8.3 %*
Female assignment, male premia (b)	0.012	*6.7 %*	0.012	*6.7 %*
including bargaining effect				
Male assignment, female premia (c)	0.000	*0 %*	0.002	*1.1 %*
Female assignment, male premia (d)	0.003	*1.7 %*	0.005	*2.8 %*
Number of observations	383,368			
Number of firms (10+ workers)	4,907			
Number of workers	68,903			

Calculation of sorting and bargaining are based on model (1) estimates. Line (a) reports sorting effect calculated using female premia: $\mathbb{E}[\psi^F_{J(i,t)} \mid g = M] - \mathbb{E}[\psi^F_{J(i,t)} \mid g = F]$. The corresponding bargaining effect is in line (c) and computed with male assignment in firms as reference: $\mathbb{E}[\psi^M_{J(i,t)} - \psi^F_{J(i,t)} \mid g = M]$. Oppositely (b) gives the estimates for the sorting effect measured with male premia: $\mathbb{E}[\psi^M_{J(i,t)} \mid g = M] - \mathbb{E}[\psi^M_{J(i,t)} \mid g = F]$ and (d) for the bargaining effect based on female assignment in firms: $\mathbb{E}[\psi^M_{J(i,t)} - \psi^F_{J(i,t)} \mid g = F]$. In both cases sorting and bargaining effects add up to the gender log wage due to firms: (a)+(c)=(b)+(d)= Gender log wage gap due to firms.

Table 14: Summary of two-way fixed effects models and model estimates for the 2005-2014 sample

	Model (1)		**Model (2)**	
	Female	*Male*	*Female*	*Male*
Age	0.021***	0.023***	0.027***	0.022***
	(0.001)	(0.001)	(0.001)	(0.001)
$\text{Age}^2/100$	-0.019***	-0.023***	-0.024***	-0.022***
	(0.001)	(0.001)	(0.001)	(0.001)
Experience	0.008***	0.012***	0.007***	0.012***
	(0.001)	(0.001)	(0.001)	(0.001)
$\text{Experience}^2/100$	0.006***	-0.002	0.006***	-0.002
	(0.001)	(0.001)	(0.001)	(0.001)
Seniority	0.010***	0.012***	0.011***	0.012***
	(0.000)	(0.000)	(0.000)	(0.000)
$\text{Seniority}^2/100$	-0.024***	-0.034***	-0.025***	-0.033***
	(0.002)	(0.002)	(0.002)	(0.002)
1 child	-	-	-0.045***	-0.028***
			(0.008)	(0.007)
2 children	-	-	-0.059***	-0.018*
			(0.009)	(0.008)
3 children	-	-	-0.090***	-0.021*
			(0.012)	(0.010)
4 children or more	-	-	-0.098***	-0.013
			(0.019)	(0.013)
18 - age of the youngest child	-	-	0.001	0.002***
			(0.000)	(0.000)
Number of observations	168,662	211,890	168,662	211,890
Adjusted R^2	0.86	0.89	0.86	0.89

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 15: Sorting and bargaining contributions to the gender wage gap, 2005-2014

	Normalization based on...			
	Accommodation and food services firms		*Lowest log VA per worker firms*	
Total gender gap	0.154	*100 %*	0.154	*100 %*
Gender log wage gap due to firms	0.010	*6.5 %*	0.009	*5.8 %*
including sorting effect				
Male assignment, female premia (a)	0.018	*11.7 %*	0.018	*11.7 %*
Female assignment, male premia (b)	0.013	*8.4 %*	0.013	*8.4 %*
including bargaining effect				
Male assignment, female premia (c)	-0.008	*-5.2 %*	-0.009	*-5.8 %*
Female assignment, male premia (d)	-0.004	*-2.6 %*	-0.004	*-2.6 %*
Number of observations	375,121			
Number of firms (10+ workers)	3,862			
Number of workers	61,371			

Calculation of sorting and bargaining are based on model (1) estimates. Line (a) reports sorting effect calculated using female premia: $\mathbb{E}[\psi^F_{J(i,t)} \mid g = M] - \mathbb{E}[\psi^F_{J(i,t)} \mid g = F]$. The corresponding bargaining effect is in line (c) and computed with male assignment in firms as reference: $\mathbb{E}[\psi^M_{J(i,t)} - \psi^F_{J(i,t)} \mid g = M]$. Oppositely (b) gives the estimates for the sorting effect measured with male premia: $\mathbb{E}[\psi^M_{J(i,t)} \mid g = M] - \mathbb{E}[\psi^M_{J(i,t)} \mid g = F]$ and (d) for the bargaining effect based on female assignment in firms: $\mathbb{E}[\psi^M_{J(i,t)} - \psi^F_{J(i,t)} \mid g = F]$. In both cases sorting and bargaining effects add up to the gender log wage due to firms: (a)+(c)=(b)+(d)= Gender log wage gap due to firms.