Family, firms and the gender wage gap in France

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Abstract

This paper explores how two main channels explaining the gender wage gap, namely the heterogeneity of firm pay policies and sex-specific wage consequences of parenthood, interact. We explore the firm heterogeneity channel by applying the model proposed by Card, Cardoso, and Kline 2016. After controlling for individual and firm heterogeneity, we show that the sorting of women into lower-paying firms accounts for 11 % of the average gender wage gap in the French private sector, whereas within-firm gender inequality does not contribute to the gap. Performing these decompositions all along workers' life cycle, we find evidence that this sorting mechanism activates shortly after birth. These gender-specific and dynamic firm choices generate wage losses all along mothers' careers, in addition to direct child wage penalties. After birth, mothers tend to favor firms with more flexible work hours and home proximity, which may be detrimental to their labor market opportunities, as, within these contexts, firms may gain relative monopsonic power.

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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I Introduction

This paper focuses on the gender wage gap and explores the relative importance of two of its main explanations: the heterogeneity of firms' pay policies, and the impact of parenthood. In spite of the increase in female education and labor market participation over the last decades, women continue to earn lower wages compared to men. In France in 2014, women's hourly wages in the private sector were on average 18.6 % lower than men's. These large wage discrepancies persist once productivity differentials are accounted for. For instance, after taking into account seniority, professional experience, age, level of education, occupation, part-time work, industry, firm size, and region of residence, an 8.4 % unexplained gap between men's and women's wages is still evident.

The literature has put forward different mechanisms to understand this enduring gender wage gap. On the one hand, key contributions have analyzed the role of heterogeneity of firms' pay policies and its consequences on gender inequalities. On the other hand, another strand of literature has focused on life cycle dynamics, especially parenthood, and their relationship to the gender wage gap. In this paper, we aim at bridging these two literatures and highlight how they may interact.

Part of the gender wage gap may indeed result from different pay policies, either within or between firms. Within-firm and within-job inequalities arise when women obtain lower wages than comparably productive male coworkers while doing the same job within the same firm. This may occur because of discrimination (Blau and Kahn, 2016 suggest that this mechanism cannot be completely dismissed) or, more subtly, because women do not bargain their wages as well as men do. Indeed, several contributions suggest that women tend to initiate wage negotiations less often than men, or perform less well than men when bargaining their own wages (see Bertrand, 2011 for a review of the literature).[1] Another part of the gender wage gap may come from between-firms and between-jobs inequalities. This mechanism would suggest that the gender wage gap is induced by gender segregation and/or

[1] See Bowles, Babcock, and Lai (2007); Bowles, Babcock, and McGinn (2005); Rigdon (2012); 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 (2010) for some mitigated field results.

sorting across industries, jobs, and firms. For example, Groshen (1991) highlights that occupational sorting accounts for a substantial part of the gender wage gap in the United States. Her contribution has motivated a large literature on the role of sorting of men and women across occupations and establishments in producing gender inequalities.[2] Recently, Card, Cardoso, and Kline (2016) proposed a framework based on linked employer-employee data to measure the bargaining and sorting effects through the within- and between- dimensions of the firm contributions to the gender gap.

Concurrently, several papers have documented how deeply child birth relates to wage losses for women. Wilner (2016) finds a large wage loss associated to motherhood in France, and a much smaller loss associated to fatherhood, even after controlling for human capital depreciation due to maternity leave, and for both individual and firm types of unobserved heterogeneity. Further, Kleven, Landais, and Søgaard (2018) estimate that around 80 % of the total gender wage gap in Denmark in recent years is attributable to child penalties. The authors do not take firm heterogeneity into account, but control for individual productivity through an event study. These wage penalties appear directly after birth, but also later throughout mothers' careers via the dynamic impacts of children on mothers' occupations, promotions, and firm choices. More generally, some influential contributions to the literature stress the importance of life cycle dynamics to understand the gender wage gap (Bertrand, Goldin, and Katz, 2010; Goldin, Kerr, Olivetti, and Barth, 2017; Goldin and Mitchell, 2017).

The main contribution of this paper is to bring together these two sets of explanations, and to assess to what extent the bargaining and sorting effects as defined by Card, Cardoso, and Kline (2016) are related to parenthood throughout workers' careers. In their paper, Card, Cardoso, and Kline (2016) document a sorting effect increasing sharply with workers' age. Here, we go a step further by relating this effect to child wage penalties. To do so, we take advantage of a rich matched

[2] Barth, Bryson, Davis, and Freeman (2016); Bayard, Hellerstein, Neumark, and Troske (2003); Nekby (2003); Trond Petersen (1995). Gobillon, Meurs, and Roux (2015) address this question with an original wage rank access approach using French data, and document that full-time executive women have lower access to the highest paid jobs compared to male counterparts.

employer-employee dataset for French private firms which also gathers information on family events. Following Card, Cardoso, and Kline (2016), we estimate sorting and bargaining effects using two-way (worker and firm) fixed effect models (Abowd, Kramarz, and Margolis, 1999; Lentz and Mortensen, 2010), with gender-specific firm-fixed effects. We then study sorting and bargaining effects according to parenthood status. Briefly, we find that the total firm contribution accounts for 8.2% of the total gender wage gap. The majority of this contribution comes from the sorting of mothers into lower-paying firms, and of fathers into higher-paying firms, relative to non-parents. The bargaining effects are close to zero for both parents and non-parents. The sorting effect among parents accounts for around 2 pp of the corresponding gender wage gap, compared to only 0.7 pp for non-parents.

A longitudinal approach shows that the sorting effect starts increasing right after birth, for both first and second births. For mothers who only have one child, this effect clearly widens from the first years after birth and starts decreasing 12 years later, which coincides with entry into secondary school for their child. For mothers who go on to have more children, the birth of the second child is also associated with a strong increase in the sorting effect, which strikingly never decreases thereafter. The bargaining effect also increases along the life cycle but at a slower pace. These findings stress the different dynamics in male and female behaviors when choosing - or being chosen by - their employers after having a child, probably attributable to mothers and fathers looking for different kinds of amenities in a firm. Whether these differences come from gender-specific preferences, social roles, or employers' attitudes, is beyond the scope of this paper. However, we highlight that compared to fathers, mothers tend to work in firms closer to their homes (and possibly to their children's kindergartens, schools, activities, etc.), and in firms where part-time work is more frequent. Flexible hours and home proximity may be at the expense of higher wages and fewer opportunities for promotion. Mothers are indeed less likely to be involved in firm-to-firm mobility compared to other workers. They also tend to work in areas where the industry-specific firm labor markets are more concentrated, so they are likely to face fewer outside options than fathers. Theses features are consistent with dynamics whereby promotions and mobilities are less profitable for women than for men, preventing them to climb up the job ladder at the same pace, and are in that respect related to "glass ceiling" and "sticky floors" phenomena.

Our paper is linked to Barth, Kerr, and Olivetti (2017), who find that the between-establishment gender wage gap component in the US is almost entirely due to married workers. Kleven, Landais, and Søgaard (2018) also show that, in addition to wage penalties, mothers are more often working in firms with high shares of women with young children. Our results are also consistent with the findings of Albrecht, Bronson, Thoursie, and Vroman (2017). Focusing on high-skilled Swedish workers, Albrecht, Bronson, Thoursie, and Vroman (2017) show that the career paths of men and women diverge at the time of the birth of their first child: women tend to work less, and in different types of firms. Their mobility rate is also affected. Compared to these papers, our contribution is original as it controls for firm level heterogeneity in the estimation, and directly links it to births.

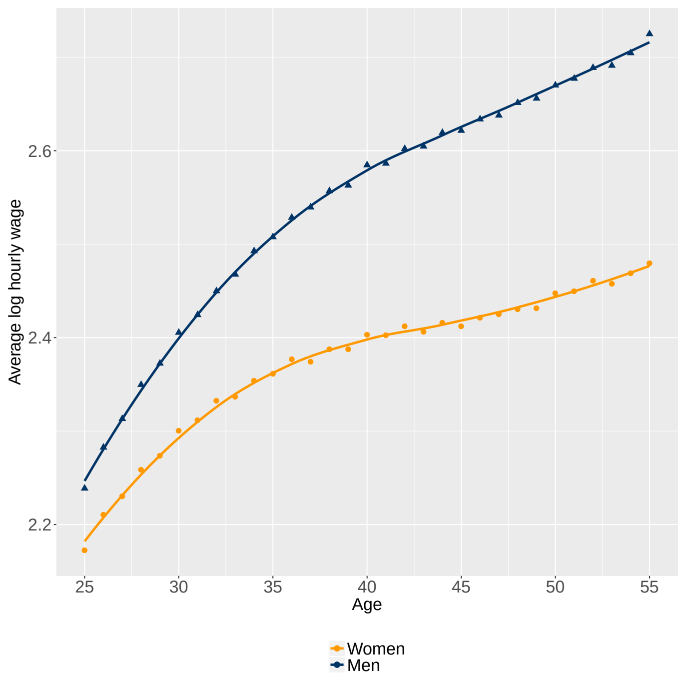
Our paper also brings new empirical elements to the literature on the decomposition proposed by Card, Cardoso, and Kline (2016). We show that the exogenous mobility and the additivity assumptions required to identify two-way-fixed effect models (Abowd, Kramarz, and Margolis, 1999) are likely to hold in the general population, but not for some occupational subgroups such as executive workers, for whom exogenous mobility is less realistic. We also provide evidence for the rent-sharing model assumption as firm-fixed effects are positively correlated with firm value-added per worker. Moreover, the comparison between our findings and those obtained by Card, Cardoso, and Kline (2016) for Portugal is of interest given the specificities of the French and the Portuguese labor markets. For France, we find that 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 the ones obtained in Portugal where firms account for 21 % of the average gender log-wage gap, with 15-20 percentage points (pp) due to the sorting channel, and 1-6 pp to the bargaining one. The discrepancies between our results and those obtained for Portugal can be related to a higher minimum wage in France, which is more than twice that of Portugal in 2016. Indeed, women in France are more likely to be paid at the minimum wage than men, and a higher minimum wage is likely to attenuate the importance of bargaining and sorting in the decomposition.

The remainder of the paper is organized as follows. Section II motivates our approach by showing how firms may impact the gender wage gap throughout workers' life cycle. Section III describes the French context and the data. The model and conditions for identification are developed in Section IV. Section V presents the results, which are further analyzed in the light of family events all along workers' life cycles in Section VI. The last section concludes.

II Children and the Gender Wage Gap

The gender hourly wage gap in France was 18.6 % in 2014. This gap increases dramatically over the life cycle as shown in Figure I. It starts at less than 5 % for individuals aged 25 to about a 20 % by age 40. From that age, the gender gap continues to increase until retirement but at a slower pace.

Figure I. Average log hourly wage by age and gender

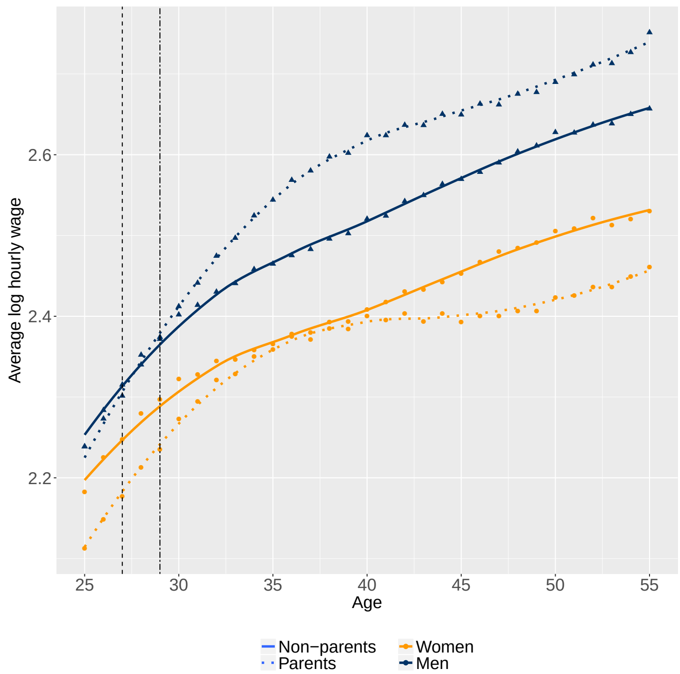


Source: DADS, *Panel Tous Salariés*. Scope: Metropolitan France. Self-employed farmers, craftsmen, shopkeepers, trainees, apprentices and private household workers are excluded. Note: average log hourly wage for forty-year-old workers is 2.40 among women, and 2.58 among men.

This age profile may be driven by parenthood as shown by Figure II: the gender wage gap deepens more for parents (dotted lines) than for non-parents (solid lines).

For the latter, the gap stabilizes after age 35, whereas it keeps increasing for parents. If parenthood leads to more family constraints for mothers than fathers, mothers will look for jobs (or stay in jobs) that are more compatible with their family lives. These jobs are likely to be concentrated in firms that offer more flexibility in working hours and other similar amenities, but where wage policies may be less generous than in other firms (see Goldin, 2014).

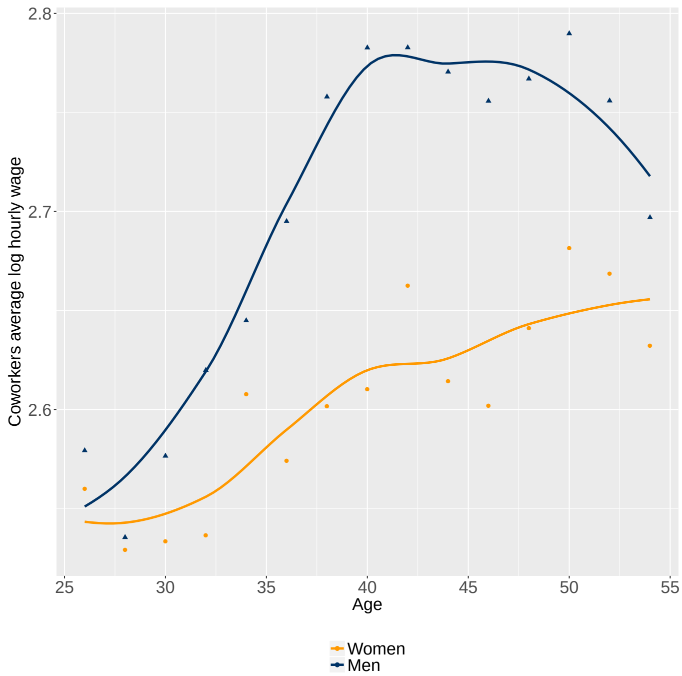
Figure II. Average log hourly wage by age, gender and family status



Source: DADS, *Panel Tous Salariés*. Note: average log hourly wage for forty-year-old workers is 2.40 among mothers, 2.41 among childless women, 2.62 among fathers, and 2.52 among childless men. Vertical dotted and dashed lines represent median age at first birth for women (27) and men (29). Before they have their first child, individuals are assigned to the non-parent group.

Figure III supports this firm-level explanation. It shows that the gap between women's and men's average coworker wages sharply increases with age. If coworkers' average wages reflects firms' pay policies, this suggests that worker sorting across firms with different pay policies may contribute to the gender gap. However, confirming a causal effect of firms on the gender wage gap requires controlling for the individual heterogeneity of workers. Moreover, the difference in coworkers' wages may reflect both a wage differentiation between men and women within the same firm, or their segregation in different firms. Our analysis thus requires to estimate the firm-specific component of wages for men and women, in order to determine to what extent the sorting and bargaining channels impact the gender wage gap, especially when workers have children, given the role of family characteristics on wages as evidenced in Figure II.

Figure III. Coworkers' average log hourly wage by age and gender



Source: DADS, *Panel Tous Salariés*, and comprehensive DADS files for coworkers' wages. Note: average coworkers' log hourly wage for forty-year-old workers is 2.61 among women and 2.78 amond men.

III Institutional setting and data sources

France shares common features and trends with other OECD countries regarding gender discrepancies in the labor market. The French female employment rate (61 % en 2015) is close to the OECD average, with a gap between male and female employment rates somewhat smaller than the OECD average (OECD, 2017). Furthermore, 30 % of employed women work part-time, as do 8 % of employed men. We find similar proportions in the dataset we use in this paper. The gender hourly wage gap we observe has slightly decreased over the period from 22.1 % in 1995 (*i.e* .175 log difference) to 18.6 % in 2014 (*i.e* .160)[3]. In particular, this decrease is due to

[3]see Figure A.1 in the appendix

hourly wage gains around the year 2000 when workers benefited from a reduction in working hours (with the introduction of the "35 hour week") while monthly wages were held constant.

III.A Institutional setting

Some specificities of the French institutional context are worth mentioning. The wage bargaining system combines a national minimum wage set by the government which applies to all workers,[4] as well as collective negotiations, at the industry and firm level. At 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 of industry-level agreements, which therefore apply to all wage-earners (whether unionized or not). At the firm level, employers and unions bargain on wage increases, provided that wages remain above the industry wage floors and above the minimum wage (see Fougère, Gautier, and Roux, 2016).

Between 1998 and 2014, the national minimum wage increased in line with the median wage, representing about 62 % of the full-time private sector median wage; this ratio remained stable between 1998 and 2014 (see Figure A.5 in the appendix).[5] The increase in the gross minimum wage did not lead to a rise in the labor costs of low-paid workers because firms benefited from substantial social security exemptions and tax credits for workers paid at the minimum wage or immediately above.[6] Hence, even though the minimum wage remained stable relative to the median between 1998 and 2014, the cost of hiring a worker paid at the minimum wage decreased from 122 % to 110 % of the gross minimum wage. The combination of a high minimum wage level with relatively low labor costs entails that low-paid workers

[4] Apprentices and some workers under 18 can be paid 80 % of the minimum wage.

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

[6] The first exemption aimed to compensate the rise in hourly wages implied by the '35 hour working week', cf *Loi n° 2003-47 du 17 janvier 2003 relative aux salaires, au temps de travail et au développement de l'emploi,* the so called *Lois Fillon.* Another decrease in the minimum wage labor cost can be attributed to the 2013 tax credit for competitiveness and employment (CICE), which applies to all workers paid up to 1.6 times the minimum wage, cf *Loi n° 2012-1510 du 29 décembre 2012 de finances rectificative pour 2012.*

are protected in two ways: they are not discarded from employment, and are paid at a relatively higher wage than their individual productivity would predict, and they are also directly protected by the rules updating the minimum wage level.

These two protections are all the more important to be considered here as women are over-represented around the minimum wage: 13.6 % of female workers in our estimation sample have a wage equal to or below 1.1 times the minimum wage, compared to 7.4 % of male workers. Therefore, the French case, combining a high minimum wage with a compulsory collective bargaining system, provides a compelling example of how a highly protected labor market relates to the gender wage gap (Blau and Kahn, 2003).

III.B Data Sources

Our main data come from the linkage of two administrative datasets, the *Déclarations annuelles de données sociales* (DADS, Annual Declarations of Social Data) database with the Census. We use the panel subsample extracted from the exhaustive DADS database, constructed by Insee (the French National Institute of Statistics) for research purposes. This panel has been updated every year since 1967 using the wage information which firms have to report annually for payroll and fiscal purposes, for each employee. This reporting is mandatory. The panel has had a linked employer-employee structure since 1975: it contains both the firm unique identifier, which comes from the French firms register, and the person unique identifier (social security number). The agricultural 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, giving a representative sample of roughly $1/24^{\text{th}}$ of the French employed population. The panel statistical unit is the worker $\times$ firm $\times$ year level: for each worker at a given year, we know the firm they worked for, their occupation, and how much they earned.

We use the same wage concept as Abowd, Kramarz, and Margolis (1999) and Postel-Vinay and Robin (2002), and Wilner (2016) more recently. We consider wage as net of social contributions (but before income tax), which corresponds to the wage information reported by firms to the fiscal services for income tax purposes. This measure is therefore of great quality and contains all wages and salaries, any paid overtime, benefits in kind, all bonuses and indemnities (including shift work), including those paid once a year and those paid after contract termination if they exceed the industry-negotiated levels. Hence, this variable is particularly appropriate for our analysis as it accurately reports the wage components which can be negotiated (bonuses) by employees. The sole limitation is that it does not completely cover profit-sharing schemes: the panel only includes remuneration which is directly paid to the employee and not saved. However, this caveat is limited as profit-sharing schemes account for only 3 % of gross earnings across all workers (and 4 % for executives), whereas bonuses represent around 17 % of earnings (resp. 17 %), according to the 2010 Structure of earnings survey. Card, Cardoso, and Kline (2016) use a hourly wage calculated as the worker's base salary plus any regular earning supplements divided by the worker's usual work hours.[7] As we have access to the number of hours worked by each worker in the private sector since 1995, our dependent variable is the hourly wage, focusing on private sector employees from 1995 to 2014.[8]

We supplement the matched employer-employee data with individual information on workers, including their education, and the number and birth dates of their children using the Permanent Demographic Sample (EDP). The EDP is a large-scale socio-demographic panel gathering information on all births, marriages and deaths since 1968 from the registry office, along with census information from 1968, 1975, 1982, 1990 and 1999. Information from the annual census surveys (which have replaced the exhaustive census since 2000) are also integrated. The sample corresponds to a 4 % survey of the population living in France. Similarly to the DADS panel, selection into the EDP is based on date of birth, and the linkage of these datasets corresponds to around 13 % of the DADS panel. We also gather

[7] Card, Heining, and Kline (2013) resort to total earnings corrected for top-coding at the Social Security maximum. They divided earnings by days worked by those in full-time jobs (they have no information on hours worked).

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

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*), which provides financial information on the firm, such as valued-added (from 2012 to 2014), surplus, income statements, and balance sheet items (sales, exports, investments). In addition, the firm level data also allows identifying the main collective agreement in force at the firm level.

Our whole sample contains 1,632,185 observations (individual $\times$ year). After taking into account sample restrictions due to the identification, which we detail below, our 'Estimation sample' used to estimate bargaining and sorting effects contains 912,784 observations, relating to 102,048 employees working in 89,908 connected firms (for 89,855 of which we have access to the number of workers and for 43,590 to financial information in 2014). Tables I and II for workers and III for firms describe the two samples. The structures of both samples are similar, however executives are slightly overrepresented in the estimation sample, as are short-term contract holders, since they are more likely to move between firms, and workers in firms with less than 10 employees, because in the estimation procedure these small firms are grouped together. High value-added firms are overrepresented in the estimation sample as they are less likely to fail.

Table I. 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 annual income (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 %

Source: DADS, *Panel Tous Salariés*. 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 II. 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 %

Source: DADS, *Panel Tous Salariés*. 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 III. 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)

Source: DADS, *Panel Tous Salariés*. Note: Firms in the entire sample have on average 78.7 workers. Firms in the after estimation sample have on average 101.7 workers.

IV Disentangling within and between-firm contributions to the gender wage gap

IV.A 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:

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

This reduced form equation is obtained 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 components, one that is fixed over time, $\bar{S}_{J(i,t)}$, a 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, which 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.

IV.B 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 firm effect average gap:

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

The first term (i) reflects the average difference between men and women firm's components 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). As for any Blinder-Oaxaca decomposition, the decomposition is not unique, and the choice of the reference group may not be inconsequential.

IV.C Identification and Firm-effect Normalization

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 for this type of model, we group together 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 restrictions, so this estimate is obtained from the set of workers employed in firms belonging to both the male and female largest connected sets. Additional exogeneity assumptions are required for the unbiasedness of equation (2). Card, Cardoso, Heining, and Kline (2017) propose a set of empirical checks to challenge these assumptions. As presented in detail in Appendix A.2, our sample satisfies these requirements.

As detailed in Abowd, Creecy, and Kramarz (2002), firm-fixed effects are identified up to a constant, which requires normalization. This must be done consistently in order to make possible the comparison between levels of male and female firm-fixed effects given that they are estimated from two different groups. Card, Cardoso, and Kline (2016) assume that female and male premia obtained from rent-sharing are null in firms when there is structurally little rent to share, and therefore little risk of sharing differentials between female and male workers.[9] Following this idea, we choose as a reference group, *i.e.* a group of firms for which both male and female firm-fixed effects are zero on average, the industry generating on average the lowest valued-added per worker; in our sample this is the hospitality and food services industry. The normalization is made after the estimation, and does not affect the estimation of the marginal impact of time-varying covariates, nor the sorting effect estimate. It impacts only the bargaining effect estimate and the total firm contribution on the gender gap.

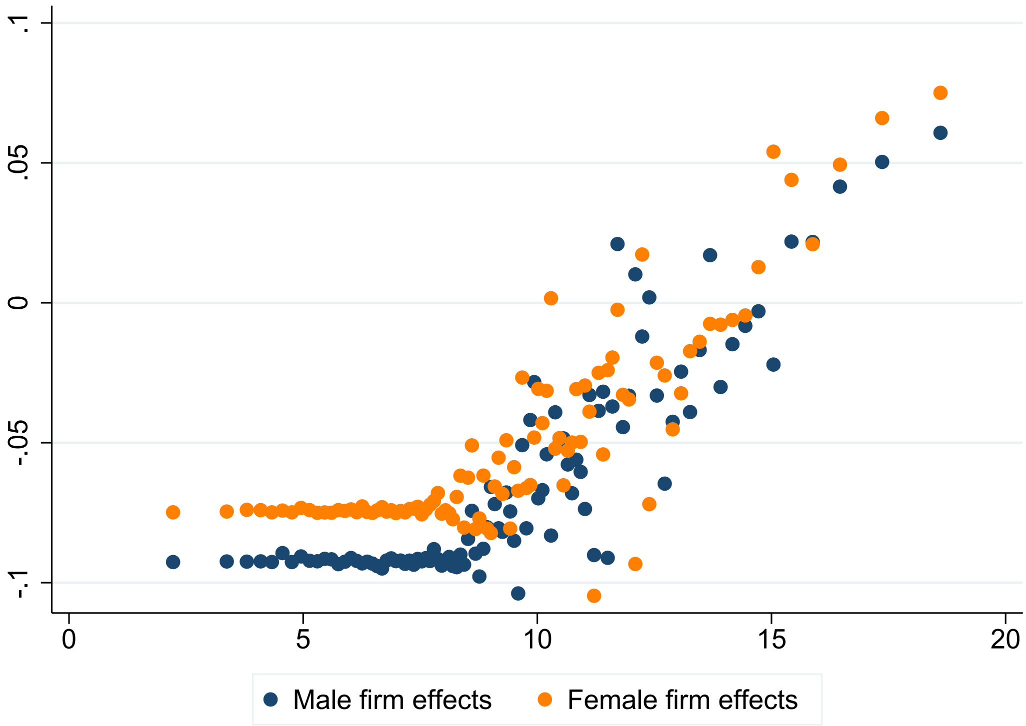
To check for the robustness of our choice, we also use the group of firms with the lowest value-added per worker as an alternative normalization, and we fix to zero the average firm effects of this group of firms. The choice of the threshold defining this group is driven by Figure IV: above a log-value-added per worker of approximately 8, there is a positive relation between the productivity of the firm and the premia female and male workers get. This result is consistent with the rent sharing theory used to derive the model.[10] The optimal level of log value-added per worker under

[9]The bargaining effect identification requires only that, in the reference group, firms tend to give similar rents to men and women.

[10]We conduct a further analysis of the estimated firm-fixed effects in Appendix A.3. It shows that the positive relationship between value added and firm-fixed effects holds after controlling for firm level variables such as firm size, composition, assets, industry, etc.

which a firm is considered in the zero fixed effect group is 8.3.[11] 26,162 firms are below this threshold, representing 29.1 % of firms and 9.4 % of worker observations.

Figure IV. Firm effects according to log per capita value-added



Source: DADS, *Panel Tous Salariés*. Lecture: Firms in the dual connected set are grouped into 100 bins according to their log value-added per worker. For each bin we plot its 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.

V Results

V.A Direct family pay gap

Table IV compares the estimates associated with time-varying worker characteristics for different specifications. In addition to worker (columns (1) and (2)) or firm and worker individual fixed effects ((3) and (4)), we control for quadratic functions of age, experience on the labor market, seniority in the firm ((1) and (3)), and family structure using the number and the age of children ((2) and (4)).

[11] The threshold is the value t^* minimizing the sum of the root mean square errors of the following model estimated for men and women: $\psi_J = \begin{cases} a & \text{if } \log \text{VA per capita} < t \\ b + c\,(\log \text{VA per capita} - t\,) & \text{otherwise.} \end{cases}$

Table IV. One-way and Two-way fixed effect model estimates

	(1)		**(2)**		**(3)**		**(4)**	
	Female	*Male*	*Female*	*Male*	*Female*	*Male*	*Female*	*Male*
Age	0.034***	0.041***	0.040***	0.039***	0.034***	0.043***	0.041***	0.041***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
$Age^2/100$	-0.035***	-0.043***	-0.041***	-0.041***	-0.032***	-0.042***	-0.038***	-0.039***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Experience	0.017***	0.020***	0.017***	0.020***	0.012***	0.013***	0.011***	0.013***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$Experience^2/100$	-0.004***	-0.009***	-0.004***	-0.009***	-0.004***	-0.006***	-0.004***	-0.006***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Seniority	0.008***	0.011***	0.008***	0.011***	0.010***	0.012***	0.010***	0.012***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$Seniority^2/100$	-0.031***	-0.043***	-0.032***	-0.042***	-0.030***	-0.041***	-0.032***	-0.040***
	(0.002)	(0.001)	(0.002)	(0.001)	(0.002)	(0.001)	(0.002)	(0.001)
1 child	-	-	-0.056***	-0.012***	-	-	-0.049***	-0.017***
			(0.004)	(0.004)			(0.004)	(0.004)
2 children	-	-	-0.069***	-0.001	-	-	-0.064***	-0.004
			(0.004)	(0.004)			(0.005)	(0.004)
3 children	-	-	-0.102***	-0.000	-	-	-0.099***	-0.005
			(0.006)	(0.005)			(0.006)	(0.005)
4 children or more	-	-	-0.114***	-0.006	-	-	-0.113***	-0.006
			(0.010)	(0.007)			(0.010)	(0.007)
18 - youngest child age	-	-	0.001***	0.002***	-	-	0.001**	0.002***
			(0.000)	(0.000)			(0.000)	(0.000)
Worker fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm fixed effects	-	-	-	-	Yes	Yes	Yes	Yes
Number of observations	403,728	515,967	403,728	515,967	403,728	515,967	403,728	515,967
Adjusted R^2	0.78	0.83	0.78	0.83	0.80	0.85	0.80	0.85

Source: DADS, *Panel Tous Salariés*. Standard errors in parentheses - * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Year dummies, individual fixed effects, firm fixed effects (for 2FE models) and interactions between child-related variables and a dummy for being born on October 2 or 3 (demographic data collection was not complete for people born those days) are included but not reported.

Returns to age are higher and more concave for men relative to women when we do not account for family structure. However, once we control for them, age parameter estimates are almost identical for men and women. This discrepancy is due to much larger wage penalties associated to motherhood for women (mothers with one child suffer a 4.9 % wage penalty relative to women with no children; the equivalent penalty for fathers is 1.7 %). Using the same dataset as us, Wilner (2016) measures the parenthood pay gap with a two-way fixed effect model from 1995 to 2011. He finds a 4.7 % wage penalty for women after their first child but no significant effect for men.[12] Accounting for children does not change the returns to professional experience nor to seniority for male and female workers, whereas one could expect the reverse if we assume that there is no human capital accumulation during maternity leave (or even a loss of human capital).

Some differences between one-way and two-way fixed effect models are also worth noting. Accounting for firm-fixed effects slightly reduces the penalty associated with children for mothers, while it increases fathers' penalties. Hence, part of the wage penalty for mothers is likely to go through the firm effect channel, but this is less clear for fathers. However, the differences between coefficients are small (0.7 percentage points per child) and not statistically significant. Whether introducing a firm effect impacts the estimation due to within or between firm differences remains to be determined.

V.B Overall sorting and bargaining effects

We now turn to the results of the decomposition of the firm contribution to the gender wage gap as detailed in equation (3). The sorting and bargaining effects are presented in Table V. Including family structure or not provides almost identical results, so we only display here the results not accounting for them (results including family characteristics are reported in Table A.2 in the appendix).[13] The gender wage

[12] Wilner (2016) computes professional experience differently from us, as he uses an experience variable corrected for maternity leave. Here, maternity leave is not excluded from experience and seniority, and interestingly we find the same child penalty as he does.

[13] As a robustness check, we also report in Table A.3 in the Appendix the results when removing the observations relating to a child's birth year, as the wage perceived by a mother at this moment may reflects more the maternity compensation than her productivity. The results of sorting and

gap amounts roughly for 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.[14] 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.

Table V. 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			

Source: DADS, *Panel Tous Salariés.* Calculation of sorting and bargaining are based on model (3) 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).

The bargaining effect is very small, and if anything negative, representing about -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. Estimates of the bargaining effect and of the total contribution of firms to the gender wage gap vary depending on the normalization choice. However, results are similar with both our reference groups: the hospitality and food services group, and firms with the lowest value added per worker. Overall, the positive role of firms on the gender wage gap comes only from the sorting of women into low-paying firms with

bargaining remain the same.

[14] We comment sorting effects computed from female premia, and the corresponding bargaining effect computed with male assignment in firms as reference. For completeness, we also report results for the opposite reference in Table V: the results show no differences.

respect to men. This is in line with the results of Card, Cardoso, and Kline (2016) who find that the firm contribution is mainly driven by sorting - which accounts for 75% of Portuguese firm contribution. However, in contrast to our results, Card, Cardoso, and Kline (2016) find also a positive contribution through the bargaining channel, which accounts for about 20 % of the Portuguese gender wage gap.[15]

V.C The role of institutional factors

The results for Portugal also differ from ours in that the overall contribution of firms to the gender gap is smaller. Altogether, national institutional settings may play an important role in interpreting results. Indeed, Blau and Kahn (2003) show from international comparisons that protective labor market institutions are negatively related to the level of the national gender wage gap. Wage settings in France are subject to a high minimum wage relative to Portugal, and more women are paid at the minimum wage than men. French workers also benefit from collective negotiations applicable to all workers of a given industry. This is likely to affect both sorting and bargaining effects: such institutions reduce the bargaining margins, and a higher minimum wage may also lead to a lower sorting effect, as low-paid workers, are also more protected *between* firms. Hence, the fact that the French minimum wage is higher than the Portuguese one could account for a lower contribution of firms to the gender wage gap in France, with both smaller sorting and bargaining effects. This explanation is in line with the bargaining effects obtained separately for periods before (1995-2004) and after (2005-2014) the rapid minimum wage growth. The French bargaining effect is higher when we estimate the model on the 1995-2004 period, when the minimum wage was lower (detailed results are provided in Appendix A.8).

Another singular aspect of the French labor market that should be taken into account in the interpretation of our results is the collective wage bargaining system. Studying the firm-specific gender gap according to firm characteristics, we see that these collective agreements explain a large share of the variance of the bargaining

[15] In a recent working paper, Bruns (2016) applies the same methodology to recent German data, and finds a total firm contribution to the gender wage gap of around 26 %, decomposed into a 25-31 pp sorting effect and a -5-0 pp bargaining effect.

effect (see Table A.4 in the Appendix). Very few other firm-level variables explain the disparities in within-firm gender gaps. In particular, we do not find any relationship between the within-firm gender gap and the firm value-added per worker, nor firm assets or size. In contrast, the firm-specific bargaining effect is positively related to the share of executives in the firm, which relates to the stronger bargaining effect for this group (see Table VI). The bargaining effect is also negatively correlated to the share of women among those executives. In addition to that, the within-firm gender gap is positively correlated with the share of women among white and blue collar workers. This suggests that a sorting mechanism is also likely to occur across jobs within firms. Finally, the within-firm gender gap is positively correlated with the share of women and men paid at the minimum wage gap. Interestingly, a one percentage point in the within-firm gender gap ((%F-%M) at the min. wage) is associated with a 0.264 percentage point increase in the firm effect gender gap. This reinforces the idea that the minimum wage policy may protect low productivity workers, and thus increases their relative bargaining power.

Last, we note that the bargaining effect is assumed by construction to be null on average for the reference group. However, if in the reference group there are more women than men at the minimum wage, and assuming that minimum wage *de facto* protects more wage-earners paid at the minimum wage, the real bargaining effect in the reference group may in contrast be negative. If so, the bargaining effect on the entire sample may be overestimated. This issue may appear if the shares of women and men at the minimum wage in the reference group differ markedly. However, this is not likely to be a key issue as in the two reference groups we use, the gender gap in minimum wage exposure is relatively smaller than in the entire sample: 24.0 % for women and 15.7 % for men (for the hospitality and food services) and 21.9 % and 14.9 % (lowest value-added firms), versus 13.6 % and 7.4 % (entire sample). Moreover, despite these differences, the two normalizations yield similar bargaining effects.

VI Life Cycle Analysis

VI.A Sorting, bargaining effects, and parenthood

Using the firm-fixed effect estimates presented above, we compute sorting and bargaining effects for different subgroups depending on occupation, education level, age, birth cohort, and parenthood status. For a group P, we can express this as:

$$\mathbb{E}\left[\psi^M_{J(i,t)} \mid g = M, P\right] - \mathbb{E}\left[\psi^F_{J(i,t)} \mid g = F,\, P\right] \tag{4}$$

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

Table VI presents the total gender gap for different subgroups, as well as the firm contribution to this gap, and how the firm contribution is decomposed between sorting and bargaining effects. Overall, the sorting effect dominates the bargaining effect for almost all population groups, with the notable exception of executive workers, for whom the bargaining effect is even larger than the total firm contribution. This group is indeed likely to have more opportunities to actively bargain wages compared to other workers, an exercise in which women may be less efficient than their male counterparts (Bowles, Babcock, and Lai, 2007; Bowles, Babcock, and McGinn, 2005; Small, Gelfand, Babcock, and Gettman, 2007). This may also reflect some vertical job segregation within firms: female executives may have less access to the top executive and highest paid jobs (Gobillon, Meurs, and Roux, 2015), which may lead to high within firm gender gaps. These two mechanisms, wage negotiation and vertical segregation, may be interlinked, as suggested by Greig (2008), who documents a correlation between one's propensity to negotiate and the rate of advancement.[16]

Further, sorting effects are roughly stable across education groups. The sorting

[16]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 A.6 to A.8 show that the exogenous mobility condition and the rent-sharing model are not credible within this group of workers: executives tend to receive better wages when they move to a different firm, whatever the productivity level of their new employers. Models with endogenous mobility (Lamadon, Manresa, and Bonhomme, 2015) would probably suit better executives' behavior. This issue should be the object of further research.

effect is somewhat smaller among more educated workers, and, since the total gender wage gap is twice as large among university graduates than among less educated workers, the relative share of the sorting effect decreases with the educational level. The difference in sorting effects between high and low-educated workers is far smaller than the difference found by Card, Cardoso, and Kline (2016) for Portugal, where the share of the sorting effect on the total gender gap is 11.5 pp larger for low-educated than for high-educated workers. The bargaining effect for low-educated workers is negative and close to the general sample estimate whereas the ones for high-school and university graduates are small but positive, probably for the same reasons as for executives.

Table VI. Sorting and bargaining effects conditional on job positions, education level, age and birth cohort

	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
College, University	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

Source: DADS, *Panel Tous Salariés*. Note: Calculation of sorting and bargaining are based on model (3) estimates and calculated using female premia and male assignment in firms as reference. We use the accommodation and food services normalization. When we compare female and male blue collars the sorting effect amounts to 0.006 and the bargaining effect is -0.003.

Age group comparisons indicate a clear increase in the sorting effect with age, confirming the pattern shown by Card, Cardoso, and Kline (2016). The estimate is twice as large for workers aged 40 or more (2.1 %) than for younger employees (1.1 %). By contrast, the bargaining effect does not seem to vary with age, and

Table VII. Sorting and bargaining effects conditional on parenthood

	N	Sorting	Bargaining	Total firm contribution	Total gender wage gap
Parents	357,031	0.020	0.000	0.020	0.214
Non parents	105,107	-0.003	-0.004	-0.007	0.024
Parents 45+	129,145	0.026	0.000	0.026	0.288
Non parents 45+	29,974	-0.014	0.002	-0.011	0.084

Source: DADS, *Panel Tous Salariés*, born on October 1st or 4th. Note: Calculation of sorting and bargaining are based on model (3) 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 0.000.

remains stable around our baseline estimate. Ultimately, the total firm contribution follows the same pattern as the sorting effect, and increases strongly with age. This finding directly relates to family structures. As shown in Table VII, the sorting effect is much larger for parents (2.0 %) than for non-parents (-0.3 %).[17] To isolate the effect of parenthood from the effect of age, we restrict our attention to parents and non-parents older than 45, supposing that most people reaching this age without a child will remain childless. For this subgroup, the difference in the sorting effect is even more pronounced: 2.6 % for parents versus -1.4 % for non-parents, which confirms that the gap is more likely to be due to parenthood than age.Differences in sorting effects between parents and non-parents (2.3 pp for all workers, and 4.0 pp for workers older than 45) come in addition to the direct effect of children on wages, which amounts to 3.2 pp difference for the first child when controlling for both individual and firm effects.[18]

VI.B Sex-specific sorting into firms appears with childbirth

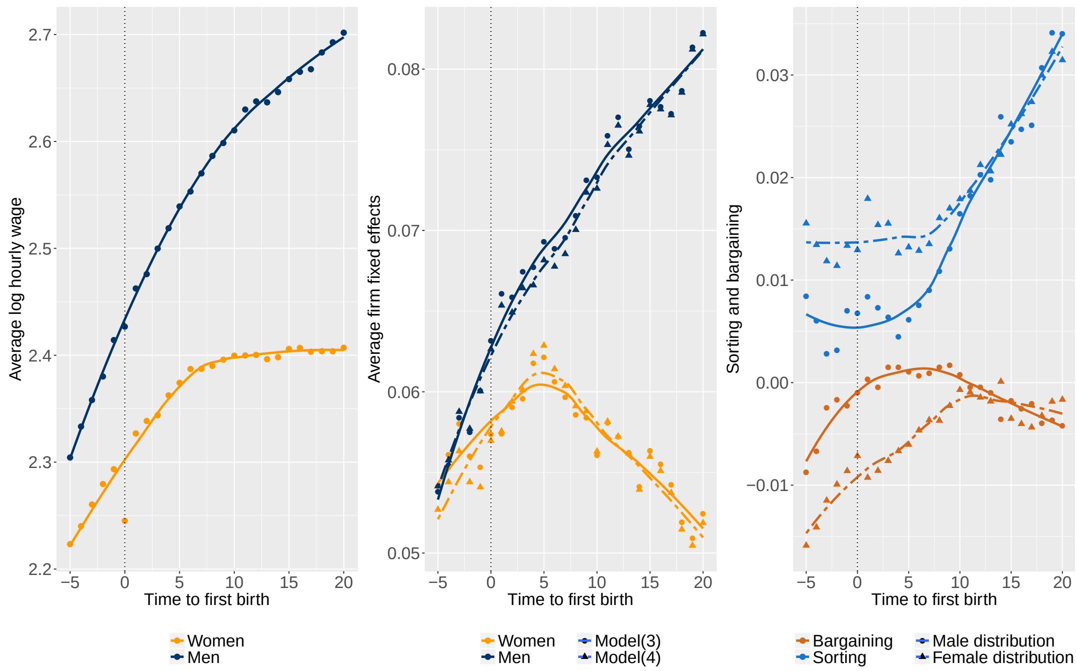
We highlighted in the previous section that the sorting of men and women across firms increases dramatically with age and with parenthood. We now analyze more precisely how inequalities grow along the family life cycle. We thus focus on workers

[17] For the rest of the analysis, we focus on people born on October 1st and 4th. As already mentioned, birth data for individuals born on the 2nd and 3rd of October was subject to data collection issues, and is likely to be measured with error. This subsample is still representative of the French population working in the private sector.

[18] The wage penalty associated with the first child is 4.9 % for women, and 1.7 % for men, translating into a 3.2 pp difference.

who eventually have children, and particularly on the birth of the first two children. Figure V shows the decomposition of the gender gap from five years before the birth of the first child to twenty years after. The left panel shows the average log hourly wage by gender, the central one plots the average firm-fixed effects for both populations, and the right panel provides the decomposition of the gap displayed in the central panel into a bargaining effect and a sorting effect. Five year before childbirth, women and men experience comparable firm premia. However, female firm effects fall rapidly behind men's after the first birth. For the first five subsequent years, women experience a slower growth rate in firm premia compared to men. Then, we observe a permanent drop in female firm premia.

Figure V. Firm premia, sorting, and bargaining effects over time to first birth



Source: DADS, *Panel Tous Salariés*, born on October 1st or 4th. Note: The average firm-fixed effect (normalized based on accommodation and food services) from model (3) for females 10 years after their first childbirth is 0.056; for males, it is 0.073. At this time to first birth, using model (3) estimates and male distribution into firms as reference gives a sorting effect of 0.016 and a bargaining effect of 0.001. The lines were obtained by smoothing the averages.

The right-hand side of the figure shows the corresponding sorting and bargaining effects (computed from Model (3) estimates).[19] The bargaining effect is relatively flat, between -0.01 and 0. It increases slightly early in the career, driving part of the

[19]Including or not family events (model (4) or (3)) in the wage equation does not change the results.

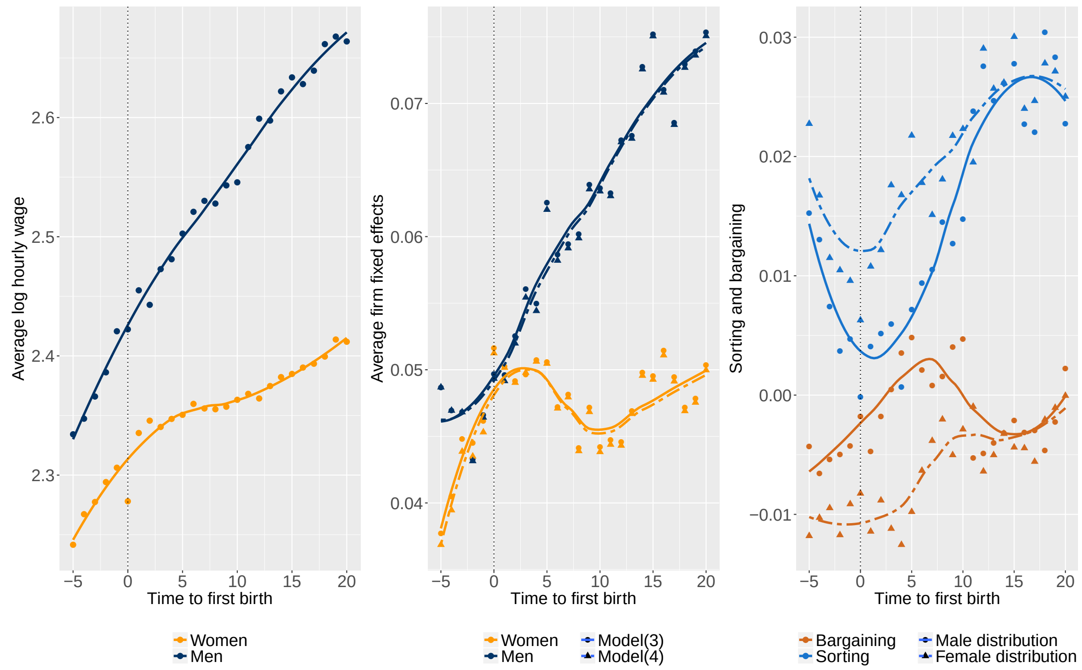
gender firm effect gap increase which occurs before the first childbirth, but it levels off about ten years later. In contrast, the sorting effect dramatically increases after the first child from approximately 0.01 in the early careers to around 0.03 twenty years after the first child, persisting at high levels until the end of careers. Compared to men, women tend to work in low-paying (and probably more "family-friendly") firms, or to move less to high-paying firms after the birth of their first child. These differences persist in the long term; this may be due to the birth of subsequent siblings. But strikingly the sorting effect never decreases, even twenty years after entry into parenthood: women do not experience upward mobility between firms 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 women in their early careers, when mobility is most profitable. Again, there is a double penalty for women having children: childbirth directly affects wages (Table IV), and accentuates a long-term gender divergence in firm effects.[20]

The sorting effect appears to be contemporaneous to the decline of the average firm effects for women several years after the birth of the first child (Figure V). We conduct similar analyses focusing on parents who will have one child only (Figure VI), from parents having two children or more (Figure VII). For parents who will only have one child, we observe a small increase in the female firm effect right after birth, followed by a rapid decrease in the following years. The combination of this jump followed by a drop can be explained by a temporary exit and then re-entry of some women working in low-paying firms, maybe due to take up of parental leave. Interestingly, the average firm effect of women increases again 11 years after birth, although it does not catch up with men's. One could associate these patterns to the entry of children into junior high school, when school hours increase and children are given more autonomy.

The decline in the average female firm effect and the increase of the sorting effect for parents of two children or more are clearly associated to the second birth. For

[20]Figures A.9 and A.10 in appendix, detailed by birth cohorts, 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. An analysis by occupation and education is provided in Figures A.11 and A.12 in the appendices.

Figure VI. Firm premia, sorting, and bargaining effects over time to first birth for parents with one child



Source: DADS, *Panel Tous Salariés*, born on October 1st or 4th. Note: The average firm-fixed effect (normalized based on accommodation and food services) from model (3) for females 10 years after their first childbirth is 0.044; for males, it is 0.064. At this time to first birth, using model (3) estimates and male distribution into firms as reference gives a sorting effect of 0.015 and a bargaining effect of 0.005. The lines were obtained by smoothing the averages.

this group, the first child does not produce very important changes in the pattern of the firm effects right after the first birth (Figure VIIa), the time at which the graph is centered. By contrast, Figure VIIb, which is centered at the birth of the second child, indicates a strong break right after this second birth. In contrast to mothers of one child, mothers of two or more children do not experience a change in the firm effect after 12 years: the decline in firm effect -which is mainly an increase in the sorting effect- continues until at least 15 years after the birth of the second child.

As shown by Figure VIII, the selection of women in the labor market may be an important factor explaining part of these trends. This figure shows the number of parents observed each year in our dataset. Any reduction in this number may be linked to individuals leaving the scope of our analysis.[21] The arrival of a child is linked to a reduction in the number of observations for both men and women, but this reduction is far sharper among women than men, so births can be associated

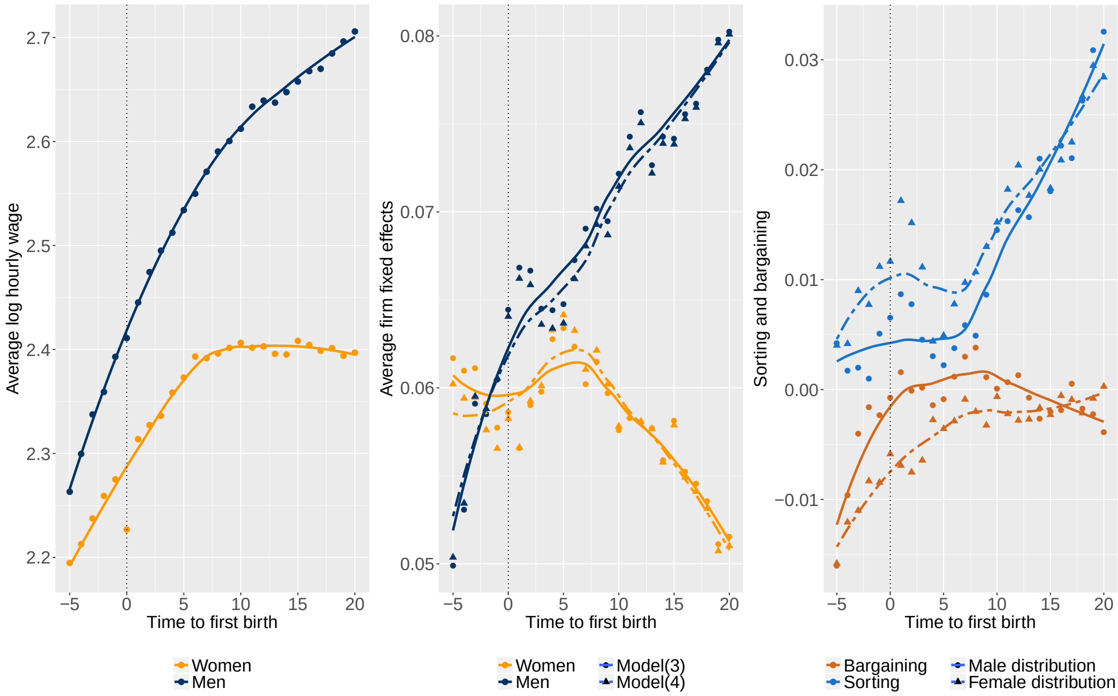
[21]They may leave employment or move to one of the sectors that we do not observe, like self-employment or the public sector.

to a peak in the difference in the number of observations.

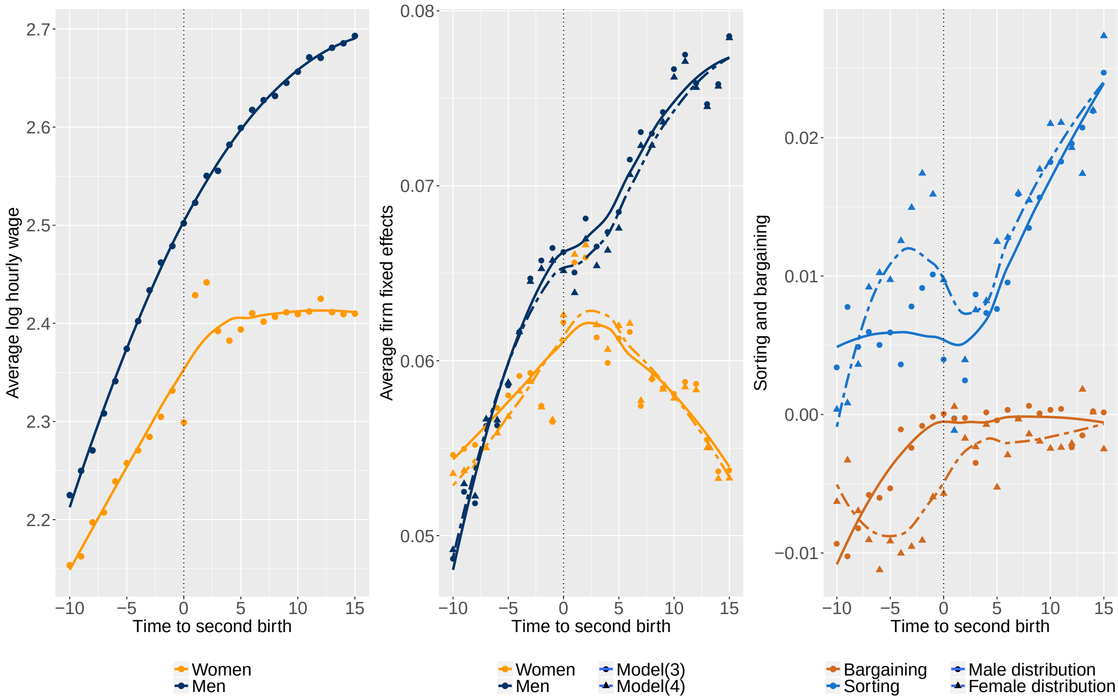
This potential selection effect has to be taken into account when interpreting the sorting effect described earlier. The decrease of the sorting effect during the couple of years following childbirth may be explained by a temporary over-representation of women working in higher paying firms, who tend to be more attached to the labor market. This would also mean that we should interpret the sorting effect as a lower bound of what we would observe if men and women were selecting into the labor market in the same way after child birth. However, as the difference between number of men and women in the labor market returns to its prebirth level also suggests that some mothers do come back to the labor market a few years later.

Figure VII. Firm premia, sorting, and bargaining effects over time to births for parents with two children or more

(a) First child arrival

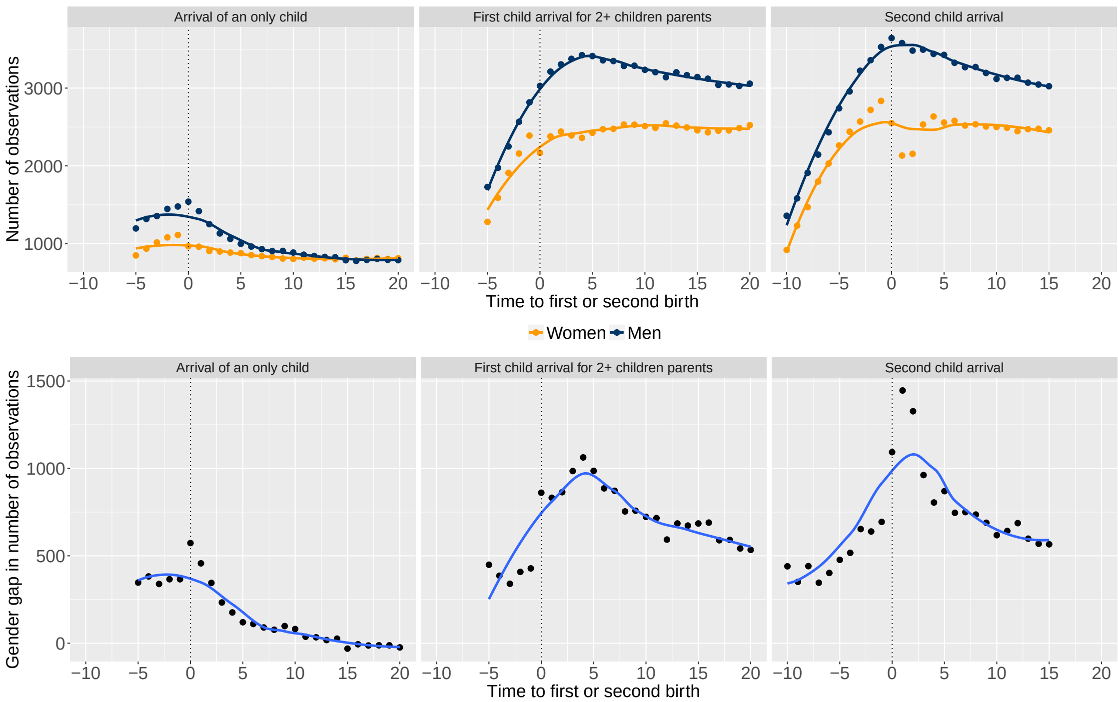


(b) Second child arrival



Source: DADS, *Panel Tous Salariés*, born on October 1st or 4th. Note: The average firm-fixed effect (normalized based on accommodation and food services) from model (3) for females 5 years after their second childbirth is 0.061; for males, it is 0.069. At this time to second birth, using model (3) estimates and male distribution into firms as reference gives a sorting effect of 0.008 and a bargaining effect of 0. The lines were obtained by smoothing the averages.

Figure VIII. Number of observations at each period

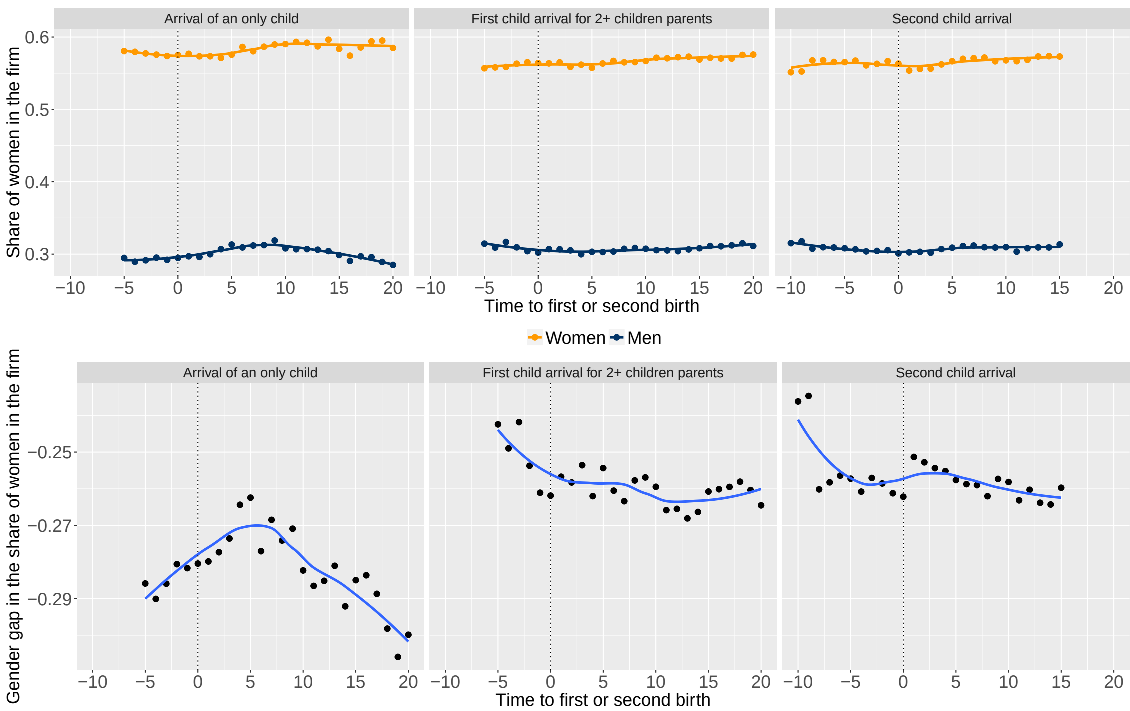


Source: DADS, *Panel Tous Salariés*, born on October 1st or 4th. Note: In our sample, we observe 967 females the year of the birth of their only child. At this time, we have 1,540 male observations, which corresponds to a 573 gender gap. The lines were obtained by smoothing the averages.

VI.C Characterizing the firms where parents work

The sorting effect highlighted before can also be described by looking at the composition of the firms where mothers and fathers work. Figure IX shows that at any moment in their careers and whatever the number of children they have, mothers work in firms where a majority of women work (between 55 and 60 % vs 30 % for fathers). This relative segregation of women increases along the career: from the beginning of their careers for parents of two or more children, and later in life for parents of one child. Segregation diminishes sharply and very temporarily in the two years following a second birth: this observation is in line with more women from low productive firms, in which women are overrepresented, exiting the labor market, compared to women in higher-paying firms with more male coworkers. The gender gap in the female coworkers' rate then goes back to its pre-birth value, as women return to work.

Figure IX. Gender Segregation Between Firms



Source: DADS, *Panel Tous Salariés*, born on October 1st or 4th, and comprehensive DADS files for coworkers' characteristics. Note: On average, five years after the birth of their only child, men are employed in firms with 31 % of female workers. For women, the share goes up to 57 %, corresponding to a -26 percentage point gap.

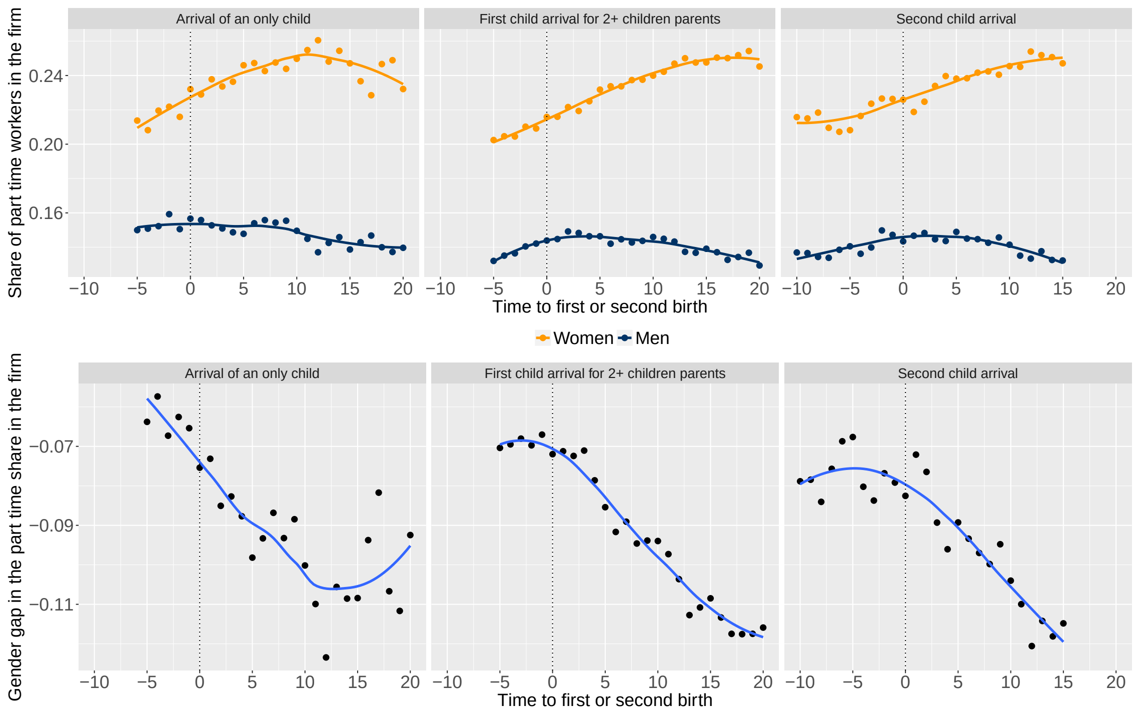
The exact mechanisms at stake remain to be determined. We are not directly able to identify to what extent this phenomenon comes from the supply or the demand side of the labor market: whether high-paying firms lay off or are reluctant to hire mothers, or whether mothers with young children look for firms offering better work conditions at the expense of higher wages. Nevertheless, gathering descriptive elements of the type of firms mothers tend to work for sheds light on the underlying mechanisms sorting mothers into low-paying firms.

The first dimension we look at is working time conditions at the firm level. Our results suggest that after birth, mothers sort or are sorted in firms more likely to offer flexible hours. Figure Xa shows the share of part-time workers in the firms where mothers and fathers work. This share remains relatively constant for men regardless of the final number of children they have and the rank of the birth: it is stable at about 14-15 %, and begins declining about 15 years after childbirth. The share of part-time workers increases over time for all mothers. The extent of the increase varies from 22 to 26 % for mothers with only one child (left panel), and from 20 to 25 % for women of more children (middle and right panels). The difference between the two curves displayed in the bottom panel of Figure Xa does not show a clear break at birth for mothers of one child, but a slightly stronger break in the case of women with more children at both the first and subsequent births. It is also interesting to note that in the case of mothers of one child only, the share of part-time workers starts to decrease after about 12 years post-birth, which coincides with the increase in the firm effects of these mothers described earlier.

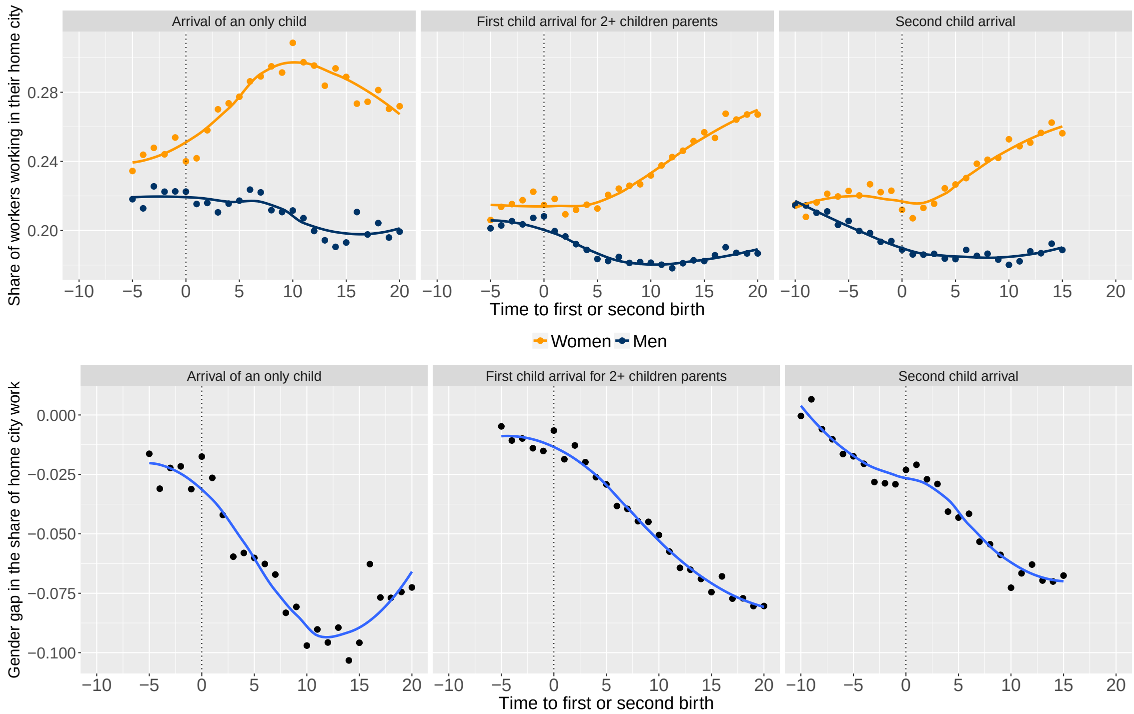
The sorting into lower-paying firms may also be due to an increased need of mothers to work close to their homes. To verify this, we look at the evolution of the share of parents working in the municipality where they live (Figure Xb). This share appears to be quite close for mothers and fathers before the birth of the first child, and this for both parents who will have one child only (left panel), and for parents of more children (central and right panels). In both cases, the share of fathers working in their town of residence continues to decrease after the arrival of a child, whereas it starts increasing strongly for mothers. For second births, the difference between mothers and fathers starts decreasing before birth, moreover, the change of the slope is stronger than for first births. Using the as-the-crow-flies home to work distance

Figure X. Firm Characteristics by gender and time to first birth

(a) Share of part time workers in the firm



(b) Share of workers employed in their home city



Source: DADS, *Panel Tous Salariés*, born on October 1st or 4th, and exhaustive DADS files for coworkers' characteristics. Note: On average, five years after the birth of their only child, 22 % of men work in the city where they live, to be compared with 28 % for women, corresponding to a -6 percentage point gap.

(measured as the distance between the centroids of the town of residence and the town of work) shows similar patterns - see Figure A.13 in the appendices.

So, mothers tend to work closer to their homes, which may ease the conciliation of professional and family life, but may bring additional constraints; in particular, this geographical restriction may give mothers fewer labor market opportunities. It may lead to less favorable wage offers if firms have a monopsony power on the local job market (Azar, Marinescu, and Steinbaum, 2017). Women could also be less likely than men to work in firms if it implies longer commuting times, or if these firms offer fewer options to reconcile work and family life. To explore these hypotheses, we plot the firm-to-firm mobility rate (Figure XIa). Firm-to-firm mobility rates decrease as workers age. The trends are similar for men and women but the latter experience a large drop at the time of birth. This pattern can be observed for all types of parents regardless of their final number of children and the rank of the child: the mobility rate gap increases from close to 0 to about 4 percentage points around birth. Subsequently, the mobility rate remains lower for mothers than for fathers for up to 10 years after the first child. Again, it is remarkable that the mobility rate of mothers of one child only catches up, and even surpasses, fathers' rates after about 12 years. For parents with more children, this rate converges to similar values at the end of their careers.

Further, we build an index of the local concentration on the labor market (Figure XIb) to investigate how far local constraints faced by women may provide monopsonic power to their potential employers. We observe 16,895 combinations of labor market areas and industries in our data. For k a labor market area, i an industry, we compute the average Herfindahl-Hirschmann index of concentration of the labor force between 2010 and 2014 using the share of workers employed in each firm j from this area $\times$ industry combination:

$$\overline{HH}_{ki} = \sum_{j \in i \times k} \left(\frac{\sum_{t=2010}^{2014} \text{number of workers}_{jt}}{\sum_{j \in i \times k} \sum_{t=2010}^{2014} \text{number of workers}_{jt}} \right)^2 .$$

We normalize it so that differences in HH indices are not driven by differences in industries. In that goal, we shift indexes by the grand average index minus the industry-average index (each industry-average normalized index is equal to the normalized index grand average): $\overline{HH}_{ki}^{\text{norm}} = \overline{HH}_{ki} + \overline{HH}_{..} - \overline{HH}_{.i}$.

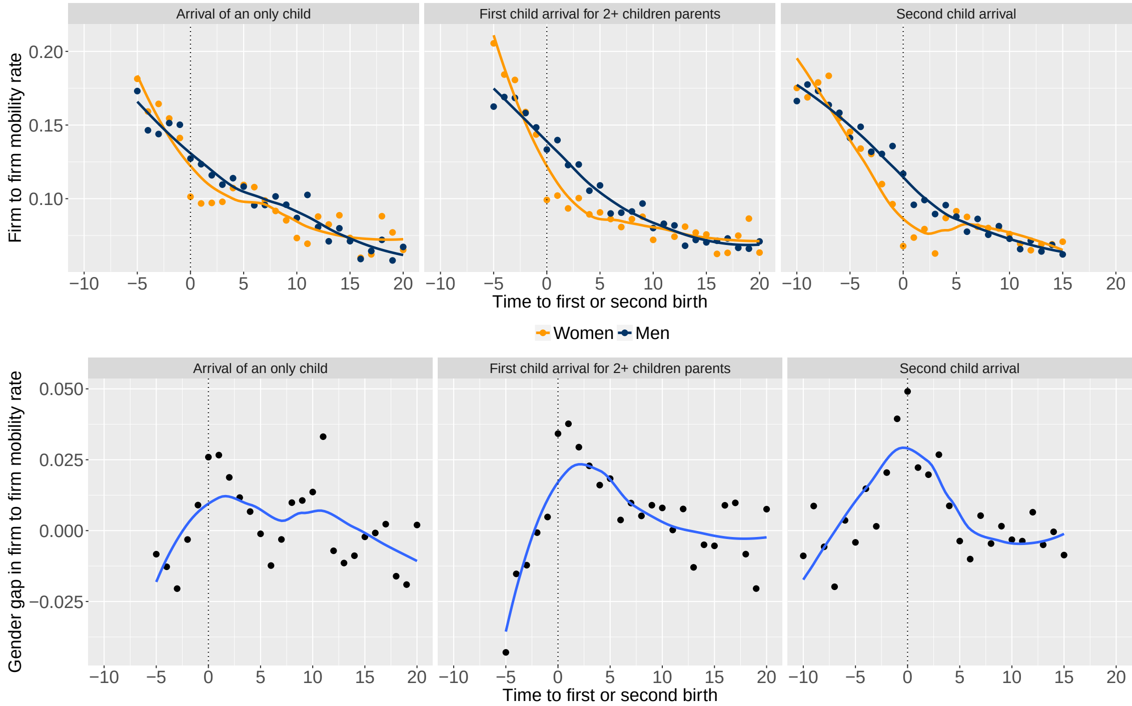
Figure XIb shows the relationship between such HH indices and the timing of births. Compared to fathers, mothers tend to work in places and industries where firms have more local monopsony positions. These concentration indices tend to decrease with age but births are associated to important breaks in this trend. The gender difference tends to diminish before births and increases afterwards. This is particularly striking for parents having one child only. The break is less sharp for other parents, although it also appears at the moment of the second birth.

Taken together, these results suggest that the sorting effect after childbirth may correspond to mothers' needs for flexible working hours and proximity of their workplace. This may lead them to work for local firms which benefit from some monopsonic power in a market defined at the area-industry level, and may thus apply less generous pay policies.[22]

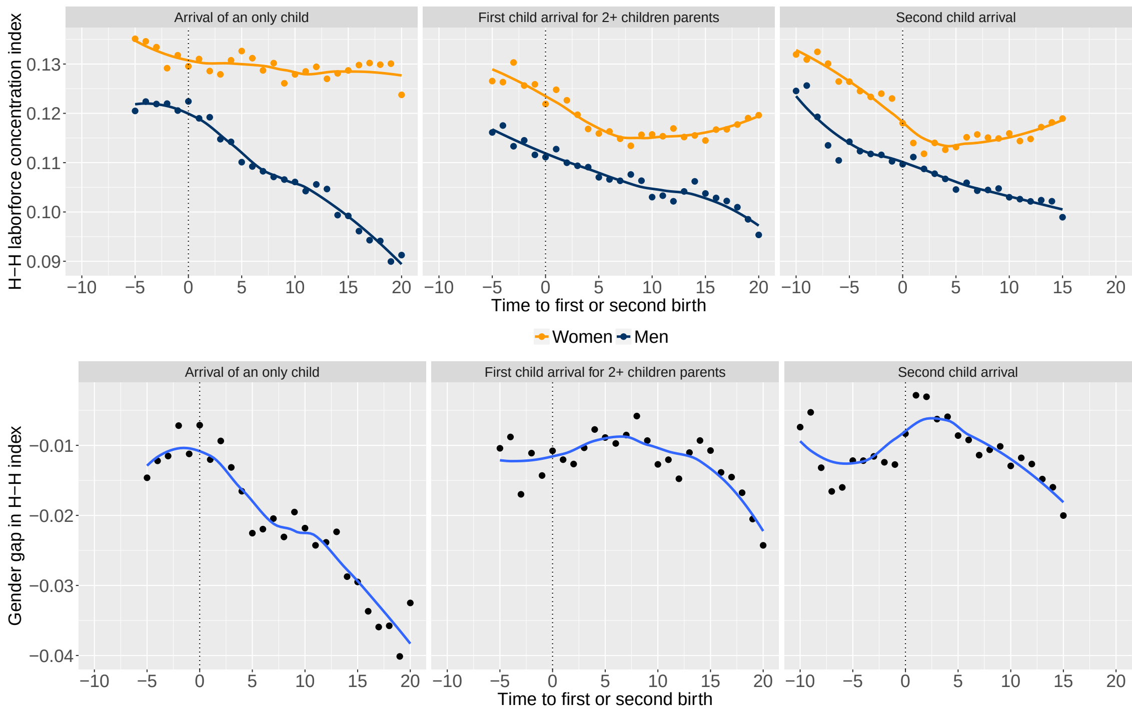
[22]We also looked at the share of minimum wage worker in the firm (see Figure A.14 in the appendix) which does not indicate clear break trends related to births.

Figure XI. Firm sorting and gender-specific labor market constraints

(a) Firm to firm mobility rate



(b) HH index on local labor markets



Source: DADS, *Panel Tous Salariés*, born on October 1st or 4th. Note: On average, five years after the birth of their only child, men are at work in areas × industries where the labor force concentration index is 0.110. The average index at that time to birth is 0.133 for women, corresponding to a -0.023 gender gap.

VII Conclusion

In this paper, we investigate the within and between-firm contributions to the gender wage gap in interaction with parenthood. 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 show that our sample fulfills with the requirements for identification of the two-way fixed effect models. 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 observed characteristics and individual heterogeneity. However, we estimate a positive sorting effect (around 11 % of the total gender gap in hourly wage), suggesting that firms contribute to the gender wage gap as women tend to be at work in firms paying lower wages than men at comparable productivity.

We find that the sorting effect is much larger for parents and older workers, showing a large heterogeneity in the firm effect over the life cycle. This implies a double child penalty: in addition to the direct child wage penalty, mothers experience wage losses through sorting between firms, and both effects are of comparable magnitudes: about 3.2 pp wage loss is due to the birth of the first child and about 1.8 pp to sorting later in the careers.

Focusing on parents, and relating the evolution of the sorting and bargaining effects to births, we show that the sorting effect clearly arises after birth, and deepens afterwards. For parents of one child only, our analysis shows an increase in the sorting effect after the birth which only declines 12 years later. For parents of more children, this effect is sharper after the birth of the second child, and increases up to 15 years after the second birth.

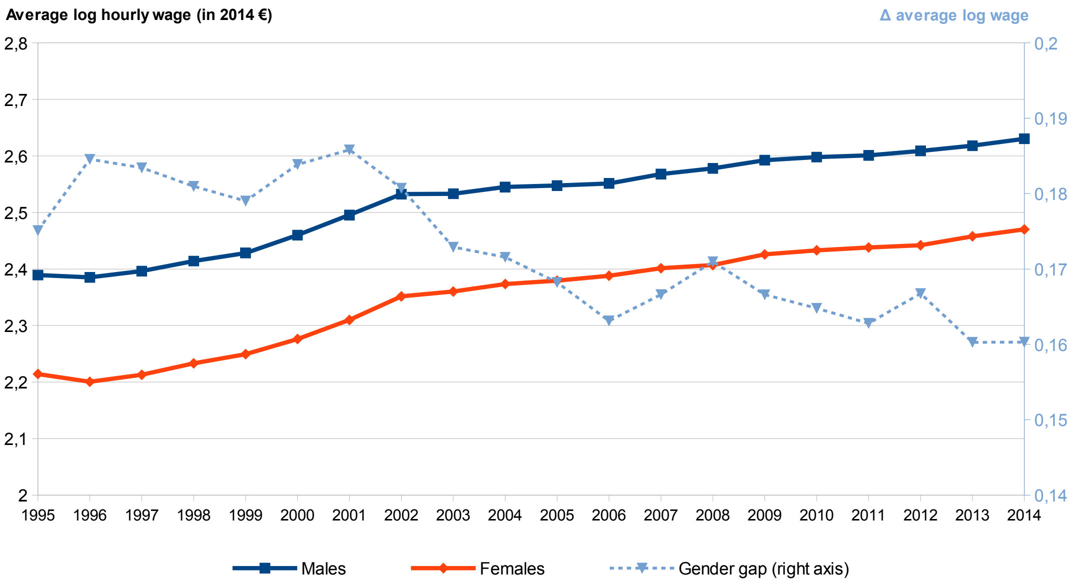
We also show that the deepening of the sorting effect coincides with important differences in the characteristics of firms where mothers and fathers work: mothers tend to work for firms with more women, allowing flexible hours - where workers more often work part-time, and which are closer to their homes. This may reflect mothers' need to combine family and work lives, their preferences, or gender-specific social attitudes of both workers and employers. Flexible hours and home proximity may be at the expense of higher pay policies, probably partly related to the monopsonic local positions of such "family-friendly" firms.

Altogether, a significant part of the gender wage gap we observe is due to gender-specific dynamics in careers and sorting between different firms after a birth, and that women's wages eventually bear most of the cost of children.

A Appendix

A.1 More trends

Figure A.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 12.9 € for females and 15.9 € for males. The gender gap corresponds to 18.6 % of the average hourly male wage.

A.2 Conditions for identification

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:

$$\mathbb{E}\left[(r_{it}-\bar{r}_i)\left(\mathbf{1}_{J(i,t)=j}-\frac{1}{T}\sum_{t=1}^{T}\mathbf{1}_{J(i,t)=j}\right)\right]=0,\quad \forall j\in\{1...J\}. \tag{5}$$

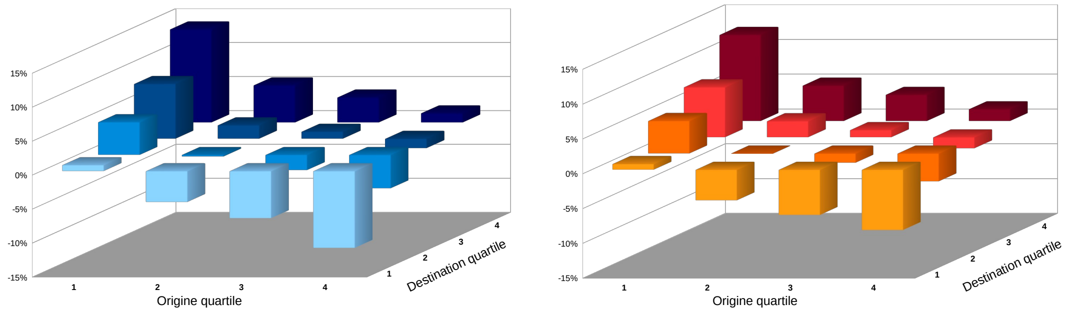
This condition must hold for each firm, j, and states that on average unobserved shocks r_{it} should not depend on the mobility of individuals (one should notice that the condition is only active for firm movers otherwise $\mathbf{1}_{J(i,t)=j}-\frac{1}{T}\sum_{t=1}^{T}\mathbf{1}_{J(i,t)=j}=0$). In 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 A.2a (men) and A.2b (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.

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 A.3a and A.3b 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

Figure A.2. Average wage change for movers conditional on origin and destination firm average wage

(a) Male Workers (b) Female Workers

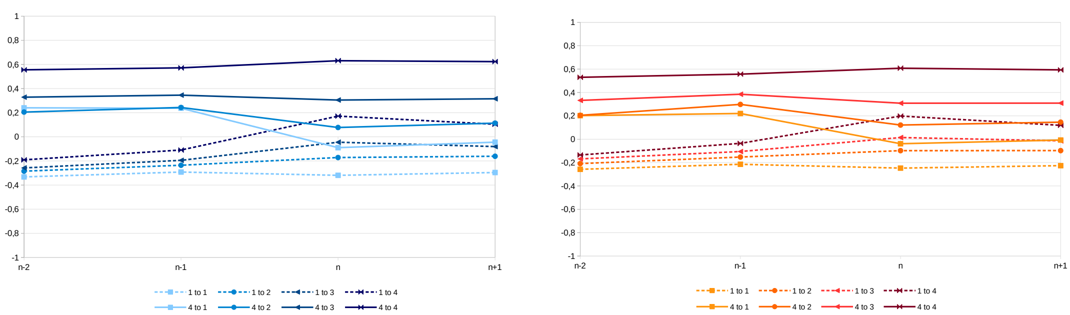


Source: DADS, *Panel Tous Salariés*. Note: A female worker moving from a firm paying average wages below the bottom quartile (Q1) of the wage distribution to a firm above top quartile (Q4) gets a average 12.4 % increase in wage. Symmetrically a female going from a Q4 firm to a Q1 firm can expect a 8.7 % drop in wage.

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 A.3. 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

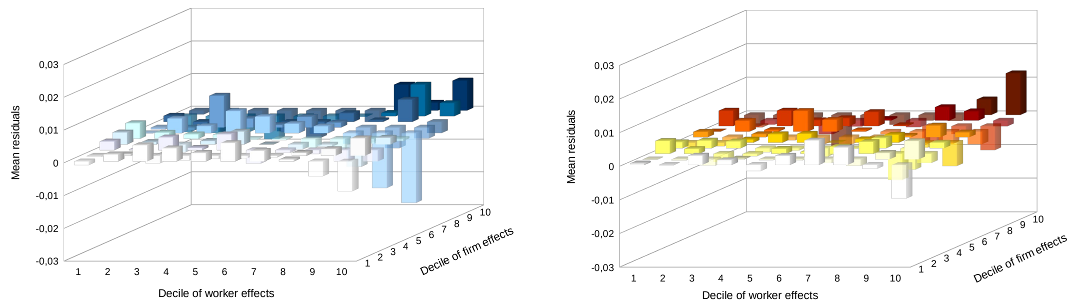


Source: DADS, *Panel Tous Salariés*. Note: A female worker moving from a firm paying average wages below the bottom quartile (Q1) of the wage distribution to a firm above top quartile (Q4) has an average residual wage of -0.13 two years before moving, -0.04 the year before her mobility, 0.20 the year she moves and 0.12 the following year. Symmetrically a female going from Q4 firm to a Q1 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 A.4a and A.4b). 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 A.4a and A.4b 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 for Portugal by Card, Cardoso, and Kline (2016) and for Germany by Card, Heining, and Kline (2013).

Figure A.4. Mean of wage residuals conditional on deciles of worker and firm-fixed effects

(a) Male Workers (b) Female Workers



Source: DADS, *Panel Tous Salariés*. 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.

A.3 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.[23] For each firm we average the estimated male and female firm-fixed effects, and we regress this variable on a

[23]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.

set of firm covariates. Findings for different specifications are presented in table A.1. 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 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 and collective agreements. 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. 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 is not significant 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.

Table A.1. Average firm premia and firm characteristics

	(1)	(2)	(3)
Constant	0.057^{***}	0.049^{**}	0.089^{**}
	(0.017)	(0.020)	(0.037)
Average age	-0.002^{***}	-0.001^{**}	-0.002^{***}
	(0.001)	(0.001)	(0.001)
Average prof. experience	0.001^{*}	0.001	0.001
	(0.001)	(0.001)	(0.001)
% part time workers	0.026^{***}	0.031^{***}	0.033^{***}
	(0.010)	(0.010)	(0.011)
% minimum wage earners	-0.243^{***}	-0.239^{***}	-0.262^{***}
	(0.013)	(0.014)	(0.015)
(%F-%M) paid at the min. wage	0.015	0.017	0.014
	(0.015)	(0.015)	(0.016)
% open end contracts	0.049^{***}	0.043^{***}	0.021^{**}
	(0.007)	(0.008)	(0.009)
% executives	0.185^{***}	0.199^{***}	0.174^{***}
	(0.012)	(0.015)	(0.016)
% clerks	0.047^{***}	0.057^{***}	0.042^{***}
	(0.011)	(0.011)	(0.012)
% white collar workers	-0.043^{***}	-0.030^{***}	-0.031^{***}
	(0.008)	(0.009)	(0.011)
% female among executives	0.022^{***}	0.019^{**}	0.014
	(0.008)	(0.008)	(0.009)
% female among clerks	0.001	0.001	0.003
	(0.006)	(0.006)	(0.006)
% female among white collars	-0.011^{**}	-0.014^{***}	-0.010^{*}
	(0.005)	(0.005)	(0.005)
%female among blue collars	0.004	0.005	0.000
	(0.008)	(0.008)	(0.008)
Value-added per worker (*ref=1*)			
Quartile 2	-0.014^{**}	-0.015^{***}	-0.008
	(0.006)	(0.006)	(0.006)
Quartile 3	0.018^{***}	0.014^{**}	0.020^{***}
	(0.006)	(0.006)	(0.007)
Quartile 4	0.036^{***}	0.033^{***}	0.038^{***}
	(0.007)	(0.007)	(0.009)
Exporting firm	0.001	0.001	0.001
	(0.005)	(0.005)	(0.006)
Assets per worker (*ref=1*)			
Quartile 2	-0.008	-0.005	0.003
	(0.006)	(0.006)	(0.007)
Quartile 3	-0.001	0.003	0.005
	(0.006)	(0.007)	(0.007)
Quartile 4	0.015^{**}	0.016^{**}	0.006
	(0.006)	(0.007)	(0.008)
Number of workers /1000	0.000^{*}	0.000	0.001
	(0.000)	(0.000)	(0.001)
Industry dummies	-	Yes	Yes
Collective agreement dummies	-	-	Yes
Adjusted R^2	0.257	0.267	0.297
Number of observations	6518	6491	6491

Source: DADS, *Panel Tous Salariés*. Standard errors in parentheses - * $p <$ 0.01, ** $p < 0.05$, *** $p < 0.01$. Accommodation and food services normalization is used.

A.4 Sorting and bargaining effects computed with model (4) estimates

Table A.2. Sorting and bargaining contributions to the gender wage gap (Model (4))

	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			

Source: DADS, *Panel Tous Salariés*. Calculation of sorting and bargaining are based on model (4) 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).

A.5 Sorting and bargaining effects computed after removing birth year wages

Table A.3. Sorting and bargaining contributions to the gender wage gap when removing wage observations on birth years for either men and women

	Normalization based on...			
	Accommodation and food services firms		*Lowest log VA per worker firms*	
Total gender gap	0.168	*100 %*	0.168	*100 %*
Gender log wage gap due to firms	0.014	*8.3 %*	0.012	*7.1 %*
including sorting effect				
Male assignment, female premia (a)	0.018	*10.7 %*	0.018	*10.7 %*
Female assignment, male premia (b)	0.019	*11.3 %*	0.019	*11.3 %*
including bargaining effect				
Male assignment, female premia (c)	-0.004	*-2.4 %*	-0.006	*-3.6 %*
Female assignment, male premia (d)	-0.005	*-3.0 %*	-0.007	*-4.2 %*
Number of observations	871,875			
Number of firms (10+ workers)	10,776			
Number of workers	101,091			

Source: DADS, *Panel Tous Salariés*. Calculation of sorting and bargaining are based on model (3) 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).

A.6 Collective agreements, firm performance and gender bargaining gap

We regress the average within firm-fixed effect gap between male and female workers on firm workers' characteristics and firm financial results. Models (1) to (3) refer to the same specifications as those in table A.1.

Table A.4. Within-firm gender gap in firm premia and firm characteristics

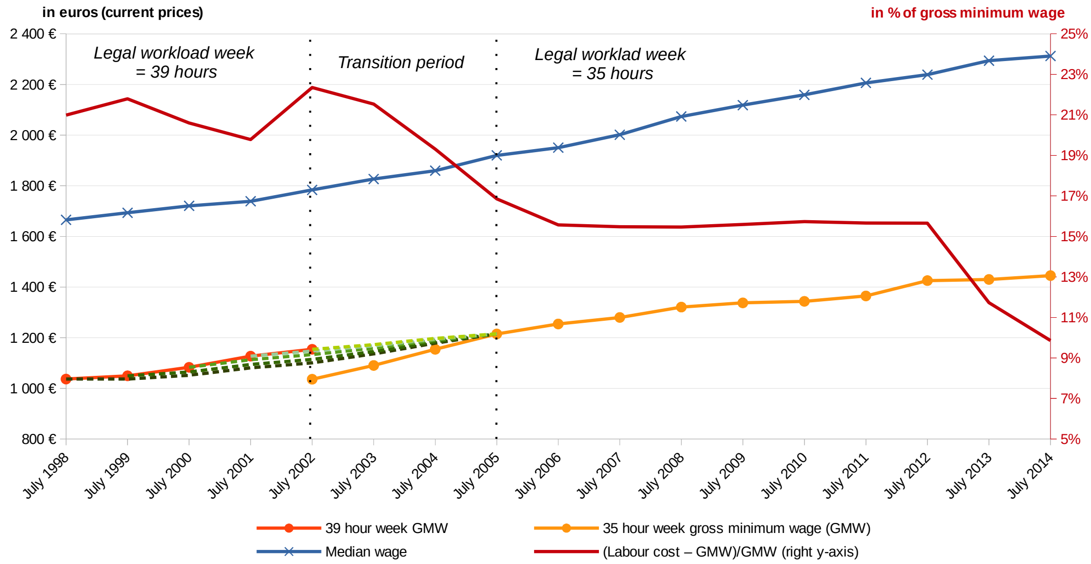
	(1)	(2)	(3)
Constant	0.011	-0.009	0.032
	(0.031)	(0.035)	(0.064)
Average age	-0.002**	-0.002	-0.002*
	(0.001)	(0.001)	(0.001)
Average prof. experience	0.002	0.001	0.001
	(0.001)	(0.001)	(0.001)
% part time workers	-0.028*	-0.014	-0.016
	(0.017)	(0.017)	(0.019)
% minimum wage earners	-0.035	-0.048*	-0.070**
	(0.024)	(0.025)	(0.028)
(%F-%M) at the min. wage	0.264***	0.257***	0.269***
	(0.028)	(0.028)	(0.030)
% open end contracts	0.010	0.001	-0.009
	(0.013)	(0.014)	(0.017)
% executives	0.125***	0.142***	0.135***
	(0.021)	(0.026)	(0.029)
% clerks	0.036*	0.049**	0.035
	(0.019)	(0.020)	(0.023)
% white collar workers	0.020	0.029*	0.033
	(0.014)	(0.016)	(0.021)
% female among executives	-0.118***	-0.115***	-0.121***
	(0.015)	(0.015)	(0.015)
% female among clerks	-0.012	-0.013	-0.013
	(0.011)	(0.011)	(0.011)
% female among white collars	0.043***	0.045***	0.042***
	(0.009)	(0.010)	(0.010)
%female among blue collars	0.075***	0.073***	0.073***
	(0.014)	(0.015)	(0.016)
Value-added per worker (*ref=1*)			
Quartile 2	0.002	0.001	0.004
	(0.010)	(0.010)	(0.011)
Quartile 3	-0.004	-0.004	-0.004
	(0.011)	(0.011)	(0.012)
Quartile 4	0.010	0.003	0.008
	(0.012)	(0.013)	(0.016)
Exporting firm	0.003	-0.008	-0.008
	(0.008)	(0.009)	(0.010)
Assets per worker (*ref=1*)			
Quartile 2	0.003	0.001	0.008
	(0.010)	(0.011)	(0.012)
Quartile 3	-0.010	-0.012	-0.005
	(0.010)	(0.012)	(0.013)
Quartile 4	-0.011	-0.007	-0.001
	(0.011)	(0.013)	(0.015)
Number of workers /1000	0.000	0.000	-0.001
	(0.000)	(0.000)	(0.001)
Industry dummies	-	Yes	Yes
Collective agreement dummies	-	-	Yes
Adjusted R^2	0.037	0.041	0.062
Number of observations	6518	6491	6491

Source: DADS, *Panel Tous Salariés*. Standard errors in parentheses - * $p < 0.01$, ** $p < 0.05$, *** $p < 0.01$. Accommodation and food services normalization is used.

A.7 More on the role of the minimum wage

Between 1998 and 2005, the *35-hour working week* laws were gradually implemented.[24] To maintain monthly earnings of workers at the bottom of the wage distribution, monthly guaranteed salaries (GMR) were enforced. As shown by figure A.5, these GMR then converged in 2005 to a unique minimum wage.

Figure A.5. Gross minimum and median wages, and relative labor cost at the minimum wage since 1998



Source: DADS, *Panel Tous Salariés*. 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.

[24] *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.*

A.8 Separate estimations for 1995-2004 and 2005-2014

We estimate our baseline model and bargaining and sorting effects separately for period 1995-2004 (Tables A.5-A.6) and 2005-2014 (Tables A.7-A.8), before and after the rapid growth of the minimum wage. The bargaining effect estimate is only negative in 2005-2014 when the minimum wage is higher.[25] The sorting effect slightly grows between the two periods. This could be due to other underlying trends, and how low-wage jobs are distributed between firms.

Table A.5. Two-way fixed effect 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

Source: DADS, *Panel Tous Salariés*. Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

[25] The financial crisis could also have contributed to a stagnation or a drop in bargaining effect, with firms having a smaller rent to distribute, either to men and women.

Table A.6. 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			

Source: DADS, *Panel Tous Salariés*. Calculation of sorting and bargaining are based on model (3) 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).

Table A.7. Two-way fixed effect 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)
$Age^2/100$	-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)
$Experience^2/100$	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)
$Seniority^2/100$	-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

Source: DADS, *Panel Tous Salariés*. Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table A.8. Sorting and bargaining contributions to the gender wage gap, 2005-2014

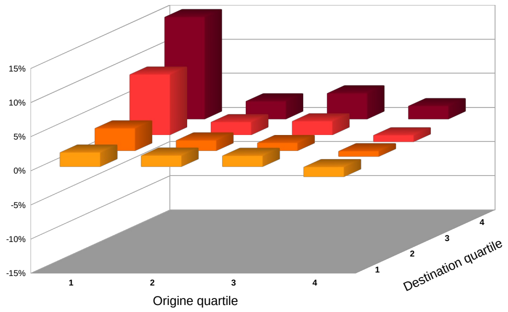
	Normalization based on...			
	Accommodation and food services firms		*Lowest log VA per worker firms*	
Total gender gap	0.154	*100 %*	0.154	*100 %*
Gender log wage gap due to firms	0.010	*6.5 %*	0.009	*5.8 %*
including sorting effect				
Male assignment, female premia (a)	0.018	*11.7 %*	0.018	*11.7 %*
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			

Source: DADS, *Panel Tous Salariés*. Calculation of sorting and bargaining are based on model (3) 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).

A.9 Firm/Wage Mobility and Additive Separability for Executives

Figure A.6. Average wage changes for executive movers conditional on origin and destination firm average wages

(a) Male Workers



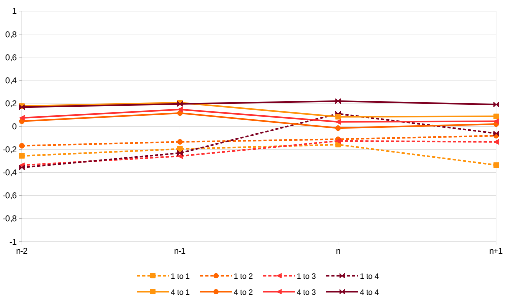
(b) Female Workers

CADRES

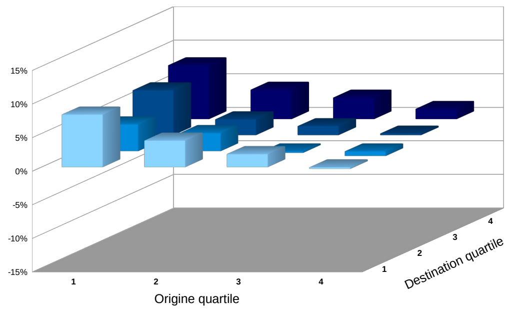
Source: DADS, *Panel Tous Salariés*. Note: Leaving a Q1 firm for a Q4 firm yields an average wage gain of 8.1 % to male executives.

Figure A.7. 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



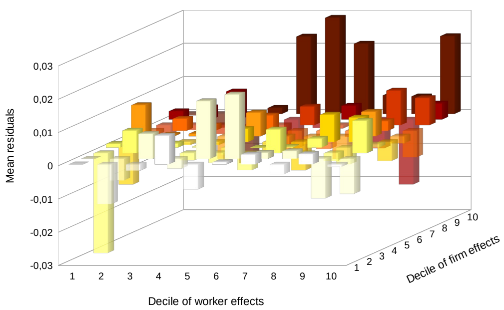
(b) Female Workers



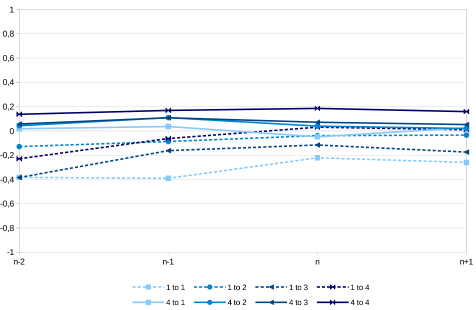
Note: Mean wage residual trend for a female executive going from a Q1 to Q4 firm is: -0.36, -0.23, 0.11 and -0.06.

Figure A.8. Mean wage residuals for executives conditional on deciles of worker and firm-fixed effects

(a) Male Workers



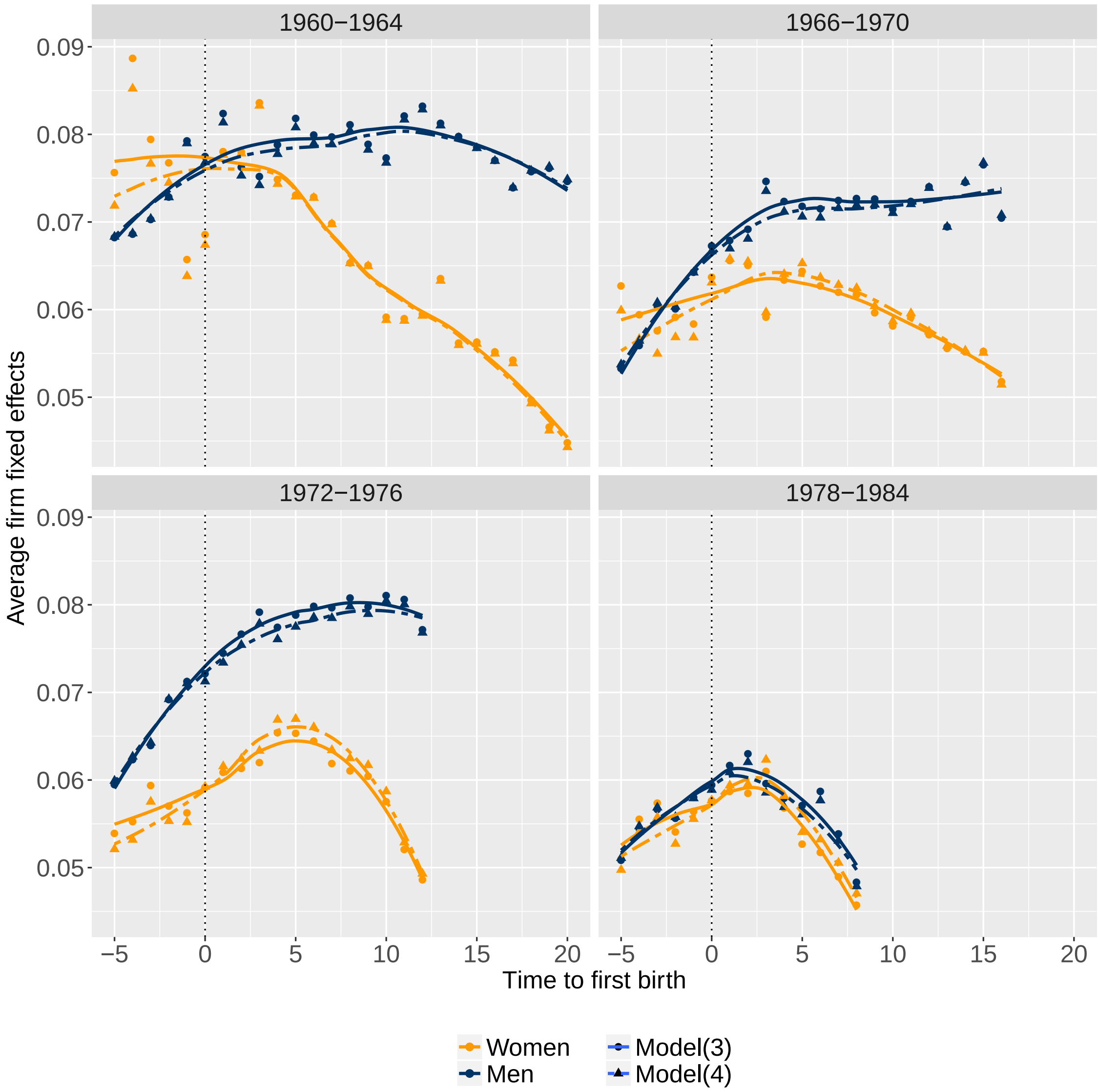
(b) Female Workers



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

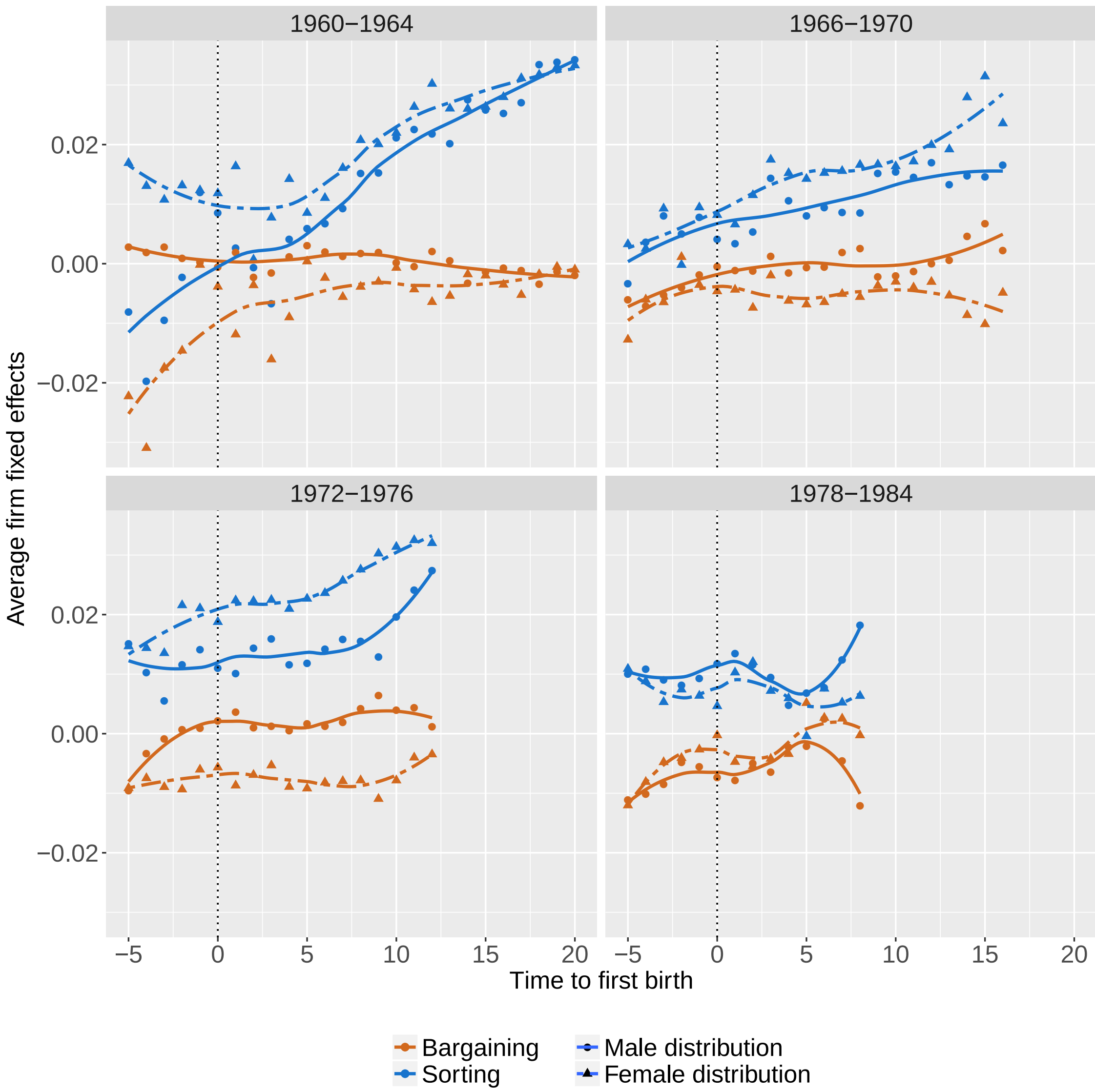
A.10 Firm effects and time to birth by cohorts

Figure A.9. Average gender gap in firm premia by cohort over time to first birth



Source: DADS, *Panel Tous Salariés*. Note: In cohorts born between 1960 and 1964, the average firm-fixed effects (normalized based on accommodation and food services) from model (3) for females 10 years after their first childbirth is 0.059; for males, it is 0.077.

Figure A.10. Sorting and bargaining effects by cohort over time to first birth

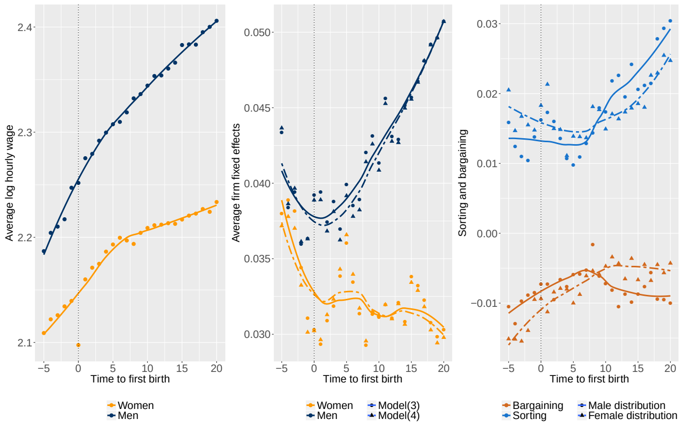


Source: DADS, *Panel Tous Salariés*. Note: In cohorts born between 1960 and 1964, we find a sorting effect of 0.021 and a bargaining effect of 0.001 10 years after workers' first childbirth, using model (3) estimates and male distribution across firms as reference.

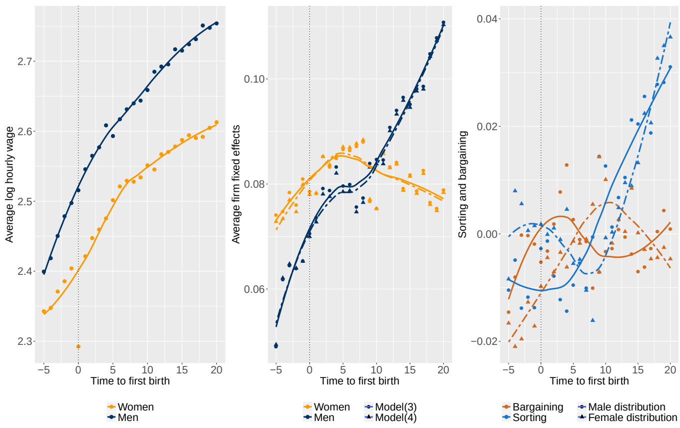
A.11 Sorting and bargaining by time to first birth for different occupations and education levels

Figure A.11. Wage gap, firm premia, sorting, and bargaining by job position

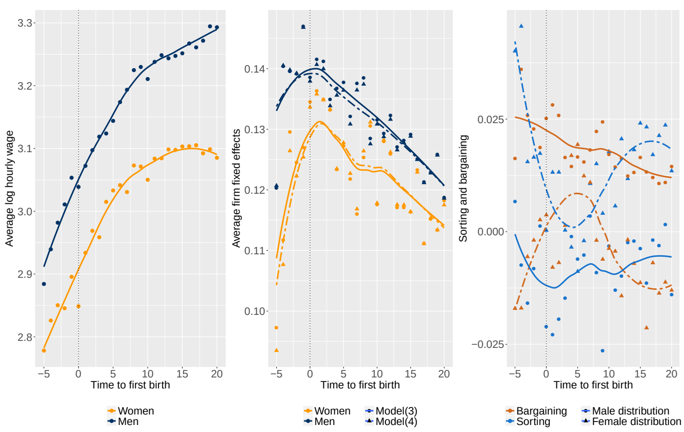
(a) Blue and white collar workers



(b) Clerks



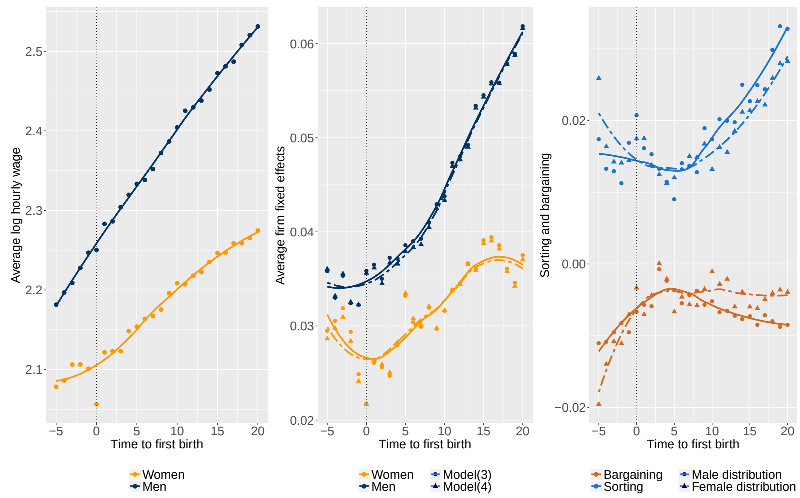
(c) Executive workers



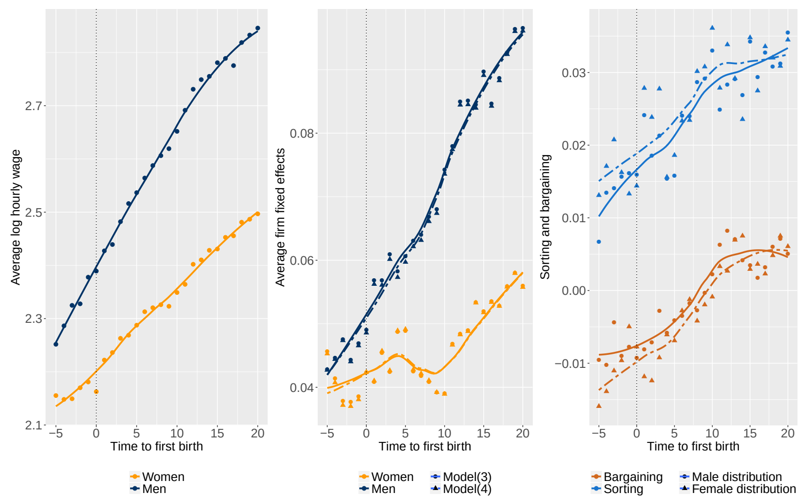
Source: DADS, *Panel Tous Salariés*, born on October 1st or 4th. Note: The average firm-fixed effect from model (3) for blue or white collar females 10 years after their first childbirth is 0.029; for males, it is 0.041. At this time to first birth, using model (3) and male distribution into firms as reference gives a sorting of 0.025 and a bargaining of -0.013.

Figure A.12. Wage gap, firm premia, sorting, and bargaining effects by education

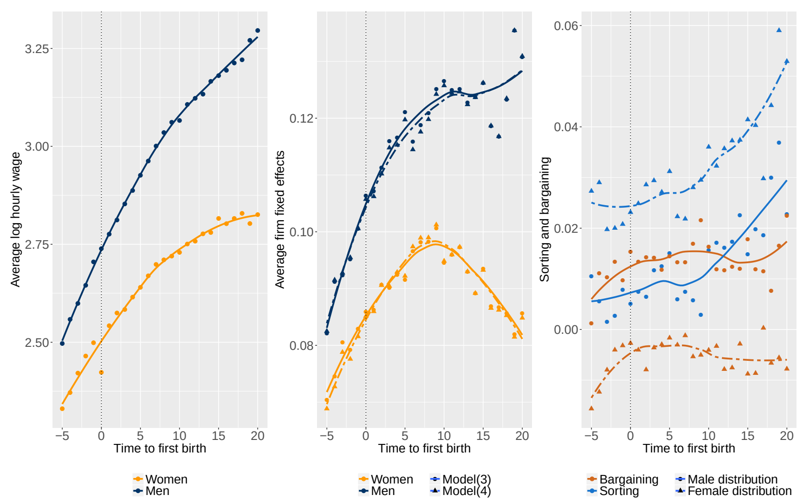
(a) Low-educated workers



(b) High-school educated workers



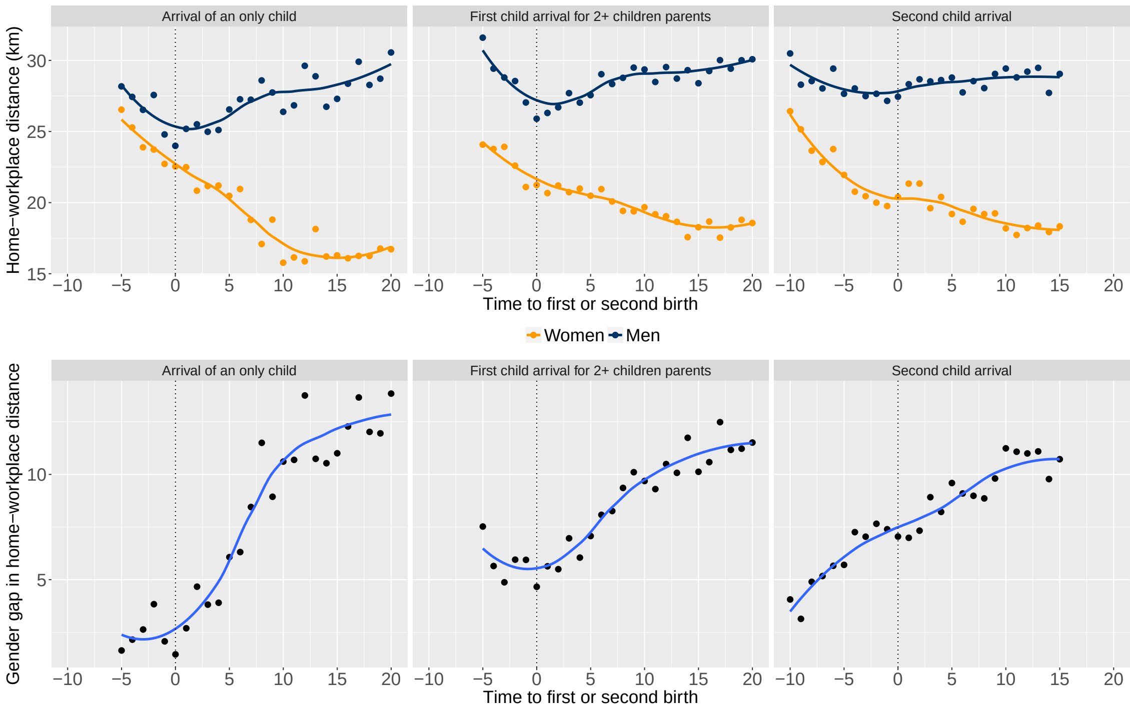
(c) College educated workers



Source: DADS, *Panel Tous Salariés*, born on October 1st or 4th. Note: The average firm-fixed effect from model (3) for low-educated females 10 years after their first childbirth is 0.029; for males, it is 0.043. At this time to first birth, using model (3) estimates and male distribution into firms as reference gives a sorting effect of 0.025 and a bargaining effect of -0.011.

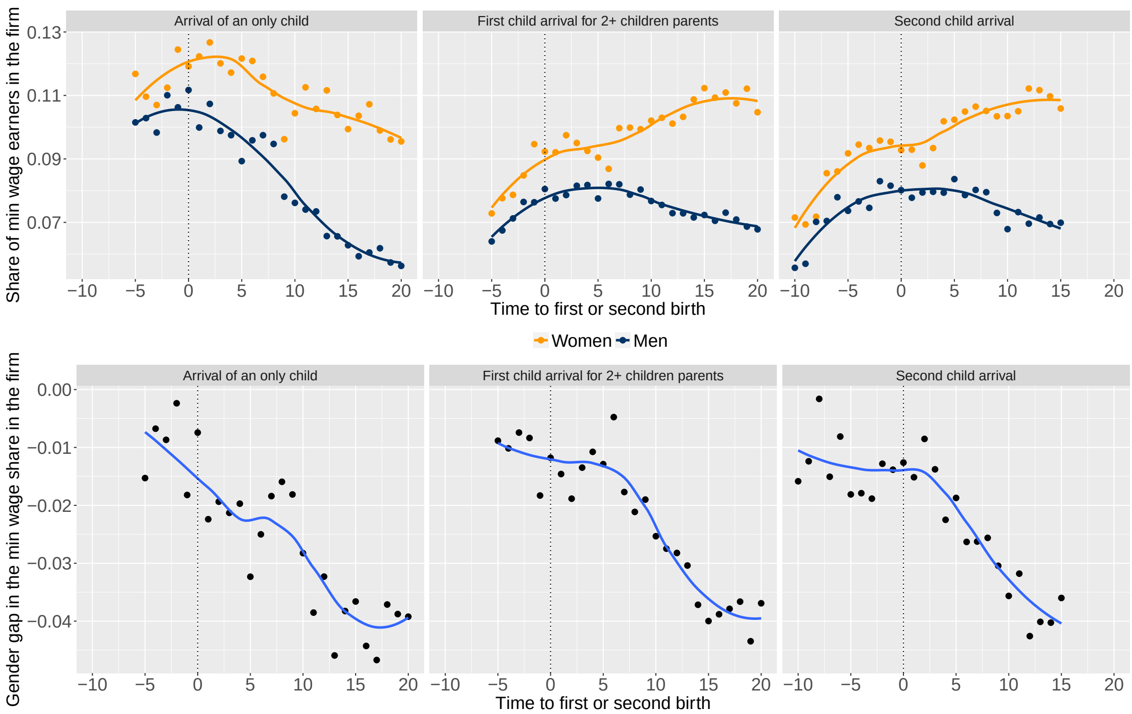
A.12 Firm characteristics at different times to the birth of workers' children

Figure A.13. Home-workplace distance



Source: DADS, *Panel Tous Salariés*, born on October 1$^{\text{st}}$ or 4$^{\text{th}}$. Note: On average, five years after the birth of their only child, men work in cities 26.5 km away from their home city, to be compared with 20.5 km for women, corresponding to a 6 km gap.

Figure A.14. Share of minimum wage-eaners in firms



Source: DADS, *Panel Tous Salariés*, born on October 1st or 4th, and comprehensive DADS files for coworkers' characteristics. Note: On average, five years after the birth of their only child, men are at work in firms paying 8.9 % of their workforce below 1.1 minimum wage. This share is of 12.2 % for women, corresponding to a -3.3 percentage point gap.

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