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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 perform Card et al. (2016) decompositions all along workers' life cycle to evidence a sorting mechanism of French private sector female workers into lower-paying firms that activates shortly after births of children. Mothers tend to work in firms with more flexible work hours and close to home. Gender-specific firm choices generate wage losses all along mothers' careers, in addition to direct child penalties. Young mothers with low wage expectations exiting the labor market after births leads to underestimate these effects.

JEL Codes: J31, J71, J16
Keywords: gender wage gap, gender inequalities, linked employer-employee data, parenthood gap

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1 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 2015, in the private sector, women's hourly wages were on average 18.5 % 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, half of the gap between men's and women's wages remains unexplained.

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 we 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). Another part of the gender wage gap may come from between-firm and between-job inequalities.

This mechanism would suggest that the gender wage gap is induced by gender segregation and/or 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.[1] Recently, Card et al. (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 births of children relate to wage losses for women. Wilner (2016) finds a large wage loss associated with motherhood in France, and a much smaller loss associated with fatherhood, even after controlling for human capital depreciation due to maternity leave, and for both individual and firm types of unobserved heterogeneity. Further, Kleven et al. (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 (See Joseph Hotz et al., 2018). More generally, some influential contributions to the literature stress the importance of life cycle dynamics to understand the gender wage gap (Adda et al., 2017; Bertrand et al., 2010; Goldin et al., 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 et al. (2016) are related to parenthood throughout

[1] For instance, Barth et al. (2016); Bayard et al. (2003); Nekby (2003); Trond Petersen (1995). Gobillon et al. (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.

workers' careers. In their paper, Card et al. (2016) document a sorting effect increasing sharply with workers' age. Here, we go a step further by relating this effect to the dates of birth of children. To do so, we take advantage of a rich matched employer-employee dataset for French private firms which also gathers information on family events. Following Card et al. (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. In addition, we control for the number of children in the household. We then study sorting and bargaining effects according to parenthood status. Briefly, we find that the sorting of women into lower-paying firms, and of men into higher-paying firms accounts for 2.0 percentage points (pp), which represents 12 %-15 % of the total gender wage gap. This sorting effect is accentuated among parents (around 2.6 pp-3.7 pp, compared to 0.8 pp-0.9 pp for non-parents once age differences are accounted for). The bargaining effects are close to zero for both parents and non-parents.

A longitudinal approach shows that the sorting effect starts increasing after births, of both first and second children. It clearly widens from 5 years after the birth of the first child (and 3 years after the arrival of the second one), and never decreases thereafter. The bargaining effect evolves along the life cycle 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 fringe benefits 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 in the few years around the birth of their child. They also tend to work in areas where the industry-specific firm labor markets are more concentrated, so they are likely to face worse outside options than fathers.

Part of the effects we find in the longitudinal approach may be linked to selection in the labor market. Indeed, we observe a relative drop in labor market participation of mothers especially in the two years after birth. We show that those women leaving the labor market after birth were previously employed in low paying firms, suggesting that the sorting we observe after birth is likely to be a lower bound of the actual effect. Imposing the selection pattern of men to be the same as the one of women conditional on pre-birth firm effects, we even find a stronger sorting effect.

Our paper is also linked to Barth et al. (2017), who find that the between-establishment gender wage gap component in the US is almost entirely due to married workers. Kleven et al. (2018) also show that, in addition to wage penalties, mothers are more often working in firms with high proportions of women with young children. Our results are also consistent with the findings of Albrecht et al. (2018). Focusing on high-skilled Swedish workers, Albrecht et al. (2018) 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 et al. (2016). We show that the exogenous mobility and the additivity assumptions required to identify two-way-fixed effect models (Abowd et al., 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. We investigate robustness of results to sample sizes. Moreover, the comparison between our findings and those obtained by Card et al. (2016) for Portugal and the ones of Bruns (2018) for Germany stresses the relationship between the firm contribution to the gender gap and national labor market specificities.[2] For France, we find that the sorting effect accounts for 12 to 15 % of the gender wage gap, whereas the bargaining effect is very small. These results differ from the ones obtained in Portugal where firms account for 21 % of the average gender log-wage gap, with 15-20 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. This explanation is in line with Bruns (2018) relating the increase in the firm contribution to the gender wage gap with the declines of union coverage and centralized wage-setting agreements between 1995-2001 and 2001-2008 in the case of Germany.

Bruns (2018) also relates the level of firm premia to motherhood, evidencing a gap in firm effects between mothers and placebo mothers after the first birth only.[3] Our analysis takes advantage of the richness of our data that links birth certificates to both mothers and fathers. This information allows us to control for the number of children in the estimation of firm premia, and to decompose the gap between mothers and fathers in terms of sorting and bargaining effects after the first and second birth.

The remainder of the paper is organized as follows. Section 2 motivates our approach by showing how firms may impact the gender wage

[2] See also Jewell et al. (2018) for a review of papers implementing CCK type decompositions.

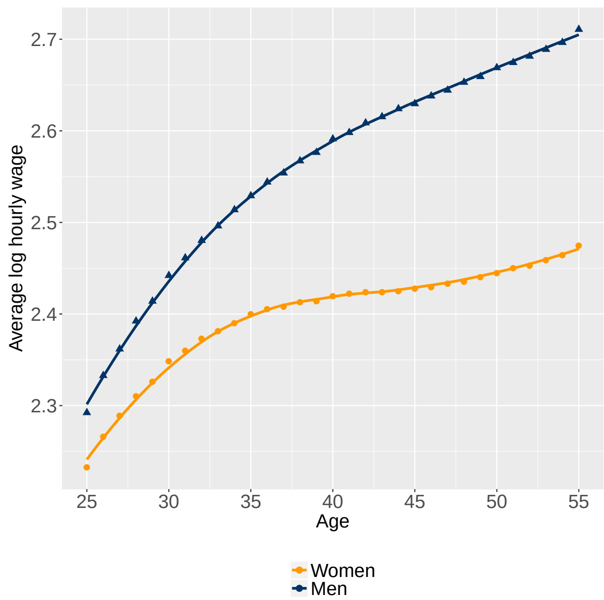
[3] Bruns (2018) imputes dates of birth from work interruptions and, when available, from administrative data on maternity leave for women only. The group of placebo mothers is a subset of women who never give birth, and the placebo date of birth of their child is drawn from a log-normal distribution fitted on the observed distribution of age of women.

gap throughout workers' life cycle. Section 3 describes the French context and the data. The model and conditions for identification are developed in Section 4. Section 5 presents the results, which are further analyzed in the light of family events and selection into the labor market issues, all along workers' life cycles in Section 6. The last section concludes.

2 Children and the Gender Wage Gap

The gender hourly wage gap in France in the private sector was 18.5 % in 2015. This gap increases dramatically over the life cycle as shown in Figure 1 which reports the average gender gap at each age between 1995 and 2015. It raises from less than 5 % for individuals aged 25 to about 20 % by age 40.

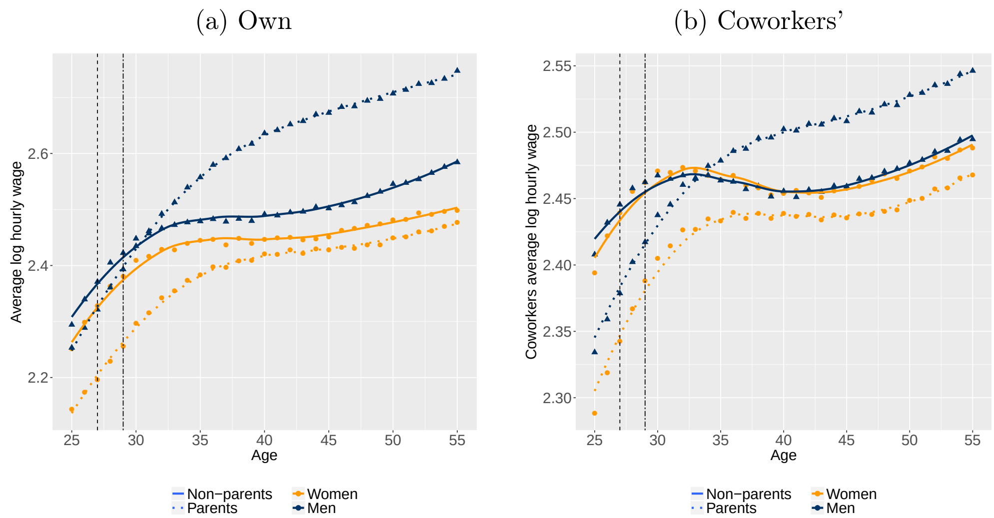
Figure 1. Average log hourly wage by age and gender



Source: DADS, *Panel Tous Salariés, 1995-2015*. 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 2.42 among women, and 2.59 among men.

This age profile may be driven by parenthood as shown by panel (a) of Figure 2: the gender wage gap deepens more for parents (dotted lines) than for non-parents (solid lines). For the latter, the gap is quite stable after age 35, whereas it keeps increasing for parents. If parenthood leads to more

Figure 2. Average log hourly wage and coworkers' one by age, gender and family status



Source: DADS, *Panel Tous Salariés, 1995-2015* and comprehensive DADS files for coworkers' wages.
Panel (a): average log hourly wage for forty-year-old workers is 2.42 among mothers, 2.45 among childless women, 2.64 among fathers, and 2.49 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.
Panel (b): average current coworkers' log hourly wage for forty-year-old workers is 2.44 for mothers, 2.45 for childless women, 2.50 for fathers and 2.46 for childness 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.

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 fringe benefits, but where wage policies may be less generous than in other firms (see Goldin, 2014).

The panel b of Figure 2 supports this firm-level explanation. The gap between fathers' and mothers' average coworker wage increases with age, whereas it remains around zero between childless men and women. If coworkers' average wage reflects firms' pay policies, this suggests that worker sorting across firms with different pay policies may contribute to the gender gap - and is especially driven by parents. 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.

3 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 (see section 3.2 for details). The gender hourly wage gap we observe decreased over the period of analysis from 21.7 % in 1995 (*i.e* 0.179 log difference) to 16.8 % in 2015 (*i.e* 0.136).[4] In particular, this decrease is due to hourly wage gains around the years 2000 when workers benefited from a reduction in working hours (with the introduction of the "35 hour week") while monthly wages were held constant.

3.1 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,[5] as well as collective negotiations, at the industry and firm level. At the industry level, employers' organizations and unions bargain on wage floors for each occupation and for each level of a productivity grid. Firms cannot opt out 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 national minimum wage (see Fougère, Gautier and Roux, 2016).

[4] See Figure 11 in the appendix.

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

Between 1998 and 2015, 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 over the period (see Figure 12).[6] 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.[7] Hence, even though the minimum wage remained stable relative to the median between 1998 and 2015, the cost of hiring a worker paid at the minimum wage decreased from 122 % to 106 % of the gross minimum wage.

The minimum wage protection is all the more important to be considered here as women are over-represented around this level of earnings: 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).

3.2 Data Sources

Our main data come from the linked employer-employee *Déclarations annuelles de données sociales* (DADS, Annual Declarations of Social Data) database. We use the panel subsample extracted from the exhaustive

[6] The minimum wage or *Smic* (Salaire minimum interprofessionnel de croissance) is adjusted to increase by i) the consumer price index (without tobacco) of the bottom quintile of the income distribution, and ii) at least half the rate of the purchasing power increase of the gross hourly wage of blue and white collar workers (SHBO). In addition to the automatic updating rule, the government can pass *coups de pouce* (boosts), such as in 2007 or 2012.

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

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 wage-earner. This reporting is mandatory. The panel has had a linked employer-employee structure since 1976: it contains both the firm unique identifier, which comes from the French register of firms, and an anonymous person unique identifier. The panel covers $1/24^{\text{th}}$ of the French wage-earners in the private sector before 2001 and $1/12^{\text{th}}$ since 2002. The panel does not cover self-employed workers who do not have to fulfill this administrative form. The public sector has been phased in the panel during the 1980s (public hospital employees in 1984; local and state public service employees in 1988).

We focus on 16-64 year old wage-earners who worked at least 15 hours in the year, in the private sector, in metropolitan France. Apprentices, trainees, private household workers (observed since 2009 only) are discarded. We also exclude self-employed farmers, craftsmen and shopkeepers who do have a wage-earner status but for whom the wages used to compute social contributions do not correspond to the full remuneration. We retain one job spell per year, the one associated with the highest earnings in the year and so, one employer per worker per year. We consider the corresponding wage, net of social contributions (but before income tax). This corresponds to the wage information reported by firms to the fiscal services for income tax purposes ("net fiscal"). This measure is therefore of great quality and contains all wages and salaries, any paid overtime, benefits in kind, all bonuses and indemnities including for shift work, those paid once a year, and those granted after contract termination if they exceed the industry-negotiated levels. This is the same wage concept as in Abowd et al. (1999), Postel-Vinay and Robin (2002), and Wilner (2016) more recently. This variable is particularly appropriate for our analysis as it accurately reports the wage components which can be negotiated, such as bonuses, by employees. A 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 and overtime represent around 13 % of gross earnings for women (resp. 15 % for men), according to the 2010 Structure of earnings survey.

Our administrative data only gather earnings paid by employers. As a consequence, it does not include maternity leave compensations paid directly by the social security system.[8] The same applies for paternity leaves, which are not compulsory and last legally 14 days. We address this potential issue by controlling for birth years in our specifications.

Our dependent variable is the hourly wage, as we observed hours worked for private sector employees from 1995 to 2015. Using hourly-wages allows us to keep part-time workers in the analysis, which is important given that we focus on the importance of children on workers career. It also contrasts with other papers: Card et al. (2013) and Bruns (2018) use daily wages and focuses on full-time workers.[9] Card et al. (2016) reconstruct a hourly wage calculated as the worker's base salary plus any regular earning supplements divided by the worker's usual work hours.

Since our main focus is on the relationship between firm pay policies and the impact of children on workers' careers, we use individual information on worker education, and the number and birth dates of their children coming from the Permanent Demographic Sample (EDP) for a subsample of the DADS panel (around 13 % of observations). 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

[8]Maternity leaves legally last 16 weeks for the first and second children, and 26 weeks for twins, third child and above. Women should take at least 8 weeks of maternity leaven among which 6, after birth. Daily compensations ("indemnités journalières") are capped, but collective agreement may provide additional compensations to maintain the usual wage during the leave ("maintien de salaire")

[9]In these two papers, the authors 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).

from 1968, 1975, 1982, 1990, and 1999. Information from the annual census surveys (which have replaced the exhaustive census since 2000) are also integrated. Selection into the EDP is based on date of birth. The EDP panel covers all individuals born the 1st, 2nd, 3rd, or 4th of October, and, since 2001, all those born one of the four first days of a quarter. For this restricted subgroup (called hereafter, the demographic sample), we know the birth dates of children and can use them in the econometric specifications, in particular to compute firm effects net of child penalties.

Further, to enhance the rest of the analysis we will supplement these data by firm level information: the composition of workers and jobs (sex, part-time, occupation), the main collective agreement in force in the firm, and financial results of the firm coming from the *File of Approximated Financial Results* (*Fichier approché des résultats d'Esane*): value-added (from 1995 to 2015), surplus, income statements, sales, exports, and investments (in 2015).

Our overall sample contain 24.5 millions of observations (person × year). As Card et al. (2016), we restrict the analysis to the dual largest connected set (DLCS) observed in the data. This sample corresponds to the largest mobility group of firms related with each other by workers' mobilities of both sexes. The DLCS associated with the overall sample covers around 70 % of its person × year and person observations, but only around 20 % of the firms (see Table 1). In addition, the proportion of firms with less than 20 workers drops from around 30 % to 9 % when restricting to the DLCS. This under-representation of small firms is quite common in this type of analysis. Indeed, firm fixed effects can only be recovered if worker mobilities in or out of the firm are observed, which depends on the size of the firm. In our case, this sample selection increases the gender wage gap on average which stands at 0.157 (in log differential) in the overall sample and 0.173 in its corresponding DLCS. Executives are also slightly over-represented and wages are somewhat higher. High value-added firms are over-represented in the DLCS as they are less likely to fail, more likely to be observed over a longer period, enabling to observe mobility.

The demographic sample covers around 13 % of the person × year observations, and around a quarter of firms. Its structure in terms of workers' and firms' characteristics follows the one of the overall sample, but small firms are under-represented in our estimation sample (3 %) compared to the DLCS obtained from the overall sample. Indeed, in this case, firm fixed effects can be estimated only in firms where mobilities of workers from the demographic sample are observed. However, restricting to the demographic sample and its DLCS does not affect the gender wage gaps in log differentials.

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

4.1 A rent-sharing model

Following Card et al. (2016) and Card et al. (2018), 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 with 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:

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

Table 1. Descriptive statistics

	Overall (1)		DLCS of (1)		Demographic sample (2)		DLCS of (2)	
	Male	Female	Male	Female	Male	Female	Male	Female
Observations	14,145,344	10,387,368	9,696,549	7,314,904	1,900,020	1,407,000	873,645	673,703
% of overall	*100 %*	*100 %*	*68.5 %*	*70.4 %*	*13.4 %*	*13.5 %*	*6.2 %*	*6.5 %*
Workers	1,593,111	1,301,728	1,086,473	870,794	202,278	168,851	106,304	85,352
% of overall	*100 %*	*100 %*	*68.2 %*	*66.9 %*	*12.7 %*	*13.0 %*	*6.7 %*	*6.6 %*
Firms	855,624	708,794	172,141	171,806	230,924	181,507	21,920	21,941
% of overall	*100 %*	*100 %*	*20.1 %*	*24.2 %*	*27.0 %*	*25.6 %*	*2.6 %*	*3.1 %*
Worker characteristics								
Net hourly log-wage	2.524	2.367	2.582	2.409	2.527	2.370	2.614	2.442
Age	38.7	38.2	38.9	38.1	38.6	38.3	38.9	38.2
Executives	0.170	0.109	0.199	0.123	0.170	0.110	0.214	0.136
White collar workers	0.145	0.516	0.150	0.467	0.141	0.517	0.150	0.438
Blue collar workers	0.486	0.160	0.431	0.178	0.486	0.157	0.406	0.181
Parents 1 child+	0.572	0.617	0.578	0.612	0.572	0.617	0.577	0.613
Parents 2 children +	0.380	0.392	0.387	0.387	0.380	0.392	0.388	0.386
Part-time job	0.127	0.336	0.125	0.301	0.125	0.335	0.130	0.296
Open-ended contracts	0.792	0.783	0.780	0.783	0.795	0.786	0.758	0.781
Paid hours	1529.6	1349.4	1552.7	1395.5	1536.7	1356.2	1555.1	1415.5
Firm characteristics								
% <20 workers	0.282	0.292	0.090	0.093	0.281	0.291	0.030	0.025
Manufacturing	0.244	0.137	0.271	0.155	0.252	0.141	0.274	0.159
Construction	0.115	0.017	0.052	0.012	0.118	0.018	0.035	0.009
Trade	0.150	0.195	0.140	0.185	0.157	0.203	0.132	0.189
Services	0.455	0.618	0.505	0.617	0.463	0.631	0.555	0.641
Value added before tax (2015)	4.394	4.256	4.513	4.302	4.397	4.262	4.663	4.427
Share of exporting firms	0.501	0.424	0.603	0.517	0.503	0.426	0.673	0.607
% of female workers	0.309	0.560	0.332	0.544	0.309	0.560	0.351	0.523
% of minimum wage earners	0.097	0.132	0.081	0.112	0.096	0.131	0.069	0.097
% of part-time workers	0.134	0.223	0.133	0.209	0.133	0.222	0.138	0.201

Source: DADS, *Panel Tous Salariés*. Lecture: The overall sample contains 14,145,344 male observations (person × year). 9,696,549 (68.5 %) of them are in the dual largest connected set. When we restrict the analysis to the demographic sample, we have 1,900,020 male observations (person × year), and 873,645 of them are also in the corresponding dual largest connected set (which represents 6.2 % of all male observations).

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.

4.2 Sorting and bargaining effects

The gender specific firm premia, $\psi^{G(i)}_{J(i,t)}$, can be recovered from equation (2). Applying a Blinder-Oaxaca decomposition to the firm effect average gap, we have:

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

The first term (i) gives the average difference between men and women firm's components if they were working in equal proportions in the same firms, it reflects differences between the gender specific bargaining parameters $\gamma^{G(i)}$ (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 be consequential.

4.3 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 already mentioned, 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 et al. (2018) propose a set of empirical checks to challenge these assumptions. As presented in details in Appendix C, our sample satisfies these requirements.

As detailed in Abowd et al. (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 et al. (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.[10] 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 accommodation 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 following the approach of Card et al. (2016) (see Appendix D for details).

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

5 Results

5.1 Direct family pay gap

Estimates associated with the time-varying worker characteristics of the worker fixed effect and the two-way fixed effect models are presented in Table 2. We control for polynomial functions of age interacted with education levels (the chosen specification is the one suggested in Card et al. 2018), and the number of children in the household. As mentioned previously, employers only pay part of the worker earnings during maternity or paternity leaves (the rest of it being paid by the social security, and not accounted for in our data), so we also control for years of birth of child.

Table 2. One way and two way fixed effect model estimates

	Worker fixed effects		Two-way fixed effects	
	Women	Men	Women	Men
$\left(\frac{age-40}{40}\right)^2$	-0.186***	-0.189***	-0.180***	-0.169***
	(0.008)	(0.006)	(0.008)	(0.006)
$\left(\frac{age-40}{40}\right)^3$	2.451***	2.527***	2.410***	2.562***
	(0.016)	(0.013)	(0.017)	(0.013)
$\left(\frac{age-40}{40}\right)^2 \times$ high school	-0.222***	-0.342***	-0.213***	-0.325***
	(0.013)	(0.011)	(0.013)	(0.011)
$\left(\frac{age-40}{40}\right)^3 \times$ high school	-2.194***	-2.137***	-2.231***	-2.146***
	(0.035)	(0.032)	(0.036)	(0.032)
$\left(\frac{age-40}{40}\right)^2 \times$ upper education	-0.748***	-0.995***	-0.691***	-0.944***
	(0.012)	(0.010)	(0.013)	(0.010)
$\left(\frac{age-40}{40}\right)^3 \times$ upper education	-1.665***	-1.470***	-1.725***	-1.556***
	(0.032)	(0.027)	(0.033)	(0.028)
1 child	-0.021***	0.038***	-0.018***	0.037***
	(0.001)	(0.001)	(0.001)	(0.001)
2 children	-0.022***	0.082***	-0.017***	0.079***
	(0.002)	(0.001)	(0.002)	(0.001)
3 children or more	-0.018***	0.133***	-0.014***	0.132***
	(0.003)	(0.002)	(0.003)	(0.002)
Adjusted R^2	0.802	0.858	0.826	0.876
Number of observations	719,338	943,000	719,338	943,000

Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Individual fixed effects, firm fixed effects, birth year dummies and interactions between year dummies and education level (including a missing education modality) are included in the regression but not reported here.

We find that mothers with one child suffer a 1.8 % wage penalty relative to women with no children, whereas fathers of one child benefit from a 3.7 % wage increase relative to men without children, entailing a 5.5 % gap between fathers and mothers of one child. Using the same dataset as ours, 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.[11] Gender gaps associated with subsequent births are increased, with a gap between fathers and mothers of 9.6 % for the second child, up to 14.6 % from the third one. Both for fathers and mothers, estimates of the year of birth dummies are negative, but stand more than 3 times bigger for mothers than fathers, in relation with the way remuneration are observed in our data during maternity/paternity leaves.

Differences between one-way and two-way fixed effect models can be noted: accounting for firm fixed effects slightly attenuates the gender gap in birth effects suggesting that part of the differential effect of the arrival of a child on wages is likely to go through the firm effect channel. However, the differences between coefficients are very small and cannot lead to a clear conclusion regarding the correlation between parenthood and firm effects.

5.2 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 3. The gender wage gap amounts roughly for 17.2 % of average male earnings. The firm contribution varies from 1.4 pp to 2.8 pp according to the chosen normalization. The sorting effect accounts for 11.6 % of the gender wage gap when using the female premia, and for 14.5 % using the male premia. In both cases, this 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.

[11] Instead of age, Wilner (2016) directly controls for experience (accounting for career interruptions) and seniority. However as noted by Card et al. (2018), actual labor market experience may raise questions of endogeneity, when carreer interruptions are related to future wages.

Table 3. Sorting and bargaining contributions to the gender wage gap

	Number of obs.	Gender wage gap	Firm contrib.	Sorting (a)	(b)	Bargaining (c)	(d)
VA	1,547,348	0.172	0.014	0.020	0.025	-0.006	-0.011
			8.1 %	*11.6 %*	*14.5 %*	*-3.5 %*	*-6.4 %*
Food	1,547,348	0.172	0.028	0.020	0.025	0.008	0.003
			16.3 %	*11.6 %*	*14.5 %*	*4.7 %*	*1.7 %*

Source: DADS, *Panel Tous Salariés*. Number of workers = 191,656; number of firms = 23,303. Column (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 column (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).

Estimates of the bargaining effect and of the total contribution of firms to the gender wage gap depend on the normalization choice. However, in both cases the bargaining effect is small. It is even negative when firm effects are normalized to zero in the group of firms having the lowest value-added per worker. It amounts to -3.5 to -6.4 % (1.7 to 4.7 % with accommodation and food services normalization) 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. Overall, the positive role of firms on the gender wage gap comes mainly from the sorting of women into low-paying firms with respect to men. This is in line with the results of Bruns (2018) who applies a similar methodology to German data, and finds a total firm contribution to the gender wage gap going from 11 % in 1995-2001 (decomposed into a 17 pp sorting effect and a -6 pp bargaining effect) to 26 % in 2001-2008 (25 and 0). Card et al. (2016) also find that the firm contribution is mainly driven by sorting - which accounts for 75 % of Portuguese firm contribution. However, they evidence a positive bargaining channel in Portugal.

5.3 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 larger than in France (21 %, vs 8 to 16 %

in our data). Altogether, national institutional settings may play an important role in interpreting those 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. Such an institution may 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 (2005-2015) the rapid minimum wage growth. The French bargaining effect is higher when we compute it on the 1995-2004 subperiod, when the minimum wage was lower (see Table 7).

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 average firm premia according to firm characteristics (see Table 8), we see that including industry dummies and then, collective agreement dummies increase significantly the adjusted R^2 of the regressions (from 0.235 to 0.277). The collective wage bargaining system explains a large proportion of the average premia paid by the firm to her employees. However, it does not seem to explain the dispersion of firm-specific gender gap in firm effects (see Table 9, which relates the firm-specific gender gap in firm effects according to firm characteristics). Moreover, we find very few firm-level variables explaining the disparities in firm specific bargaining effects. In particular, we evidence no relationship between the within-firm gender gap and the firm value-added per worker when controlling for collective agreement dummies, nor with firm assets or investments. In contrast, the firm-specific bargaining effect is positively related to the proportion of women among blue collar workers. This suggests that a sorting mechanism is also likely to occur across jobs within firms.

5.4 Robustness to sample size

As shown by Table 1, our analysis requires large sample restrictions, and one could wonder whether our main results are robust to these. The problem with those restrictions is twofold: first, the resulting estimation sample is a selected sample, which could bias the results; second, estimates of firm premia are less precise as the sample size is smaller. Indeed, Andrews et al. (2008) shows that the precision of firm premia estimates depend on the number of between firm mobilities observed in the sample, and the merge of our matched employers-employees sample with the demographic sample decreases the sample size, and *a fortiori* the number of mobilities. However, the precision problem raised by Andrews et al. (2008) matters for the analysis of non-additive combinations of individual and firm fixed effects, like the correlation between individual and firm fixed effects. Sorting and bargaining effects are linear combinations of firm fixed effects, so the precision of the estimate will not be biased by the lack of precision of firm effect estimates.

Nevertheless, our data allows us to conduct an alternative estimation on the overall sample of the DADS panel ($1/24^{\text{th}}$ of the population before 2002 and $1/12^{\text{th}}$ after). The corresponding DLCS is much larger than the demographic sample (more than 17 millions observations compared to 1.5 millions), its drawback is that it does not contain information on birth dates of children, neither on education. To deal with the lack of information on education level, we use an alternative specification interacting age dummies with job positions, but we cannot control for the direct child penalty when estimating firm premia when estimating the model on the overall sample, which may bias sorting and bargaining effects.

In line with these first results, we find that sorting and bargaining effects are also close whether we consider the demographic sample or the overall sample (see Table 4).[12] In particular, sorting effect estimates using female premia (a) always range between 0.019 and 0.024 and between 0.025 and 0.029 when using male premia (b). Both the sample size reduction and the change of specification (using education instead of job position and including children variables) slightly reduce sorting effects. The effect of the sample size on the bargaining effect is very small regardless of the normalization. The effect of the change of specification on the bargaining effect is stronger. but in all cases the bargaining effect remains small in absolute value.

Altogether, these alternative specifications reinforce our previous conclusion: the gap due to firms is mainly due to sorting, and the bargaining effect is very close to zero. In the following, we focus on results based on our main specification that control for ages and year interacted with education and kids estimated on the demographic DLCS.

Table 4. Sorting and bargaining contributions to the gender wage gap: demographic vs. overall sample

Specification	Number of obs.	Gender wage gap	Total firm contribution	Sorting (a)	Sorting (b)	Bargaining (c)	Bargaining (d)
Small value-added per worker firms							
(Age + year) × position	17,011,453	0.173	0.024	0.024	0.029	0.000	-0.005
			13.9 %	*13.9 %*	*16.8 %*	*0 %*	*-2.9 %*
	Demographic sample						
(Age + year) × position	1,547,348	0.172	0.024	0.022	0.027	0.002	-0.003
			14.0 %	*12.8 %*	*15.7 %*	*1.2 %*	*-1.7 %*
(Age + year) × education + kids	1,547,348	0.172	0.014	0.020	0.025	-0.006	-0.011
			8.1 %	*11.6 %*	*14.5 %*	*-3.5 %*	*-6.4 %*
Accommodation and food services							
(Age + year) × position	17,011,453	0.173	0.025	0.024	0.029	0.001	-0.004
			14.5 %	*13.9 %*	*16.8 %*	*0.6 %*	*-2.3 %*
	Demographic sample						
(Age + year) × position	1,547,348	0.172	0.023	0.022	0.027	0.001	-0.004
			13.4 %	*12.8 %*	*15.7 %*	*0.6 %*	*-2.3 %*
(Age + year) × education + kids	1,547,348	0.172	0.028	0.020	0.025	0.008	0.003
			16.3 %	*11.6 %*	*14.5 %*	*4.7 %*	*1.7 %*

See Table 3 for further details on sorting and bargaining computation. Note: When we estimate the two-way fixed effect model in the overall sample using (age + year) × position as control variables, we find a total firm contribution of 0.024 corresponding to 13.9 % of the total gender gap (lowest value-added per worker firms normalization). If we keep the same specification but estimate it on the demographic sample, we find a firm contribution of 0.024 (14.0 % of the gender gap in this sample). Now, if we use education dummies and number of children as controls in the model, firm contribution becomes 0.014 (8.1 % of the gender gap).

[12] For comparison of the two firm effect distributions see figure 20 in the appendix.

6 Life Cycle Analysis

6.1 Sorting, bargaining effects, and parenthood

Using the firm fixed effect estimates obtained from the demographic sample, we compute sorting and bargaining effects for different subgroups depending on education, 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 5 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. 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 et al., 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.[13]

[13] For these reasons, together with the fact that gender gap is particularly acute among executives (Bertrand et al., 2010), we considered estimating separately the model for executives. However, figures 17 to 19 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 et al., 2015) would probably suit better executives' behavior. This issue should be the object of further research.

Table 5. Sorting and bargaining contributions for different subgroups

	Obs.	Gender gap	Firm contribution	Sorting (a)	Sorting (b)	Bargaining (c)	Bargaining (d)
All - Lowest VA per worker	1,547,348	0.172	0.014	0.020	0.025	-0.006	-0.011
Parents	917,259	0.254	0.022	0.024	0.032	-0.003	-0.011
Non parents	630,089	0.070	0.004	0.014	0.015	-0.010	-0.012
Parents 45+	410,576	0.290	0.023	0.026	0.037	-0.003	-0.014
Non parents 45+	103,741	0.082	-0.005	0.008	0.009	-0.013	-0.014
< High school	720,225	0.163	0.015	0.026	0.028	-0.011	-0.013
High school	265,585	0.184	0.017	0.019	0.028	-0.002	-0.010
College, University	400,390	0.292	0.019	0.016	0.027	0.003	-0.008
<30 year old	400,886	0.068	0.010	0.021	0.021	-0.010	-0.010
30-39	429,105	0.135	0.010	0.014	0.018	-0.004	-0.008
40-49	397,412	0.208	0.016	0.018	0.029	-0.002	-0.013
>49	319,945	0.261	0.017	0.024	0.031	-0.007	-0.014
Blue collar workers	476,807	0.169	0.015	0.028	0.025	-0.014	-0.010
White collar workers	426,126	0.044	-0.016	0.006	-0.006	-0.021	-0.009
Clerks	364,038	0.099	0.004	0.005	0.017	-0.001	-0.013
Executives	275,187	0.167	-0.000	-0.016	0.012	0.016	-0.013
All- Food and housing	1,547,348	0.172	0.028	0.020	0.025	0.008	0.003
Parents	917,259	0.254	0.036	0.024	0.032	0.012	0.004
Non parents	630,089	0.070	0.018	0.014	0.015	0.005	0.003
Parents 45+	410,576	0.290	0.037	0.026	0.037	0.011	0.001
Non parents 45+	103,741	0.082	0.010	0.008	0.009	0.002	0.001
< High school	720,225	0.163	0.029	0.026	0.028	0.004	0.001
High school	265,585	0.184	0.031	0.019	0.028	0.012	0.004
College, University	400,390	0.292	0.033	0.016	0.027	0.017	0.006
<30 year old	400,886	0.068	0.025	0.021	0.021	0.004	0.004
30-39	429,105	0.135	0.024	0.014	0.018	0.010	0.006
40-49	397,412	0.208	0.030	0.018	0.029	0.012	0.002
>49	319,945	0.261	0.032	0.024	0.031	0.007	0.001
Blue collar workers	476,807	0.169	0.029	0.028	0.025	0.001	0.004
White collar workers	426,126	0.044	-0.001	0.006	-0.006	-0.007	0.005
Clerks	364,038	0.099	0.018	0.005	0.017	0.013	0.001
Executives	275,187	0.167	0.014	-0.016	0.012	0.030	0.002

See Table 3 for further details on sorting and bargaining computation. Note: When we use the lowest VA per worker firms as the normalization group of the firm fixed effects, we find a firm contribution to the gender wage gap among parents of 0.018 decomposed into a 0.023 sorting effect (respectively 0.032 when measured with male premia) and a -0.005 bargaining effect (resp. -0.014).

Further, sorting effects are roughly stable across education groups. The sorting 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 (between 0.028/0.163 - 0.027/0.292 = 7.9 and 10.5 pp) is of the same magnitude than the one found by Card et al. (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 increases with the level of education probably for the same reasons as with position.

Age group comparisons indicate sorting effect is relatively high before 30, then diminishes, and finally increases after age 40. This pattern is in line with the one evidenced by Card et al. (2016). By contrast, the bargaining effect does not seem to vary with age, and remains stable around our baseline estimate. Ultimately, the total firm contribution follows the same pattern as the sorting effect, and increases strongly after 40. This finding can be related to family structures. As shown in Table 5, the sorting effect is also larger for parents (0.024-0.032) than for non-parents (0.014-0.015). 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: 0.026-0.037 for parents versus 0.008-0.009 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 (either 0.024-0.014= 1.0 pp or 0.032-0.015= 1.7 pp depending on the reference chosen) come in addition to the direct effect of children on wages, which amounts to 5.5 pp difference for the first child when controlling for both individual and firm effects.[14]

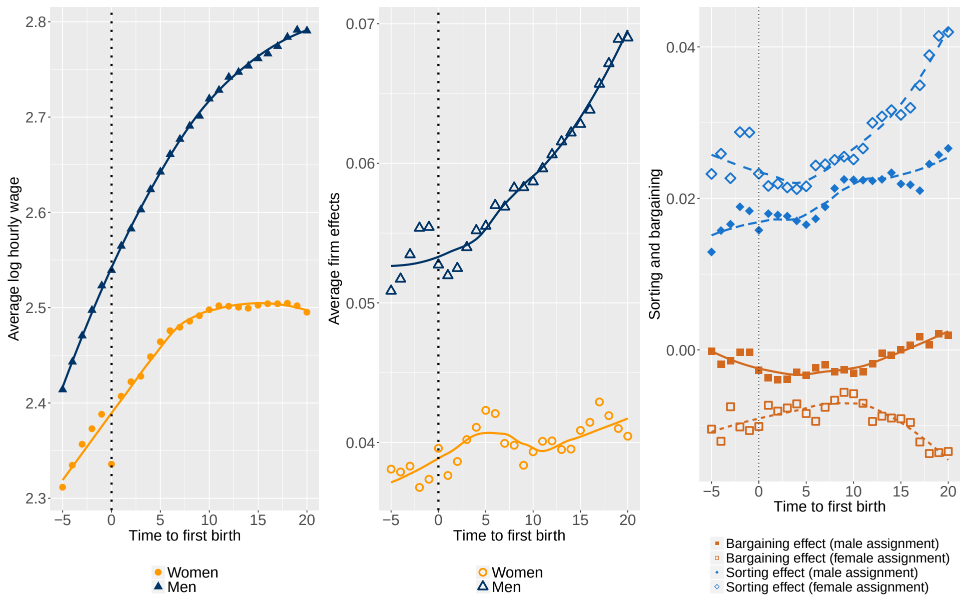
6.2 Sex-specific sorting and births of children

We highlighted in the previous section that the sorting of men and women across firms increases with age and parenthood. We now analyze more precisely how inequalities grow along the family life cycle. We thus focus on workers who eventually have children, and particularly on the birth of the first two children. Figure 3 shows the decomposition of the gender gap from five years before the birth of the first child to twenty years after.

[14] The wage penalty associated with the first child is -1.8 % for women, and 3.7 % for men, translating into a 5.5 pp difference.

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 years before birth, women and men experience comparable firm premia evolution. However, average female firm effects rapidly stop increasing a few years after the first birth and plateau until 20 years after the first child was born. Oppositely, male firm effects keep increasing over the entire career of fathers.[15]

Figure 3. Firm premia, sorting, and bargaining effects over time to first birth



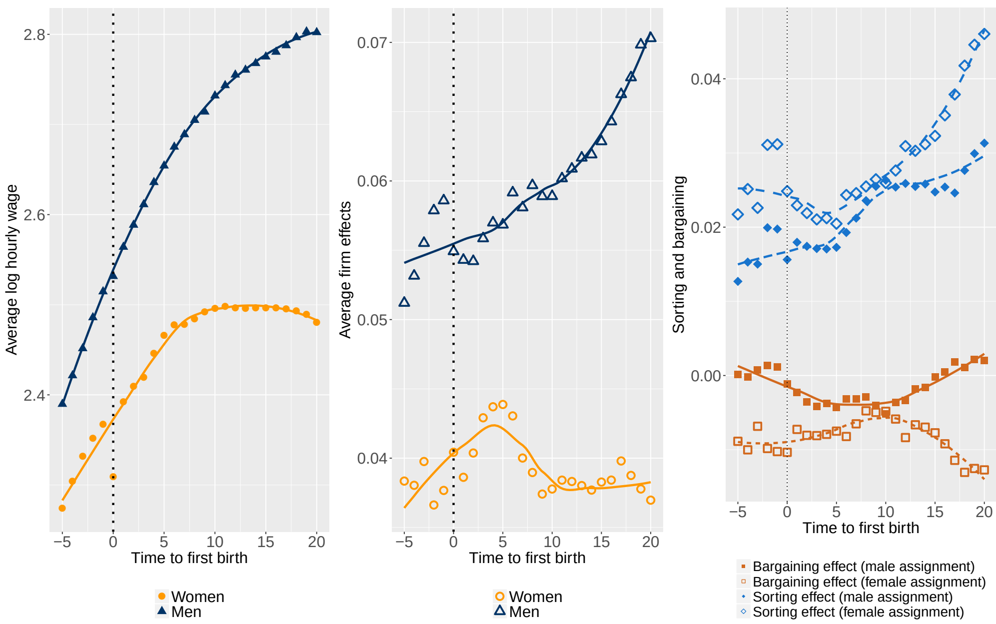
Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1st to 4th. Note: The average log hourly wage for females 5 years after birth of their first child is 2.46, to be compared with 2.64 for males. The average firm fixed effect (normalized based on low value-added) for females at this time is 0.042; for males, it is 0.055. Using male distribution into firms as reference gives a sorting effect of 0.016 and a bargaining effect of -0.003. The lines were obtained by smoothing the averages.

The right-hand side of the figure shows the corresponding sorting and bargaining effects. The bargaining effect is relatively flat, between -0.015 and 0. In contrast, the sorting effect increases a few years after the first birth from approximately 0.02 in the early careers to around 0.03-0.04

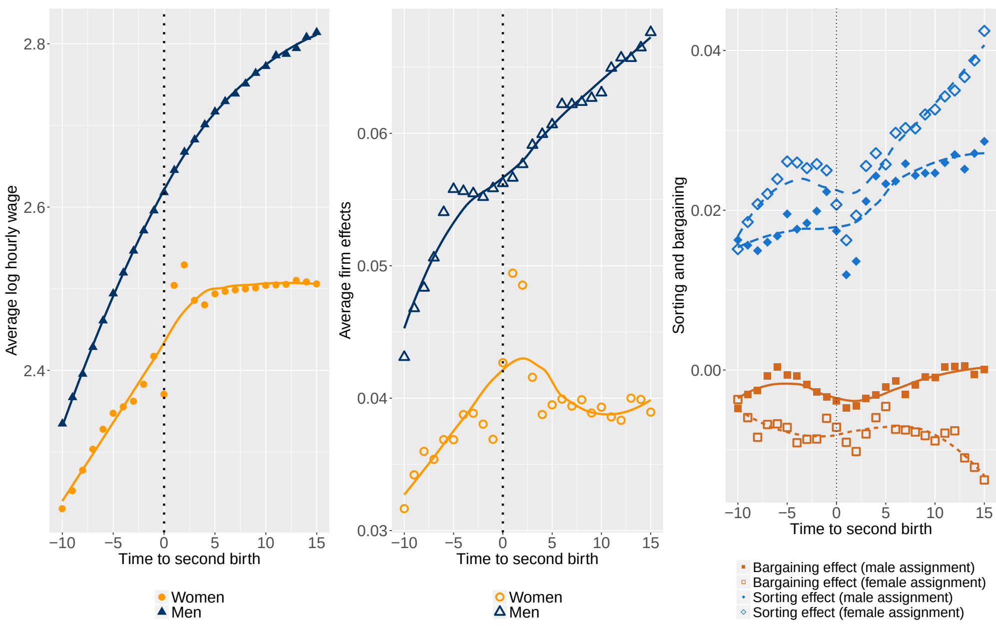
[15] We check in Figure 21 that the trends in firm premia, sorting and bargaining effects are unaffected by considering the estimates obtained in the overall sample instead of the demographic subsample, with a simpler specification.

Figure 4. Firm premia, sorting, and bargaining effects over time to births for parents of 2 or more

(a) First birth



(b) Second birth



Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1^{st} to 4^{th}. Note: The average log hourly wage for females 5 years after birth of their second child is 2.49, to be compared with 2.72 for males. At this time, the average firm fixed effect (normalized based on low value-added) for females is 0.039; for males, it is 0.061. At this time to second birth, using male distribution into firms as reference gives a sorting effect of 0.023 and a bargaining effect of -0.002. The lines were obtained by smoothing the averages.

(depending on the reference) 20 years after the first birth. Compared to men, women tend to work in low-paying 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. This observation is in line with Albrecht et al. (2018), 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: birth directly affects wages (Table 2), and accentuates a long-term gender divergence in firm effects.[16]

The decline in the average female firm effect and the increase of the sorting effect for parents of two children or more is even more pronounced at the time of the second birth. Figure 4b, which is centered at the birth of the second child, indicates a strong break right after this second birth.

6.3 Selection into the labor Market

Part of these sorting and bargaining effects may be due to differences in labor market selection between men and women. As shown by Figure 5, the arrival of a child is linked to a reduction in the number of observations for both men and women employed, but this reduction is far sharper among women.[17]

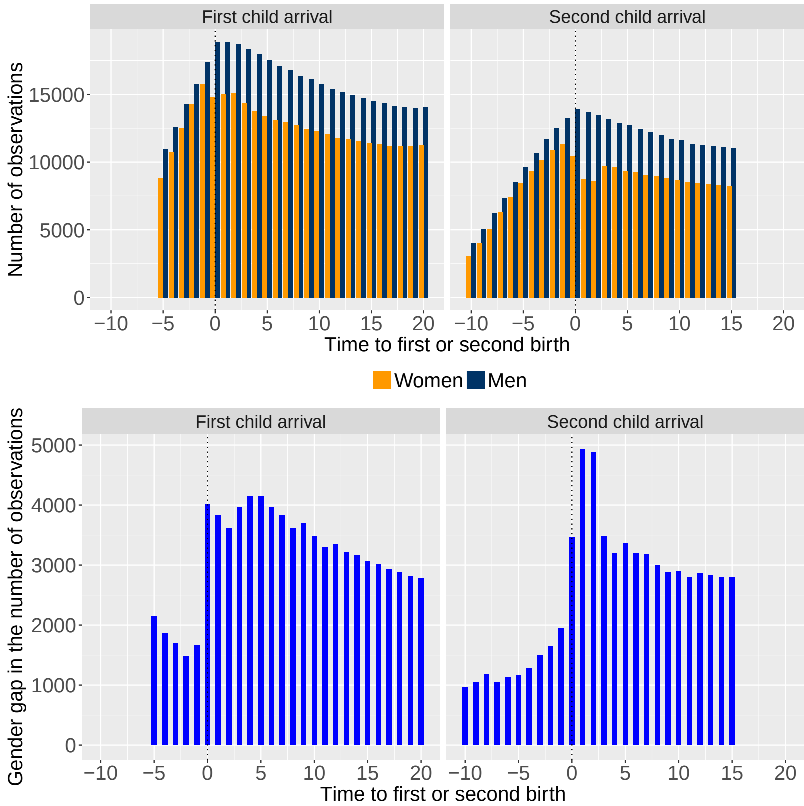
The effect of differentiated selection by gender is theoretically ambiguous depending on whether it is more privileged women who tend to leave the labor market due to social norms or mothers for whom child care would cost more than they could earn (Neal, 2004). As documented in the literature the selection pattern may also vary along the life cycle (see for instance Machado, 2017; Mulligan and Rubinstein, 2008).

In our context, the selection issue does not affect the estimates of firm

[16] Figure 22, detailed by birth cohorts, shows that the increase in sorting effect is observed within generations, and is not only due to differences across generations. An analysis by education is provided in Figure 23.

[17] These individuals may leave employment or move to one of the sectors that we do not observe, *i.e.* self-employment or the public sector.

Figure 5. Number of observations at each period



Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1st to 4th. Note: In our sample, we observe 14,821 females the year of the birth of their first child. At this time, we have 18,840 male observations, which corresponds to a 4,019 gender gap. The lines were obtained by smoothing the averages.

effects, but it is likely to impact the interpretation of the decomposition across time. To document it, we first compare the groups of mothers and fathers who do not leave the labor market in the year of the first birth or the year after (high commitment) to parents who leave the labor market for at least one year in this period after the birth (low commitment). Table 6 compares these two groups in terms of average wage and firm effects.[18] Women who are at least one year out of the labor market the year of birth or the subsequent one (low-commitment group) are clearly a selected sample compared to women who work during those two years (high-commitment group): they have lower pre-birth wages (2.433 - 2.318 = 0.115 log hourly

[18] Table 10 in the appendix provides more details about these two groups of parents in terms of age, education, job position, and number of children.

wage difference on average). The wage difference between these two groups of women increases with time relative to the first birth up to a 0.150 log difference on average in the ten years following the birth. Interestingly, this difference partly reflects differences between average firm effects observed for these two groups of women that increases from 0.017 to 0.024 from the pre-birth period to the later period we observe.

Table 6. Descriptive statistics in the high/low-commitment subgroups to the labor market after first birth

	Year - 1 (before first birth)			Years 1 to 10		
	Low commitment	High commitment	Δ	Low commitment	High commitment	Δ
Females						
Average Log hourly wage	2.318	2.433	0.115	2.368	2.518	0.150
Average Firm effect	0.031	0.048	0.017	0.028	0.052	0.024
Nb of obs.	21,479	53,165	-	12,328	65,913	-
Males						
Average Log hourly wage	2.466	2.540	0.074	2.574	2.689	0.115
Average Firm effect	0.052	0.059	0.007	0.047	0.065	0.018
Nb of obs.	20,722	70,686	-	10,401	85,755	-

Note: Workers the year before birth are considered in the low or high commitment groups if they experienced work interruption the year of birth or the following one or not. A women in the low commitment group receives an average log hourly wage of 2.318 the year before birth, to be compared with 2.433 for a women more committed to the labor market. The difference between the 2 groups amounts to 0.115. In the ten years following the first birth, the discrepancy across these two types of workers goes up to 0.150.

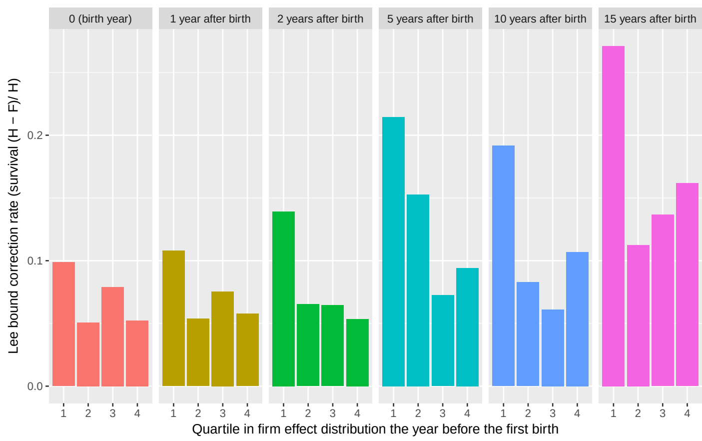
Women staying on the labor market are thus a selected sample of women with higher potential wages and firm effects. Regarding men, the difference in log wages between high and low committed men ranges from 0.074 to 0.115, and in firm effect from 0.007 to 0.018. Hence, the difference between low and high committed women is larger at all times than between men.

These observations indicate that in the years following birth the selection process is stronger for women. So, women who leave the labor market after birth would have lower firm fixed effect than their male counterpart, which would increase the gender gap in firm premia.

In order to provide more elements about the potential effect of the selection, we also propose an alternative decomposition that corrects for the differentiated selection between men and women. Assuming that conditional on pre-birth firm premia there is no selection on future potential firm premia, we simulate a counterfactual group of men for whom the selection pattern is similar to the one of women observed after the birth.

To do so, we restrict our sample to those individuals who are observed the year before the birth, for whom we are actually able to establish the position in the pre-birth firm effect distribution. For each year after the birth, we are also able to characterize the selection pattern conditional on the pre-birth firm premium. For instance, Figure 6 shows that the relative differential in survival rates between men and women who were belonging to the first quartile of the year before birth sex-specific distribution of firm premia amounts to 13.9 % two years after the first birth. Based on this information, we are able to draw a counterfactual population of men applying the same selection pattern conditional on the pre-birth quartile of firm premium as the one of women.[19]

Figure 6. Gender gap in survival rate on the labor market for workers observed in employment the year before the birth of the first child, conditional on their position in pre-birth firm fixed effect distribution

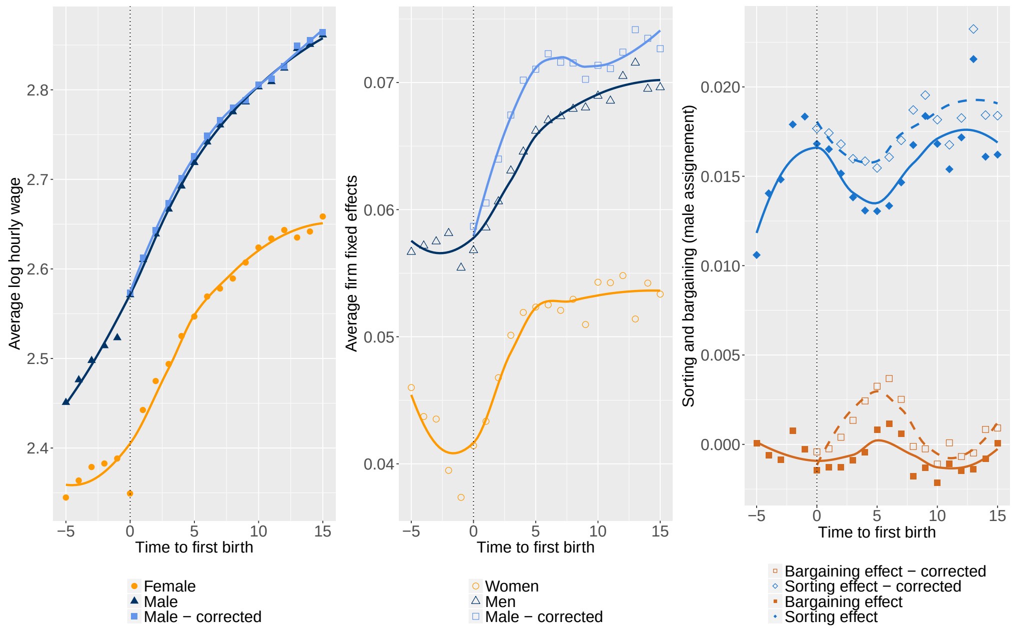


Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1^{st} to 4^{th}. Note: We consider the workers observed the year before the birth of their first child, and we recover to which inter-quartile interval they belong at this date in the gender-specific firm effect distribution. Two years after the birth of the first child, the relative differential of survival rate between males and females in the bottom quartile is 13.9 %.

[19] We also apply classical Lee's approach to build Figure 25. However, the lower bound is not informative about the gender gap due to firms (middle panel). The sorting effect is also weakened once we base our computation on the altered population of men. Nevertheless, the upward slope of the sorting effect remains and is even stronger.

Figure 7 shows the results of the decomposition based on this approach. The two left panels show log wage gap and firm effect evolutions across the lifecycle for women, for men, and for a corrected subgroup of men.[20] The right panel plots the sorting and bargaining effects computed with male assignment to firms.[21] The plain lines represent the two effects before correction, while the dotted curves correspond to the effects computed after applying our correction by quartile of firm effect distribution before birth.

Figure 7. Gender wage gap, firm effects, sorting and bargaining with Lee bounds correction on wages - based on quartile of average firm effects the year before the first birth



Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1st to 4th. Note: We consider the workers observed the year before the birth of their first child. The average corrected male firm effects 5 years after the first birth is 0.071, to be compared with a 0.066 effect for males before correction, and a 0.052 firm effect among women. At this time to first birth, the corrected sorting effect amounts to 0.016 (instead of 0.013 without correction) and the bargaining effects is 0.003 (instead of 0.001).

The central panel indicates that the average corrected male firm effects are higher at all times after the first birth than their uncorrected counterparts. That is to say, when we mimic the female selection process in employment

[20] The corrected value we rely on is the average of 1,000 replications of the correction process described above.

[21] The female assignment to firms is more difficult to obtain since we need to compute an average firm effect of male coworkers *after having applied the correction*.

after a first birth, and apply it to the male population, we find higher labor market outcomes (which is consistent with a positive selection process). Both sorting and bargaining effects are somewhat increased by the male distribution correction (right panel). We report in the appendices the same analysis for the second birth (Figures 26 and 27).

Finally, gender differentials in selection is likely to decrease the observed gap between men's and women's firm premia, and thus lead to an under-estimation of the sorting and bargaining effects. The final simulation that we perform shows that this bias is likely to be limited.

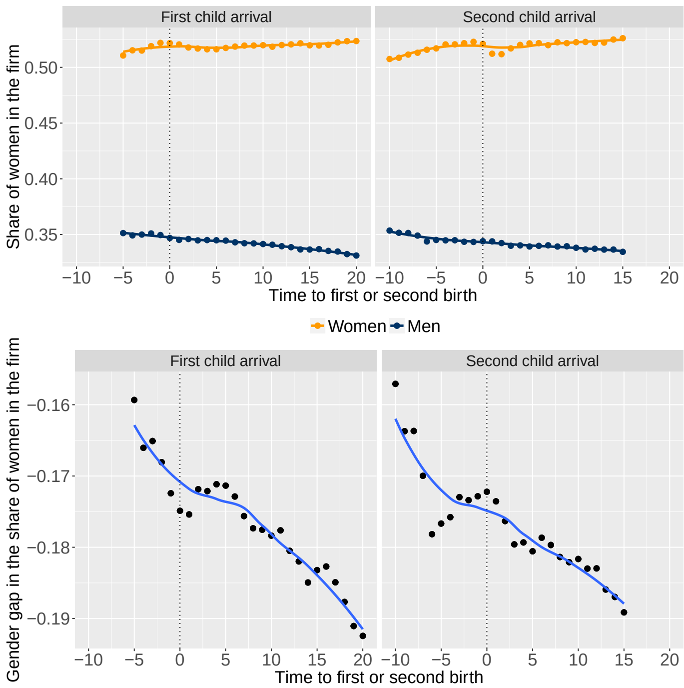
6.4 Characterizing the Firms Where Parents Work

The sorting effect highlighted before can also be described by looking at the composition of the firms where mothers and fathers work. Figure 8 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 50 and 55 % vs. 35 % for fathers). This relative segregation of women increases slowly along the career, except temporarily in the years around birth (especially the second one): 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.

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 on the type of firms mothers tend to work for sheds light on the underlying mechanisms contributing to mothers' sorting into low-paying firms.

The first dimension we look at is working time conditions at the firm level.

Figure 8. Gender Segregation Between Firms



Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1st to 4th, and comprehensive DADS files for coworkers' characteristics. Note: On average, the year of the birth of their first child, men are employed in firms with 34.7 % of female workers. For women, the proportion goes up to 52.2 %, corresponding to a -17.5 percentage point gap.

Our results suggest that after birth, mothers sort or are sorted in firms more likely to offer flexible hours. Figure 9a shows the proportion of part-time workers in the firms where mothers and fathers work. This proportion decreases slowly for men regardless of the final number of children they have and the rank of the birth from about 14 to 12 %. The proportion of part-time workers in the firm is roughly stable before and after the birth of the first child for mothers. However, the proportion of part-time colleagues increases after the second birth by approximately 1 pp. The difference between the two curves (men and women) displayed in the bottom panel of Figure 9a shows a break after birth.

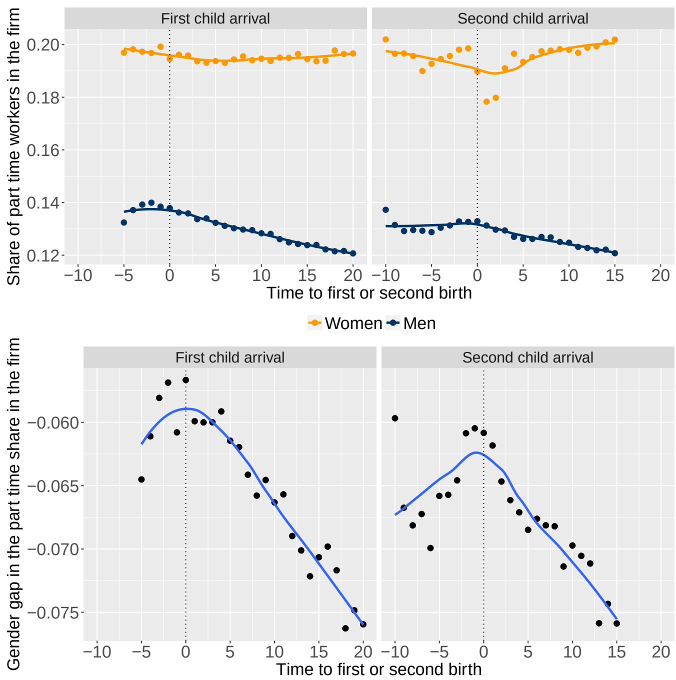
The sorting into lower-paying firms may also be due to an increased need of mothers to work close to their homes. To verify this, we look at the evolution of the proportion of parents working in the municipality where they live (Figure 9b). This proportion appears to be quite close for mothers and fathers before both the first and second birth. In both cases, the proportion of fathers working in their town of residence continues to decrease after the arrival of a child (at least for 10 years), whereas it starts increasing strongly for mothers. For second births, the difference between mothers and fathers starts decreasing shortly before birth, moreover, the change of the slope is stronger than for first births.[22]

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 et al., 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 10a). Firm-to-firm mobility rates decrease as workers age. The trends are similar for men and women but the latter experience a large drop after births of children. This pattern can be observed for all types of parents regardless of the rank of the child: the mobility rate gap increases from close to 0 to about 4 pp around birth. Subsequently, the mobility rate remains lower for mothers than for fathers for up to 10 years after the first birth. Again, it is remarkable that the mobility rate of mothers of one child only catches up, and even surpasses, fathers' one about 12 years after the first birth or 8 after the second one.

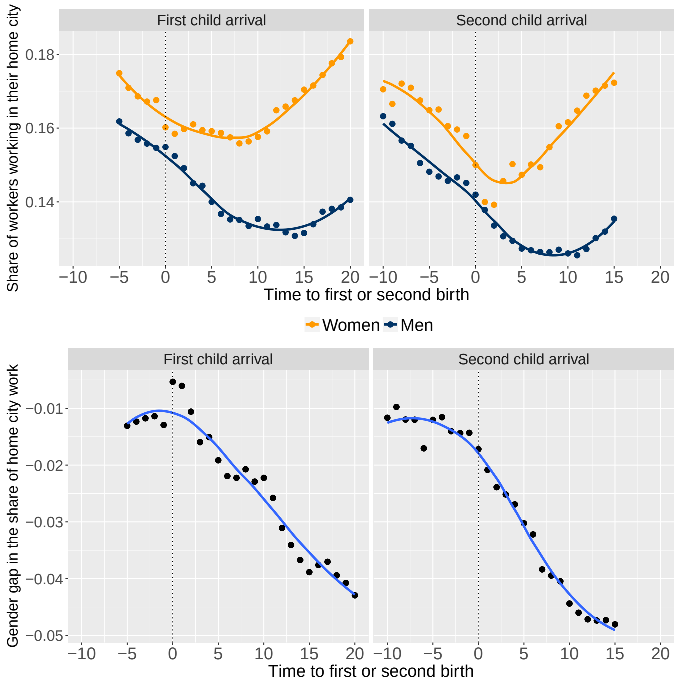
[22]Using the as-the-crow-flies home to work distance (measured as the distance between the centroids of the town of residence and the town of work) shows similar patterns - see Figure 28 in Appendix.

Figure 9. Firm Characteristics by gender and time to first birth

(a) Proportion of part time workers in the firm



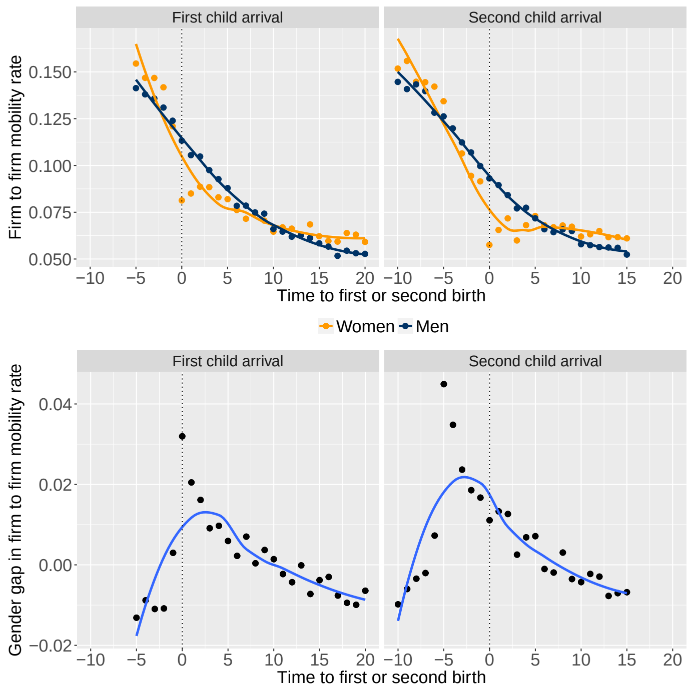
(b) Proportion of workers employed in their home city



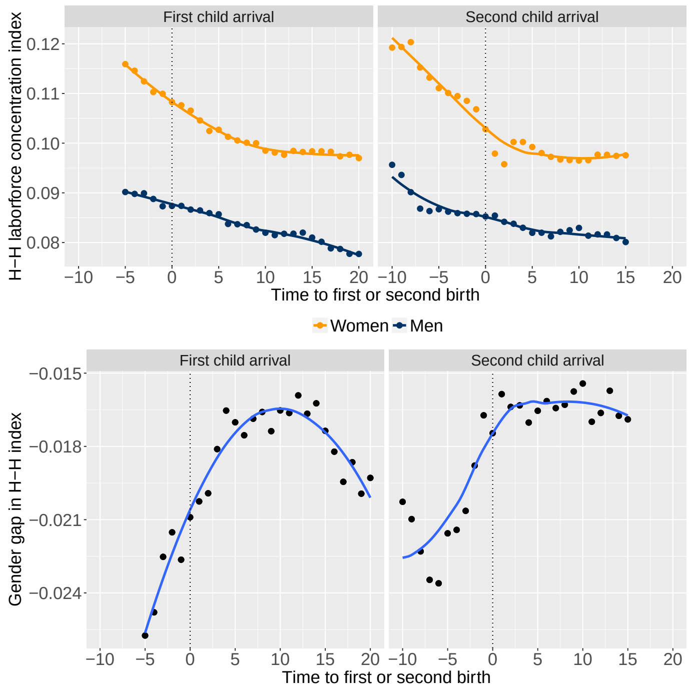
Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1st to 4th, and exhaustive DADS files for coworkers' characteristics. Note: On average, the year of the birth of their first child, 15.5 % of men work in the city where they live, to be compared with 16.0 % for women, corresponding to a -0.5 percentage point gap.

Figure 10. Firm sorting and labor market constraints

(a) Firm to firm mobility rate



(b) HH index on local labor markets



Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1^{st} to 4^{th}. Note: On average, the year of the birth of their first child, men are at work in areas × industries where the labor force concentration index is 0.087. The average index at that time to birth is 0.108 for women, corresponding to a -0.021 gender gap.

Further, we build an index of the local concentration on the labor market (Figure 10b) to investigate how far local constraints faced by women may provide monopsonic power to their potential employers.[23] 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 with important breaks in this trend. The gender difference tends to diminish before birth and increases afterwards for first child or plateaus for second ones. Taken together, these results suggest that the sorting effect increase after birth 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.[24]

7 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 et al. (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. The bargaining effect is close to zero: on average women tend to be paid as well as their male coworkers after controlling for observed characteristics and individual heterogeneity. We estimate a positive sorting

[23] 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 proportion 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 each industry-average normalized index is equal to the normalized index grand average: $\overline{HH}_{ki}^{\text{norm}} = \overline{HH}_{ki} + \overline{HH}_{..} - \overline{HH}_{.i}$.

[24] We also looked at the proportion workers paid at the minimum wage in the firm (see Figure 29) which is also in line with the previous findings.

effect (12 to 15 % of the total gender gap in hourly wage), suggesting that firms contribute to the gender wage gap as women are at work in firms paying lower wages than men of comparable productivity.

We find that the sorting effect is much larger for parents, showing a significant heterogeneity in the firm effect over the life cycle. 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 at least two children, this pattern is sharper after the birth of the second child.

Gender differentials in selection into the labor market affect the progress of these trends along the careers, especially just after birth, both through the sorting and the bargaining effects. We bring pieces of evidence that selection differentials are due to selection in the labour market of mothers who have higher expected firm effects, which at the end entails an under estimation of the gender gap and of its trends through the different channels.

We also show that the deepening of the sorting effect coincides with notable 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 wages, 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 sorting between different firms during careers and after a birth. There is a double child penalty: in addition to the direct wage cost within the firm, mothers experience wage losses through sorting between firms: about 5.5 pp wage loss is due to the birth of the first child and 2.0 to 2.5 pp to sorting later in the careers.

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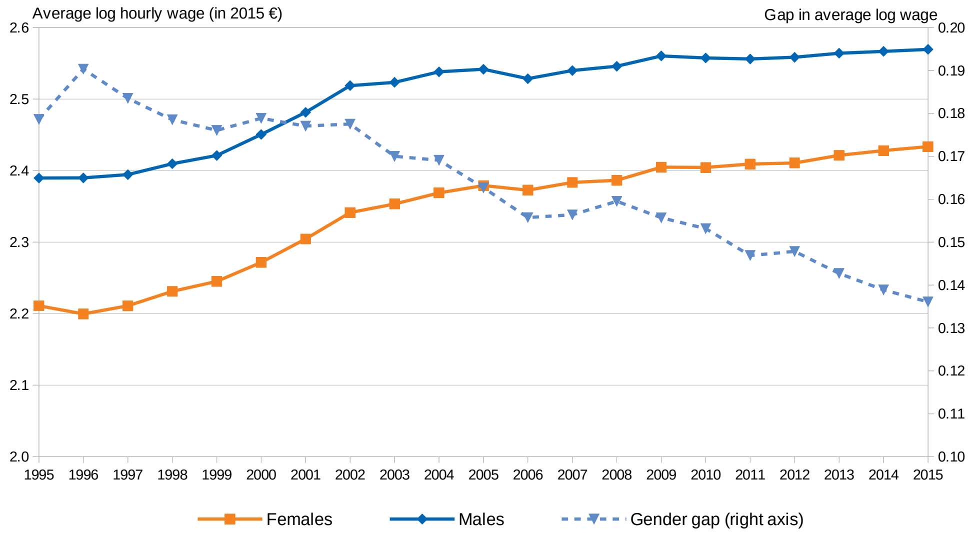
APPENDIX: FOR ONLINE PUBLICATION

A Wage evolutions for men and women since 1995

Figure 11. Average hourly wage for women and men since 1995



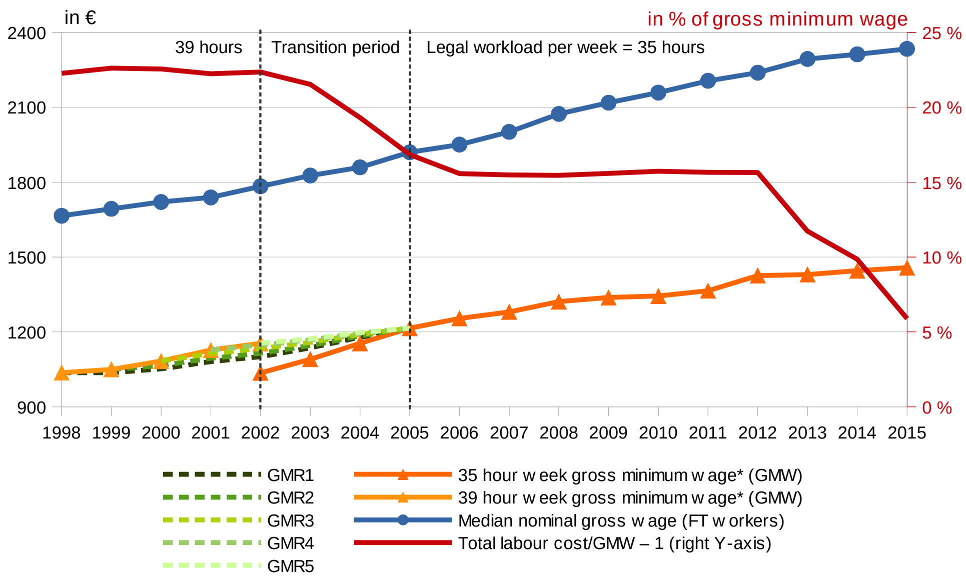
Source: DADS, *Panel Tous Salariés, 1995-2015*. Scope: Metropolitan France. Workers aged less than 16 or more than 65, self-employed farmers, craftsmen, shopkeepers, trainees, apprentices and private household workers are excluded. Note: in 2015, the average hourly wage for private sector employees is 12.4 € for females (2.43 in log) and 14.9 € for males (2.57 in log). The gender gap corresponds to 16.8 % of the average hourly male wage or a 0.14 log gap.

B More on the role of the minimum wage

Between 1998 and 2005, the *35-hour working week* laws were gradually implemented.[25] To maintain monthly earnings of workers at the bottom of the wage distribution, monthly guaranteed salaries (GMR) were enforced. As shown by figure 12, these GMR then converged in 2005 to a unique minimum wage.

[25] *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.*

Figure 12. Gross minimum and median wages, and relative labor cost at the minimum wage since 1998



Source: DADS, *Panel Tous Salariés*. Note: In 2015, the gross median wage for full-time workers is 2,334 €; the gross minimum wage for a full-time job (35 hours a week) is 1,458 €; the total labor cost at the minimum wage (gross minimum wage and legal social contributions paid by the employer, net of social exemptions and of the tax credit for competitiveness and employment that applies for workers paid less than 1.6 minimum wage since 2013) amounts to 5.88 % of the gross minimum wage.

C Conditions for identification

Card et al. (2018) point that the OLS estimates of equation (2) are unbiased provided that:

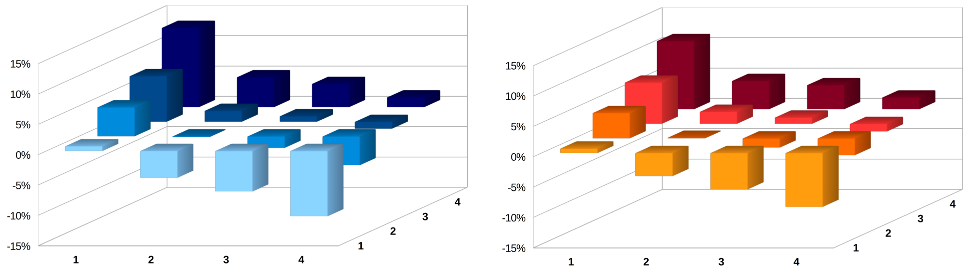
$$\mathbb{E}\left[(r_{it} - \bar{r}_i)\left(\mathbf{1}_{J(i,t)=j} - \frac{1}{T}\sum_{t=1}^{T}\mathbf{1}_{J(i,t)=j}\right)\right] = 0, \quad \forall j \in \{1...J\}. \qquad (5)$$

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 ($\bar{r}_i$). Since r_{it} encompasses shocks on worker, firm or worker-firm match 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 et al. (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 13a (men) and 13b (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 13.1 % wage increase on average. Symmetrically a man going from a high paying firm (Q4) to a low paying one (Q1) can expect a 10.8 % wage drop.

Figure 13. Average wage change for movers conditional on origin (X-axis) and destination (Z-axis) firm average wage

(a) Male Workers (b) Female Workers

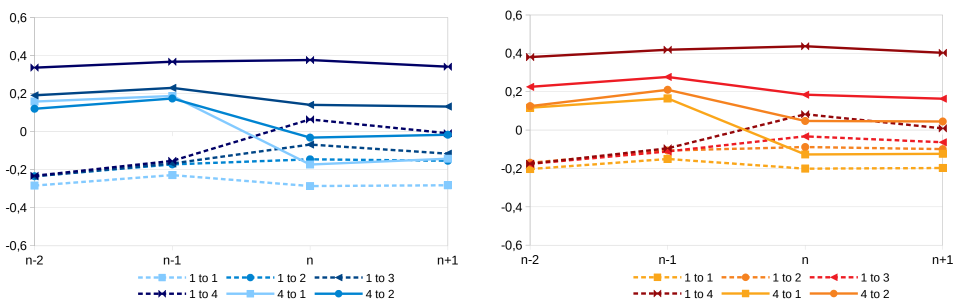


Source: DADS, *Panel Tous Salariés, 1995-2015*. Note: A male worker moving from a firm paying average wages below the bottom quartile (Q1) of the wage distribution to a firm above top quartile (Q4) gets a average 13.1 % increase in wage. Symmetrically a male going from a Q4 firm to a Q1 firm can expect a 10.8 % drop in wage.

Besides symmetry, the exogenous mobility condition implies absence of transitory wage shocks 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 (those used in the two-way fixed effect model) for movers. Figures 14a and 14b 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. We find no evidence of such shocks that help predict before or after mobility wage residuals. Based on these elements, exogenous mobility seems to be a reasonable assumption in our dataset.

Figure 14. 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

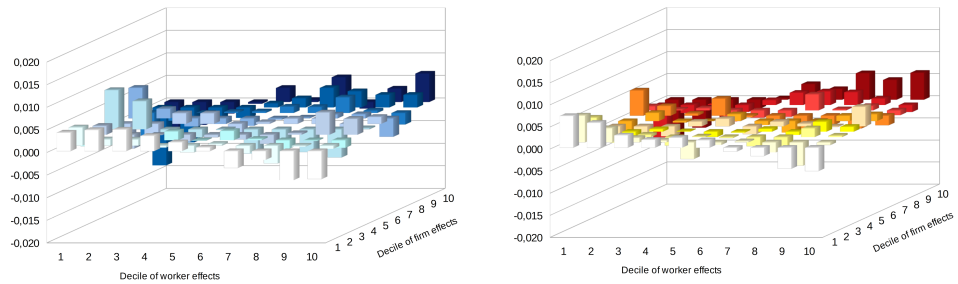


Source: DADS, *Panel Tous Salariés, 1995-2015*. Note: A male worker moving from a firm paying average wages below the bottom quartile (Q1) of the wage distribution to a firm above top quartile (Q4) has an average residual wage of -0.23 two years before moving, -0.15 the year before mobility, 0.06 the year he moves and -0.01 the following year. Symmetrically a male going from Q4 firm to a Q1 firm can expect a residual wage of 0.16 two years before moving, 0.19 the year before mobility, -0.17 the year he moves and -0.14 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 15a and 15b). If wages depend not only on worker and firm productivity, but also on the *interaction* of the two factors, the residuals should follow 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 15a and 15b show no such 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 -0.01 and 0.01 which corresponds to less than +/-0.5 % of the average hourly log wage for either men and women. The order of magnitude of the residuals is comparable to the one obtained for Portugal by Card et al. (2016) and for Germany by Card et al. (2013).

Figure 15. Mean of wage residuals conditional on deciles of worker and firm fixed effects

(a) Male Workers (b) Female Workers



Source: DADS, *Panel Tous Salariés, 1995-2015*. 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 0.006.

D Lowest-VA normalization

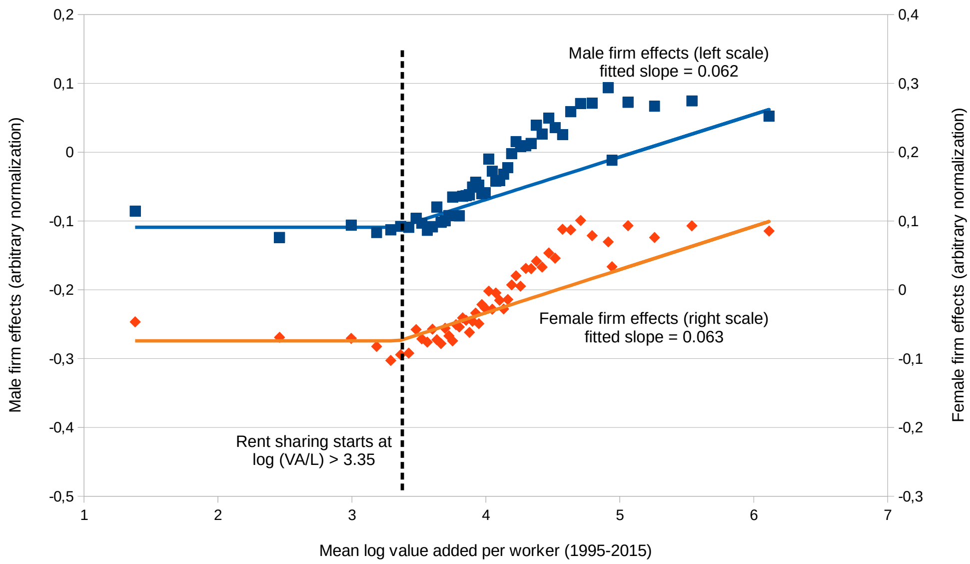
To check for the robustness of our choice, we also use the group of firms with the lowest value-added per worker as an alternative normalization following the approach of Card et al. (2016) and we fix to zero the average firm effects of this group of firms.

The choice of the threshold defining this group is based on Figure 16: above a log value-added per worker of approximately 3, there is a positive relationship 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.[26] The optimal level of log value-added per worker

[26] We conduct a further analysis of the estimated firm fixed effects in Appendix F. 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.

under which a firm is considered in the zero fixed effect group is 3.3.[27] 1,709 firms are below this threshold, representing 7.3 % of firms and 6.2 % of worker observations.

Figure 16. Firm effects according to log per capita value-added



Source: DADS, *Panel Tous Salariés, 1995-2015.* Lecture: Firms in the dual connected set are grouped into 50 bins according to their average log value-added per worker over the period. 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 two percentiles (average VA = 6.11) female premia before normalization are equal to 0.085 and male premia are equal to 0.052.

E Separate estimations for 1995-2004 and 2005-2014

We compare sorting and bargaining effects with our baseline model for period 1995-2004 and 2005-2015 (Table 7), before and after the rapid growth of the minimum wage. Setting to zero either the average firm effect in the accommodation and food services industry or among firms with a low value-added per worker we find a larger bargaining effect (in absolute and

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

relative terms) between 1995-2004 than between 2005-2015, when the minimum wage is smaller. 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. The sorting effect slightly grows between the two periods relative to total gender wage gap. This could be due to other underlying trends, and how low-wage jobs are distributed between firms.

Table 7. Sorting and bargaining contributions to the gender wage gap in 1995-2004 and 2005-2015

	Gender wage gap	Firm contribution	Sorting (a)	Sorting (b)	Bargaining (c)	Bargaining (d)
1995-2015 (VA)	0.172	0.014	0.020	0.025	-0.006	-0.011
		8.1 %	*11.6 %*	*14.5 %*	*-3.5 %*	*-6.4 %*
1995-2004	0.192	0.015	0.018	0.023	-0.003	-0.008
		7.8 %	*9.4 %*	*12.0 %*	*-1.6 %*	*-4.2 %*
2005-2015	0.166	0.013	0.020	0.026	-0.007	-0.013
		7.8 %	*12.0 %*	*15.7 %*	*-4.2 %*	*-7.8 %*
1995-2015 (Food)	0.172	0.028	0.020	0.025	0.008	0.003
		16.3 %	*11.6 %*	*14.5 %*	*4.7 %*	*1.7 %*
1995-2004	0.192	0.029	0.018	0.023	0.011	0.006
		15.1 %	*9.4 %*	*12.0 %*	*5.7 %*	*3.1 %*
2005-2015	0.166	0.028	0.020	0.026	0.008	0.002
		16.9 %	*12.0 %*	*15.7 %*	*4.8 %*	*1.2 %*

Source: DADS, *Panel Tous Salariés*. Total number of observations = 1,547,348; between 1995 and 2004 = 512,135; between 2005 and 2015 = 1,035,213. See for instance Table 3 for further details on sorting and bargaining computation.

F Firm fixed effects, within firm gender gap in firm effects, and firm characteristics

To check whether our results are consistent with the model, we analyze the firm fixed effects - either the average for both male and female workers, or the within firm gender gap - as functions of several firm characteristics.[28] For each firm we average the estimated male and female firm fixed

[28] Here we use the fixed effects obtained from the lowest value-added 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.

effects, and we regress this variable on a set of firm covariates. Findings for different specifications are presented in Table 8. Table 9 gives estimates when regressing the within firm gender gap in firm effects on the same firm characteristics. 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, the investments 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, average 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 premia about 5 to 6 % higher than firms of the first quartile. Firms in the top quartile of assets per worker distribution also pay on average higher firm premia. Furthermore, firm fixed effects are higher in firms employing a large proportion of executives and clerks (relative to blue collars). In contrast, fixed effects are dramatically lower in firms with a high proportion of workers paid at the minimum wage level, and of white collars. A higher proportion of women among white collars is related to somewhat lower firm fixed effects. This finding may indicate that firms where occupations are segregated by gender pay lower wages than others. Finally, we do not find significant differences between exporting and non-exporting firms, nor relationship between the level of investments per worker and firm premia.

Contrary to average firm fixed effects, within firm gender gaps in firm effects are very little explained by firm level variables (Table 9). In particular, there is no evidence of a relationship between firm productivity indicators (either value-added, assets or investments) and the gender gap in firm effects. After controlling for industries and collective agreements, only the proportion of females among white and blue collars and the fact that the firm exports products are significantly related to higher gender gap in firm effects.

Table 8. Average firm premia and firm characteristics

	(1)	(2)	(3)
Constant	-0.015**	-0.005	0.009
	(0.007)	(0.009)	(0.026)
Average (age-40)/40	0.020**	0.022***	0.016*
	(0.008)	(0.008)	(0.009)
% part time workers	0.011	0.013	-0.007
	(0.010)	(0.010)	(0.013)
% minimum wage earners	-0.211***	-0.196***	-0.188***
	(0.015)	(0.015)	(0.017)
(%F-%M) paid at the min. wage	0.027	0.020	0.013
	(0.020)	(0.020)	(0.021)
% managers	0.214***	0.212***	0.208***
	(0.009)	(0.012)	(0.014)
% clerks	0.040***	0.042***	0.046***
	(0.011)	(0.012)	(0.014)
% white collar workers	-0.049***	-0.029***	-0.004
	(0.006)	(0.007)	(0.010)
% female among managers	0.011*	0.012*	0.012*
	(0.006)	(0.006)	(0.006)
% female among clerks	0.009	0.013*	0.008
	(0.007)	(0.007)	(0.007)
% female among white collars	-0.011	-0.017**	-0.016**
	(0.007)	(0.007)	(0.007)
% female among blue collars	0.010	0.014**	0.012*
	(0.006)	(0.007)	(0.007)
Number of workers/1000	0.000	0.000	0.001
	(0.000)	(0.000)	(0.001)
Exporting firm	0.002	0.004	0.006
	(0.004)	(0.004)	(0.004)
Value-added per worker (*ref=1*)			
Quartile 2	-0.004	-0.003	-0.002
	(0.004)	(0.004)	(0.005)
Quartile 3	0.018***	0.021***	0.022***
	(0.005)	(0.005)	(0.005)
Quartile 4	0.055***	0.061***	0.053***
	(0.005)	(0.005)	(0.007)
Assets per worker (*ref=1*)			
Quartile 2	0.005	0.002	0.005
	(0.005)	(0.005)	(0.005)
Quartile 3	0.006	0.005	0.005
	(0.005)	(0.006)	(0.006)
Quartile 4	0.020***	0.014**	0.012*
	(0.006)	(0.006)	(0.007)
Investments per worker (*ref=1*)			
Quartile 2	-0.000	-0.002	0.000
	(0.005)	(0.005)	(0.005)
Quartile 3	0.001	-0.001	0.003
	(0.005)	(0.005)	(0.005)
Quartile 4	-0.001	-0.003	-0.002
	(0.006)	(0.006)	(0.006)
Industry dummies	-	yes	yes
Collective agreements dummies	-	-	yes
Adjusted R^2	0.237	0.256	0.279
Number of observations	10,859	10,856	10,856

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

Table 9. Within firm gender gap in firm premia

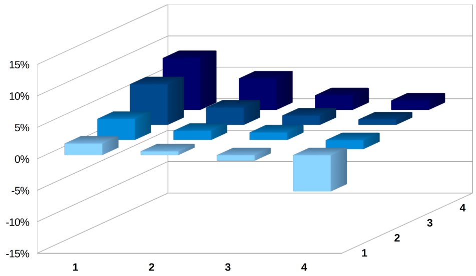
	(1)	(2)	(3)
Constant	-0.061***	-0.058***	-0.010
	(0.012)	(0.015)	(0.047)
Average (age-40)/40	0.001	0.001	-0.003
	(0.014)	(0.014)	(0.015)
% part time workers	-0.001	0.002	0.014
	(0.017)	(0.017)	(0.023)
% minimum wage earners	0.019	0.011	0.011
	(0.027)	(0.027)	(0.030)
(%F-%M) paid at the min. wage	0.055	0.059*	0.052
	(0.035)	(0.035)	(0.037)
% managers	0.034**	0.023	0.015
	(0.015)	(0.020)	(0.025)
% clerks	0.033*	0.037*	0.036
	(0.020)	(0.022)	(0.026)
% white collar workers	0.015	0.023*	0.032*
	(0.011)	(0.013)	(0.018)
% female among managers	-0.014	-0.007	-0.003
	(0.011)	(0.011)	(0.012)
% female among clerks	-0.012	-0.004	-0.006
	(0.012)	(0.012)	(0.013)
% female among white collars	0.033***	0.028**	0.022*
	(0.012)	(0.013)	(0.013)
% female among blue collars	0.023*	0.027**	0.031**
	(0.012)	(0.012)	(0.013)
Number of workers/1000	-0.000	-0.000	-0.001
	(0.000)	(0.000)	(0.001)
Exporting firm	0.020***	0.016**	0.018**
	(0.006)	(0.007)	(0.008)
Value-added per worker (*ref=1*)			
Quartile 2	0.010	0.007	0.006
	(0.008)	(0.008)	(0.008)
Quartile 3	0.013	0.009	0.007
	(0.009)	(0.009)	(0.010)
Quartile 4	0.023**	0.020**	0.019
	(0.009)	(0.010)	(0.012)
Assets per worker (*ref=1*)			
Quartile 2	0.003	0.002	-0.001
	(0.009)	(0.009)	(0.010)
Quartile 3	-0.005	-0.007	-0.011
	(0.009)	(0.010)	(0.011)
Quartile 4	-0.013	-0.014	-0.018
	(0.010)	(0.011)	(0.012)
Investments per worker (*ref=1*)			
Quartile 2	-0.007	-0.006	-0.003
	(0.008)	(0.008)	(0.009)
Quartile 3	-0.002	-0.002	0.001
	(0.009)	(0.009)	(0.009)
Quartile 4	0.005	0.005	0.009
	(0.010)	(0.010)	(0.011)
Industry dummies	-	yes	yes
Collective bargaining dummies	-	-	yes
Adjusted R^2	0.003	0.005	0.004
Number of observations	10,859	10,856	10,856

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

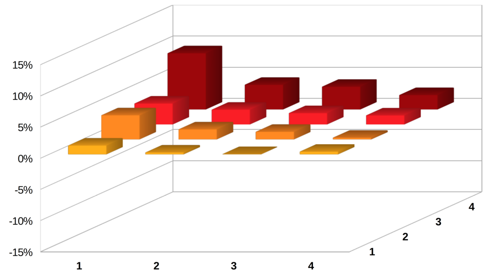
G Mobility and additive separability (Executives)

Figure 17. Average wage changes for executive movers conditional on origin and destination firm average wages

(a) Male Workers



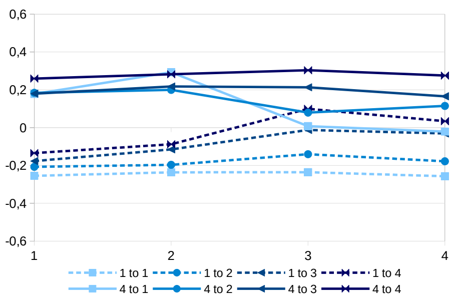
(b) Female Workers



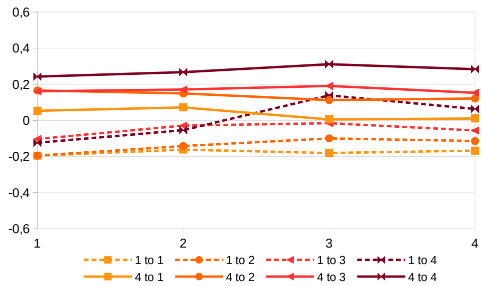
Source: DADS, *Panel Tous Salariés, 1995-2015.* Note: Leaving a Q1 firm for a Q4 firm yields an average wage gain of 8.3 % to male executives.

Figure 18. 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



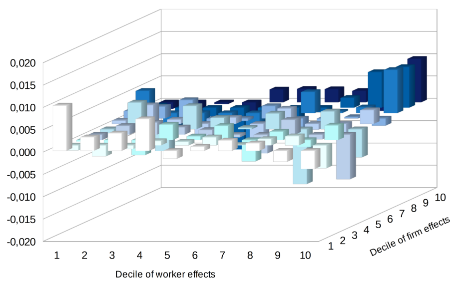
(b) Female Workers



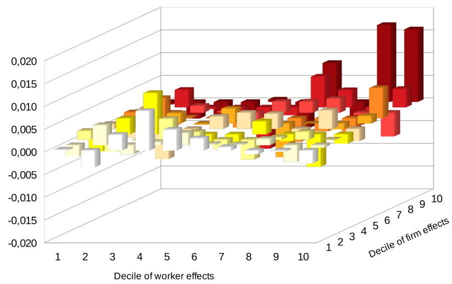
Note: Mean wage residual trend for a male executive going from a Q1 to Q4 firm is: -0.13, -0.09, 0.10 and 0.03.

Figure 19. Mean wage residuals for executives conditional on deciles of worker and firm fixed effects

(a) Male Workers



(b) Female Workers



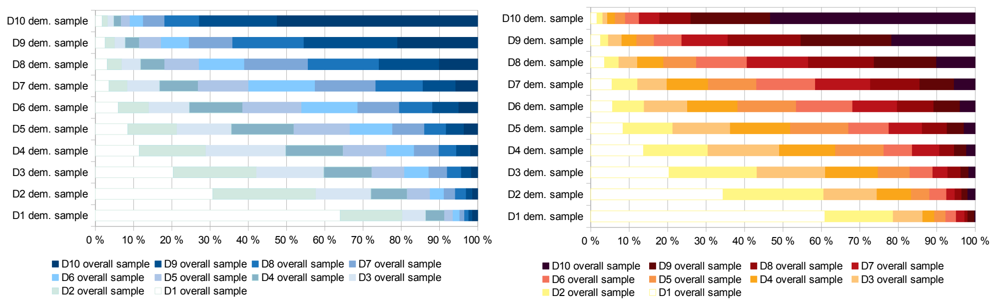
Note: Mean wage residuals for male executives in D10 working in a D10 firm is 0.010.

H Robustness to sample size

To show the similarities between estimations from the two samples (whole sample and demographic sample), we check that the rank of a given firm, in the distribution of male/female fixed effect, remains roughly the same in both sets of estimates. This is true in particular when workers are located at the tails of the distribution: figure 20 illustrates the consistency of deciles of firm fixed effects obtained from both overall (Y-axis) and demographic (X-axis) samples).

Figure 20. Distribution of firms per deciles of the overall sample firm effect distribution according to their position in the demographic sample firm effect distribution

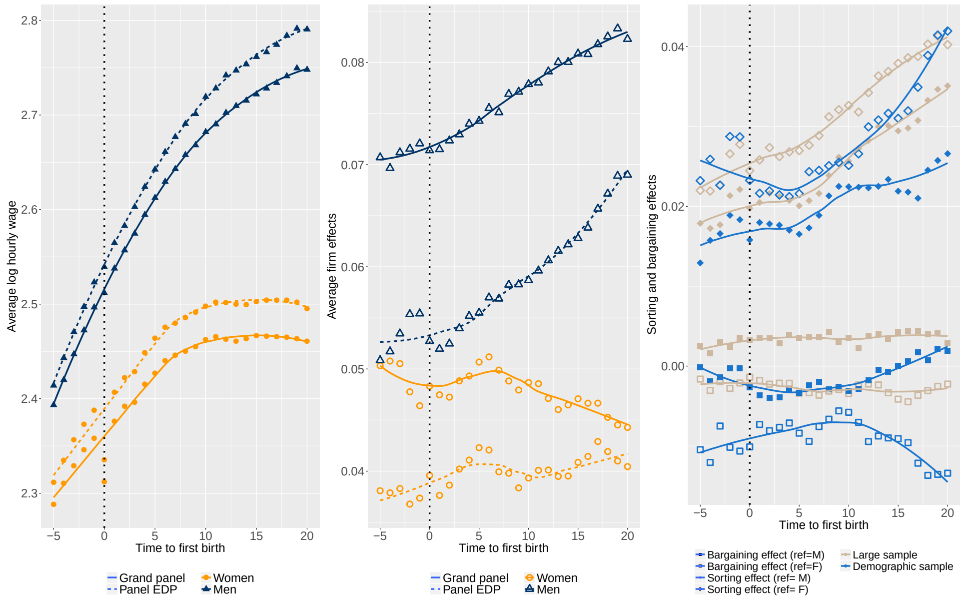
(a) Male firm effects (b) Female firm effects



Note: 64.0 % of firms whose male firm effect is in the bottom decile of the firm distribution in the demographic sample estimation also have a male firm effect in the bottom decile in the overall sample estimation.

I Event studies with large sample vs demographic subsample estimates

Figure 21. Firm premia, sorting, and bargaining effects over time to first birth in large sample vs. demographic sample

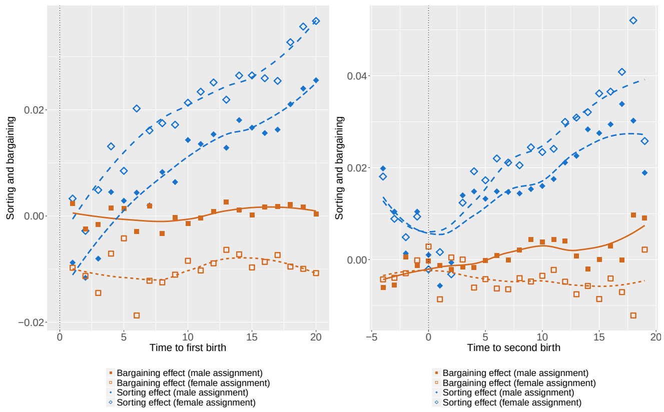


Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1st to 4th. Note: The average firm fixed effect estimated in the overall sample (respectively in the demographic sample) for females 5 years after their first child was born is 0.051 (resp. 0.042); for males, it is 0.074 (resp. 0.055). At this time to first birth, using male distribution into firms as reference gives a sorting effect of 0.020 (resp. 0.017) and a bargaining effect of 0.003 (resp. -0.003).

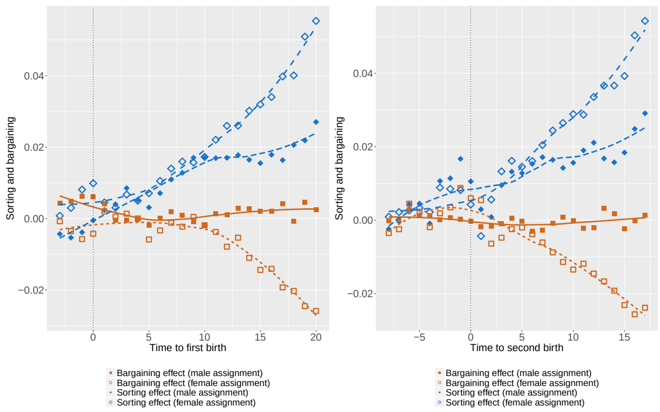
J Firm effects and time to birth by cohorts

Figure 22. Sorting and bargaining effects by cohort and time to births

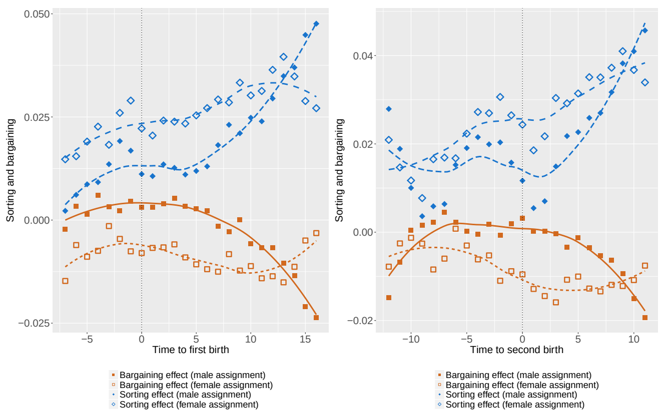
(a) Born 1960-1964



(b) Born 1966-1970



(c) Born 1972-1976

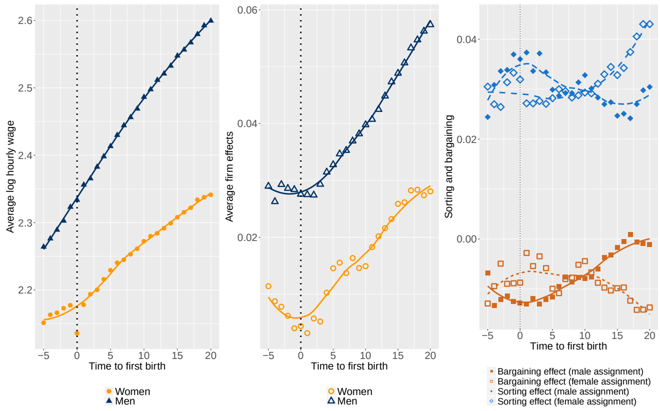


Source: DADS, *Panel Tous Salariés*. Note: In cohorts born from 1972 to 1976, five years after the first child was born using male distribution into firms as reference gives a sorting effect of 0.012 and a bargaining effect of 0.003.

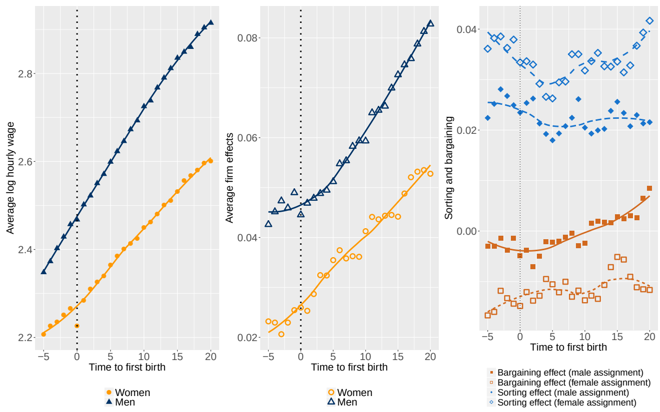
K Sorting and bargaining effects by education

Figure 23. Wage gap, firm premia, sorting, and bargaining effects by education

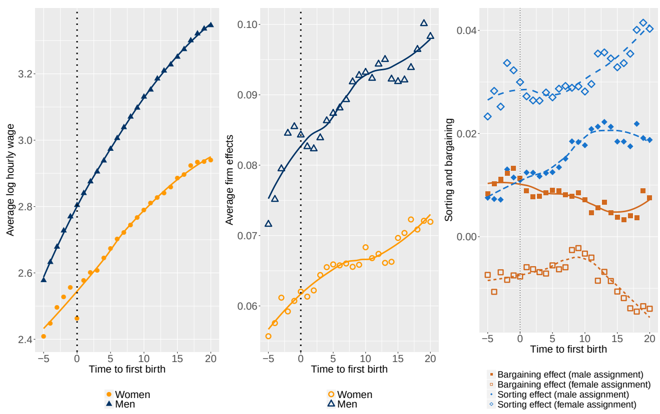
(a) Low-educated workers



(b) High-school educated workers



(c) College educated workers



Source: DADS, *Panel Tous Salariés, 1995-2015*, born between October 1^{st} and 4^{th}. Note: The average firm fixed effect for college educated females 5 years after their first child was born is 0.066; for males, it is 0.087. At this time to first birth, using male distribution into firms as reference gives a sorting effect of 0.013 and a bargaining effect of 0.009.

L Selection

L.1 Details on the high/low committed groups to the labor market

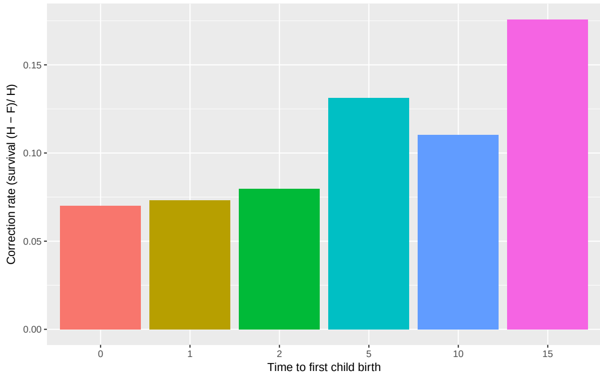
Table 10. Detailed descriptive statistics in the high/low-commitment subgroups to the labor market after first birth

	Year -1 (before first birth)			Years 1 to 10		
	Little committed	Highly committed	Δ	Little committed	Highly committed	Δ
Females						
Log hourly wage	2.318	2.433	0.115	2.368	2.518	0.150
Firm effect	0.031	0.048	0.017	0.028	0.052	0.024
Age	27.9	29.2	1.4	32.1	33.1	1.0
< high school	34.9 %	20.4 %	-14.5 %	36.6 %	20.9 %	-15.7 %
High school	27.7 %	24.0 %	-3.8 %	28.4 %	24.2 %	-4.2 %
> high school	37.4 %	55.7 %	18.3 %	35.0 %	54.9 %	19.9 %
Executives	9.7 %	17.4 %	7.7 %	10.5 %	21.2 %	10.7 %
Clerks	19.1 %	27.5 %	8.3 %	19.4 %	29.1 %	9.7 %
White collar workers	51.6 %	44.4 %	-7.2 %	50.8 %	40.0 %	-10.8 %
Blue collar workers	19.4 %	10.6 %	-8.7 %	19.2 %	9.6 %	-9.6 %
No child	89.8 %	85.5 %	-4.3 %	0.0 %	0.0 %	0.0 %
1 child	3.0 %	3.7 %	0.7 %	65.5 %	62.5 %	-3.0 %
2 children	5.1 %	8.0 %	2.9 %	30.1 %	33.7 %	3.6 %
3 children or more	2.1 %	2.7 %	0.7 %	4.5 %	3.9 %	-0.6 %
Number of observations	21,479	53,165	-	12,328	65,913	-
Males						
Log hourly wage	2.466	2.540	0.074	2.574	2.689	0.115
Firm effect	0.052	0.059	0.007	0.047	0.065	0.018
Age	29.3	30.9	1.6	34.2	35.0	0.8
< high school	44.2 %	37.3 %	-6.9 %	50.0 %	38.5 %	-11.5 %
High school	21.1 %	22.1 %	0.9 %	18.6 %	21.1 %	2.5 %
> high school	34.6 %	40.6 %	6.0 %	31.4 %	40.3 %	8.9 %
Executives	17.4 %	21.8 %	4.4 %	19.1 %	26.6 %	7.5 %
Clerks	19.7 %	24.0 %	4.3 %	20.1 %	24.7 %	4.6 %
White collar workers	17.7 %	15.6 %	-2.1 %	14.2 %	12.5 %	-1.7 %
Blue collar workers	45.0 %	38.4 %	-6.6 %	46.4 %	35.9 %	-10.5 %
No child	91.5 %	84.4 %	-7.1 %	0.0 %	0.0 %	0.0 %
1 child	2.2 %	3.6 %	1.4 %	60.4 %	59.6 %	-0.9 %
2 children	3.7 %	8.5 %	4.8 %	31.4 %	34.8 %	3.5 %
3 children or more	2.6 %	3.5 %	0.9 %	8.2 %	5.6 %	-2.6 %
Number of observations	20,722	70,686	-	10,401	85,755	-

Note: Workers the year before birth are considered as little or highly committed to the labor market whether they experienced work interruption the year of birth or the later or not. A women in the low commitment group receives an average log hourly wage of 2.318 the year before birth, to be compared with 2.433 for a women more committed to the labor market. The difference between the 2 groups amounts to 0.115. In the ten years following the first birth, the discrepancy across these two types of workers goes up to 0.150.

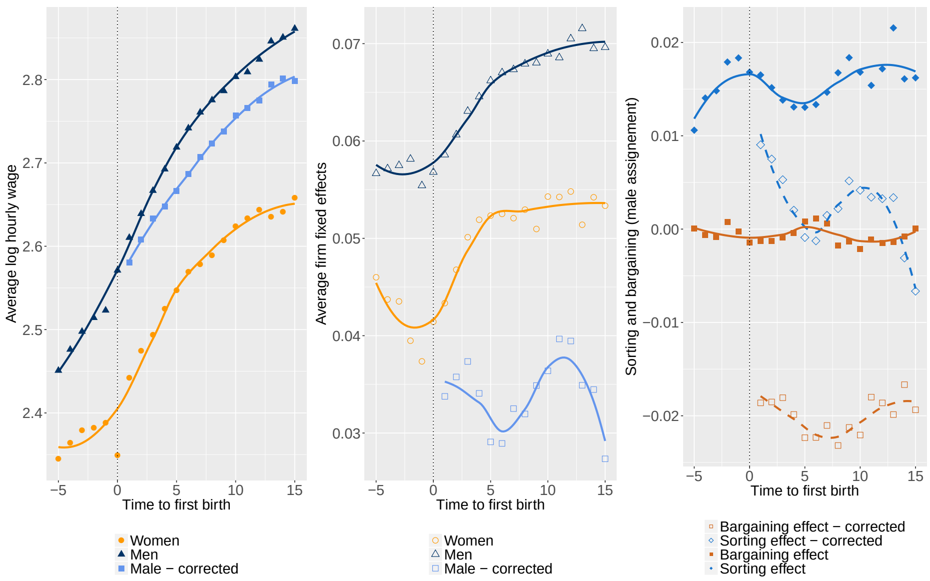
L.2 Lee bound approach

Figure 24. Gender gap in survival rate on the labor market for workers observed in employment the year before their first child



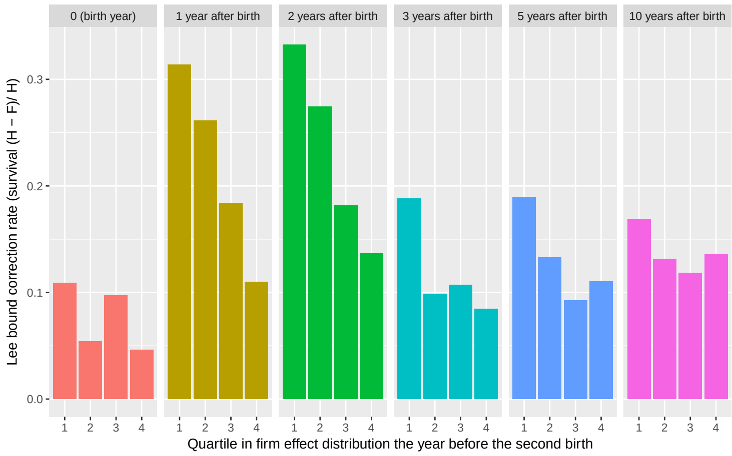
Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1st to 4th. Note: We consider the population of males and females who are observed in the sample the year before the birth of their first child. Two years after the birth of the first child, the relative differential of survival rate between males and females is 8.0 %.

Figure 25. Gender wage gap, firm effects, sorting and bargaining with Lee bounds correction on firm effects



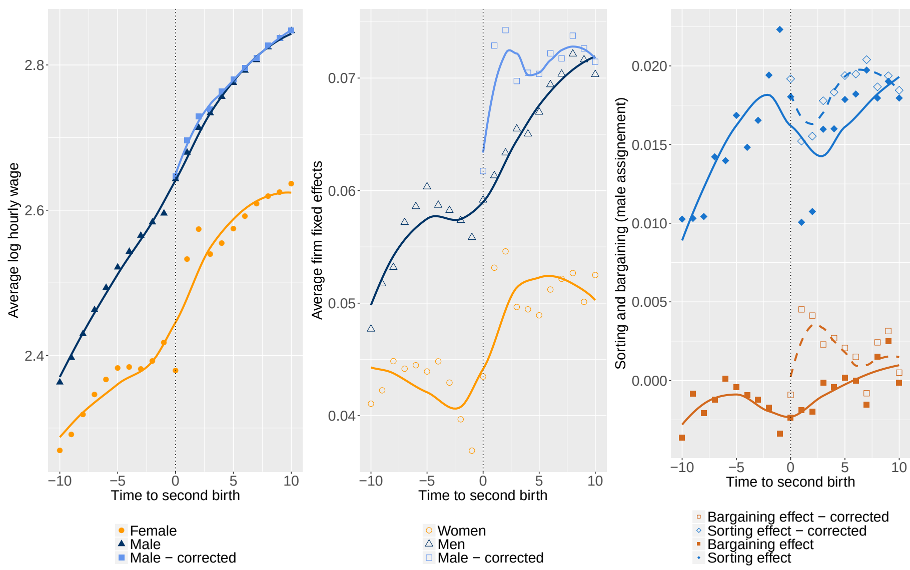
Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1st to 4th. Note: We consider the population of males and females who are observed in the sample the year before the birth of their first child. The average corrected male firm effects 5 years after the first birth is 0.029, to be compared with a 0.066 effect for males before correction, and a 0.052 firm effect among women. At this time to first birth, the corrected sorting effect amounts to -0.001 (instead of 0.013 without Lee correction) and the bargaining effects is -0.022 (instead of 0.001).

Figure 26. Gender gap in survival rate on the labor market for workers observed in employment the year before the birth of the second child, conditional on their position in before-birth firm fixed effect distribution



Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1st to 4th. Note: We consider the population of males and females who are observed in the sample the year before the birth of their second child, and we recover to which inter-quartile interval they belong at this date. Two years after the birth of the second child, the relative differential of survival rate between males and females who were in the bottom 25 % of their gender-specific firm effect distribution before birth is 33.2 %.

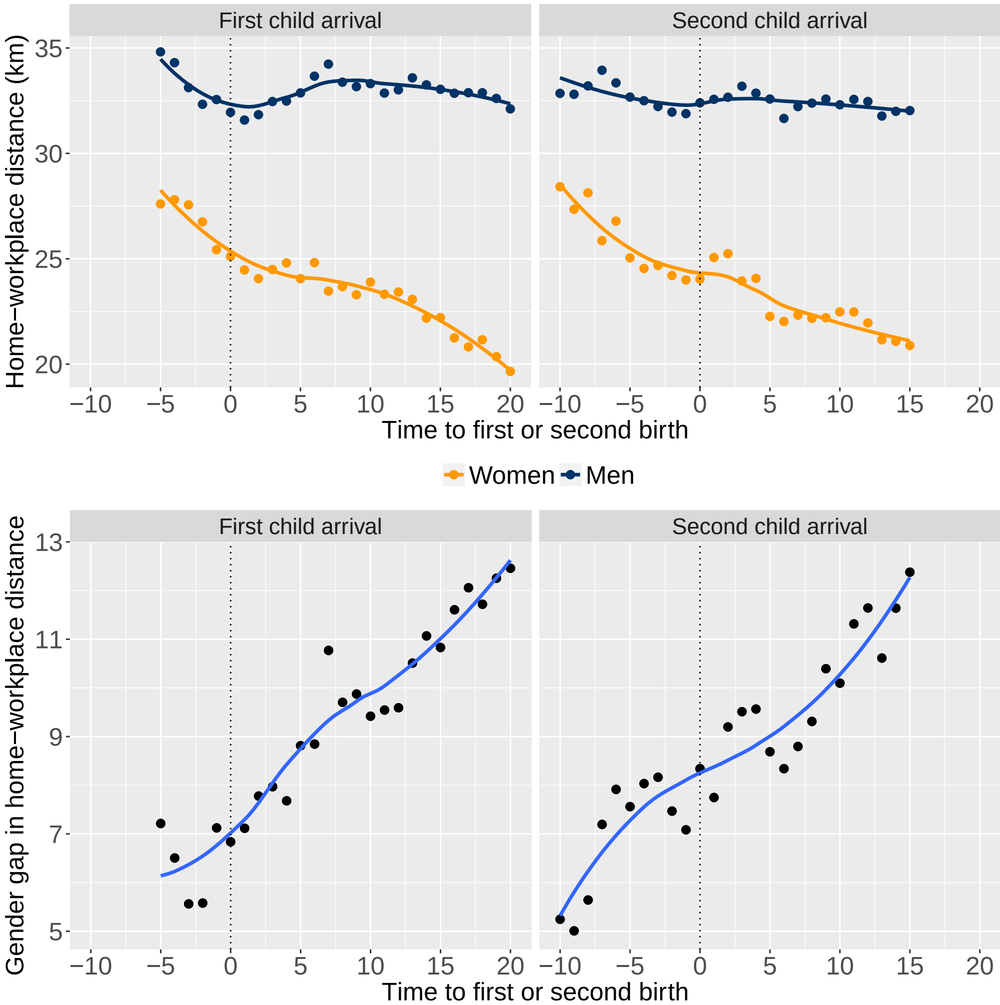
Figure 27. Gender wage gap, firm effects, sorting and bargaining with Lee bounds correction around second birth- based on quartile of average firm effects the year before second child was born



Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1st to 4th. Note: We consider the population of males and females who are observed in the sample the year before the birth of their second child. The average corrected male firm effects 5 years after the second birth is 0.070, to be compared with a 0.067 effect for males before correction, and a 0.049 firm effect among women. At this time to second birth, the corrected sorting effect amounts to 0.019 (instead of 0.018 without Lee correction) and the bargaining effects is 0.002 (instead of 0.000).

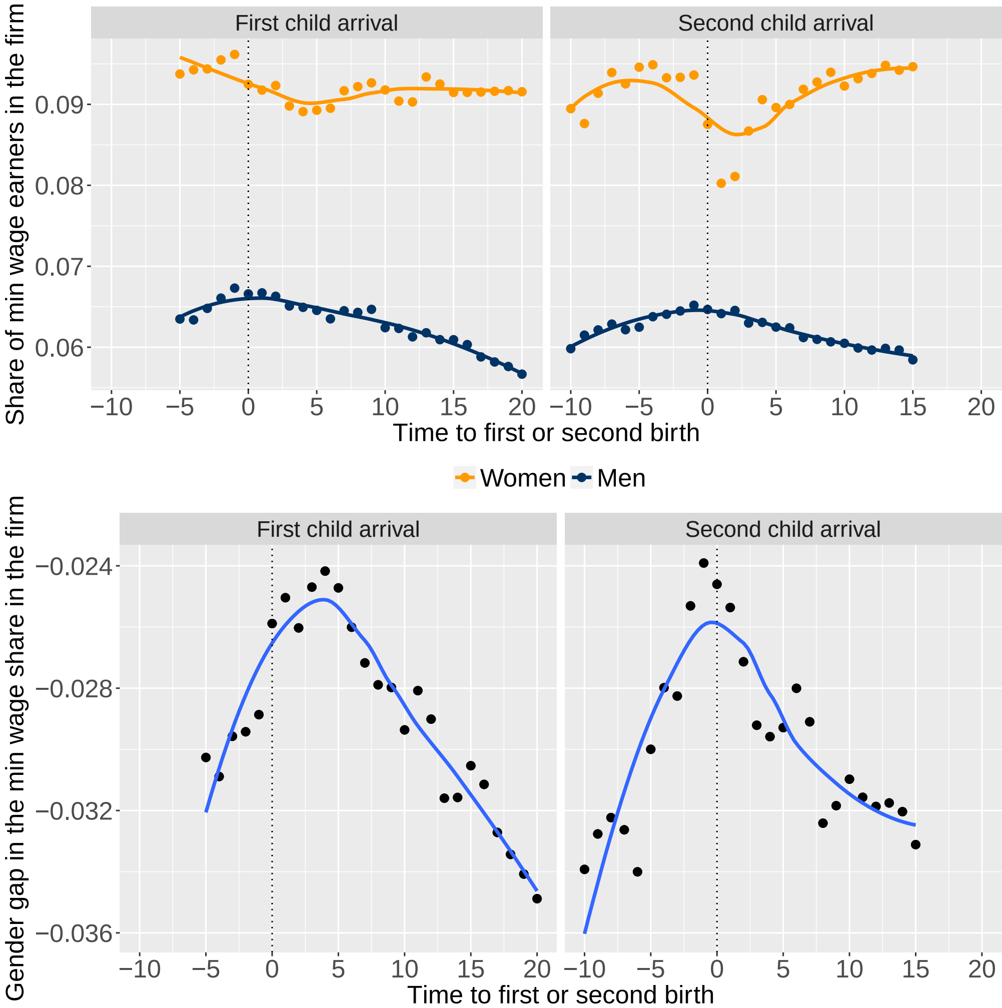
M Firm characteristics at different times to the birth of workers' children

Figure 28. Home-workplace distance



Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1st to 4th. Note: On average, five years after the birth of their first child, men work in cities 32.9 km away from their home city, to be compared with 24.1 km for women, corresponding to a 8.8 km gap.

Figure 29. Proportion of minimum wage-earners in firms



Source: DADS, *Panel Tous Salariés, 1995-2015*, born on October 1st to 4th, and comprehensive DADS files for coworkers' characteristics. Note: On average, five years after the birth of their first child, men are at work in firms paying 6.5 % of their workforce below 1.1 minimum wage. This proportion is of 8.9 % for women, corresponding to a -2.5 percentage point gap.