Gender Gaps in Performance Pay: New Evidence from Spain *

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ABSTRACT

In this paper we use micro data from a large wage survey carried out in Spain during 2006 to analyze the magnitude of the gender gap in the performance–pay (PP) component of hourly wages. Under the presumption that PP is determined in a more competitive fashion than the other wage components, we argue that there should be less room for gender discrimination in PP. Accordingly, the gender PP gap be lower than in total pay. However, our finding is just the opposite. After controlling for observable characteristics, non-random selection into performance-pay jobs and segregation into different firms and occupations, the estimated adjusted gap in favour of men remains fairly high (around 30 log points). Further, we document a "glass ceiling" pattern in this gap throughout the distribution of PP. After examining alternative ways of rationalizing these findings, we conjecture that monopsonistic exploitation exerted by employers might be the one more consistent with our evidence.

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Keywords: *performance pay*, *gender gaps, selection bias, quantile regressions, glass-ceiling*

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1. Introduction

In contrast to a vast body of literature which focuses on explaining differences in the total pay received by male and female workers, there are few papers that analyze gender gaps in one particular portion of wages that is becoming increasingly relevant, namely *performance-related pay* (henceforth PP).[1] This wage component is particularly interesting because it is generally considered as a better proxy of the conventional "wage equals marginal revenue product" textbook condition than other components (e.g. the base wage) that often do not depend so closely on individual performance. This is especially the case in countries where wages are set at semi- or fully-centralized collective bargaining agreements rather than in a decentralized fashion.

Following this intuitive reasoning, Lemieux *et al*. (2009) have analyzed the impact of PP on wage inequality in the US. Their basic hypothesis is that, through a widespread reduction in the cost of gathering and processing information, a growing incidence of PP may have contributed to an increase in inequality, mainly at the top of the wage distribution. Indeed, their finding that PP accounts for 25% of male wage inequality between the late 1970s and early 1980s supports this conjecture.

In this paper, we contribute to this literature by presenting new evidence drawn from a large dataset regarding pay and other working conditions of employees in Spain in 2006. The dataset contains a detailed breakdown of their total wages into different components. We re-examine the hypothesis posed by Lemieux *et al*. (2009), but from a different angle. Specifically, our interest lies first in estimating *gender gaps* in PP remuneration, to then interpret the main findings in the light of the different implications of several relevant theories about how these gaps may arise.

To our knowledge, there are only three studies in the relevant literature on gender gaps which are clear forerunners to ours. The first study is by Chauvin & Ash (1994), who use wage micro data drawn from a survey of business school graduates in the US to analyze how the gender gap structure changes across different pay components. Unlike our sample, which covers a large fraction of employees in Spain, theirs is rather small and does not allow to identify firm fixed effects, as we do here.

[1] Most papers in the literature on this topic have dealt mainly with analyzing the incentive effects of PP on productivity: see e.g. Dohmen and Falk (2009), Lazear (2000) and Lavy (2009), and the references therein. A gender perspective on this issue, albeit one related to the education system, can be found in Lavy (2011), whereas Bertrand and Hallock (2001) and Gayle *et al*. (2012) deal with gender gaps in CEO compensation.

The second one is by Booth and and Frank (1999), who use information drawn from the 1991 wave of the British Household Panel Survey (BHPS) on the extent of PP in the UK. They test for workers' selection into jobs offfering PP (henceforth, PP jobs) and for its effects on wages, distinguishing by gender. Unlike our case, their dataset does not contain the specific amount of PP remuneration but only its incidence. Their main finding is that women are 8% less likely to be on PP and that this type of remuneration raises male wages by 3 percentage points (pp.) more than female wages, a result which is attributed to some sort of gender discrimination.

The final forerunner, and the closest to our own work, is a recent paper by Manning and Saidi (2010), who use the two most recent waves of the Workplace Employee Relations Survey (WERS) in the UK (a matched employer-employee establishment-based survey) to focus on PP availability as an indicator of competition in the workplace. Their goal is to check whether the finding in laboratory experiments that men and women exhibit different attitudes to competition can be mapped into the real world (see, e.g., Gneezy et al., 2003, and Niederle and Vesterlund, 2007). Their main conclusion is that gender gaps in PP and even in work effort are modest. Hence, the ability of such theories to explain the gender pay gap is limited. Our findings differ drastically from theirs, being more in line with those of Booth and Frank (1999). In particular, for our sample of workers we estimate a much larger gender gap in PP than in the other wage components. We argue that, since the UK and Spain have quite different regulations regarding industrial relations and collective bargaining, our results illustrate the consequences that these differences may have gender gaps in PP.

From a theoretical perspective, there are several hypotheses about how PP availability could affect gender gaps in terms of both its magnitude and the selection of male and female workers into PP jobs:

- If PP is determined in a more competitive fashion than the other wage components, equally-performing men and women (with similar unobservable traits and equal preferences towards risk) should receive similar remuneration in terms of this wage component where firms. Hence, gender gaps in PP should be lower than in the other wage components which are less sensitive to meritocracy. This implication could be particularly relevant in countries with rigid labour markets where non-PP wages are set by collective bargaining, as is the case of Spain. As documented in Arumpalam *et al.* (2007), unions may represent more stongly the interests of male employees, e.g. due to their higher membership rates or because they work more frequently full-time. Moreover, if women were to perceive some form of (taste and/or statistical) discrimination against them operating in non-PP jobs, then it is likely that, to make up for these disadvantages, they would prefer to choose PP jobs.

- Nonetheless, even in the case of equal risk aversion, the assumption of equally-performing men and women may be a controversial one. In effect, insofar as effort in the marketplace may be negatively affected by housework, PP could also provide a relevant channel through which women's greater involvement in home production hinders their returns in the labor market. Therefore, gender differences in effort unrelated to the workplace may still entail gender differences in PP, even if this wage component is determined in a more competitive fashion than the remaining components.
- Another source of gender differences in PP could be occupational segregation, which may arise from several sources originating on either the worker's or the employer's side. As the regards the worker's side, women might sort themselves into non-PP jobs (e.g., most public sector jobs) because either they anticipate that these positions are more compatible with their greater household responsibilities, or they have different preferences with regards to the pecuniary and non-pecuniary elements of the job, or they are more risk averse than men. Thus, in line with the so-called *mommy track* hypothesis (see Mincer and Polacheck, 1977), they may willingly opt for jobs that entail steadier and, possibly, lower pay in exchange for less penalties in case of career interruptions. In a similar vein, another reason for why women may select themselves into non-PP jobs is that, as discussed earlier, they dislike competing with men in highly competitive jobs, which often entail merit-related pay. As for the firm's side, occupational segregation may arise from statistical discrimination: if employers expect lower female work attachment, they might be more reluctant to place women (with the same observable skills as men) in fast-track jobs, which are likely to involve PP (see Lazear and Rosen, 1990). Moreover, anticipation by women of some sort of statistical discrimination may discourage them from applying for these jobs, leading to self-fulfilling equilibria in both sides of the labour market (see Dolado et al., 2013). At any rate, irrespective of the underlying source of occupational segregation, the main implication of this hypothesis is a lower incidence of women in PP jobs where pay and working conditions are more risky. Yet, once selection into PP jobs is accounted for and the comparison takes place between similarly-skilled male and female employees in the same PP jobs (same firm and occupation), it is much more arguable that, everything else equal, women would prefer lower PP than men. In sum, our view is that, if gender differences in attitudes toward risk matter, they should mostly affect selection into PP jobs rather than differences in PP for those male and female employees who have freely chosen these jobs.
- Firms with monopsonistic power in frictional labor markets may discriminate against women in the PP component. This may be the case if employers perceive

women as having lower job mobility or lacking alternative job offers (see Booth et al, 2003 and Manning, 2003)

Given all these considerations, this paper seeks to dig deeper into the extent and the determinants of gender gaps in PP in Spain with the goal of identifying which of the above-mentioned theories of gender gaps fit best with our findings. Our data comes from the 2006 wave of the Spanish Earnings Structure Survey, which contains detailed micro-data information on the various components of the wage, such as the base wage, overtime pay and other wage complements. In comparison to the longitudinal dataset used by Manning and Saidi (2010), the cross-sectional nature of our dataset has the drawback of not being able to control for workers´ fixed effects. However, in exchange, it has the advantage of providing information on how PP is disaggregated by particular occupations within firms, whereas these authors only have information on the amount PP within firms and not by occupations. As will become clear below, the more disaggregated nature of our wage information becomes crucial in unraveling the alternative rationalizations of the gender PP gap. Furthermore, rather than concentrating exclusively on the gender PP gap at the mean, as the other papers do, we also analyze how gaps evolve throughout the PP distribution. Notice that this analysis is relevant, since the theories listed above differ in their implications on this issue.

The rest of the paper is structured as follows. Section 2 describes the wage survey and provides some basic descriptive statistics regarding the whole sample, the distribution and extent of PP, and the differences between the observable characteristics of workers in PP and non-PP jobs. In Section 3 we test our key hypothesis on whether PP is set in a more competitive way than other wage components. Section 4 reports adjusted PP gender gaps, once differences in personal and job characteristics across genders and non-random selection of workers into PP jobs are accounted for. Section 5 discusses which of the previous theories about PP gender gaps fits best with our findings. Finally, Section 6 draws some conclusions.

2. Data and descriptive statistics

The data source is the third (2006) wave of the Spanish Earnings Structure Survey (*Encuesta de Estructura Salarial* or ESS for short), which was the latest available wave at the time of writing this paper.[2] The survey is based on two-stage random

[2] The previous waves correspond to 1995 and 2002. In October 2012 another wave, corresponding to wages in 2010 was launched but unfortunately the sample size is much smaller than in the 2006 wave, due to budgetary restrictions. Another noticeable feature of ESS is that age brackets start with workers aged less than 30 and therefore include individuals aged 16-25. Yet, given that the participation rate of young people in this age cohort is

samples of workers from establishments in the manufacturing, construction and service industries, including small firms with less than 10 employees. First, establishments are randomly selected from the General Register of Payments records of the Social Security system, which are stratified by region and establishment size. In the second stage, samples of workers from each of the establishments selected are again randomly drawn. Overall, not only are sample sizes much larger than those provided by any other Spanish wage survey but, aside from wage remuneration, ESS also collects individual information on workers' demographics (such as age and educational attainment) and job characteristics (including industry, occupation, contract type, type of collective bargaining, export activity, establishment size, and region).

The main advantages of the EES (2006) is that it is the first wave to include a module where employers report detailed information on the breakdown of the total annual wages paid to their workers into fixed and variable components. Besides total monthly gross wages and effective (weekly) working hours, the ESS (2006) also provides information on both the ordinary (base wage and other complements due to shift-work, tenure, job risks, etc.) and non-ordinary components of annual gross earnings. Regarding the latter, two different types of payment are distinguished:

• ***Fixed Annual Non-ordinary Payments***. This payment "basically corresponds to extraordinary remuneration at Christmas and summer vacation time (known in Spanish as *pagas por navidad y verano*)[3], the standard rates for overtime work and participation in firms´ ordinary and extraordinary profits". It is specifically stated that these payments are known in advance by the employee, typically established at the collective bargaining level, and that they do not depend on the performance of either workers or firms.

• ***Variable Annual Non-ordinary Payments***. In contrast to the previous category, these are payments related to workers´ individual performance as well as to firms' overall performance. The amount is not known in advance by the worker and its precise definition changes from firm to firm: it is determined as a function of production/revenue targets, quality and quantity of sales, profits, etc. These payments are not received periodically, and they lump together bonuses, merit-related remuneration and piece rates.

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very low, since many of them are still in post-compulsory education (after 16 years of age), this shortcoming is unlikely to change our main results.

[3] This implies that the fixed part of the total annual gross wage is distributed into 12 ordinary installments and 2 extraordinary ones in June and December. This tradition dates back to the industrial relations setting of the Franco dictatorship period in Spain from 1939 to 1977.

Given this breakdown of total wage compensation, in what follows we identify the PP component as *Variable Annual Non-ordinary Payments.* Conversely, the non-PP component is identified as the sum of the ordinary wage and the *Fixed Annual Non-ordinary Payments.* To avoid potential differences in pay due to differences in hours worked, we also use available information on weekly hours in EES, to compute total pay as well as non-PP and PP remunerations in hourly terms.

One rather important issue to consider at this stage is that the *Variable Annual Non-ordinary Payments* component not only includes payment for PP but also profit-sharing schemes, as is also the case in the BHPS dataset used by Booth and Frank (1999). Thus, interpreting PP exclusively in terms of workers' merit pay may result in a non-negligible measurement error. Yet, since most of our analysis focuses on adjusted gender gaps in PP for men and women working in the same firm and occupation (18 occupational categories), differences in profit-sharing payments are likely to be small. As a result, there is some justification for interpreting the chosen definition of the PP component as mostly reflecting merit pay.

2.1. Description of the dataset

Our sample consists of full-time workers aged 18-65 for whom the interview month (October) was an ordinary period in terms of their labour status. Table 1 displays the weighted descriptive statistics for the male and female samples. Our sample covers 195,153 employees in almost 18,000 firms, out of which 129,930 (66.6%) are men while 65,233 (33.4%) are women. Although in 2006 retirement age was 65 for both genders, it is noteworthy that the share of female workers aged over 50 is 7 pp. lower than the male one, while the proportion of younger women is higher. This reflects the fact that the strong increase in female labour force participation in Spain is a fairly recent phenomenon which dates back only to the eighties.

The first finding to highlight is that the incidence of women in PP jobs is only slightly lower than males' (17.7% against 19.4%). At first sight, this preliminary evidence is not consistent with those theories which predict strong gender differences in workers' selection into PP jobs. However, this is an issue which deserves further scrutiny further below, once we adjust gaps for gender differences in observable characteristics.

Next, the demographics of workers reveal three distinctive features: (i) on average women have significantly higher educational attainment levels than men (e.g. the percentage of female workers with a university degree (32%) is almost twice that of men (18%), whereas the fraction of women with no more than primary education is 10 pp. lower (18% vs. 28%); (ii) the average age of women is about two years less than that of men (from interpolation of the mid-points of the different age brackets); (iii) female job tenure is about 1.5 years shorter than male tenure. As regards the characteristics of employers, we find that women are 9% more likely to work in larger firms (> 200 employees), and that they are 3% less likely to be covered by firm-level bargaining agreements.

Lastly, in terms of total gross hourly wages, the raw gender gap in favour of men is 21 log points while the corresponding gap for the subsample of PP workers reaches 25.6 log points. Among the latter, the most interesting finding is that the gender gap in the PP component (46 log points) doubles the gap in the fixed wage component (23.4 log points). Notice that this gender PP gap is strikingly higher than the one reported by Manning and Saidi (2010) for British workers, despite the fact that gender differences in participation in PP jobs are small in both countries. It is therefore interesting to explore the reasons for such contrasting results.

2.2. Characterizing PP remuneration

Table 2a compares the sample characteristics of workers and firms in the PP and non-PP samples, distinguishing by gender.[4] The main finding is that workers who receive PP are more skilled (40% of women and 28% of men in the PP sample have a college degree compared to 29% and 15% in the non-PP sample). Likewise, they are older (by about 10% in the 41-50 age cohort), have longer tenure (about 2.5 years longer for women and 4 years for men), are more likely to have an indefinite contract, and work in larger establishments (typically less subject to centralized bargaining).

Table 2b presents the incidence of PP jobs per sector and occupation. The sectors where PP is most and least prevalent are Financial Intermediation (60%) and Education (9%), respectively. The results per occupation confirm that the incidence of PP is much higher for the high-wage categories: 50% for Managers and 30% for Professionals and Technicians.

Finally, Table 2c reports the share of female workers who receive PP throughout the distribution of this wage component (i.e. the proportion of women among workers receiving PP in each of the deciles), which can be compared to the above-mentioned average share of women who receive PP in our sample (17.7%). As can be observed, there is a sharp decline in this proportion as we move upwards in the PP distribution - from 41% at the bottom to 16% at the top.

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[4] The non-PP sample includes those workers who do not report any positive variable annual non-ordinary payments. Some of these workers may have the right to receive PP remuneration in their employment contracts but, for one reason or another, they did not get it. Yet it is not possible to disentangle these two types of workers within the sample of non-PP workers. Given that the focus of the paper is not on PP per se but rather on gender gaps in PP, our implicit assumption here is that the distribution of this potential measurement error is similar among men and women and hence that it will not affect our results in a significant way.

2.3. Raw gender PP gaps

We next analyze the fraction of total hourly wages accounted for by PP and study more closely the size of the gender gap in this specific wage component. The first four columns in Table 3 present the total hourly wages for workers with PP schemes (expressed in €) across genders as well as the corresponding shares of total wages accounted by the PP component.

As can be observed, workers who receive PP earn on average much more than those who do not receive PP (about 64% and 50% higher wages for men and women, respectively), in line with the evidence offered in Table 2a about their higher skills and longer job tenure. Yet the share of total wages accounted for by the PP component turns out to be rather low (on average 7% for women and 9% for men). Yet, it increases throughout the wage distribution, reaching 22% (men) and 17% (women) at the 90th percentile (P90th). Taking both features together, de la Rica *et al*. (2010) report that the contribution of the gender gap in PP to the overall gender pay gap for the whole sample of workers is rather small: about 7% on average and 12% at the top of the wage distribution. Notwithstanding this, when the analysis is restricted to the PP sample, these contributions become larger: 18% on average and almost 25% at the top of the distribution.

Summing up, three main lessons can be drawn from the preliminary evidence presented so far: (I) there are no major gender differences in the incidence of PP, (II)) the gender gap in the (hourly) PP component is much higher than in the total hourly wage, particularly at the top of the wage distribution, and (III) PP makes its mark in higher wages, in agreement with the previous evidence on the higher observable skills of workers in the PP sample.

As stressed earlier, the finding that the gender gap in PP is much higher than in total hourly wages, even within firms and/or occupations (see below), constrasts sharply with the evidence for the UK by Manning and Saidi (2010). One plausible rationalization of these different patterns is suggested by the arguments provided by Dolado *et al*. (1997), who point out that employers in Spain improve the pay of high-skilled workers above the compressed wages (especially base wages) agreed with unions in the collective bargaining. This is done through formal and informal agreements involving PP schemes which are are set on a more discretionary basis by employers. Insofar as unions compress the base wage distribution and that these wage components are determined more by occupational categories and tenure than by individual characteristics, it is plausible that these arrangement lead to a consideraly lower (in raw terms) non-PP gender gap than the PP gender gap. This contrasts sharply with the situation in the UK, where union coverage is much lower than in Spain. This conjecture is supported by the fact that the standard deviation of the (logged) fixed component of total hourly wages in Spain (0.61 and 0.60 for men and women, respectively) is less than one-half of the standard deviation of the (logged) PP component (1.41 and 1.34, respectively).

However, before exploring potential explanations for the gender PP gap in Spain, several preliminary steps must be taken. First, in line with Lemieux *et al.* (2009), it needs to be shown that, even in a much more regulated labour market that the US one, PP depends more on workers´ characteristics than the other components of the hourly wage do. Next, we need to check if the above-mentioned patterns of the raw PP gender gap remain similar once differences in observed individual and job characteristics across genders are adjusted for, non-random selection into the PP sample of workers is corrected for, and comparisons are restricted to men and women working in the same firm and occupation. In other words, it is only under the competitive labor market paradigm and under similar observable characteristics that the documented gender PP gap can be considered to be "strikingly high", as indicated above. These issues are sequentially addressed in the next two sections.

3. Is PP determined in a competitive fashion?

This section first analyzes whether the PP component is "more attached to the worker" while the non-PP component of the wage is more "attached to the job". Following Lemieux *et al*. (2009), the basic insight is that, if PP responds mainly to workers´ productivity, then human capital variables – basically age, education and job tenure- should have higher market returns in PP jobs. Conversely, returns to job characteristics- such as firm size, sector, and tenure in the firm- should receive a higher market reward in non-PP component.

To address this issue, Table 4 reports standard *Mincerian* (logged) total hourly wage regressions estimated by OLS where the returns (estimated coefficients) to job and human capital variables are displayed separately in the first two columns for PP and non-PP samples, respectively. The last column, in turn, shows the results from a pooled regression where interactions of human capital and job characteristics with an indicator of receiving PP are added as covariates to test for statistically significant differences between returns in the two samples. Thus, denoting the hourly wage of worker *i* in firm *j* as W_{ij}, individual and job characteristics as X_i and X_j, respectively, and an indicator (1/0) for receiving PP as D_i , the estimated model is:

$$\ln W_{ij} = \beta_0 + \beta_1 D_i + X_i \beta_2 + X_j \beta_3 + D_i X_i \phi_1 + D_i X_j \phi_2 + \varepsilon_{ij}$$

where, from the previous considerations, it is expected that $\phi_1 > 0$ and $\phi_2 < 0$.

Our results are similar to those reported in Lemieux *et al.* (2009) for the US. For example, in the PP sample the returns to college and secondary education are 41% (0.304 vs. 0.215) and 60% (0.09 vs. 0.06) larger, respectively, than in the non-PP sample. Likewise, the returns to age, as a proxy for potential experience and, to a lesser extent, job tenure follow the same pattern. By contrast, the returns to firm size and other job characteristics are significantly higher in the non-PP sample. This is also the case for the coefficients estimated on the industry and occupational dummies, not reported here to save space. Overall, we interpret this evidence as supporting the view that PP is more closely linked to workers' productivity than the other wage components. Yet, the fact that, in general, the estimated returns on firms characteristics are statistically significant indicates that workers tend to be categorized by firms into jobs albeit less so in the PP sample.

4. Adjusted gender PP gaps

The next step is to compute gender PP gaps adjusting for differences in observed individual and job characteristics. However, the fact that slightly less than one-fifth of workers in the whole sample are subject to PP schemes and that they have different personal and job characteristics than non-PP workers leads us to consider that non-random selection of workers into the PP sample may be a relevant issue to address. This is particularly important if the selection process into PP is not exactly the same for males and females because ignoring these differences in selection may lead to biased estimates of the adjusted gender PP gaps.

4.1. Selectivity issues

Finding instruments suitable for addressing a potential selectivity bias is a very difficult task given that our dataset lacks information on family issues, such as civil status or number/age of children, which are the traditional instruments used in this context. Instead, we use the availability of wage bargaining at firm level (*Firm Agreement*) as the identifying variable in the participation equation. For given individual and other job characteristics, workers who end up in jobs with this type of decentralized wage agreement are more likely to receive PP than those in other jobs where unions play a prominent role in determining wages and often limit the use of PP schemes.

The choice of this indicator as an instrument could be criticized on the grounds that it could affect total wages. Yet we cannot find strong arguments as to why it should affect the magnitude of the PP component in the PP sample, since the amount of PP depends mainly on workers' performance and it is not clear why effort should be higher when bargaining takes place at the firm leve (e.g. its estimated coefficient is statistically insignificant in the first column of Table 4). This makes us inclined to trust the validity of this exclusion restriction despite the fact that there may not be strong conceptual arguments in its favour. At any rate, the results do not differ qualitatively from those achieved when we omit the restriction and rely exclusively on non-linearities to achieve identification. In what follows we present estimates with and without controlling for sample selection because, given that PP recipients show higher observable skills than the rest, not controlling for sample selection could lead to downward biased estimates of the actual gaps which must be taken into consideration when interpreting the results.

4.2. Gender gaps in participation in PP schemes

Table 5 presents the results of a probit model estimated to explain participation in the PP sample (PP=1, non-PP=0). This model is later used to compute the inverse Mills ratio in a conventional two-stage Heckit approach to control for selection in the estimation of (logged) hourly PP *Mincerian* regressions. In the first column we present the estimates of the coefficients in the probit using the standard explanatory variables, with a *Female* indicator capturing gender differences in the probability of receiving PP remuneration. As can be observed, women are less likely to get PP than comparable men with the same observable characteristics working in identical occupations. From those estimates, one can compute the corresponding estimated marginal effect of being a woman rather than a man (with the other covariates evaluated at their means) on the probability of participating in the PP sample. Though statistically significant, the estimated marginal effect is fairly small, namely the probability of a female worker receiving PP is only 1.6 pp. lower than that of a male worker. Notice that this small effect is line with our previous finding of rather similar participation rates by men and women in PP jobs when differences in observable characteristics had not been adjusted for. The remaining estimates are in line with the evidence presented in Table 2a: higher educational attainment, longer tenure and belonging to the 31-50 age group are covariates that also increase this probability.

4.3. Adjusted gender PP gaps within firms and occupations

We next estimate gender PP gaps adjusting for gender differences in observed characteristics in the restricted PP sample. Furthermore, as will be explained further below, we estimate these gaps in different setups regarding combinations of firms and ocupations of these workers. As before, we use a *Mincerian* log wage specification with a *Female* intercept and where the remaining returns to individual and job characteristics are assumed to be equal across genders. Our focus lies on comparing the estimated coefficient on the *Female* indicator in a regression (augmented by the inverse Mills ratio obtained from the participation equation reported in the second column of Table 5) under four different specifications: (i) a pooled regression (*P*); (ii) within- occupations (*WO*);[5] (iii) within-firms (*WF*); and (iv) within-firms & occupations (*WFO*).

Table 6 reports the estimates obtained under these alternative specifications. The OLS results (without correction for selection) are also included in the first column for purposes of comparison. The following findings stand out:

- First, the adjusted gender PP gap in the OLS pooled specification is about 41 log points (compared to a raw gap of 46 log points). It is noteworthy that this gap is much larger than the adjusted gap of 19.6 log points estimated by de la Rica (2010) for the fixed wage component of workers in the PP sample.
- Second, once selection bias is controlled for in this specification, the gap increases to 45 log points. The fact that this gap is larger than in the OLS specification is explained by the positive sign of the highly significant coefficient of Heckman´s lambda, which reflects strongly favourable selection of workers receiving PP. Since in our sample women have higher educational attainment than men (despite having lower tenure), this leads to a larger gap when selection is taken into account.
- Third, controlling once more for selection biases, the estimate of the gap in the within-firm specification (34 log points) is considerably smaller than the one in the within-occupation specification (43 log-points) which, in turn, is quite close to the estimated gap in the pooled specification (41 log points).
- Finally, the gap in the joint within-firm and occupation model (29 log-points) is slightly lower than the gaps in the within-firm and within-occupation models.

[Table 6 about here]

4.5 Quantile regressions

Further evidence on the gender PP gap can be obtained from a comparison of its pattern throughout the distribution of this wage component. To that end we use quantile regressions (QR) that account for corrections for selectivity under the within-firm & occupation specification. Following Buchinsky' s (1998) approach, the correction for selectivity for workers who receive PP is based on a two-stage approach. First a two-term series expansion of the inverse of the Mills ratio in Table 5 is used to obtain an estimate of a latent index that approximates the unknown quantile functions of the truncated bivariate distribution for the error terms in the wage and participation

[5] We use the most disaggregated occupational classification available for our dataset: 18 occupational categories

equations. Then, the covariance matrix for the two-stage QR and the selectivity-corrected estimates is obtained by bootstrapping the design matrix with 100 replications.

Table 7 reports the QR estimates of the coefficient on the *Female* indicator for a few relevant percentiles of the PP distribution. A clear "glass ceiling" (i.e., increasing) pattern emerges, with the gap growing from 20 log-points in the bottom deciles to 43 log points at the top of the distribution.

[Table 7 about here]

5. Reconciling the evidence with the potential explanations

The main findings regarding gender PP gaps for our sample of workers in Spain, once individual and job caracteristics and non-random sample selection (labeled as "similar men and women" in what follows) are adjusted for, can be summarized as follows:

1. The female incidence PP schemes is only slightly lower than the male incidence.
2. The adjusted gender PP gap for similar men and women working in the same firm and occupation is around 2/3 of the total raw gap.
3. There is a clear glass-ceiling pattern in the gender PP gap between similar men and women working in the same firm and occupation.

Given this evidence, some of the potential explanations discussed in the Introduction about our findings on the gender PP gap in Spain can be ruled out. In particular:

- The fact that the incidence of women in PP schemes is only slightly lower than men's does not support explanations of the gap based on gender differences in risk aversion since this hypothesis would predict much lower incidence of women in PP jobs. Moreover, it also goes against the hypothesis stating that, if women were to anticipate discrimination in non- PP jobs, they should be more prevalent in PP jobs.

- The fact that the adjusted gender PP gap within occupations is nearly the same as the one across occupations, and that it remains at almost 30 log-points (two-thirds of the total PP gap) when we focus on similar men and women within the same firm and occupation, implies that occupational segregation in its different formats is not consistent with this sizeable differential.

By contrast, notice that our findings about similar male and female incidence rates in PP jobs would be consistent with the hypothesis stating that, under a competitive labour market paradigm, there should be no differences in participation in PP jobs among equally productive men and women. However, this leaves still open the issue of why the PP gender gap is so high.

Having ruled out the previous hypotheses, this leaves two explanations for our findings: one based on supply and the other on demand considerations. Regarding the supply side, women may exert less effort in the workplace due to their heavier burden in housework. This mechanism would lead to a large gender PP gap between similar men and women working in almost identical jobs. As for the demand side, the gender PP gap may be due to discrimination by employers with monopsonistic power who find it optimal to pay women less in terms of PP than equally productive men. Identifying which is the more appropriate among the two explanations is quite a complicated task.

Although the information available in our dataset does not allow us to test for gender gaps in preferences regarding PP, there is a feature that does not seem to support this hypothesis, namely the glass-ceiling pattern observed in these gaps. In principle, these differences might be expected to lead to similar gaps throughout the distribution, unless women's preferences with respect to the pecuniary components of PP are assumed to decrease with qualification and skill levels, a contention for which there is, to our knowledge, no empirical support.

Despite the absence of family-related information in our dataset, we can also provide some strong evidence against the possibility that women may exert less effort in PP jobs because of their greater commitment at home. To do so, it is important to consider that in our sample we only considers full-time workers, that we control for age, education and tenure – all related to productivity- and that the number of overtime hours reported are similar for men and women who receive PP (60.2 and 59.8 per year, respectively). All these features together lead us to think that gender differences in effort do not play an essential role in explaining the gender PP gap. Moreover, an indirect test for gender differences in effort can be implemented by checking whether the proportion of the total hourly wage accounted for by PP is lower for similar men and women in the same jobs and occupations. The insight is simply that greater effort should lead to a higher proportion of PP in the total wage. However, as mentioned above, recall that, on average, these proportions are 9% for men and 7.2% for women, which do not look so different. Yet, before reaching a firm conclusion on this issue, we should once more adjust for observable characteristics. Though not reported here for the sake of brevity, we have run a *Mincerian* regression similar to the one reported in column (5) of Table 6, where now the dependent variable is the *logit* transformation of the percentage that PP represents in total pay.[6] We obtain that the Female indicator explains one-fourth (0.45 pp.) of the 1.8 p.p. gap (=9.0-7.2) in this proportion. Thus, although there is some evidence that differences in effort exerted may play a role in explaining the gap, the contribution estimated is not large enough for this hypothesis to be considered as the key explanation.

Additionally, it is not easy to reconcile the glass-ceiling pattern found for the gender PP gap with this hypothesis since it is hard to provide plausible reasons why these gender differences in effort should be larger among the highest qualified workers. If anything, the opposite pattern (i.e., larger gender gaps for low-qualified workers) should be expected, as the opportunity costs of exerting lower effort on the part of less-skilled female workers are likely to be lower.

Having ruled out the previous rationalizations, the only hypothesis which seems to be consistent with our empirical findings is one involving some sort of discrimination on the employer's side. In particular, assuming that labour mobility is lower among women than among men, employers seem to exploit their monopsonistic power in frictional labour markets by paying less PP to women than to similar men working in the same firm and occupation. Furthermore, recall that the rate of exploitation under monopsony (i.e., the relative difference between marginal revenue and wage) is the inverse of the elasticity of labor supply (Alshenfelter et al. 2010). Since this elasticity is likely to be lower for more highly-skilled workers (see Hirsch et al., 2010),[7] the exercising of monopsonistic power by employers implies that the PP gender gap increases over the course of the PP distribution, which is in line with our findings.

Obviously, the above conclusion is just a conjecture since the lack of family-related information in our dataset prevents us from formally testing whether family conditions affecting women (e.g. being married or having children/elderly dependents in their charge) play a major role in explaining the findings. It is high on our future research agenda to explore whether the 2006 ESS can be merged with other datasets where such information is available.

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[6] The *logit* transformation, $\ln(R/1-R) \in (-\infty, +\infty)$, achieves congruencyy with the support of the distribution of the error tem in the regression, where $R \in (0, 1)$ is the proportion of PP in the total hourly wage. Denoting by b the estimated coefficient in the regression, the effect of the *Female* dummy, D, on R becomes $\delta R / \delta D = bR(1-R)$.

[7] For example, this would be the case if the income effect is strong at higher wages, likely to be related to high skills, and therefore the labor supply schedule becomes either vertical or even backward bending.

6. Concluding remarks

We have used a large cross-sectional wage survey for workers in Spain to examine whether the gender PP gap differs from the gaps in the other components of total wage remuneration. We have found evidence that: (i) PP is linked more with workers´ performance, (ii) there are minor gender differences in selection of workers into PP jobs, and (iii) women who receive PP have several observable characteristics which are better than those of men (e.g. educational attainment). Yet our main finding is that the gender gap in PP is much higher-- both in raw terms and after adjusting for observable characteristics and for segregation into different firms and occupations-- than the gap in non-PP remuneration, and that there are clear signs of a "glass ceiling" effect (wider gaps and lower participation of women in the upper parts of the PP distribution).

We argue that, in principle, these findings taken together cannot be reconciled with some hypotheses, such as occupational segregation, less effort by women in the workplace, the competitive labour market paradigm, and gender differences in attitudes toward competition or in risk aversion. Our preferred explanation relies instead on some sort of monopsonistic discrimination by employers against women, due to their lower job mobility or lack of potential job offers. Furthermore, this rationalization of PP gender gaps would be consistent with their glass-ceiling pattern. This is so since the rate of exploitation under monopsony is the inverse of the elasticity of labor supply and this elasticity tends to be lower for more highly-skilled workers.

Yet, it is important to highlight that this conclusion is just a mere conjecture since the lack of family-related information in our dataset prevents us from formally testing whether specific family conditions affecting women (e.g. being married or having children/elderly dependents in their charge) play a major role in explaining the findings. It is high in our future research agenda to check whether 2006 ESS can be merged with other datasets where such missing information is available.

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List of Tables

Table 1 – Sample characteristics (Full-time workers aged 18-65)

Variables	Women (65,233)	Men (129,930)
	Mean	Mean
Individual Characteristics		
Education		
Primary or less	0.176	0.275
Secondary	0.508	0.545
University	0.316	0.180
Age		
Less than 30	0.257	0.200
31-40	0.354	0.323
41-50	0.245	0.265
>50	0.143	0.212
Tenure (years)	7.410	8.867
Indefinite Contract	0.727	0.768
Wages		
Total Hourly Wage (logs)	2.185	2.391
Performance Pay (only PP workers)		
% PP wokers	0.177	0.194
Total Hourly Wage (logs)	2.508	2.764
Fixed Hourly Wage (logs)	2.430	2.664
PP Hourly Wage (logs)	-0.663	-0.203
Firm Characteristics		
Size		
<50 employees	0.339	0.403
51-200 employees	0.265	0.288
>200 employees	0.396	0.309
Firm Bargaining Agreement.	0.133	0.198
Exporting firms	0.183	0.194

Source: ESS (2006)
Note: The null of equal means across genders is always rejected.

Table 2a: Characteristics of workers and firms by type of job and gender

	PP sample		Non-PP sample	
	Women (14.789 obs.)	Men (29.460 obs.)	Women (50.444 obs.)	Men (100.470 obs.)
Variables	Mean	Mean	Mean	Mean
Education				
Primary or less	0.107	0.178	0.196	0.304
Secondary	0.494	0.545	0.512	0.545
University	0.399	0.277	0.292	0.151
Age				
Less than 30	0.204	0.149	0.273	0.215
31-40	0.313	0.385	0.346	0.325
41-50	0.265	0.294	0.239	0.257
>50	0.150	0.244	0.141	0.203
Tenure (years)	9.281	12.037	6.861	7.938
Indefinite Contract	0.814	0.862	0.741	0.741
Firm Characteristics				
Size				
<50 employees	0.201	0.235	0.380	0.452
51-200 employees	0.239	0.297	0.272	0.285
>200 employees	0.560	0.467	0.348	0.262
Collective Bargaining (ref: Industry level)				
Firm Collective Bargaining	0.193	0.288	0.154	0.167
Firm Market (ref: International Market)				
Local or Nat. Market	0.181	0.239	0.135	0.153

Source: ESS (2006)
Note: The null of equal means across is always rejected.

Table 2b: Incidence of PP per industry and occupation

	Mean	Std. Dev.	No. Obs.
Industries			
Financial Intermediation	0.598	0.49	10475
Energy	0.324	0.468	4627
Transport	0.324	0.468	12710
Health	0.287	0.452	14178
Retail trade	0.241	0.427	17131
Manufacturing	0.205	0.404	74332
Real Estate and Res. Serv.	0.194	0.395	16342
Mining & Extractive Ind.	0.188	0.391	2919
Other Services	0.146	0.353	9040
Construction	0.127	0.333	17096
Hotels and Restaurants	0.123	0.328	8315
Education	0.092	0.289	7998
Occupations			
Managers	0.497	0.5	6190
Technicians	0.326	0.469	30184
Professionals	0.288	0.453	20295
Clerks	0.257	0.437	24761
Personal Services	0.196	0.397	17528
Operators and Assemblers	0.18	0.384	34822
Craftsmen	0.169	0.375	37918
Agriculture and Fisheries	0.146	0.353	542
Laborers, unskilled workers	0.127	0.333	22923

Source: ESS (2006)

Table 2c: Share of women throughout PP distribution

	% Women in percentiles
[P1th-P10th]	32.7%
[P11th-P25th]	32.4%
[P26th-P50th]	31.6%
[P51th-P75th]	25.6%
[P76th-P90th]	19.8%
[P91th-P95th]	14.9%
[P95th-P100th]	13.7%

Source: ESS (2006)

Table 3: Hourly wages in PP and non-PP samples

	PP sample				Non-PP sample	
	Women		Men		Women	Men
	Total Hourly Wage(€)	Ratio PP/Total Wage (%)	Total Hourly Wage(€)	Ratio PP/Total Wage(%)	Total Hourly Wage(€)	Total Hourly Wage(€)
Average	14.503	7.164	19.144	9.012	9.678	11.665
$P10^{th}$	6.060	0.976	7.801	0.932	3.721	4.689
$P25^{th}$	8.577	2.087	10.804	2.563	5.884	7.308
$P50^{th}$	12.479	4.657	16.051	6.073	8.126	9.826
$P75^{th}$	18.800	9.491	23.546	12.751	12.048	14.192
$P90^{th}$	24.842	16.684	33.127	21. 743	17.795	20.162

Source: ESS (2006)

Table 4: Log hourly wage regressions

Dependent Variable: (Log) Total Hourly Wage

	(1) PP sample	(2) Non-PP sample	(3) Pooled sample
PP Indicator			0.208***
			(0.009)
Female	-0.223***	-0.212***	-0.219***
	(0.004)	(0.003)	(0.004)
Age 30-39 (ref.:<30)	0.167***	0.098***	0.095***
	(0.006)	(0.003)	(0.004)
Age 41-49	0.218***	0.116***	0.114***
	(0.007)	(0.004)	(0.004)
Age 50-59	0.235***	0.161***	0.161***
	(0.008)	(0.004)	(0.005)
Age >60	0.262***	0.155***	0.158***
	(0.014)	(0.007)	(0.008)
College	0.277***	0.223***	0.215***
(ref: Primary)	(0.007)	(0.005)	(0.004)
Secondary	0.077***	0.063***	0.060***
	(0.006)	(0.003)	(0.003)
Tenure	0.044***	0.042***	0.043***
	(0.001)	(0.000)	(0.000)
Tenure sq.	-0.001***	-0.001***	-0.001***
	(0.001)	(0.000)	(0.000)
Indefinite Contract	0.282***	0.313***	0.312***
	(0.006)	(0.003)	(0.003)
Firm Size: 50-199	0.067***	0.095***	0.094***
(Ref: <50)	(0.005)	(0.003)	(0.003)
Firm Size: >199	0.118***	0.166***	0.164***
	(0.005)	(0.003)	(0.003)
Firm Agreement	0.011	0.014*	0.013
	(0.012)	(0.008)	(0.008)
Export market	0.027***	0.035***	0.045***
	(0.005)	(0.003)	(0.003)
Interactions with PP			
Female*PP			-0.023
			(0.005)
Age 30-39*PP (ref:<30)			0.059***
			(0.007)
Age 41-49*PP			0.103***
			(0.008)
Age 50-59*PP			0.089***
			(0.010)
Age >60*PP			0.127***
			(0.016)
College*PP			0.100***
(ref: Primary)			(0.007)

Secondary*PP			0.040***
			(0.006)
Tenure*PP			0.011**
			(0.005)
Indefinite Contract *PP			-0.025***
			(0.007)
Firm Size: (ref<50)			-0.027***
50-199*PP			(0.006)
Firm Size: >199*PP			-0.042***
			(0.006)
Firm Agreement*PP			-0.006
			(0.005)
Export. Firm *PP			-0.014***
			(0.006)
No. Obs.	44249	150914	195163
R sq.	0.605	0.511	0.573

Note: s.e´ s in parentheses. Estimations also control for industry, regional dummies and occupational dummies; (*), (**) and (***) denote statistically significant at 10, 5 and 1 percent, respectively.

Table 5: Probit estimation
Dependent Variable: Receiving Performance Pay (1/0)
(Estimated coefficients)

Female	-0.047***
	(0.008)
Age 30-39 (ref:<30)	0.052***
	(0.010)
Age 40-49	0.032***
	(0.011)
Age 50-59	0.015
	(0.013)
Age >60	-0.076***
	(0.023)
University	0.260***
(ref: Primary)	(0.013)
Secondary	0.164***
	(0.009)
Tenure	0.030***
	(0.001)
Tenure square	-0.001***
	(0.000)
Indefinite Contract	0.037***
	(0.010)
Firm Size: 50-199	0.295***
(Ref: <50)	(0.009)
Firm Size: >199	0.485***
	(0.008)
Firm Collective Agreement	0.096**
	(0.009)
Exporting firm	0.122***
	(0.009)
No. Observations	195163
Pseudo R^2	0.111

Note: s.e´s in parentheses. (*), (**) and (***) denote statistically significant at 10, 5 and 1 percent, respectively.

Table 6: Estimates of log. PP hourly wage equation
Dependent Variable: log PP hourly wage component

	(OLS)	(IV)	(WO)	(WF)	(WFO)
Female	-0.407***	-0.453***	-0.432***	-0.361***	-0.298***
	(0.014)	(0.015)	(0.016)	(0.015)	(0.019)
% Fem. rate in Firm			-0.103***		
			(0.033)		
% Fem. rate in Occupation				-0.200***	
				(0.037)	
% Fem. rate in Firm & Occ.					-0.295***
					(0.028)
Inv. Mills Ratio		1.628***	1.693***	1.690***	1.984***
		(0.170)	(0.170)	(0.198)	(0.141)
Personal Characteristics		Yes	Yes	Yes	Yes
Job Characteristics		Yes	Yes	Yes	Yes
No. obs.		44,249	44,249	44,249	44,249
R-sq.		0.160	0.165	0.165	0.167

Note: s.e´s in parentheses. Coefficients in (1) are derived from an OLS regression over the overall sample of workers. Coefficients in (2) are derived from a Heckit estimation, performed to correct for selection into PP jobs. Coefficients in (3) to (5) also control for the femaleness rate within firms, within occupations and within firms and occupations, respectively. Inverse Mills ratios derived from estimates in (2) are included in the last three columns as an additional covariate to correct for selectivity.

Table 7: Adjusted Gender Gaps in PP – Quantile Regressions
(with selection correction and with firm and occupation fixed effects)

Dependent Variable: Log PP Hourly Wage

	(1) **P10th**	(2) **P25th**	(3) **P50th**	(4) **P75th**	(5) **P90th**
Female (WFO)	-0.226***	-0.281***	-0.318***	-0.357***	-0.366***
	(0.034)	(0.023)	(0.022)	(0.026)	(0.027)

Note: s.e´s in parentheses. Estimations also control for the whole set of covariates in Table 6.