The role of non-base compensation in explaining the motherhood wage gap: Evidence from Italy

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Abstract

This paper underlines the importance of accounting for non-base compensation in explaining the motherhood wage gap. We consider two alternative measures of hourly wage using Italian EU-SILC data from 2007 to 2019: the *base-wage* and the *full-wage*. The former refers to the contractual base wage, while the latter includes performance-based bonuses, productivity bonuses, commissions, pay incentives, and other extra payments. We address the endogeneity issues of motherhood and examine the effect of motherhood status across various quantiles of the wage distribution for the two hourly wage measures. Empirical findings provide evidence of a motherhood *base-wage* premium, which becomes non-significant when using the *full-wage* measure, suggesting that non-base compensation is a source of inequality for mothers. These findings are consistent across the wage distribution. Exploring potential heterogeneity across macro-regions and periods, we find no notable regional disparities except minor distinctions for the Southern regions, alongside a decline in the *base-wage* premium over time and the emergence of a *full-wage* penalty in recent years. A comparative analysis with a sample of men reveals that fathers enjoy a premium with both wage measures. Nevertheless, fatherhood is also associated with reduced extra remunerations, yet to a lesser extent than motherhood.

Keywords: Motherhood wage gap, non-base compensation, Italy.

JEL codes: J31, J16, C21.

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1 Introduction

Despite significant progress in terms of gender convergence in recent decades (Goldin 2014), gender inequality in labor market earnings remains a persistent and widespread issue worldwide (Olivetti and Petrongolo 2016, Blau and Kahn 2017). The empirical evidence indicates that a substantial fraction of these gender disparities stems from the motherhood penalty, which refers to the decline in women's earnings due to motherhood (Kleven, Landais and Søgaard 2019). While the literature extensively covers a broad spectrum of mechanisms responsible for the motherhood wage gap, there is a notable absence of evidence regarding the role of non-base compensation in explaining the wage differential between mothers and non-mothers. This void in research is worth emphasizing. We are pioneering in our article by being the first to thoroughly investigate the disparities in non-base compensation between mothers and childless women.

Researchers have identified multiple mechanisms to explain the motherhood wage gap and its varying magnitude across countries (Cukrowska-Torzewska and Matysiak 2020). Because of their family responsibilities, mothers often opt for part-time and flexible working hours and actively look for family-friendly jobs that accommodate their child-rearing needs, even though associated with lower earnings (Matteazzi et al. 2014, Adda et al. 2017, Shure 2019). Mothers are less likely to be employed in greedy jobs that require working during the weekend, holidays, dinner hours, and that tend to be disproportionately rewarded. This can widen disparities in the labor market associated with motherhood (Bertrand et al. 2010, Goldin 2021). Employers discrimination against mothers is another explanation (Correll et al. 2007, Ishizuka 2021). Furthermore, a selection effect might be at play, where women who prioritize less their careers tend to have more children or have them earlier (Korenman and Neumark 1992, Amuedo-Dorantes and Kimmel 2005). Lastly, the social context, including country-specific policies, degree of market competitiveness, prevailing gender norms, and cultural factors that affect labor market conditions for mothers, might also play a role in influencing the motherhood wage gap (Budig et al. 2016, Kleven, Landais, Posch, Steinhauer and Zweimüller 2019, Christafore and Leguizamon 2019, Casarico and Lattanzio 2023).

However, the wage gap between mothers and non-mothers can stem from disparities in extra payments granted to employees on top of their base or contractual wage. While collective agreements usually set the base wage, extra payments, such as performance-based bonuses, productivity bonuses, commissions, pay incentives, are often negotiated individually between the employer and the employee. We argue that non-base compensation adversely affects the motherhood wage gap, as mothers are less likely to obtain these extra payments, or more likely to receive lower amounts. By analyzing the literature, we can identify different mechanisms that rationalize our hypothesis. In particular, due to family responsibilities mothers often reduce their work effort (Anderson et al. 2003, Harkness and Waldfogel 2003), and, in response, employers may grant mothers reduced discretionary extra payments. In addition, the literature provides evidence of gender differences in negotiation attitudes, i.e., women are less likely to negotiate for promotions, better job opportunities, or extra remunerations (Dittrich et al. 2014, Exley et al. 2020, Biasi and Sarsons 2022, Sin et al. 2022), as well as in wage expectations and negotiation strategies (Kiessling et al. 2024), and willingness to compete (Niederle and Vesterlund 2007). Remarkably, these differences can also arise between mothers and non-mothers for the following reasons: (i) mothers are aware of the negative consequences of motherhood on labor outcomes (Luhr 2020); (ii) mothers have different preferences for non-wage job attributes, prioritizing non-monetary benefits, such as flexible work schedules, and trading wages for these job amenities that help them balance their professional and family lives (Felfe 2012); and (iii) mothers are more likely to conform to social norms regarding their expected roles, appropriate behaviors, and work-related activities (Schmidt et al. 2023).

Empirical evidence regarding the importance of employers' discretionary allocation of extra remunerations is scanty due to limited data on variable pay practices and earnings. Existing research has predominantly focused on their effect on the gender wage gap (Casarico and Lattanzio 2024, Hirsch and Lentge 2022). However, to the best of our knowledge, their influence on the motherhood wage gap remains unexplored. Therefore, this paper aims to fill this gap in the literature on the motherhood wage gap.

To this end, we use Italian data from the European Union Statistics on Income and Living Conditions (EU-SILC) covering the period from 2007 to 2019. To assess the role of non-base compensation in explaining the motherhood wage gap, we rely on two distinct hourly wage measures: the *base-wage*, which is derived from monthly earnings and measures the base compensation, and the *full-wage*, calculated from annual earnings and including extra remunerations. Our analysis unfolds in several steps. First, we explore the role of the motherhood status in explaining the two different hourly wage measures, i.e., the *base-wage* and the *full-wage*, and we compare the effects. Second, to address endogeneity issues, we employ Oster (2019)'s approach to compute bounding values for the motherhood effect on wages, accounting for unobservable factors that affect motherhood status and wages. We also employ matching techniques to evaluate the motherhood pay differentials within a sample of mothers and childless women who share similar observable characteristics. Subsequently, we conduct a decomposition analysis to examine the contribution of composition and wage-structure effects to the wage gap between mothers and non-mothers, when allowing returns to observable characteristics to be different for the two groups of women. Furthermore, we investigate the effect of the motherhood status across different quantiles of the wage distribution using recent advancements in quantile regression analysis accounting for the selection into motherhood (Rios-Avila and Maroto 2022, Borgen et al. 2023). We replicate the analysis across different time periods and Italian macro-regions to evaluate whether any regularities or notable differences emerge over the years in a country marked by relevant territorial disparities. To contextualize our findings within the gender wage gap literature, we provide empirical evidence on a sample of men to assess the broader influence of non-base compensation on the parenthood wage gap.

Our results reveal novel insights into the importance of the discretionary allocation of additional payments to mothers and childless women in correctly assessing the effect of motherhood on wages. We find that mothers enjoy a 4.4 percent *base-wage* premium. However, this premium becomes not statistically significant when considering the *full-wage*, highlighting that extra remunerations are an important source of inequality for mothers. Therefore, neglecting these additional payments could lead to a misevaluation of the motherhood wage gap because mothers may be less likely to obtain these additional payments or more likely to receive lower amounts than childless women. Interestingly, this evidence applies also to men, although fathers generally enjoy a wage premium regardless of the wage measure. Overall, our findings suggest a negative role of parenthood for both men and women in terms of extra payments, albeit to a lesser extent for men. Lastly, we find that parenthood plays a minor role in explaining the overall gender wage gap in Italy.

The remainder of the paper is organized as follows. Section 2 introduces the data and the empirical strategy. Results are discussed in Section 3, and Section 4 concludes.

2 Data and estimation strategy

2.1 Data and sample selection

We use Italian EU-SILC cross-sectional data that provide information on poverty, income, social exclusion, and living conditions from a nationally representative sample of households. We pool survey waves from 2007 to 2019.[1] We restrict our sample to include women aged 25-40 (prime years for childbearing)[2] who hold a single job at the time of the interview and has been employed at least one month during the income reference period (IRP, hereafter),

[1]While the EU-SILC four-year rotating panel would allow us to control for time-invariant unobserved characteristics, it enables us to investigate the short-term (i.e., in the first two or three years after childbirth) impact of motherhood on earnings (see, for instance, Cukrowska-Torzewska 2017). However, it is important to note that this is outside the scope of the present study, which focuses on the motherhood wage gap regardless of the year of the last childbirth.

[2]Later in the paper, we provide a robustness analysis of this age-based selection.

i.e., the calendar year preceding the interview.[3] In our empirical analysis, we rely on two different wage measures, one referring to the IRP and the other to the interview. Given that EU-SILC data do not provide information on the characteristics of the job held during the IRP, we restrict our sample to women who have experienced no changes to either their job (change of employer) or their contract (for instance, from part-time to full-time, or vice-versa) with the same employer since the last interview to make sure that changes in wages are not due to job changes.[4] To define the motherhood status, we use a dummy variable equal to one if at least one child is residing in the household at the time of the interview and zero otherwise. The final sample consists of 18,832 observations (8,760 mothers and 10,072 childless women) distributed across Italian macro-regions as follows: 5,012 in the North-West (2,264 mothers and 2,748 childless women), 5,544 in the North-East (2,729 mothers and 2,815 childless women), 4,959 in the Centre (2,252 mothers and 2,707 childless women) and 3,317 in the South and Islands (1,515 mothers and 1,802 childless women).[5]

Respondents report two different earnings measures: the annual employee cash or near cash income earned during the IRP and the gross monthly earnings (for employees) at the time of the interview. While the latter refers to the gross income obtained in a typical month, including overtime and pro-rata the 13^{th} or 14^{th} month payments, the former includes also additional payments such as profit sharing, cash-paid performance-based bonuses, and payments based on productivity, all earned during the IRP. Furthermore, respondents provide information about their work schedule, by indicating the number of months worked during the IRP and their typical weekly working hours at the time of the interview. Using this information, we compute the yearly and monthly worked hours,[6] which we use to derive two different measures of the log-hourly wages (expressed in 2015 euros): log-hourly *full-wage* and log-hourly *base-wage*, respectively.

It is noteworthy that the *base-wage* measure, computed from monthly earnings, largely aligns with contractual provisions outlined in collective agreements, which have extensive coverage in Italy (between 80 and 90 percent, see OECD 2017). On the other hand, the *full-wage* derived from annual earnings includes additional remunerations, which employers may award to employees on a discretionary basis. Comparing the role of the motherhood status on both outcomes is crucial to understanding the motherhood wage gap because differences

[3]The incidence of women employed during the IRP and not at time of the interview is 0.4 percent, whereas the share of women employed at the interview and not in the IRP is 4 percent.

[4]This selection results in the exclusion of approximately 8 percent of observations.

[5]The share of employed women is 0.58 throughout the period considered, albeit notable heterogeneity across macro-regions is evident. Specifically, this share ranges from 0.71 and 0.73 in the North-West and North-East regions, respectively, to 0.32 in the Southern regions. Meanwhile, the value stands at 0.65 in the Centre of Italy.

[6]We consider 4.3 working weeks in a month.

in access to additional payments may exist between mothers and non-mothers.

Table A1 in the Appendix provides descriptive statistics for the total sample and subsamples categorized by macro-region. In Italy and across all macro-regions, on average, the (log) *full-wage* exceeds the (log) *base-wage*, and mothers earn more than childless women regardless the wage measure. Notice that the motherhood wage gap, i.e., the difference in the average log-hourly wage of mothers and childless women, tends to be smaller when using the *full-wage* measure. This is because, when comparing the *full-wage* and the *base-wage* by motherhood status, the increase in wages from the *base-wage* to the *full-wage* measure is greater for childless women compared to mothers. This descriptive evidence suggests that women without children are more likely than mothers to obtain extra remunerations or to receive higher amounts.

The higher wages of mothers can reflect the greater labor market experience of women with children, who are generally older than women without children, although having lower levels of education. In terms of job characteristics, mothers are more likely to be employed in economic sectors such as education, health care, and public administration (denoted as "public sector" in Table A1), where public employment is more prevalent than in other economic sectors, with this evidence being more pronounced moving from Northern to Southern regions. The prevalence of permanent contracts and part-time jobs is higher among mothers compared to childless women across all macro-regions. In addition, the share of women holding managerial positions or employed in high-status occupations (i.e., occupation 1 in Table A1) is larger among childless women (mothers) in Northern and Central (Southern) regions.

2.2 Empirical strategy

Descriptive statistics show a positive raw motherhood pay gap. In other words, on average mothers have higher hourly wages compared to women without children. The econometric analysis aims to estimate whether mothers enjoy a wage premium or undergo a wage penalty, *ceteris paribus*, and to assess whether differences emerge according to the definition of hourly wage adopted. Let Y_{irt} denote the outcome of woman i, in macro-region r, reported for year t. As explained in the previous section, we consider two different measures of the log-hourly wage as outcome variables, i.e., Y_{irt} can correspond to either *base-wage* or *full-wage*. Let M_{irt} be the motherhood status. Our main specification is:

$$Y_{irt} = \beta_0 + \beta_1 M_{irt} + \beta_2 X_{irt} + \theta_t + \gamma_r + u_{irt}, \tag{1}$$

where β_1 is the parameter of interest, measuring the motherhood wage premium ($\beta_1 > 0$) or penalty ($\beta_1 < 0$), and X_{irt} is a set of individual attributes and job-related characteristics listed in Table A1. Since the model is estimated on a pooled dataset of multiple cross-sectional yearly waves, we include period fixed effects θ_t. For the *base-wage* measure, periods correspond to the following groups of years: pre-crisis (2007-2008), crisis (2009-2011), recovery (2012-2014), and pre-COVID-19 (2015-2019). Due to a one-year lag of the IRP with respect to the interview year, the definition of sub-periods slightly differs for the *full-wage* measure, for which the pre-crisis period covers the years 2006-2008, and the pre-COVID-19 period the years 2015-2018. We also add macro-region fixed effects γ_r.

Regressing log-hourly wages on motherhood status will most likely produce biased estimates due to the endogeneity of motherhood decisions with respect to unobservable individual characteristics, such as career aspirations, abilities, and attitudes, as well as unobservable characteristics of the firm where the woman has chosen to work, such as company's gender culture, variable pay practices, and criteria for career progressions. To deal with endogeneity, we rely on two approaches. First, we adopt Oster (2019)'s technique, which establishes upper and lower bounds for the OLS estimates of β_1. Second, we apply Propensity Score Matching (PSM) and Coarsened Exact Matching (CEM) methods to account for self-selection into motherhood.

More specifically, Oster's approach requires comparing changes in β_1 estimates and R-squared (R^2) of the "controlled" specification of Equation (1) with a more parsimonious specification, referred to as "uncontrolled", including only the motherhood status. Let $\tilde{\beta}_1$ and $\dot{\beta}_1$ ($\tilde{R}^2$ and $\dot{R}^2$) denote the motherhood wage effect (R-squared) of the controlled and uncontrolled specification, respectively. The identification of bounding values also hinges on the choice of two unknown parameters: the coefficient of proportionality (δ) and the overall R-squared (R^{max}) of a comprehensive model controlling for all variables, both observed and unobserved, with $\tilde{R}^2 \leq R^{max} \leq 1$. In particular, the coefficient δ measures the relative importance of unobservables compared to observables in estimating β_1. Thus, for $|\delta| = 1$, both observable and unobservable factors have equal importance in influencing β_1 estimates. Conversely, $|\delta| < 1$ ($|\delta| > 1$) means that unobservable factors are less (more) important than observable factors. Furthermore, positive (negative) values of δ indicate that observed and unobserved factors impact the estimate of β_1 in the same (opposite) directions. Following Oster (2019), we compute the bounding values for $\tilde{\beta}_1$ assuming $R^{max} = 1.3 \times \tilde{R}^2$ and either $\delta = 1$ or $\delta = -1$. The estimated coefficient $\tilde{\beta}_1$ of the controlled specification refers to the case $\delta = 0$.

The PSM technique allows the comparison of mothers and childless women who share similar characteristics, differing solely in their motherhood status. This requires estimating the likelihood of being a mother conditional on observables that impact the choice of motherhood (Rosenbaum and Rubin 1983). Hence, the motherhood wage effect corresponds to the Average Treatment effect on the Treated (ATT) of being mother which is defined as

$$ATT = E\left(Y^{m} \mid K, M = 1\right) - E\left(Y^{c} \mid K, M = 0\right), \tag{2}$$

where Y^j is the log-hourly wage of women in group j, with $j = m$ for mothers and $j = c$ for childless women, respectively, and $K = (X, \theta_t, \gamma_r)$ is the full set of controls included in Equation (1) and listed in Table A1. The term $E\left(Y^c|K, M = 0\right)$ in Equation (2) measures the counterfactual outcome. After conditioning on the chosen set of observable characteristics, mean outcomes are conditionally independent on treatment (i.e., being a mother), that is $(Y^m, Y^c) \perp M = 1 \mid K$. Under this assumption, it follows that $(Y^m, Y^c) \perp M = 1 \mid p$, with p denoting the propensity score (i.e., the probability of being a mother, deriving from a probit model) used to create a sample of matched similar women.[7]

We further rely on the CEM method proposed by Iacus et al. (2012) to improve causal inference by reducing imbalance on a set of pre-treatment control variables between mothers and childless women. More precisely, the method consists of coarsening relevant control variables into subgroups and then identifying strata. Mothers and childless women within the same stratum have identical values for all the coarsened variables. The central challenge lies in achieving a balance that refines the definition of the control group as much as possible - with a greater number of variables used to define the strata and a greater number of subcategories within each variable - while still securing a sufficient sample of childless women within the same stratum. The variables we use to define the strata are the following: experience in years, indicators of university degree, being in a couple, occupation, and firm size. This coarsening of variables allows identifying 201 strata and obtaining a matching rate of 87 percent of women in the total sample. Using the matched samples, we implement our OLS analysis on the log-hourly *base-wage* or *full-wage* with motherhood status and other characteristics listed in Table A1. Given that there are not equal numbers of treated and control units within strata, OLS estimators require weighting observations according to the size of their strata. Like the PSM, the CEM procedure does not improve the balance between mothers and childless women across unobserved characteristics. However, compared to PSM, there is evidence showing that CEM may outperform PSM with respect to covariate balance and effect bias (Iacus et al. 2012, King and Nielsen 2019). This advantage has been

[7] We choose the caliper matching method, with a caliper of 0.2 of the standard deviation of the propensity score, and allow for replacement, i.e., a woman in the control group can be matched to more than one woman in the treated group if she is also similar to another treated woman (Rosenbaum and Rubin 1985, Stuart 2010). The standard errors for the PSM estimation are computed using bootstrap based on 100 replications (Lechner 2002, Smith and Todd 2005).

recently mitigated (Ripollone et al. 2019).

It is worth noting that the wage gap between mothers and non-mothers can stem from either differences in observable characteristics or disparities in the return associated with those characteristics between the two groups of women. To investigate this issue, we resort to the Kitagawa-Oaxaca-Blinder (KOB, hereafter) decomposition analysis (Kitagawa 1955, Oaxaca 1973, Blinder 1973), which requires estimating Equation (1) separately for mothers and childless women to assess the group-specific returns to observed characteristics. Let $\bar{Y}^j$ be the mean of the log-hourly wage of women in group j, with $j = m, c$. The KOB decomposition analysis divides the difference $\bar{Y}^m - \bar{Y}^c$ into the sum of two components, one measuring the composition effect and the other associated with the wage-structure effect. More specifically, the composition effect captures the contribution to the earnings gap of groups' differences in the mean of predictors. Thus, it measures the gap that would be present if the returns of the observed characteristics were the same for all women and aligned with those of childless women. On the other hand, the wage-structure effect measures the contribution to the wage gap of groups' differences in the return to characteristics, thus reflecting a motherhood premium (if positive) or penalty (if negative).

Furthermore, we investigate how the impact of the motherhood status changes over the earnings distribution by estimating a Quantile Treatment Effect (QTE) model based on recent econometric advances in quantile analysis accounting for selection into motherhood (Rios-Avila and Maroto 2022, Borgen et al. 2023).[8]

To explore heterogeneity across Italian macro-regions and time-periods, we replicate the analysis separately for the four Italian macro-regions and sub-periods as defined above. Lastly, we enrich the analysis by considering a sample of same-aged Italian men. This allows us to compare the role of non-base compensation in the motherhood and fatherhood wage gaps and to contextualize the results within the existing literature on the effect of parenthood on wages and the overall gender wage gap.

3 Results

3.1 The motherhood wage differential

Table 1 provides the estimates of the motherhood wage effect based on the different methods outlined in the previous section. OLS estimates reveal the existence of a motherhood *base-*

[8]We use estimated propensity scores p to derive a set of individual inverse probability weights that are employed in a weighted least square regression with the Recentered Influence Function (RIF) as dependent variable. We use the same set of controls in both the first and the second stage, and bootstrap standard errors (1000 bootstraps) to account for uncertainty in the estimation of the RIF.

wage premium in Italy, which amounts to 7.6 percent in the uncontrolled specification, and decreases to 4.4 percent once individual and job-related variables are controlled for. When using the *full-wage*, the motherhood pay premium is not statistically significant in the controlled specification.[9] These results suggest that, ceteris paribus, mothers enjoy a wage premium when considering the regular or contractual earnings, i.e., the base compensation defined by collective agreements. However, this premium disappears when we consider the hourly wage measure including extra remunerations, indicating that mothers are less likely than childless women to earn additional payments on the top of their contractual wage or receive lower amounts of these extra payments.

One potential explanation for the premium can be related to the timing of the first childbirth. Italy is among the European countries with the highest average age at the birth of the first child, possibly suggesting that women may opt to delay their first childbirth until they have established their careers.[10] In line with this argumentation, Glauber (2018) and Kwak (2022) assert that the motherhood wage gap has narrowed over time, leading to a motherhood premium, because women delay childbearing (which allows women to accumulate more work experience and establish careers before becoming mothers) and tend to work longer hours.

The results based on Oster (2019) approach is consistent with the OLS-based evidence. When using the *base-wage*, the bounding set is relatively narrow and does not include zero. Additionally, the coefficient of proportionality is 1.6 suggesting that unobservable factors would need to have roughly twice the influence of observables to make the motherhood premium negligible. This implies that omitted variable bias is a minor issue. Differently, the bounding set of the motherhood pay premium includes zero when using the *full-wage* measure.

The analysis based on the PSM and CEM is also consistent with the OLS results, possibly suggesting that selection bias related to motherhood is only a minor issue.[11] As previously

[9] Table A2 in the Appendix reports the complete set of results. As a sensitivity check, we estimated a Heckman (1979) model to account for self-selection into employment. As exclusion restrictions, explaining the working decision but not the log-hourly wage, we use the household non-labor income and the partner's labor income. The results, reported in Table A2, are overall consistent with OLS estimates.

[10] In our data we observe that women who became mothers before the age of 25 have lower wages compared to both childless women (who are approximately the same age) and those who delayed motherhood until after 30, or even more significantly, until after 35 years old. This highlights the correlation between the timing of motherhood and wage levels, with earlier motherhood associated with lower wages.

[11] To assess our success in matching, we calculated and compared the standardized bias of the propensity scores for our overall and matched samples (Rosenbaum and Rubin 1985). We find a substantial reduction in percentage bias due to the matching process. Comparing the pseudo R-squared of the unmatched and the matched sample, we observe a significant reduction after the matching procedure, that effectively produced a sample in which, in terms of our explanatory variables, much of the motherhood decision remains random (Sianesi 2004). The results are available from the authors upon request.

argued, the timing of childbirth may be the driver of the premium. In other words, it is not a matter of being or not a mother, but when a woman chooses to become a mother.

Differently from the OLS analysis, the KOB decomposition analysis enables us to assess the contributions of differences in observable characteristics and variations in the returns to these characteristics between mothers and childless women in explaining the motherhood wage gap. Therefore, it is valuable to compare the magnitude of the OLS estimate of β_1 with the size of the wage-structure effect, which considers the heterogeneous impact of all explanatory variables on wages between the two groups of women. The estimate of the wage-structure effect is positive and significant when using the *base-wage*, indicating a pay premium for mothers, and only slightly smaller than $\tilde{\beta}_1$. Regarding the *full-wage* measure, the wage-structure effect is negligible, indicating the absence of a premium for mothers when extra remunerations are considered.

The detailed analysis of the composition effect[12] reveals that the motherhood gap is mainly explained by the longer labor market experience of mothers compared to childless women. Indeed, mothers are, on average, older than women without children. Additionally, job-related variables play a role in explaining the gap, as mothers are more likely to work in sectors like education, health, and public administration, to have a permanent job contract, and to be employed in large firms that pay higher wages. Interestingly, other individual characteristics, including Italian nationality and living in a couple, contribute to explain the gap. The contribution of Italian nationality is negative due to a positive return of being Italian on childless women's wages and a lower share of Italian women among mothers than among childless women. Differently, living in a couple, which means either being married or cohabiting, has a positive return on childless women's wages.[13] Given that the share of married/cohabiting women is higher among mothers than among childless women, the composition effect associated with this variable is positive. Education tends to reduce the gap, as childless women are, on average, more educated than mothers, and higher education is associated with higher wages.

Lastly, Figure 1 shows the estimates of the motherhood effect along the wage distribution. When using the *base-wage* measure (left panel), the pay premium is only slightly increasing along the wage distribution (over the period 2006-2018). We find no evidence of a premium nor a penalty for mothers along the *full-wage* distribution (over the period 2007-2019).

[12] The composition effect is given by the product of childless women's returns to characteristics and the difference in the average characteristics of mothers and non-mothers. To compare the results with OLS estimates, we assume that the non-discriminatory wage structure is the one estimated for childless women (the reference group).

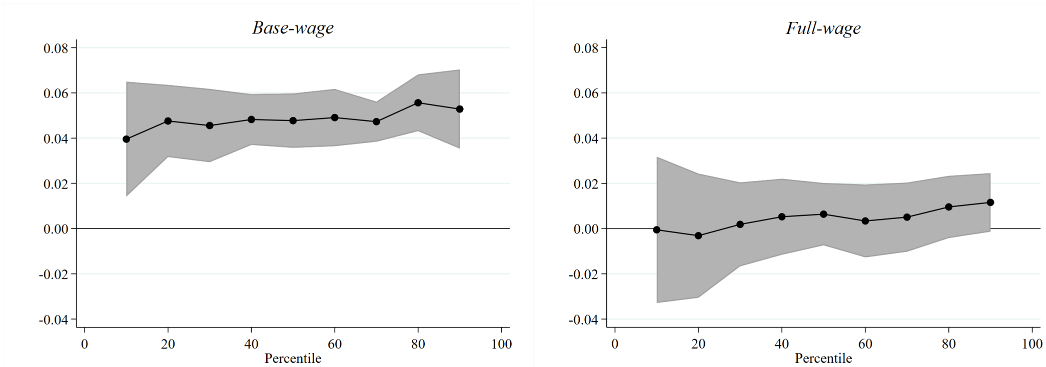
[13] Budig and Lim (2016) find a marriage or cohabiting premium also among women.

Table 1: Results by model specification

	base-wage	*full-wage*
OLS model		
Uncontrolled specification $\dot{\beta}_1$	0.076***	0.035***
	(0.004)	(0.005)
Controlled specification $\tilde{\beta}_1$	0.044***	0.000
	(0.005)	(0.005)
Oster (2019)'s analysis		
Bounding values	[0.024; 0.057]	[-0.019;0.015]
Coefficient of proportionality $(\delta \mid \tilde{\beta}_1 = 0)$	1.788	0.026
PSM		
ATT	0.048***	0.007
	(0.007)	(0.008)
CEM		
ATT	0.037***	-0.002
	(0.005)	(0.006)
KOB decomposition analysis		
Mean of log-hourly wage of mothers	2.368***	2.422***
	(0.003)	(0.003)
Mean of log-hourly wage of childless women	2.292***	2.387***
	(0.003)	(0.003)
Motherhood wage gap	0.076***	0.035***
	(0.004)	(0.005)
Composition effect	0.044***	0.044***
	(0.004)	(0.005)
Education	-0.007***	-0.007***
	(0.001)	(0.001)
Experience[a]	0.017***	0.025***
	(0.002)	(0.002)
Job-related characteristics[b]	0.019***	0.017***
	-0.002	-0.002
Other individual characteristics[c]	0.016***	0.009**
	-0.003	(0.002)
Sub-period and macro-region fixed effects	-0.001	0.001
	(0.001)	(0.001)
Wage-structure effect	0.032***	-0.009
	(0.005)	(0.006)
N	18,832	18,832

Note: (i) In the OLS model, the uncontrolled specification includes only the dummy for motherhood status, whereas the controlled specification adds the set of individual and job-related characteristics listed in Table A1, as well as sub-period and macro-region fixed effects. (ii) PSM estimates include the full set of controls of the controlled specification. (iii) The coarsening variables in the CEM estimates include the experience in years and indicators of university degree, being in a couple, occupation, and firm size. The following weighted OLS estimations of the effect of motherhood on the log-hourly wage include the full set of controls. (iv) In the KOB analysis, groups of variables are the following: *a* second-order polynomial degree of real labor market experience; *b* dummy variables for occupation, public sector, managerial position, part-time, firm size, and permanent job contract; *c* dummy variables for Italian nationality and living in a couple. (v) Standard errors are in brackets. (vi) Significance levels: $^{**}p < 0.05$, $^{***}p < 0.01$.

Figure 1: QTE estimates by wage measure



Note: (i) The control variables are the set of individual and job-related characteristics listed in Table A1, as well as sub-period and macro-region fixed effects, specified as dummy variables. (ii) The figure shows confidence intervals at 95% level.

Robustness analysis. Our study sample includes women aged 25-40, who were employed in the same job and with the same employer during the IRP and at the interview. These sample restrictions are due to EU-SILC survey design.

More specifically, EU-SILC data provide information on the total number of children residing with the mother at the time of the interview rather than the total number of children a woman has given birth to. Including women older than 40 poses the risk of misclassifying some as childless, particularly those who had children at younger ages and whose children have already left the parental home. This age-based sample selection is supported by the data, which reveal a substantial decline in the share of mothers among women aged 42 and older. As a robustness check, we enlarge the sample to include women aged 20-24 (the upper secondary education lasting five years in Italy, between the ages of 14 and 19) and women aged 41. We estimate our OLS model on three different samples of women aged 20-40, 20-41, and 25-41, respectively. In line with the baseline estimates, the results indicate a *base-wage* premium of 0.045 in all samples, and a negligible (not statistically significant) *full-wage* premium (Table A3 - Panel 1).

Furthermore, we relaxed the selection criteria that previously enabled us to consider only women who held the same job during the IRP and at the time of the interview. This narrows down the analysis to focus solely on the *base-wage* measure, for which the entire set of information (earnings, working hours, and job-related variables) is reported at the interview. The results reveal a *base-wage* premium of 0.052 (Table A3 - Panel 2).

Lastly, we assess the robustness of our results considering alternative wage measures. The *full-wage* is obtained by dividing the annual income earned during the IRP by the typical working hours reported at the time of the interview. EU-SILC data do not report information on the hours worked during the IRP. To address this data limitation, we rely on monthly earnings expressed in full-time equivalent, which we obtain using the (in-sample) year-specific average hours worked by full-timers and part-timers, instead of relying on self-reported working hours. The results reveal a motherhood premium of 0.050 when using the full-time equivalent monthly *base-wage*, whereas no significant effect appears when using the full-time equivalent monthly *full-wage* (Table A3 - Panel 3). Finally, considering the potential for mothers to work fewer hours than what is reported for a 'typical' week due to childcare responsibilities, we estimate an OLS model with the monthly *base-wage* reported at the interview as the dependent variable. Our findings reveal that mothers experience a premium of 0.035 (Table A3 - Panel 4). Overall, the robustness analyses demonstrate that our baseline estimates of the motherhood wage effect are robust to sample and wage measure definitions.

3.2 Heterogeneity at macro-region and sub-period levels

Macro-regions analysis. Table 2 provides the estimates of the motherhood wage effect for the uncontrolled and controlled specifications by Italian macro-region (Panels A-D).[14] When considering the *base-wage*, we observe a motherhood wage premium in all macro-regions ranging from 2.7 percent in the South and Islands to 4.9 percent in the North-Western regions (controlled specification). The analysis based on the *full-wage* reveals neither a premium nor a penalty in all macro-regions. Interestingly, while Italy does exhibit significant regional differences in various aspects, such differences are not as prominent when it comes to the motherhood earnings gap. Our data suggest that the wage gap between mothers and non-mothers and the motherhood pay premium are fairly consistent across Italian macro-regions. This evidence may be explained by the characteristics of the Italian wage bargaining system, which sets wages based on nationwide sectoral contracts with limited local wage adjustments, thus ensuring wage equalization across Italian macro-regions (Boeri et al. 2021).

The analysis based on Oster (2019)'s approach demonstrates that unobservable factors should play a very relevant role to nullify the *base-wage* premium observed in the Northern and Central regions. Specifically, for these regions the coefficient δ is almost 2 (Table 2, column 6), and the bounding set is notably narrow, with 0 not falling within it. These results reveal the robustness of our estimates of the motherhood wage premium to omitted variable bias for these macro-regions. Somewhat different considerations apply to the Southern regions, where the bounding set includes 0 and to cancel out the premium the unobservable variables should play almost the same role as the observables. This evidence aligns with

[14] The full set of OLS results is reported in Table S1 in the *Supporting Information* file. Table S2 in the same file shows the Heckman model estimates by macro-region.

the ATT estimates based on PSM and CEM presented in Table 3.[15] Indeed, obtained results show a wage premium for mothers in the Northern and Central macro-regions when considering the *base-wage*.

However, the motherhood premium disappears when we consider the *full-wage*. These results suggest that, compared with mothers, childless women are likely to benefit more from additional payments, eliminating the estimated motherhood premium associated with the base remuneration. Thus, neglecting extra payments leads to a misevaluation of the motherhood wage differential.

The KOB decomposition analysis (Table 4) shows that the motherhood wage gap is significant in all macro-regions for both wage measures, except for the North-East macro-region when using the *full-wage*. Regarding the *base-wage*, we find a positive and significant wage-structure effect, implying the presence of a wage premium for mothers in all macro-regions except the South.[16] When considering the *full-wage* measure, the composition effect almost entirely explains the gap, leaving the wage-structure effect non-significant in all macro-regions. Hence, KOB results aligns with the PSM and CEM estimates, as well as the analysis based on the Oster (2019)'s approach.

All results discussed up to this point pertain to average effects. Differently, Figures 2A and 2B illustrate the QTE estimates for *base-wage* and *full-wage*, respectively. QTE estimates based on the *base-wage* show a significant wage premium for mothers in the North-West macro-region, which remains rather constant across all percentiles. In the North-East, the motherhood premium is positive but not significant in the lower and upper tail of the distribution. In the other segments of the distribution it is rather constant. The QTE estimates exhibit an increasing pattern in the Centre of Italy, with mothers at the top of the wage distribution enjoying a premium of approximately 10 percent. This evidence is consistent with studies highlighting that the motherhood wage premium observed on average is a high-quantile phenomenon (Glauber 2018, Kwak 2022). Differently, in the Southern macro-region, the pattern of the effect of motherhood appears U-shaped. However, the premium is positive and significant only in the bottom part of the distribution, characterized by women predominantly employed in economic sectors where public employment is not prevalent. When looking at the *full-wage*, the effect of motherhood status remains non-significant throughout the wage distribution in all macro-regions. By comparing the results

[15] Regarding the balancing property, we find a substantial reduction in percentage bias across all macro-regions due to the matching process. We also observe a significant reduction in pseudo R-squared after matching. This makes us confident that the matching procedure effectively produced a sample in which much of the motherhood decision remains random.

[16] The detailed decomposition shows a similar pattern across macro-regions similar to that for Italy. Results are available from the authors upon request.

Table 2: OLS estimates, bounding values and the role of unobservables by macro-region and sub-period

	Uncontrolled specification		Controlled specification		Bounding values	Coefficient of proportionality (δ)
	$\dot{\beta}_1$	$\dot{R}^2$	$\tilde{\beta}_1$	$\tilde{R}^2$	$\tilde{\beta}_1$	$\delta \vert \tilde{\beta}_1 = 0$
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A - North-West						
base-wage	0.079***	0.020	0.049***	0.212	[0.032;0.062]	2.209
	(0.008)		(0.008)			
full-wage	0.028***	0.002	-0.004	0.227	[-0.020;0.008]	-0.310
	(0.009)		(0.010)			
Panel B - North-East						
base-wage	0.070***	0.017	0.043***	0.234	[0.027; 0.055]	1.981
	(0.007)		(0.008)			
full-wage	0.009	0.000	-0.007	0.247	[-0.016;-0.000]	-1.005
	(0.008)		(0.009)			
Panel C - Centre						
base-wage	0.070***	0.015	0.047***	0.230	[0.033;0.056]	2.284
	(0.008)		(0.009)			
full-wage	0.049***	0.006	0.012	0.253	[-0.008;0.027]	0.642
	(0.009)		(0.010)			
Panel D - South and Islands						
base-wage	0.085***	0.016	0.027**	0.249	[-0.020;0.055]	0.651
	(0.011)		(0.014)			
full-wage	0.064***	0.007	-0.002	0.239	[-0.053;0.029]	-0.045
	(0.013)		(0.016)			
Panel E - Pre-crisis						
base-wage	0.093***	0.026	0.043***	0.289	[0.010;0.065]	1.211
	(0.010)		(0.011)			
full-wage	0.055***	0.007	0.024**	0.279	[0.005;0.037]	1.182
	(0.009)		(0.010)			
Panel F - Crisis						
base-wage	0.083***	0.021	0.060***	0.274	[0.045; 0.070]	2.443
	(0.008)		(0.009)			
full-wage	0.019	0.001	0.004	0.275	[-0.005;0.010]	0.444
	(0.010)		(0.010)			
Panel G - Recovery						
base-wage	0.069***	0.015	0.038***	0.278	[0.020;0.051]	1.718
	(0.009)		(0.009)			
full-wage	0.041***	0.004	0.007	0.257	[-0.013;0.021]	0.387
	(0.010)		(0.011)			
Panel H - Pre-COVID-19						
base-wage	0.070***	0.014	0.036***	0.201	[0.016;0.050]	1.574
	(0.007)		(0.008)			
full-wage	0.019**	0.001	-0.030***	0.219	[-0.053;0.029]	-1.665
	(0.009)		(0.010)			

Note: (i) The uncontrolled specification (columns 1-2) includes only the dummy for motherhood status. (ii) The controlled specification (columns 3-4) adds the set of individual and job-related variables listed in Table A1 as well as sub-period (Panels A-D) or macro-region (Panels E-H) fixed effects. (iii) Bounding values (column 5) for the OLS estimates (column 3) are computed assuming $\delta = +/-1$ and $R^{max} = 1.3 \times \tilde{R}^2$. (iv) Standard errors are in brackets. (v) Significance levels: $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 3: PSM and CEM estimates by macro-region and sub-period

	base-wage		*full-wage*	
	PSM	CEM	PSM	CEM
Macro-region				
North-West	0.063***	0.051***	0.010	0.003
	(0.011)	(0.009)	(0.014)	(0.011)
North-East	0.069***	0.037***	0.013	-0.010
	(0.012)	(0.009)	(0.013)	(0.011)
Centre	0.054***	0.037***	0.018	0.010
	(0.015)	(0.010)	(0.016)	(0.012)
South and Islands	0.036	0.002	-0.006	-0.034
	(0.024)	(0.019)	(0.026)	(0.020)
Sub-period				
Pre-crisis	0.071***	0.039***	0.044**	0.028**
	(0.019)	(0.014)	(0.018)	(0.012)
Crisis	0.057***	0.044***	0.003	-0.001
	(0.013)	(0.010)	(0.017)	(0.012)
Recovery	0.034**	0.032***	-0.015	0.012
	(0.013)	(0.010)	(0.017)	(0.013)
Pre-COVID-19	0.042***	0.030***	-0.018	-0.038***
	(0.013)	(0.010)	(0.016)	(0.012)

Note: (i) PSM estimates the set of individual and job-related characteristics listed in Table A1, as well as macro-region and sub-period fixed effects. (ii) The coarsening variables in the CEM estimates include the experience in years and indicators of university degree, being in a couple, occupation, and firm size. The following weighted OLS estimations of the effect of motherhood on the log-hourly wage include the set of individual and job-related characteristics listed in Table A1, as well as macro-region and sub-period fixed effects. (iii) Bootstrapped standard errors are in brackets in PSM. (iv) Robust standard errors are in brackets in CEM. (v) Significance levels: $^{**}p < 0.05$, $^{***}p < 0.01$.

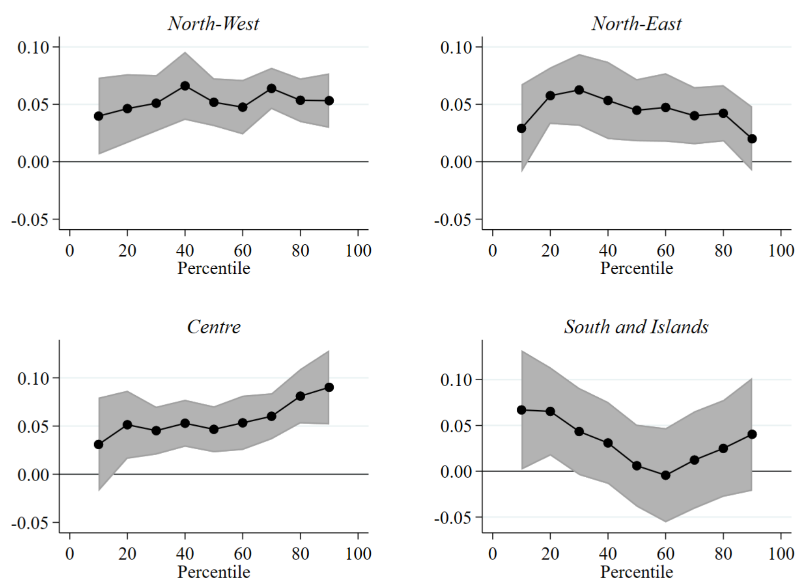
Table 4: Detailed KOB decomposition results

	base-wage	*full-wage*	*base-wage*	*full-wage*	*base-wage*	*full-wage*	*base-wage*	*full-wage*
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Macro-region	North-West		North-East		Centre		South and Islands	
Mean log-hourly wage of mothers	2.399***	2.455***	2.395***	2.446***	2.361***	2.409***	2.283***	2.351***
	(0.006)	(0.007)	(0.005)	(0.006)	(0.006)	(0.007)	(0.009)	(0.010)
Mean log-hourly wage of childless women	2.320***	2.428***	2.325***	2.437***	2.291***	2.360***	2.199***	2.287***
	(0.005)	(0.006)	(0.005)	(0.006)	(0.005)	(0.006)	(0.008)	(0.009)
Motherhood wage gap	0.079***	0.028***	0.070***	0.009	0.070***	0.049***	0.085***	0.064***
	(0.008)	(0.009)	(0.007)	(0.008)	(0.008)	(0.009)	(0.011)	(0.013)
Composition effect	0.034***	0.028***	0.040***	0.026***	0.044***	0.054***	0.082***	0.095***
	(0.007)	(0.008)	(0.008)	(0.009)	(0.008)	(0.009)	(0.014)	(0.017)
Wage-structure effect	0.045***	0.000	0.030***	-0.017	0.026***	-0.006	0.003	-0.031
	(0.009)	(0.010)	(0.009)	(0.011)	(0.010)	(0.011)	(0.017)	(0.019)
Sub-period	Pre-crisis		Crisis		Recovery		Pre-COVID-19	
Mean log-hourly wage of mothers	2.371***	2.467***	2.364***	2.440***	2.341***	2.418***	2.387***	2.365***
	(0.007)	(0.007)	(0.006)	(0.007)	(0.006)	(0.007)	(0.005)	(0.007)
Mean log-hourly wage of childless women	2.278***	2.412***	2.282***	2.421***	2.272***	2.377***	2.317***	2.347***
	(0.006)	(0.006)	(0.006)	(0.007)	(0.006)	(0.008)	(0.005)	(0.006)
Motherhood wage gap	0.093***	0.055***	0.083***	0.019	0.069***	0.041***	0.070***	0.019**
	(0.010)	(0.009)	(0.008)	(0.010)	(0.009)	(0.010)	(0.007)	(0.009)
Composition effect	0.063***	0.034***	0.033***	0.021**	0.043***	0.054***	0.046***	0.061***
	(0.010)	(0.009)	(0.008)	(0.010)	(0.008)	(0.011)	(0.007)	(0.009)
Wage-structure effect	0.030**	0.021	0.050***	-0.002	0.026***	-0.013	0.024***	-0.043***
	(0.013)	(0.011)	(0.010)	(0.011)	(0.010)	(0.012)	(0.009)	(0.011)

Note: (i) The control variables are the set of individual and job-related characteristics listed in Table A1. (ii) In the analysis by macro-region (sub-period), sub-period (macro-region) fixed effects are controlled for. (iii) Standard errors are in brackets. (iv) Significance levels: $^{**}p < 0.05$, $^{***}p < 0.01$.

of the quantile analysis for the two measures of hourly wages, it appears that in the Central regions of Italy, mothers in the upper part of the distribution are considerably less likely to receive extra remunerations, or they receive lower amounts of these additional payments. Indeed, while the pattern based on the *base-wage* is increasing, the one related to the *full-wage* remains almost flat. Similar considerations also apply to mothers in the lower segment of the labor market in the Southern regions.[17]

Figure 2A: QTE estimates by macro-region, *base-wage*

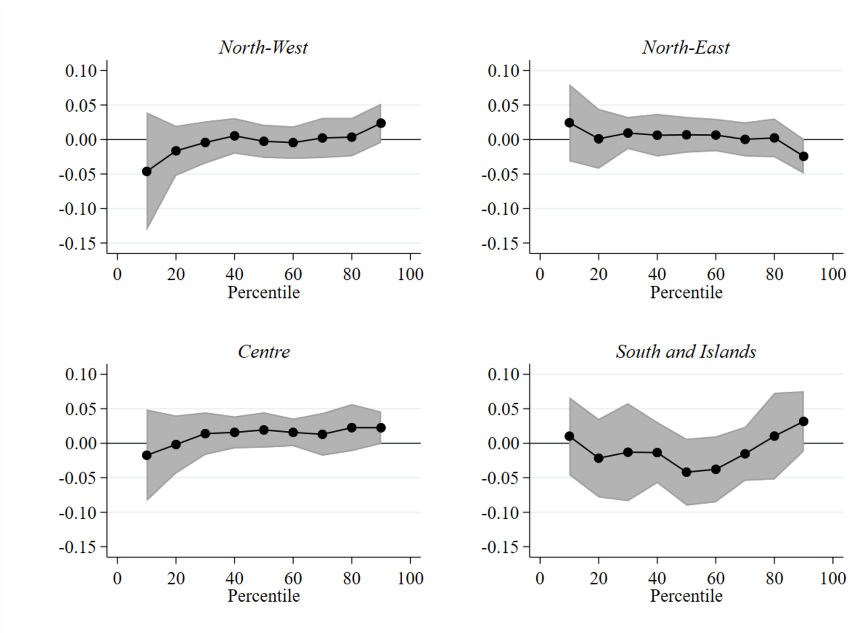


Sub-periods analysis. To investigate whether the motherhood wage differential has changed over time, we replicate the entire analysis by sub-period. Table 2 shows the estimates of the motherhood wage effect for the uncontrolled and controlled specification by sub-period (Panels E-H).[18] The OLS results show a *base-wage* premium in all sub-periods, slightly smaller in the recovery and the pre-COVID-19 periods. We find evidence of a *full-wage* premium

[17]To check this, we computed the difference in the levels of the *full-* and the *base-wage*. This difference, which serves as an approximation for the hourly extra remunerations, is used as dependent variable of a QTE model. Figure A1 in the Appendix shows the results. As suspected, in the Central regions, the effect of motherhood on hourly extra remunerations is somewhat stronger in the upper part of the distribution. Conversely, it is rather constant along the distribution in the Northern regions, except for the upper and lower tails of the North-East, where the effect of motherhood on hourly extra remunerations is higher. In the Southern regions, mothers in the lowest segment of the labor market have to cope with lower extra remunerations.

[18]The full set of OLS results are reported in Table S3 in the *Supporting Information* file. Table S4 in the same file shows the Heckman model estimates by sub-period.

Figure 2B: QTE estimates by macro-region, *full-wage*



Note: (i) The control variables are the set of individual and job-related characteristics listed in Table A1, as well as sub-period fixed effects, specified as dummy variables. (ii) The figure shows confidence intervals at 95% level.

in the pre-crisis period and a *full-wage* penalty in the pre-COVID-19 period. Oster, PSM, and CEM estimates are overall consistent with the OLS findings (Tables 2 and 3). The results of the decomposition analysis (Table 4) suggest that the motherhood wage gap has slightly decreased over time for both wage measures. The size of the wage-structure effect has also diminished across the years, revealing a decreasing motherhood *base-wage* premium and the emergence of a motherhood *full-wage* penalty in the pre-COVID-19 period. Indeed, we observe that the *full-wage* is reducing over time for both mothers and childless women, though to a larger extent for the former group of women.[19]

The quantile analysis on the *base-wage* (Figure 3A) indicates a decline in the motherhood premium over time, as shown by the downward movement of the quantile curve across the various periods. Interestingly, this reduction is more pronounced in the bottom part of the distribution. Conversely, the motherhood premium has exhibited a relatively stable pattern over time in the upper part. Regarding the quantile analysis on the *full-wage* measure (Figure 3B), we observe a downward movement of the quantile curve over time, though the

[19]Table S5 in the *Supporting Information* file provides the results of the KOB analysis by macro-region and sub-period. Results indicate a positive and significant, albeit decreasing, motherhood *base-wage* premium in the Northern macro-regions, where the premium is not significant in the more recent sub-period. The analysis shows the onset of a *full-wage* penalty in the North-West, North-East, and Central regions in the pre-COVID-19 period.

premiums or penalties are almost negligible along the entire distribution.

Figure 3A: QTE estimates by sub-period, *base-wage*

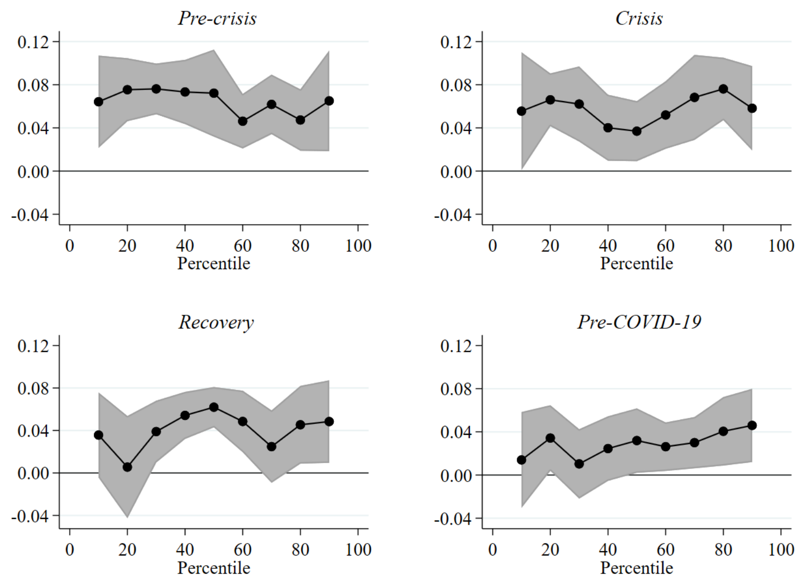
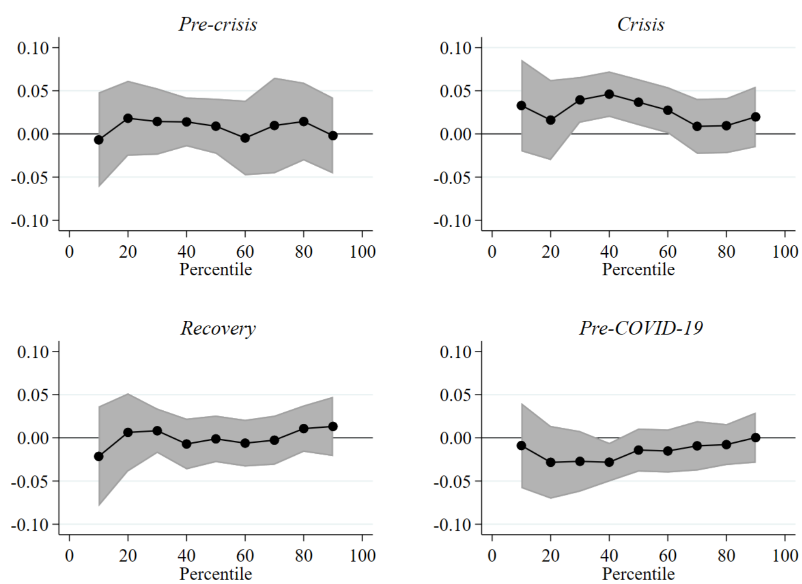


Figure 3B: QTE estimates by sub-period, *full-wage*



Note: (i) The control variables are the set of individual and job-related characteristics listed in Table A1, as well as macro-region fixed effects, specified as dummy variables. (ii) The figure shows confidence intervals at 95% level.

3.3 Fatherhood and the gender wage gap

In this section, we extend the analysis to a sample of men. Our objective is twofold. Firstly, we assess whether the findings regarding the impact of parenthood on earnings, derived from the study of women, apply to men. Secondly, we aim to offer insights into the contribution of both motherhood and fatherhood pay gaps to the overall gender wage gap in Italy. Given that parenthood generally appears to have a positive impact on men's wages while negatively affecting women's wages, it may widen the disparity between men's and women's average wages, thereby exacerbating the gender wage gap (Angelov et al. 2016, Cukrowska-Torzewska and Lovasz 2020).

From a theoretical standpoint, the positive effect of fatherhood on men's wages, the so-called fatherhood wage premium, can be attributed to several channels. Specifically, fatherhood is associated with greater job tenure, which may account for the premium (Millimet 2000). Moreover, men tend to move to higher-paying jobs or increase their working hours upon the arrival of a child (Pollmann-Schult 2011). Within-household specialization further contributes to explaining the fatherhood premium (Lundberg and Rose 2000, Glauber 2008, Killewald and Gough 2013, Angelov et al. 2016). Accordingly, fathers often view their family responsibilities as motivating factors for increasing their work effort (Eggebeen and Knoester 2001, Kmec 2011). In turn, employers may reward more fathers because they are expected to be more productive than childless men (Correll et al. 2007).

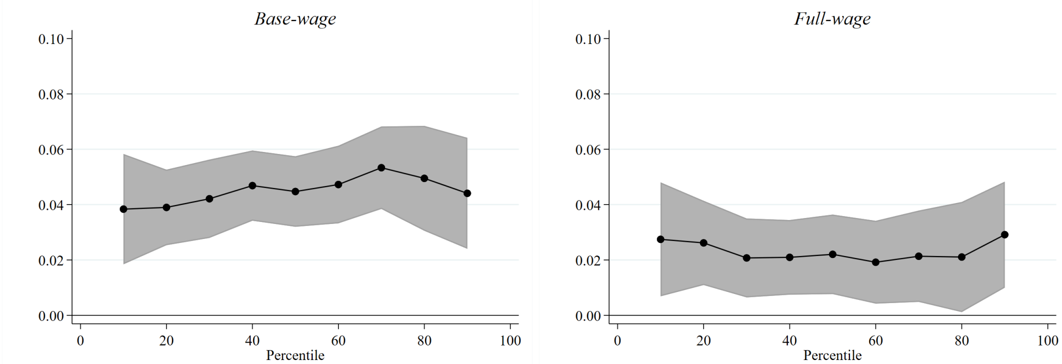
OLS estimates, based on a sample of men aged 25-40 in full-time jobs,[20] reveal the existence of a fatherhood *base-wage* premium of 0.043 and a *full-wage* premium of 0.021 (Table A4 in the Appendix).[21] Hence, differently from mothers, a premium holds for fathers, even when we consider the *full-wage* measure, which is in line with the existing related literature.[22] However, this fatherhood premium is smaller in magnitude, suggesting that fathers have lower access to extra payments than childless men. This leads us to a generalizable finding applicable to both mothers and fathers. Furthermore, the quantile analysis shows that the fatherhood *base-wage* and *full-wage* premiums are relatively constant along the distribution (Figure 4).

[20] We define the fatherhood status with a dummy variable equal to one if at least one child is residing in the household at the time of the interview and zero otherwise. Given the low proportion of men in part-time employment (3.4%), we have reduced our sample to men in full-time jobs. As for women, we select men with no job changes between the IRP and the time of the interview. The final sample consists of 22,182 observations (8,291 fathers and 13,891 childless men). Table S6 in the *Supporting Information* file shows descriptive statistics for the sample of men.

[21] Our data show a substantial decline in the share of fathers among men aged 45 and older. As a robustness check, we estimate our baseline model on a sample of men aged 25-44. The results indicate a significant fatherhood *base-wage* premium of 0.046 and a *full-wage* premium of 0.021.

[22] In Italy, Cukrowska-Torzewska and Lovasz (2020) find evidence of a fatherhood (*full-wage*) premium of 0.038 while a motherhood (*full-wage*) penalty of -0.040 (not significant).

Figure 4: QTE estimates by wage measures - *Men*



Note: (i) The control variables are the set of individual and job-related characteristics listed in Table S6 in the *Supporting Information* file, as well as sub-period and macro-region fixed effects, specified as dummy variables. (ii) The figure shows confidence intervals at 95% level.

As mentioned above, parenthood may contribute to widening the overall gender wage gap. Indeed, the gender wage gap can be decomposed into a gender wage gap among childless individuals, augmented by a fatherhood wage gap, subtracted by the motherhood wage gap, where the latter two gaps are weighted by the share of fathers and mothers, respectively, among men and women (Cukrowska-Torzewska and Lovasz 2020). Our data indicate that the gender *base-wage* and *full-wage* gaps in Italy are 0.061 and 0.118, respectively. The *base-wage* (*full-wage*) gap is decomposed as follows: a childless gender wage gap of 0.064 (0.111), a fatherhood wage gap of 0.086 (0.061), and a motherhood wage gap of 0.076 (0.035). However, given the low shares of fathers and mothers in Italy (0.37 and 0.47, respectively), the contribution of parenthood to the overall gender wage gap is marginal. In Italy, the fatherhood premium contributes to widening the gender wage gap, while the motherhood premium helps reduce it. However, the two effects offset each other, indicating that parenthood is not the main driver of the gender wage gap in Italy. It is rather a gender issue.

4 Conclusions

In this paper, we explore how the non-base compensation contributes to understanding the motherhood wage gap in Italy. We use Italian EU-SILC data from 2007 to 2019, and employ two different measures of hourly wage: the *base-wage* derived from monthly earnings and referring to the base compensation, and the *full-wage* calculated from annual earnings and inclusive of additional remunerations, such as bonuses, performance-based premia, and commissions. The obtained results show that these extra remunerations play a crucial role in explaining the motherhood wage gap, emphasizing them as a factor contributing to inequality between mothers and childless women. Failure to account for them could result in an inaccurate assessment of the motherhood wage gap, thereby impeding the implementation of effective policies.

When considering the base compensation, mothers enjoy a wage premium, which becomes not statistically significant when extra remunerations are considered. This evidence, robust to selection and omitted variable bias and rather constant across quantiles of the wage distribution and macro-regions, reflects the likelihood of different allocation of these additional remunerations between mothers and non-mothers. These additional remunerations are not generally set by collective agreements and are often negotiated individually between employer and employee.

Two potential mechanisms may contribute to explaining this result. Firstly, employers might allocate extra remunerations based on the perceived employee's effort and commitment in a discretionary manner that can penalize mothers. Secondly, a possible lack of negotiation attitudes among mothers, due to deeply ingrained societal gender norms and expectations that perceive mothers as less competitive and profit-oriented, can reinforce the adverse effect of motherhood in terms of extra remunerations. The effect could also stem from distinct negotiation strategies between mothers and non-mothers. Mothers may prefer to negotiate more flexibility over traditional salary increases, seeking options like reduced hours, ease of entering or leaving work, or shorter lunch breaks. Therefore, it is crucial to understand whether any discrimination exists regarding access to non-base compensation, potentially linked to stereotypes among human resources personnel, or informal criteria used to grant extra remunerations that are not gender and motherhood-neutral. Future research should delve into the dynamics of women's wage negotiation, identifying any barriers that may impede this process. This will enable to propose appropriate interventions to strengthen mothers' negotiation strategies.

Upon comparing the outcomes of women and men, it becomes apparent that parenthood tends to penalize access to extra payments for both men and women, though to a lower extent for men. Indeed, fathers still enjoy a wage premium regardless of the wage measure. In addition, results reveal that parenthood only marginally contributes to the overall gender wage gap in Italy. Hence, it is rather a gender issue.

From a policy perspective, our work calls for more transparency in workplace practices used to determine the base and variable components of the pay. The EU Directive 2023/970 aligns with this aim, although it targets employers with at least 100 workers, mandating regular reporting on pay. Regarding the Italian case, where small family enterprises are prevalent, our results sustain extending this requirement to encompass firms with fewer than 100 workers. Such reporting enables more accurate measurement of the gender wage gap within enterprises and drives employers to take action by implementing gender-neutral job evaluation criteria to address inequalities. Another policy option to address the gender wage gap could involve enhancing the recently introduced Gender Equality Certification System under the Italian National Recovery and Resilience Plan. This system sets guidelines for gender equality and offers incentives to companies that complete the certification process. At present, certification is voluntary. Making certification mandatory or increasing incentives for certified companies could effectively reduce the gender wage gap.

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Appendix - Supplementary materials

Table A1: Descriptive statistics

	Italy				Macro-regions															
					North-West				North-East				Centre				South and Islands			
	Childless women		Mothers		Childless women		Mothers		Childless women		Mothers		Childless women		Mothers		Childless women		Mothers	
	mean	sd	mean	sd	mean	sd	mean	sd	mean	sd	mean	sd	mean	sd	mean	sd	mean	sd	mean	sd
Outcomes of interest:																				
Log-hourly *base-wage*	2.292	0.284	2.368	0.287	2.320	0.270	2.399	0.272	2.325	0.260	2.395	0.267	2.291	0.281	2.361	0.284	2.199	0.323	2.283	0.331
Log-hourly *full-wage*	2.387	0.329	2.422	0.327	2.428	0.312	2.455	0.319	2.437	0.301	2.446	0.300	2.360	0.326	2.409	0.321	2.287	0.370	2.351	0.377
Individual characteristics:																				
Real labor market experience	8.556	4.954	11.630	4.990	8.863	4.975	12.152	4.869	9.474	5.345	12.434	4.936	8.323	4.698	11.304	4.938	7.006	4.213	9.883	4.865
Italian nationality	0.898	0.303	0.863	0.344	0.908	0.289	0.858	0.349	0.899	0.301	0.844	0.363	0.875	0.331	0.856	0.351	0.915	0.280	0.915	0.279
University degree	0.312	0.463	0.230	0.421	0.308	0.462	0.204	0.403	0.310	0.463	0.234	0.424	0.320	0.466	0.239	0.427	0.307	0.461	0.248	0.432
Living in a couple	0.293	0.455	0.856	0.352	0.348	0.476	0.845	0.362	0.339	0.474	0.883	0.321	0.269	0.444	0.837	0.369	0.171	0.377	0.850	0.358
Job-related characteristics:																				
Occupation 1	0.147	0.354	0.126	0.332	0.159	0.366	0.116	0.320	0.141	0.348	0.118	0.323	0.139	0.346	0.127	0.333	0.147	0.354	0.155	0.362
Occupation 2	0.236	0.425	0.245	0.430	0.242	0.429	0.250	0.433	0.249	0.433	0.248	0.432	0.226	0.419	0.225	0.418	0.220	0.415	0.261	0.440
Occupation 3	0.238	0.426	0.218	0.413	0.243	0.429	0.242	0.428	0.254	0.435	0.220	0.414	0.248	0.432	0.216	0.412	0.191	0.393	0.182	0.386
Occupation 4	0.237	0.425	0.216	0.411	0.226	0.418	0.201	0.401	0.205	0.404	0.216	0.412	0.238	0.426	0.230	0.421	0.300	0.458	0.216	0.412
Occupation 5	0.143	0.350	0.195	0.396	0.130	0.336	0.190	0.392	0.151	0.358	0.198	0.398	0.149	0.356	0.202	0.402	0.142	0.349	0.186	0.389
Public sector	0.233	0.423	0.283	0.450	0.219	0.414	0.263	0.440	0.244	0.430	0.289	0.454	0.220	0.414	0.242	0.429	0.255	0.436	0.360	0.480
Firm size	0.264	0.441	0.304	0.460	0.297	0.457	0.341	0.474	0.290	0.454	0.319	0.466	0.256	0.437	0.297	0.457	0.183	0.387	0.232	0.422
Permanent job contract	0.802	0.398	0.879	0.326	0.847	0.360	0.920	0.271	0.823	0.381	0.890	0.313	0.782	0.413	0.876	0.330	0.733	0.443	0.805	0.397
Managerial position	0.158	0.365	0.144	0.351	0.168	0.374	0.154	0.361	0.199	0.399	0.174	0.379	0.135	0.342	0.111	0.314	0.115	0.320	0.125	0.331
Part-time	0.175	0.380	0.341	0.474	0.164	0.370	0.334	0.472	0.118	0.323	0.368	0.482	0.171	0.376	0.317	0.466	0.289	0.453	0.337	0.473
N	10,072		8,760		2,748		2,264		2,815		2,729		2,707		2,252		1,802		1,515	

Note: (i) Real labor market experience is measured in years. (ii) Occupation dummies from 1 to 5 refer, respectively, to the following categories: (1) legislators, senior officials, and managers, professionals; (2) technicians and associate professionals; (3) clerks; (4) service workers and shop and market sales workers; and (5) skilled agricultural and fishery workers. (iii) Public sector is a dummy variable that refers to education, health, and public administration; (iv) Firm size is a dummy variable equal to one for firms with more than 50 employees and zero otherwise. (v) Part-time is a dummy variable equal to one if the employee works less than 30 hours per week. (vi) sd stands for standard deviation.

Table A2: OLS and Heckman estimates

	OLS		Heckman	
	base-wage	*full-wage*	*base-wage*	*full-wage*
	(1)	(2)	(3)	(4)
Mother	0.044***	0.000	0.041***	-0.007
	(0.005)	(0.005)	(0.005)	(0.005)
Experience	0.011***	0.016***	0.012***	0.020***
	(0.001)	(0.002)	(0.002)	(0.002)
Experience (squared)	-0.000***	-0.000***	-0.000***	-0.001***
	(0.000)	(0.000)	(0.000)	(0.000)
Italian nationality	0.065***	0.103***	0.067***	0.107***
	(0.006)	(0.007)	(0.007)	(0.007)
University degree	0.080***	0.084***	0.082***	0.090***
	(0.005)	(0.006)	(0.005)	(0.006)
Living in couple	0.016***	0.015***	0.014***	0.009
	(0.004)	(0.005)	(0.005)	(0.005)
Occupation: 2	-0.032***	-0.011	-0.032***	-0.011
	(0.006)	(0.007)	(0.006)	(0.007)
Occupation: 3	-0.066***	-0.041***	-0.066***	-0.041***
	(0.007)	(0.008)	(0.007)	(0.008)
Occupation: 4	-0.143***	-0.139***	-0.142***	-0.139***
	(0.007)	(0.008)	(0.007)	(0.008)
Occupation: 5	-0.191***	-0.212***	-0.191***	-0.212***
	(0.008)	(0.009)	(0.008)	(0.009)
Public sector	0.078***	0.066***	0.078***	0.066***
	(0.005)	(0.005)	(0.005)	(0.005)
Firm size	0.082***	0.112***	0.082***	0.112***
	(0.004)	(0.005)	(0.004)	(0.005)
Permanent job contract	0.065***	0.109***	0.065***	0.109***
	(0.006)	(0.007)	(0.006)	(0.007)
Managerial position	0.018***	0.031***	0.018***	0.031***
	(0.005)	(0.006)	(0.005)	(0.006)
Part-time	0.085***	0.037***	0.085***	0.037***
	(0.005)	(0.006)	(0.005)	(0.006)
Constant	2.136***	2.161***	2.127***	2.132***
	(0.013)	(0.014)	(0.015)	(0.016)
N	18,832	18,832	18,832	18,832
R-squared	0.243	0.253		

Note: (i) Sub-period and macro-region fixed effects are controlled for. (ii) Exclusion restrictions explaining employment decisions, but not wages, are household non-labor income and partner's labor income. (iii) Standard errors are in brackets. (iv) Significance levels: $^{**}p < 0.05$, $^{***}p < 0.01$.

Table A3: Results of the robustness analysis

	base-wage (1)	*full-wage* (2)
Panel 1 - Age selection		
Mother (20-40 years old)	0.044***	0.004
	(0.004)	(0.005)
N	20,489	20,489
R-squared	0.258	0.265
Mother (20-41 years old)	0.046***	0.006
	(0.004)	(0.005)
N	22,310	22,310
R-squared	0.264	0.273
Mother (25-41 years old)	0.046***	0.005
	(0.004)	(0.005)
N	20,798	20,798
R-squared	0.251	0.262
Panel 2 - No additional sample restrictions		
Mother (25-40 years old)	0.052***	
	(0.004)	
N	28,159	
R-squared	0.261	
Panel 3 - Full-time equivalent monthly wage		
Mother (25-40 years old)	0.050***	0.007
	(0.005)	(0.005)
N	18,832	18,832
R-squared	0.203	0.239
Panel 4 - Monthly wage		
Mother (25-40 years old)	0.035***	
	(0.005)	
N	18,832	
R-squared	0.500	

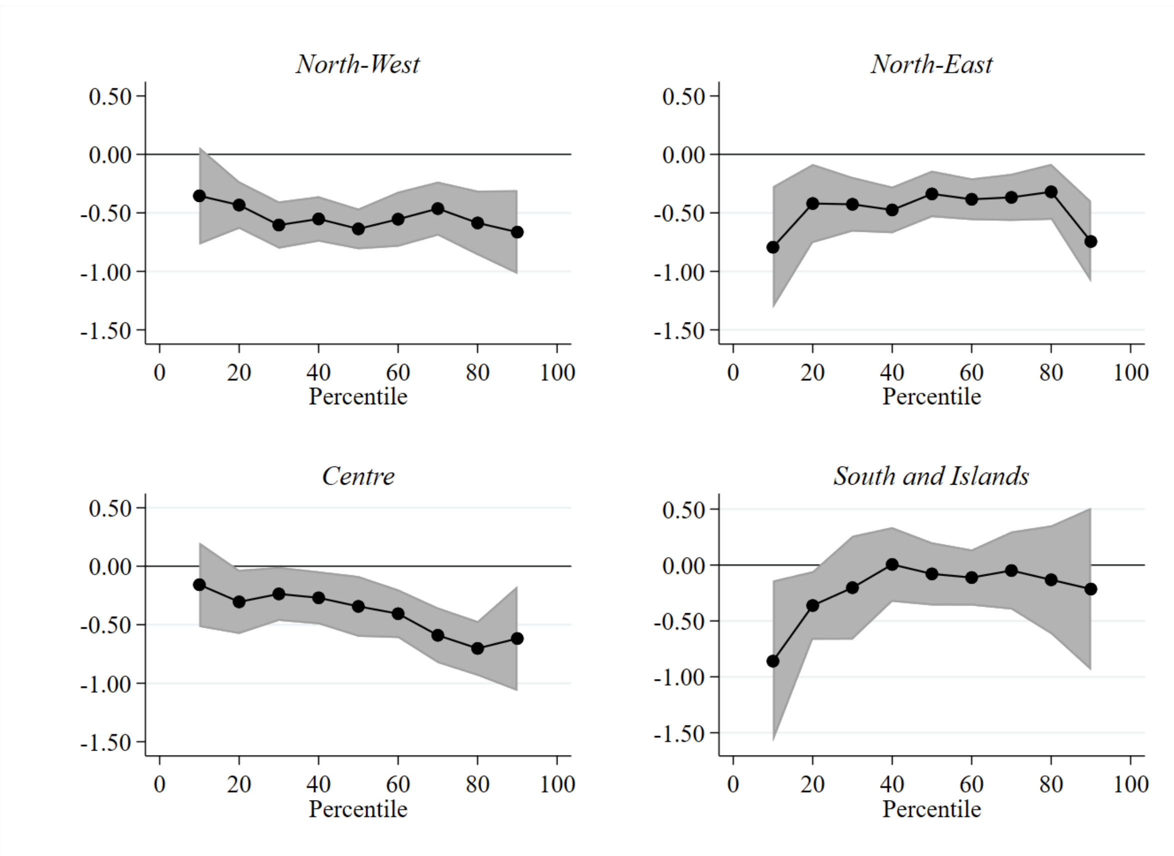
Note: (i) The full-set of results is available from the authors upon request. (ii) The control variables are the set of individual and job-related characteristics listed in Table A1, as well as macro-region and sub-period fixed effects. (iii) Significance levels: $^{**}p < 0.05$, $^{***}p < 0.01$.

Table A4: OLS estimates – Sample of men

	base-wage	*full-wage*
	(1)	(2)
Father	0.043***	0.021***
	(0.005)	(0.005)
Experience	0.008***	0.016***
	(0.001)	(0.001)
Experience (squared)	-0.000***	-0.000***
	(0.000)	(0.000)
Italian nationality	0.044***	0.107***
	(0.005)	(0.006)
University degree	0.081***	0.086***
	(0.005)	(0.006)
Living in couple	0.042***	0.032***
	(0.005)	(0.005)
Occupation: 2	-0.031***	-0.014*
	(0.007)	(0.008)
Occupation: 3	-0.057***	-0.035***
	(0.008)	(0.008)
Occupation: 4	-0.090***	-0.100***
	(0.008)	(0.009)
Occupation: 5	-0.115***	-0.122***
	(0.007)	(0.008)
Public sector	0.082***	0.106***
	(0.005)	(0.006)
Firm size	0.065***	0.111***
	(0.003)	(0.004)
Permanent job contract	0.094***	0.137***
	(0.006)	(0.007)
Managerial position	0.059***	0.066***
	(0.004)	(0.004)
Constant	2.195***	2.204***
	(0.012)	(0.013)
N	22,182	22,182
R-squared	0.210	0.260

Note: (i) Sub-period and macro-region fixed effects are controlled for. (ii) Standard errors are in brackets. (iii) Significance levels: $^{**}p < 0.05$, $^{***}p < 0.01$.

Figure A1: QTE estimates of the hourly extra remunerations by macro-region



Note: (i) The dependent variable is the difference between the level of the *full-wage* and the level of the *base-wage*. (ii) The control variables are the set of individual and job-related characteristics listed in Table A1, as well as sub-period fixed effects, specified as dummy variables. (iii) The figure shows confidence intervals at 95% level.