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THE PUBLIC-PRIVATE PAY GAP:
A ROBUST QUANTILE APPROACH

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Abstract

This paper investigates whether a public sector premium exists even after controlling for observable characteristics and for additional motivations, other than monetary, that may induce workers to prefer employment in the public sector. We do this by studying the entire conditional wage distribution on Italian micro data, covering the period 1998-2008. The evidence under random sampling shows the existence of a wage differential averaging at about 14% for women and 4% for men, generally lower at the high tail of the wage distribution and in the Northern regions. The premium significantly increases when possible sorting is considered; the correction is particularly large above the median of the wage distribution, therefore suggesting that the additional motivations may play an important role above all at higher wage levels. When we restrict our comparison to large private firms, a differential is confirmed for women but not for men.

JEL Classification: H50, J31, J45, J50.
Keywords: public employment, wage differentials, wage determination.

Contents

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* Bank of Italy, Economics, Research and International Relations.

1 Introduction [1]

Wage and employment characteristics in the public sector differ significantly from those in the private sector. Analyses comparing average earnings outcomes in public and private sectors generally find positive wage differentials for public sector employees. These analyses either use aggregate data on average wages, which are easily obtained from published sources, or use individual data in an attempt to control for the quality and the composition of public employees relative to private employees. When controlling for employment characteristics, the wage differential typically decreases with the number of controls and varies with the econometric specification and across countries.

Several types of econometric techniques have been adopted to investigate the wage gap between sectors using individual data. One approach estimates a single earning equation augmented with a dummy variable indicating whether the worker is employed in the public sector or not. The coefficient of such a variable provides a measure of the wage differential after having controlled for other characteristics of the worker. In the same vein, another econometric specification allows the coefficients to vary across sectors. It consists of two wage equations, one for each sector, in order to capture different returns to worker characteristics. The wage gap is obtained by taking the difference between the worker's actual wage and the estimate of what the worker would earn in the sector in which he is not employed. Through standard decomposition techniques it is then possible to disentangle the impact on the wage gap of differences in worker endowments from that associated with unexplained factors (usually interpreted as the 'rent to public sector'). More recently, quantile regression techniques have also been used to compare wages in the public and the private sectors along the entire wage distribution. These analyses usually find higher differentials at the bottom end of the distribution and lower or even negative premia at the top deciles.

The results obtained by all the above techniques should be interpreted with caution. Indeed, there may be a sample selection bias due to the possibility that sorting of employees between sectors is not random but occurs on the basis of unobserved characteristics. This problem has typically been addressed by jointly estimating equations for the worker's sector of employment and for earnings (relying on a set of exclusion restrictions), or by using longitudinal data. To choose appropriate instruments the literature has tended to follow some studies into whether certain characteristics of individual employees increase the probability that they will seek employment in the public sector. For example, Bellante - Link (1981) find that workers employed in the public sector are more risk averse than private sector employees. Perry - Wise (1990) argue that public sector jobs are likely to provide a greater scope than in the private sector to engage in community or welfare-type activities that will affect individual workers' labour supply decisions; finally, imitation of parents and the presence of social networks, as summarized by variables related to family background, can

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be important channels to enter one sector or the other. A somewhat related argument is proposed by Heyes (2005) and Barigozzi - Turati (2011), who explore the role of intrinsic motivation in the market for nurses. They claim that workers, if intrinsically motivated to perform their job, receive a non-pecuniary benefit (a vocational premium) in addition to the wage rate. This implies that a wage increase, by attracting workers with low intrinsic motivation under certain conditions, may have adverse effects on the average vocation and hence on the average quality of services. Estimated wage gaps obtained by means of sample selection corrections are generally found to be greater than those not conditioned on these corrections.

Despite the large body of empirical literature on the public sector pay gap, the theory of why such a wage differential exists is scant.[2] Most arguments invoked to justify the evidence of a wage differential are related to the peculiarity of the public employer's objective function, which encompasses lobbying (Gunderson, 1978), electoral motives (Fogel - Lewin, 1974) or budget-maximization (Niskanen, 1975; Reder, 1975). Other arguments focus on the relatively inelastic labour demand curve in the public sector that the unions would exploit, to the extent allowed by the government budget constraint, to gain higher wages for public sector workers. In a country where tax evasion is deemed to be significant and largely confined to the private sector (up to 27% in Italy according to Schneider, 2002) a further reason for the pay gap can be that, for a given gross salary, tax evasion leads to different net salaries across sectors; thus, the only way to make the public sector attractive for workers is to offer the same *net* salary rate (Levaggi, 2007).[3]

In this paper we contribute to the empirical analysis. We investigate the public-private pay gap using Italian micro data covering the period 1998-2008. In contrast to most of the existing literature, we analyse the entire distribution of (log) wage, by using conditional quantile regression techniques, and properly consider the possible endogeneity of the sector choice. To our knowledge, the only other attempt to jointly address these issues is provided by Melly (2006). As exclusion restrictions, in addition to variables related to parents' occupational status as in the existing literature, we investigate the roles of risk aversion, inter-temporal preferences and pro-social attitude. We check the robustness of our findings distinguishing between small and large firms or using different measures of job compensation.

The paper is organized as follows. Section 2 presents the existing empirical evidence, focussing in particular on Italy; Section 3 briefly describes our empirical strategy; Section 4 presents the data and some descriptive statistics; Section 5 discusses our findings; Section 6 offers some concluding remarks.

2 Related literature

Most of the early research on this issue focussed on the US; only few studies based on macro data were carried on non-US countries. At the beginning of the '90s a number of studies started

[2]The only theoretical paper on the sectoral wage differential is Holmlund (1993). In his paper the public and private sector unions bargain over wages with a utilitarian government. The model predicts a wage premium only if the unions act uncooperatively. In this case, in fact, the unions do not internalize two sources of externalities: the first involves the increase in taxes needed to pay for the increase in public sector wages, which are spread across all workers; the second is due to the fact that increases in public sector wages will cause employment in the public sector to decrease, and therefore consumption of goods produced by the public sector to decrease, but again for all workers and not just for public sector workers.

[3] Notice however that papers that explicitly tackle the issue of tax evasion in Italy use the same dataset that we use here as the benchmark for the true income (see Cannari - Ceriani - D'Alessio, 1995, and Fiorio - D'Amuri, 2005).

addressing wage differentials in Europe, Australia and some developing countries. Bender (1998) and Gregory - Borland (1999) provide extensive surveys of such analyses in a range of countries.

Among more recent studies focussing on European countries, Portugal - Centeno (2001) use the 1995 wave of the European Community Household Panel (ECHP) to compare wage differentials between the general government and the private sector in the European Union member states. Considering identical worker's characteristics, they find that the wage gap is wider in Portugal, Ireland, Luxembourg, Spain and Italy; it is narrower in Austria, Belgium, Germany; in Denmark the differential turns out slightly negative.

Cross-country analyses can be useful to assess the impact on wage differentials of different institutions, wage setting schemes, macro-economic and labour market conditions or culture. However, heterogeneity in data collection and, in particular, in the definition of the public sector may yield spurious results.

Within EU studies, a single country perspective is adopted by Disney - Gosling (1998, 2003) who focus on British data in the nineties. Using a panel data fixed effects approach robust to possible sector sorting, they find that women in the public sector earn more than in the private sector, whilst on average men do not. Melly (2005, 2006) analyses the public sector pay gap in Germany. In his 2005 paper he measures and decomposes the differences in wage distributions between public sector and private sector employees for the years 1984-2001, and obtains higher conditional wages in the public sector only for the women. Furthermore, wage premia are found only at the low end of the wage distribution, whereas differences in workers' characteristics explain more than the raw wage gap at high wages. These features appear quite stable over both decades. Using data on Germany in 2003, Melly (2006) addresses the issue of possible sorting between the two sectors, along the entire conditional distribution. He finds a positive premium at the low end of the conditional wage distribution and a significant negative premium at the upper tail under the assumption of exogenous sector choice. When endogenous sorting is considered, the premium increases by roughly 40 percentage points. Postel-Vinay - Turon (2007) estimate a model of income and employment dynamics for the UK in the period 1996–2003. They allow for unobserved heterogeneity in the propensity to be unemployed or employed in either job sector, exploiting the panel structure of their data. They find a positive average public premium both in income flows and in the present discounted sum of future income flows. Bargain - Melly (2008) estimate the public sector wage premium in France over the period 1990-2002, both at the mean and at different quantiles of the wage distribution. They account for unobserved heterogeneity by using fixed effects estimations on panel data. Contrary to common results, they find that wages do not substantially differ across sectors after controlling for unobserved heterogeneity. They explain these results by arguing that the public sector manages to attract better workers in the lower part of the distribution, in part due to non-monetary gains, but fails to retain the most productive ones at the top.

2.1 The wage gap in Italy

In Italy the difference in pay between the public and the private sectors in recent decades has always been sizeable. If we look at aggregate data from national accounts the gap was about 20 percent in 1980 and reached almost 40 percent in 1990, following a series of particularly favourable wage renewal contracts in the public sector; it decreased to 22 percent in 1995, reflecting the overall fiscal consolidation effort required under the Maastricht Treaty to join the European Monetary Union; the differential started increasing again at the beginning of the last decade, to reach 33 percent in

2010 (Figure 1).

This evolution seems to contrast with the introduction in 1993 of the current legislation, which aimed at a "privatization" of employment relations in the public sector, that is at making pay and employment condition determination mechanisms in the public sector closer to those in the private sector, by envisaging a greater role for negotiation, imposing tighter constraints on wage growth, and replacing the automatic component for wage increases with schemes based on merit. The 1993 reform assigned a larger role to collective bargaining and created an independent agency (Agenzia per la Rappresentanza Negoziale nella Pubblica Amministrazione - ARAN) to be responsible for negotiating pay levels and working conditions for most public employees. In principle, closer cooperation between unions in the public and the private sectors during the wage determination process (which could have followed the greater role assigned to collective bargaining in the public sector) together with a more centralized bargaining process (due to the creation of a single agency representing a vast majority of public employees) should have fostered wage moderation in the public sector or, at least, determined wage dynamics in the public sector more in line with those in the private sector. (Forni - Giordano, 2003, present a theoretical model that delivers these results.) Indeed, many analysts agree that the reform failed to achieve its main targets (see Dell'Aringa, 2006).[4]

Possible reasons for the gap are that public sector workers are older, are better educated and take managerial positions more frequently than in the private sector. In fact, taking into account the significant heterogeneity in the composition of the labour force (by age, gender, education and occupational level) , the pay gap turns out to be lower but still sizeable.

Most of the papers that analyse the public-private pay gap in Italy using micro data are based on the Bank of Italy Survey of Household Income and Wealth (SHIW), which contains information about personal and occupational characteristics, wages (net of income and payroll taxes) and type of economic activity.

Among them, Bardasi (1996) looks at 1991 incomes. She uses a two-stage approach to take into account the possibility that the distribution of people between sectors may not be random, but rather result from self-selection. She finds a wage premium in the public sector of 9 and 35 per cent for men and women, respectively. Although differences in observable human capital characteristics explain a small fraction of the overall gap, most of it is due to a residual component, which captures differences in compensation of given individual characteristics and the effect of geographical and occupational dummies. Furthermore, self-selection seems to negatively affect the wage differential, above all for men. If workers were distributed randomly across sectors, the average wage in the public sector would be 25 and 17 per cent higher for men and women, respectively.

Brunello - Dustmann (1997) compare public-private wage differentials in Italy and in Germany. As for Italian data, they use the 1993 wave of the SHIW; German data are from the German Socio-Economic Panel and refer to 1989 incomes. They find a positive wage gap in both countries, higher in Italy than in Germany (21 and 7 per cent, respectively). Moreover, by decomposing the differential into two factors, one associated with individual characteristics and the other with price differentials, they find that the positive difference in Germany is entirely explained by the first factor, which more than offsets a negative price differential that is larger for workers with higher education. In contrast, in Italy the compensation of different characteristics adds to a positive price differential that is higher for less educated workers. Therefore, they conclude that, for given

[4]Analyses on both the motivations and the effects of the 1993 reform are provided, among others, by Dell'Aringa (1997) and Lucifora (1999).

individual characteristics, public sector workers are penalized in Germany and rewarded in Italy. According to the authors, their preliminary findings do not reveal significant sample selection biases.

Comi - Ghinetti - Lucifora (2002) analyse the gap over the period 1977-1998 using data from the SHIW. They also find a positive wage gap for the public sector that is higher for women and for low-income workers. They observe that this difference, after peaking in 1995 (20 per cent for women, after controlling for individual characteristics), started decreasing in 1998. They also explain the wage moderation in the public sector observed in the last part of their sample period by mentioning the reform in the pay determination mechanisms introduced in 1993 to "privatize" the employment relationship for most public sector employees.

Although gender is one key determinant of the public sector pay gap, the geographical heterogeneity also turns out to be substantial. An analysis of pay differentials at regional level is provided by Dell'Aringa - Lucifora - Origo (2007). They show that such differentials, which vary significantly across Italian regions, are partly explained by local labour market conditions that influence wages in the private sector and only marginally affect wages in the public sector. In the years 1991-2002 in most regions in the North the gap, after controlling for individual and occupational characteristics, turns out to be below 10 per cent, about 4-6 per cent in the largest and most industrialized areas (Piemonte, Lombardia and Emilia Romagna). In contrast, in the southern regions the differential almost always exceeds 15 per cent, reaching 20-25 per cent in the regions characterized by high unemployment and where public employment is widespread (Calabria and Sicilia).

Finally, Lucifora - Meurs (2006) investigate the public sector pay gap in 1998 in France, Great Britain and Italy. They use the quantile regression (QR) method to analyse wage differentials along the wage distribution, and find that the pay gap significantly declines along it. The estimated wage premium from the pooled model ranges between 10 and 12 per cent (lower when more control variables are included in the specification) in the lowest deciles of the distribution, and is about zero in the highest deciles (in the case of women the gap remains positive even at the top deciles, while for men it turns negative in the upper part of the distribution). Quite surprisingly, when conditioning on a larger set of variables they find the highest premium in Great Britain (about 6 per cent on average), where the private sector wage is systematically used as a reference for pay determination in the public sector. Estimates for France and Italy, where no application of this comparability principle is in place, are lower and do not differ considerably from one another (approximately 5 per cent).

To summarize, the presence of a wage gap in favour of the public sector in Italy is largely documented. The existing analyses, most of which focus on a particular year, show a wage differential generally exceeding 20 per cent for all employees, which decreases but remains quantitatively large and significant after controlling for the individual characteristics of the workers. The gap varies a lot between men and women, workers with different educational and professional qualifications and across geographical areas. With the exception of Lucifora - Meurs (2006), these studies focus on averages rather than on the entire distribution of wages and, with the exception of Bardasi (1996) and Brunello - Dustmann (1997), they take the decision between private and public sector as random. In contrast, we relax this assumption and consider the entire wage distribution.

3 The empirical strategy

In our analysis, in addition to Ordinary Least Square (OLS), we use quantile regression (QR) to assess the value of the pay gap along the entire wage distribution.

Indeed, there is in general no guarantee that the mean of a distribution successfully summarizes all its relevant features. This is particularly true in the presence of outliers, as the results obtained using standard approaches are very sensitive to them. Moreover, QR is useful when the impact of explicative variables is different across different parts of the distribution as, in contrast to OLS, it does not impose that covariates have the same effects over the entire distribution. In particular, we employ the quantile regression methodology developed by Koenker - Bassett (1978). It consists of a semi-parametric estimate that combines two parts: one is parametric and refers to the deterministic part, i.e. the covariates; the other is non parametric and concerns the random residuals about which no distributional assumption is imposed. In a simple model $y_i = x_i\beta + u_i$ with $i = 1, \ldots, N$, when we are interested in the τ-th conditional quantile of y_t, β is obtained from the minimization

$$\underbrace{\min}_{\beta \in R} \left[\sum_{i \in \{i: y_i \geq x_i\beta\}} \tau |y_i - x_i\beta| + \sum_{i \in \{i: y_i < x_i\beta\}} (1-\tau)|y_i - x_i\beta| \right].$$

Under some regularity conditions, the solution to the minimization problem gives a consistent and asymptotically normal estimator.[5] Moreover, it is shown that, whereas for a symmetric distribution function median and mean regressions are equivalent, for a more complicated distribution function the median regression is uniformly more efficient than the mean regression. For the present study all these aspects are very useful and will be exploited throughout.

An additional concern that is related to our specific goal regards the exogeneity of the decision to work in the public sector. Suppose for example that workers in the public sector are more risk averse (Bellante - Link, 1981) and that managers in the private sector have more responsibility: if we don't control for the amount of responsibility in the OLS and QR, we would estimate as a "private premium" what is in fact compensation for responsibility. Of course, the opposite is also true: if, for example, private companies give fringe benefits to their workers as part of salary, we would conclude that there exists a "public premium" if we don't control for the amount of non monetary benefits. Hence, endogeneity of sector choice can bias our estimates and can eventually lead to a misunderstanding of the driving forces of the data generating process. It follows that the assumption of exogeneity must be checked and cannot be imposed.

To our knowledge, the first systematic treatment of endogeneity with median regression is due to Amemiya (1982) who proposed a consistent estimator, whose asymptotic properties were further generalized in Powell (1983). The approach is a two stage least absolute deviations (2SLAD) and is analogous to the two stage least square (2SLS) for the mean.

More recently in a series of papers Chernozhukov - Hansen (2005) suggest a dual approach. They consider a model

$$\begin{cases} y &= D\alpha(U) + X\beta(U) & \text{U |X,Z} \sim \text{Uniform(0,1)} \\ D &= \delta(X, Z, V) & \text{V is statistically dependent on U} \\ \tau &\mapsto D'\alpha(\tau) + X\beta(U) & \text{is strictly increasing in } \tau \end{cases}$$

where D is statistically dependent on the scalar random variable U that aggregates all the unobserved factors affecting the structural outcome equation, X is a set of exogenous covariates and Z is a set of instruments related to D, but independent on U. Thus, the true structural parameters α

[5] The main conditions are that i) the distribution F and its density f are continuous and ii) the matrix $T^{-1}X'X$ is positive definite.

cannot be estimated with the usual quantile regression, but with an Instrumental Variable Quantile Regression (IVQR). The method is a GMM on

$$Q_n(\tau, \alpha, \beta, \gamma) = \frac{1}{n}\sum_{i=1}^{n} \rho_\tau(Y - D\alpha - X\beta + Z\gamma).$$

The spirit of the approach is that if we knew the true value of α we could obtain consistent estimates using the ordinary quantile regression of $y - D\alpha$ on $X\beta$. Thus we can run a series of quantile regressions for *given* values of α (grid search) to obtain

$$(\beta(\hat{\alpha, \tau}), \gamma(\hat{\alpha, \tau})) = \underbrace{\arg\min}_{\beta,\gamma} Q_n(\tau, \alpha, \beta, \gamma)$$

and select $\alpha(\tau)$ that makes γ as close to 0 as possible. More formally, we must find a value of α that satisfies

$$\alpha(\hat{\tau}) = \underbrace{\arg\inf}_{\alpha \in A}[W_n(\alpha_n)], \qquad W_n(\alpha) = n[\gamma(\hat{\alpha, \tau})' \hat{A}(\alpha) \gamma(\hat{\alpha, \tau})],$$

with $A(\alpha)$ set equal to the asymptotic covariance matrix of $\sqrt{n}(\gamma(\hat{\alpha, \tau})' - \gamma(\alpha, \tau)')$, so that $W_n(\alpha)$ can be employed as the Wald statistic to test $\gamma(\alpha, \tau) = 0$ (Chernozhukov - Hansen - Jansson, 2007). The estimator is shown to be consistent and asymptotically normal as

$$\sqrt{n}([\alpha(\hat{\tau}), \beta(\hat{\tau})]' - [\alpha(\tau), \beta(\tau)]') \rightarrow_{d.} N(0, \Omega_\tau),$$

for Ω_τ as given in Chernozhukov - Hansen (2005). We conclude this section with a summary of the four steps required to estimate the coefficients:

1. define a set of (j=1,...,J) values for α;
2. save the covariance matrix to be used for $W_n(\alpha_j)$;
3. choose the $\alpha(\tau)$ that minimizes $W_n(\alpha)$ and obtain an estimate for β;
4. obtain the inference on α using Ω_τ.

Finally, note that this procedure works when the number of exogenous covariates is large, but the number of endogenous covariates is small (typically one, as in our case, or two), because a grid search is employed.

4 The data

Our data are taken from SHIW and refer to the period 1998–2008. Its target population is representative of the Italian population; the sample consists of a small part followed over time (i.e., panel) and a larger part refreshed at each wave. To avoid a large drop in observations, we do not exploit the distinction throughout the paper. The data contain information about a wide range of personal and occupational characteristics (age, gender, marital status, educational level, region of residency, sector of economic activity and occupational level), wages (net of income and payroll taxes) and type of activity (firm size, part-time status, number of months worked in the year, average number of hours worked in a week). As for all survey data, the quality of SHIW is affected by unit and item non-response. Unit non-response is sizeable (it amounts to about 50 per cent of the selected households), whereas item non-response depends on the variable under study. The first issue is tackled by a post stratification weighting process, whereas the second issue is irrelevant for employees and so the present study does not suffer from these drawbacks. Furthermore, the data quality is satisfactory when monetary aggregates are compared to Italian National Accounts (Banca d'Italia, 2008).

One critical aspect related to this study concerns the identification of the public sector, which can be defined using two different questions, one regarding the sector of economic activity and the other categorized as firm size. In particular, a public worker can be identified when *i)* his/her sector of activity is "public administration, defence, education, health and other public services" or, up to 2006, when *ii)* firm size "is not applicable, because public employee". We checked the robustness of our results using both definitions, but the rest of the paper is based on sector of activity, as in Lucifora - Meurs (2006).

Another issue is related to the choice of the appropriate variable for wage comparison. Indeed, as the average number of hours worked in a week in the public sector (35) is lower than that reported by workers in the private sector (39), using the monthly wage could underestimate the wage differential across sectors (we return to this issue in Section 5.5). Thus in our benchmark specification we approximate the hourly wage as $(YW/M)/(4 * \bar{H})$, where YW is the yearly wage, M the number of months worked in the year, and $\bar{H}$ the average number of hours worked in a week.

4.1 Descriptive statistics

In this section we illustrate some descriptive statistics for key wage related characteristics of the sample at hand. The statistics refer to employees aged 15–65, as is usual in this literature.

In Table 1 we tabulate the share of public sector workers in total employment by wave, gender and area. We consider five different areas: North West (NW), North East (NE), Centre (C), South (S) and Islands (Is). Public sector employment in Italy is sizeable. It is more widespread among women than among men, in the South and in the Islands than in the rest of the country and was declining from 1998 to 2008. In our sample the percentage of public sector employment was 43.5% for women and 28.5% for men in 1998, and went down to 35.4% and 20.9%, respectively, in 2008.[6] The reduction occurred in all areas although in the North to a lesser extent, bringing about a "convergence" towards a less dispersed range across areas. The gender gap remained constant over time at about 14%, with again substantial differences across areas.

In Tables 2–3 we report some descriptive statistics for individual characteristics for the pooled 1998–2008 sample (the same evidence emerges in each single year). In Table 2 we present summary statistics for the entire distribution of the log of hourly wage and age. On average in Italy the hourly wage in the private sector is higher in the sample of men than in the sample of women by 12% (12% in the North East, 15% in North West, 16% in the Centre, 21% in the South and 3% in the Islands). If we focus on selected quantiles, the picture is even more differentiated over the territory: for example, in the South the gender gap is 25% at the 25th quantile, 21% at the median and 19% at the 75th, i.e. decreasing over quantiles, whereas in the Centre it does not

[6] In National Accounts data the fraction of public sector employees in total employment was equal to 20% in 2008. The difference with our sample is likely due to the way in which the public sector is identified, in particular the sample may include employees working outside general government, e.g. in education and health services.

vary substantially along the wage distribution (15% at the 25th and 75th quantiles and 16% at the median). In the public sector the gender gap is smaller than in the private sector: 4% as the national average, positive in the North and in the Islands, and negative or zero in the Centre and in the South.

More important for the purposes of our analysis is the comparison of wages across sectors. On average the wage gap is in favour of the public sector by 34% for women and 25% for men; it is 30% and 36% at the 25th and the 75th quantile for women and 21% and 26% respectively for men. In the sample of women it is less than 30% in the North and much higher in the Centre (36%), in the South (55%) and in the Islands (41%). In the sample of men these differences are smaller: 20–25% in the North and in the Centre, slightly higher than 30% in the South and 40% in the Islands. Thus not only the gender gap, but also the private-public pay gap differs significantly across regions.

A graphical summary of these characteristics is presented in Figure 2. In panel *a)* we show the Kernel density estimation of the (log of) hourly distribution by gender and sector; the lower panel reports the cumulative distribution functions. In panel *b)* we fix the gender to compare the (unconditional) sector wage differential: it is quite clear that the horizontal distance between the two curves in the sample of women is larger than in the sample of men, with the public sector cumulative distribution function lying to the right of the private sector cumulative distribution function (i.e. at each quantile the wage is higher in the public sector than in the private sector); moreover, the difference tends to be larger at higher quantiles than at others. In panel *c)* we fix the sector: it is clear that the horizontal difference is larger in the private sector than in the public sector, with the curve for men lying to the right of the curve referring to women in the private sector (i.e., at each quantile men earn more than women), and virtually overlapping in the public sector.

Turning to workers' characteristics, the age of men and women is similar within each sector (with men being on average older by 1–2 years), but the difference across sectors is sizeable: on average public sector employees are older by about 5–7 years. Other workers' characteristics are reported in Table 3. The share of married individuals is higher amongst public sector employees than that in the private sector. More relevant to the wage process are education and job position. Public employees are better educated.[7] The fraction of women with a graduate or post-graduate degree is 32% in the public sector as against 8% in the private sector; for men it is 23% and 7%, respectively. The educational gap is observed in all geographical areas, but is wider in Southern regions. As for job position, the vast majority of both men and women are white collar workers in the public sector (up to 92% of women and 81% of men in the South), and blue collar workers in the private sector, with relative frequencies changing widely across areas. As expected, the fraction of managers is the lowest in both sectors, both for men and for women: it is three times as much in the public sector than in the private sector for men (7% as against 2%); it is 3% in the public sector and negligible in the private sector for women. In general, in the public sector occupational positions seem more favourable than in the private sector in relation to educational attainments, above all for men and in the Southern regions. In Italy the fraction of men with intermediate or high education is 71%, the fraction of white collar workers and managers is 85% in the public sector; the shares are 49% and 32%, respectively, in the private sector. For women the differences across sectors are less striking: the share of workers with intermediate and high education is 86%, that of white collar workers and managers is 87% in the public sector, as opposed to shares equal

[7] In the table the figures for education do not sum up to 1 because they do not include the share of individuals with minimal education.

to 57% and 48% in the private sector. Finally, part time jobs are observed more frequently for women than for men, who almost never work part-time, and in the private sector more than in the public sector.

Given the high variability in job position across areas, gender and sectors, in Table 4 we enrich the analysis of hourly wage by also looking at the position. Table 4 shows that the greatest pay differential between men and women is for private sector managers, whereas for private sector blue and white collar workers the difference is about 13–15%. Such differences are much smaller in the public sector (for example, the gender difference for white collar workers is virtually zero). As for the public-private pay gap, it is highest for white collar women workers, in particular those living in the South and in the Islands. For men the premium is generally lower and negative for managers.

So far we have documented a public sector pay gap by gender, job position and geographical area; the interesting question is now whether it persists even after controlling for relevant characteristics determining the wage process. We do this in the rest of the paper.

5 The results

In this section we present the model specification (Section 5.1) and our empirical findings on the wage gap in Italy. In Section 5.2 we report the results under random sampling, based on quantile and ordinary least square regressions; in Section 5.3 we remove the assumption of random sampling and control for possible endogeneity of the decision to work in the public sector with the earning process.

5.1 The model

To investigate the public sector pay gap we estimate a standard wage equation, extensively used in previous studies:

$$\ln W_{it} = \alpha X_{it} + \beta PUB_{it} + \epsilon_{it}, \tag{1}$$

where $\ln W_{it}$ is the log of hourly wages (net of income and payroll taxes) of individual i at time t, X_{it} is a vector of individual characteristics, PUB is a dummy variable, which takes the value one if the individual works in the public sector, α and β are vectors of parameters, and ϵ_{it} is an error term.

We focus on the coefficient attached to PUB, which is the "return on public sector". Other things being equal, if the coefficient is positive there exists a wage gap in favour of the public sector.

Estimating equation 1 using OLS, or QR, yields consistent estimates if and only if the distribution of workers between the two sectors is random and the vector X consists of exogenous variables. In fact, the distribution of people between the two sectors may not be random, but rather depend on the unobserved individual characteristics of the workers, such as different aptitude for risk or different preference for time flexibility; workers might also be heterogeneous across sectors with respect to the gratification that they derive from different types of jobs (e.g., they may desire to be a civil servant or work in the non-profit sector) and thus self-select according to those features; finally, there may be other factors, both monetary and non-monetary, such as fringe benefits, occupational pension benefits, job tenure or wage variability, that can contribute to explain wage differentials. In these situations, interpreting the wage gap as a "premium" or a "penalty" may not be appropriate, unless a correction is made.

Hence, we begin our analysis by making the assumption of random sampling, as most of the literature has done so far. We then relax this assumption and address the possibility of non-random sampling. From the previous section it should be clear that the wage differential may differ along the wage distribution; thus the following analysis is on selected quantiles as well as on the mean.

5.2 Random sampling

In Table 5 we present the estimates obtained with QR and OLS using data for all employees, and distinguishing by gender. We report only the coefficients attached to the public sector, although the regressions are conditional on a set of characteristics, namely: age (a second degree polynomial); marital status; educational level (a set of dummies for basic education, i.e. primary school, lower secondary school, which corresponds to compulsory school under Italian law, higher secondary school, and BA or post-graduate degree); job position (a set of dummies for blue collar workers, white collar workers and managers); whether part-time or not; geographical area; and year.[8] The reference individual is 40 years old, with a BA, working in the private sector as a full-time white collar worker in central Italy and observed in 2008.

The average differential is 14% for women and much lower for men, about 4%. However, further insights can be gained from the investigation of the complete distribution of the conditional wages. Indeed, the differential varies along the wage distribution and a formal test of equality of slope coefficients across quantiles rejects the null hypothesis at standard confidence levels (the F-statistics are 3.03 with a P-value of 0.002 for women and 9.75 with a P-value of 0.000 for men). As usual, much of the heterogeneity is on the tails of the distribution. For women, the differential ranges between 17% at the 10th quantile and 12% at the 90th: it declines fast between the 10th and the 30th quantile (from 17% to less than 12%) then it remains constant between the 40th and the 90th quantiles at about 12–13%, such that a formal statistical test does not reject the null hypothesis of the equality of slope coefficients between these quantiles. Unlike for women, the differential for men exhibits a clear decreasing pattern from lower to higher quantiles: it is 10% at the 10th quantile, less than 5% at the median, 2.5% at the 60th quantile and not really significant from the 70th onward.

The gender difference is on average 10% with important variation across quantiles. Although working in the public sector is more rewarding for women than for men at all quantiles, the gender difference increases along the wage distribution: it is about 5% at the 10th quantile, and is above 10% from the 60th quantile onward. This appears clearly in Figure 3 *a)*, where the maximum (vertical) distance between the two lines is at the right-hand side of the graph.

By splitting the sample between different geographical areas and occupational qualifications, other interesting features emerge from the analysis.

By geographical area, in relation to the sample of women, there is a differential in favour of the public sector in all areas that generally increases going from the North to the Centre and to the South of Italy. On average in the last decade the differential for women has been about 10% in the North, 15% in the Centre, 29% in the South and 18% in the Islands. Along the wage distribution, the pattern in each area resembles that observed in Italy as a whole. For men, the pay gap is almost always positive along the wage distribution in the North East, in the South and in the Islands (on average equal to 4%, 7% and 13%, respectively). In the North West and in the Centre it is not

[8] All complete estimates described in this and the following sections are available from the authors upon request.

significant on average; in the North West at high quantiles working in the public sector entails a negative rather than a positive differential.

The other important conditioning is on position, for both blue collar and white collar jobs, where interesting differences emerge by gender.[9] For blue collar women workers the conditional earning differential in favour of the public sector is decreasing from lower to upper quantiles, and generally much smaller than for white collar workers (the mean averaging at 9%, as against 16% for white collar workers). For men, instead, while the decreasing pattern along the wage distribution is common to both white collars and blue collars, the wage differentials in favour of the public sector is higher for blue collar workers than for white collar workers (on average 8% and 4%, respectively); at higher quantiles, the pay gap is not significantly different from zero at standard confidence level.

According to these results, there is an advantage to working in the public sector that is greater for women than for men. For women it is positive in all geographical areas and generally higher in the Centre and in the South. For men the pay gap is virtually zero in the North and the Centre of Italy, whereas it is positive in the South and in the Islands; the pay gap decreases along the wage distribution: at high quantiles no pay advantage, and in some cases even negative differentials are found.

However, these results must be evaluated with care because of a possible endogeneity between the decision to work in the public sector decision and the earning process. We try to resolve this issue in the rest of the paper.

5.3 Non random sampling

In this section we specifically address the possibility of non random sampling across sectors. We first focus on the mean; then, making use of the techniques discussed in Section 3, we analyse the quantiles adopting a two steps procedure. The first step purges the endogeneity between public sector choice and the earning process, and the second step uses the projection of the sector decision obtained in the first step to estimate the premium, along with a grid search. The grid ranges between -1 and +1 (i.e., 100% penalty/premium) with a step of 0.01. Consistent estimates are obtained based on valid exclusion restrictions, i.e. *i)* restrictions that influence the probability of working in the public sector, and *ii)* restrictions that are uncorrelated with the earning process.

Almost all the existing literature uses only variables related to parents' sector of occupation as instruments for sector sorting, mainly due to data limitations. There are several reasons to control for the parents' sector of occupation, beside the fact that it is uncorrelated with the individual's wage level: *i)* the family background may shape the preferences of an individual to work in one sector or another; *ii)* the opportunity for some workers to leave their own job and be replaced by their children may be relevant, especially in the Italian context; *iii)* the connections built up by parents during their working career may be important channels to enter one sector or the other. Among the studies focussing on Italian data, Brunello - Dustmann (1997) argue that the job of the fathers is key for the job of the children. We investigated also the role played by the job of the mothers, jointly with that of the fathers. The results (available upon request) do not change significantly; however, due to a large number of missing data on the job of the mothers, we decided not to consider this information. On the other hand, Bardasi (1996) argues that public sector workers are more risk averse and value stability more and thus chooses home ownership and the

[9] We do not compare the public sector pay gap for managers along the quantiles because very few of them are in the lower part of the distribution.

presence of children as exclusion restrictions.

Undoubtedly family background, risk aversion and preference for stability are all convincing exclusion restrictions. Thus, we exploit a detailed set of information available in SHIW regarding the father's job, home ownership together with the presence of children (hereafter home-related variables), as well as a direct measure of the attitude to taking financial risks, which was never taken into account in previous studies.[10] Unfortunately while the latter may be a better proxy of risk aversion than home-related variables, it has only been available since 2004 and, due to questionnaire design, only for a smaller subset of respondents. As a consequence, when we use this information the sample size shrinks dramatically.

Furthermore, our dataset allows us to control for two other possible motivations not yet explored by the existing literature, although only for wave 2004. One is the rate of time preference, that may affect people's choice to join one sector or the other. We expect that a lower discount rate increases the probability to be a public sector worker, as deferred salaries, in the form of pension benefits, are in Italy generally more generous in the public sector. The other instrument measures the propensity of the individual to be engaged in active pro-social activity and is meant to capture the importance of a pro-social motivation to work in the public sector.[11] Although controlling for such additional motivations limits our analysis to a single year (2004), the exercise is important for two reasons: on the one hand including other possible motivations allows us to evaluate whether the existing practice to control for family background alone is valid or not in the case of Italy (we *test* whether it is a good assumption or not, rather than *assume* that it is a good assumption); on the other hand we can sort the importance of motivations.

5.3.1 Sorting the sector of activity

In the first step we regress the public sector indicator on exogenous covariates and exclusion restrictions: fathers' sector of occupation, risk attitude measured by both home-related variables and our direct measure of risk aversion, discounting preferences and pro-social vocation. In Table 6 we report the coefficients for the exclusion restrictions estimated using the whole sample: columns "Father" and "Home" present the results obtained when family background or home-related variables are considered one at a time; column "Father & Home" when family background and home-related variables are included together. Similarly, in Table 7 we report the coefficients for the exclusion restrictions for year 2004; we also include columns "Discount", "Risk" and "Social", which present the results for discount rate, risk aversion and pro-social motivation, respectively. The column FHDRS reports the results obtained when all possible exclusion restrictions are accounted for, but since the inclusion of R implies a large loss of observations, we consider FHDS as the most inclusive.

[10] The exact wording of the question for risk aversion is: "When managing your financial investments, would you describe yourself as someone who looks for:" and four possible responses are offered: "VERY HIGH returns, regardless of a HIGH risk of losing part of your capital" to "LOW returns, WITHOUT any RISK of losing your capital", which we recoded in high vs low risk aversion.

[11] The exact wording for the rate of time preference is "Imagine you were told you had won on the lottery the equivalent of your households net annual income. The sum will be paid to you in a year's time. However, if you give up part of the sum you can have the rest immediately", and five possible answers regarding the fraction the respondent is willing to give up are envisaged. The exact wording for pro-social activity is "In the last year, have you taken an active part in gatherings of any of the following groups or associations: associations/groups involved in social, environmental, union policy, religious, cultural, sports or recreational, professional, or voluntary activities?", and a YES/NO answer.

In what follows, we focus our comments on the exogeneity and relevance of instruments, which are the two essential requisites for the IV technique.[12] As for the former, given that in general we have more exclusion restrictions than endogenous variables, we can test the over-identifying restrictions for the validity of our instruments (J-test; Hansen, 1982). For the relevance of instruments we take advantage of the recent contribution by Stock - Yogo (2005).

When we consider all the years, the test rejects the orthogonality condition for "Home" and "Father & Home", for both men and women. When using "Father" we have one endogenous variable and one instrument and, thus, we cannot test for the over-identifying restrictions (Hansen, 1982); nevertheless, to test how good family background is as an instrument, we have computed the J-test for the case in which the mother's sector of occupation is also considered. The result of the J-test, not shown to avoid confusion, supports the idea that family background is a good set of instruments because it meets the orthogonality condition (the J-test for women is 1.346 with a P-value of 0.246, while for men is 1.797 with a P-value of 0.180).

To evaluate the relevance of our instruments we use the F-statistic as proposed by Stock - Yogo (2005). As a rule of thumb, with one endogenous variable the F-statistic for joint significance of exclusion restrictions is above 10 if instruments are relevant, below if they are weak. The F-statistic for "Father" is always the highest and well above the threshold. "Home" on its own is a weak indicator for sorting in the public sector for men, as the test is below 10. However, when we consider both "Father" and "Home" the test for the joint significance of the instruments is always larger than 10.

The results clearly show that "Father" is an appropriate instrument: it satisfies all the properties required by IV technique for consistent estimation. The coefficients reported in Table 6 suggest that having a father in the public sector significantly increases the probability of working in the same sector for both men and women.

Finally, in order to evaluate whether adding the other instruments improves the efficiency of the estimator, we perform the redundancy test (Breusch - Qian - Schmidt - Wyhowski, 1999), which basically consists of the comparison of the variances of the estimators based on different sets of instruments: a particular set of instruments is redundant if its inclusion does not reduce the variance of the estimator. The entries in Table 6 are calculated exclusively on the column "Father & Home", because "the potential ambiguity [in ascertain redundant instruments] encountered in a start small approach does not occur in a start big approach" (Breusch - Qian - Schmidt - Wyhowski, 1999, p. 99). Thus, in the column "Father" the redundancy of the family background instrument based on estimates from the column "Father & Home" is tested; in the column "Home" the redundancy of home ownership and children based on estimates from the column "Father & Home" is tested. Neither "Father", nor "Home" are redundant, with the statistic from "Father" much larger than that from "Home".

The redundancy test confirms that "Father" is a valid instrument. At the same time, home-related variables, although non redundant, fail to pass the other tests (exogeneity and relevance).

We conclude the section with a focus on the year 2004, which aims at evaluating whether additional information can be exploited using other instruments. Indeed, for year 2004 we can control for a larger set of motivations. Except for FHDRS for women and FHDS for men, all our instruments are valid as they meet the requisite for consistent estimates. Focussing on the additional

[12] We performed these analyses on conditional mean, as this theory for quantiles is not yet established and extending it is beyond the scope of the paper. Notice however that this first step is common to both IV and IVQR techniques.

exclusion restrictions only, the discount rate variables (in the "Discount" column) do not have much explanatory power (individually they are not significant not even at the 10% confidence level) and come with the wrong sign. The risk attitude when investing in financial markets (col. "Risk") is important for men but not for women. Although in this case the sample size is reduced due to questionnaire design, we will borrow from this evidence to suggest a possible interpretation of the results that we obtain under the hypothesis of endogeneity. Finally, pro-social motivation (in the "Social" column) is significant for women at the 5% confidence level, but largely insignificant for men. Moreover, for women home-related and pro-social instruments are not redundant at the 10% confidence level (note however that the size of the sample including the pro-social variable is small). For men, the null hypothesis of redundancy is strongly rejected for "Father", "Father & Home" and "Risk", whereas "Discount" and "Social" are redundant other than weak. By and large these robustness checks support the standard practice of using only the family background as exclusion restrictions. Interestingly, however, pro-social indicators appear to increase the probability to be employed in the public sector for women only, whereas risk aversion helps in explaining the sector choice only for men.

Summing up, according to our tests, the relevant set of instruments is confirmed to be family background as proxied by the sector of activity of the father: leaving aside this restriction brings to weak identification and prevents all the asymptotic properties for a correct identification to be matched.

For these reasons we now study the public-private sector pay gap focussing on the family background ("Father") as our instrument.

5.4 The pay gap

Based on this background we can now study the public sector pay gap. In Table 8 we report the results using "Father" and "Father & Home" to instrument the worker's sector of activity. We investigate averages as well as all quantiles as we did for the exogenous case. In what follows, we first present the results and then we comment on them. For women the conditional pay differential is on average about 40% when we consider "Father" and 60% when we consider "Father & Home". For men, the conditional pay gap is 34% when we instrument the sector of activity using "Father" and 39% when we consider "Father & Home": in both cases it is lower than that estimated for the sample of women. Also over quantiles, notwithstanding the lack of precision at the very end of the wage distribution,[13] the differentials in favour of the public sector are larger when we control for additional motivations than under the exogenous case. When using "Father", for both women and men the pay gap is larger by approximately 20 percentage points up to the median, and by something between 30 and 45 percentage points between the 60th and the 80th quantile. The same pattern along the wage distribution, shifted up, is found when using "Father & Home" together.

The main message that we can draw from this analysis is that, whatever the set of instruments, the assumption of random sampling tends to underestimate the true differential, both at averages and over quantiles, suggesting that the results obtained assuming exogeneity of sector choice must be viewed with caution. More formally, based on the classical Hausman (1978) test, the coefficients estimated under the assumption of exogeneity are found to be statistically different from those obtained under non random sampling (at the standard confidence level).

[13] Larger standard errors derive from the presence of the density function in the variance formula, which tends to make the objective function less precise in the regions of low density (Koenker, 2005, p. 72).

As an attempt to further understand these results, we investigate where, along the wage distribution, the endogeneity bias exerts its main power. Following Angrist - Chernozhukov - Fernandez-Val (2006), the bias of the QR estimates due to omitted variables, like family background in our case, depends on the (weighted) correlation between the public sector choice and a remainder term, which incorporates family background. As can be seen from a comparison of the estimated coefficients, with and without correction for sector choice endogeneity, in general the bias increases from lower to higher quantiles, suggesting that, other things being equal, the sorting process is more important at high quantiles than at low quantiles, i.e. precisely at those quantiles where previous studies found lower or negative differentials. To provide an overall picture of the bias, in Figure 4 we plot the differences between the coefficients obtained controlling and without controlling for sorting. The bias for men is similar in size to that for women, and clearly increases along the quantiles, except at the very right end.

The size of the correction and its increasing pattern along the wage distribution raise two questions. Is the size of the correction and, consequently, of the differential credible? What is the source of its increasing pattern?[14] Obviously the robustness of coefficient estimates to different sets of instruments is a strength for the sample of men; on the other hand, the variability of the coefficients for women suggests that the estimates for this group have to be viewed with caution. Indeed, for women other forms of sorting in addition to the private/public sector decision might be relevant, in particular the preliminary decision to enter the work force or not. However, taking this into account requires a different model, as our results are valid conditional on individuals being employed. The literature either removed this problem by focussing on men, as in Dustmann - Van Soest (1998) who raised the issue explicitly, or completely ignored it. This said, even a wage differential in the order of 30–40% as estimated for the sample of men may be too large. However, in the following section we will present the estimates based on different measures of job compensation and firm sizes that, on the one hand, confirm as a general result the underestimation bias under the assumption of exogeneity and, on the other hand, point to lower conditional pay differentials.

As for the second question, we suggest the following interpretation for our findings. At the low quantiles of the earning distribution what really matters is the monetary motivation, so that the bias for not controlling for the sorting mechanism is relatively small. In contrast, at the high earning quantiles other factors, related to family background, also matter and the bias from not controlling for them leads to quite imprecise conclusions. A possible explanation for these findings might hinge on the results presented in Table 7 that suggest that risk aversion for men and pro-social attitude for women matter for sorting in the public sector. In particular, we found that for men the risk aversion motivation is more important at the higher wage levels where one is more likely to have to take decisions, particularly in the private sector (the results are available upon request). Such a higher responsibility might explain the larger correction at the higher quantiles for men. Also for women the impact of pro-social propensity on the choice of sector increases with the wage level. However, in this case, we could not find any convincing argument to explain the evidence. Of course, given that our analysis is based on the variable "Father" as the only instrument for sector

[14] Although there is evidence for the US that blue collar workers are willing to queue to obtain public sector jobs, whereas highly skilled workers are hard to recruit and retain in the public sector (Katz - Krueger, 1991), for Italy we are not aware of any quantitative analysis exploring this issue. However, informal observation seems to suggest that workers are willing to queue to obtain public sector jobs above all in Southern Italy, where educational choices at graduate level also appear to be influenced, among other things, by their relative ability to yield a position in the public sector.

choice, this interpretation holds as long as family background, by shaping individual preferences, is also able to capture risk aversion and pro-social attitudes.

Unfortunately, we cannot investigate the differential by geographical area and/or by job position with the IVQR technique. Indeed this technique is appropriate only when the number of endogenous variables is one or two at most, and the alternative solution of splitting the sample would substantially reduce the number of observations.

5.5 Robustness checks

Our previous results are persuasive, but can be questioned on several counts. First of all, the economic theory beyond the Mincerian equation uses experience, not age, as proxy for the human capital accumulation. We should thus also control for experience instead of age in our estimation of the wage equation. Second, public sector wage differentials may only be found compared with small and medium enterprises (SMEs), that are more likely to adopt bad practices in violation of contractual rules, such as gender discrimination and use of undeclared employment. This problem may be addressed by controlling for firm size. Third, we have performed our analysis using the hourly wage as a measure of workers' compensation, while one can argue that, as long as the hours worked are not a choice variable of the worker, monthly earnings better represents the remuneration that is possibly more important for the employees.

Although our choices are common in the existing literature, in this section we investigate the robustness of our previous conclusions to these potential drawbacks. In the spirit of a robustness check, in Table 9 we present only the results of the analysis on the means (those obtained for quantiles are available upon request).

The upper panel of the table is built under the assumption of exogenous sector decision, whereas the lower panel reports the results obtained when this assumption is removed and the sector of activity of the fathers is used as the instrument for the public sector dummy.

As is clear from the first columns (denoted by "Exp."), for both women and men using experience instead of age as proxy for the human capital accumulation does not change the public sector differential significantly, reflecting the fact that labour experience and age are highly correlated. Under exogeneity there exists a positive wage differential for public sector workers of about 13.6% for women and 3.9% for men (14.1% for women and 4.2% for men, respectively, using age); when we allow for sector sorting, the differential increases by about 25 percentage points for women and men, to 0.38 and 0.31, respectively (the coefficient is 0.39 for women and 0.34 for men using age).

The other two issues are somewhat peculiar to the Italian case. To address the first issue, related to firm size, we have included in our specification two sector indicators: one for small private firms, one for large private firms. The coefficient of the former (latter) will measure the pay gap between small (large) firms and the public sector, thus a negative coefficient will be estimated if the pay gap is in favour of the public sector workers.[15] We have defined the firm as large if it employs more than 100 workers. The results reported in the second column ("Big Ent.") are

[15] If the public sector is more directly comparable to large private firms than to the private sector as a whole, another approach to estimate the pay gap is to split the sample. Using this approach, under random sampling we find no (significant) differential for women and a significative *negative* differential for men, thus supporting the insight that our results are in some sense "driven" by the SMEs. However, when we correct for possible sorting, for women the point estimate (0.35) is smaller but close to the benchmark presented in Section 5.4 (0.39), although it is estimated rather imprecisely; for men we obtain a lower but still largely significant differential in favour of the public sector workers (the point estimate is now 0.23).

interesting. Indeed workers of large private firms are better off than workers of small private firms in all cases, but public sector workers are better off than both of them in the sample of women, under the assumption of exogeneity. When we control for possible sorting, for women the pay gap in favour of the public sector becomes larger, irrespective of firm size, although the magnitude of the correction is smaller than in the benchmark case. When we consider men, the correction makes the pay gap larger, only if compared with small firms, whereas men working in large private firms are better off than those working in the public sector. However if we allow for a more flexible model specification, where each coefficient is interacted with the firm size dummies, the pay gap for men is again estimated in favour of public sector workers, irrespective of whether they are compared with small or large firms' workers. As a consequence, the evidence for men working in large firms is not clear cut.

Finally, in the last columns of Table 9 we report the results obtained when monthly, instead of hourly, wage is our dependent variable controlling and not controlling for the number of hours usually worked per week ("Monthly" and "Monthly - No Hours", respectively). In both cases, all estimated differentials are lower than in our benchmark specifications. While for women the gap always remains always positive and significant, for men a differential in favour of private sector workers emerges under exogeneity. When we consider the possible sorting in the employment sector, all our qualitative previous conclusions are re-established. As expected, the differential in favour of the public sector is larger when we control for the number of hours worked per week.[16]

6 Conclusions and final remarks

In this paper we studied the public private sector pay gap in Italy in the period 1998–2008, using micro data from the Bank of Italy Survey of Household Income and Wealth. Compared to previous studies on Italy that control for relevant characteristics of employment, this paper provided an updated analysis that evaluates the wage differential in the last decade.

Unlike most of the existing literature that focuses on the average differential, we analysed the entire distribution of (log) wage, by using conditional quantile regression techniques, and properly considered the possible endogeneity of the sector choice. Because the performance and the validity of IV estimators crucially depends on the exogeneity and relevance of instruments, we investigated several possible motivations invoked in the literature as key to enter the public sector. By exploiting a rich set of information provided in the 2004 wave of the Survey, we were able to discuss whether certain characteristics of employees, such as family background, risk aversion, degree of forward-lookingness and preferences for pro-social activities increase the probability that they will seek employment in the public sector. As a measure of risk aversion we considered the presence of children and home ownership, as in the existing literature, as well as the attitude of the respondent to taking financial risks; to assess the degree of forward-lookingness or impatience we relied on direct measures of discounting; as an indicator of altruistic vocation we controlled for the engagement of the respondent into pro-social voluntary activities; finally, as for family background we accounted for the parents' sector of occupation.

[16]The issue here is the well-known omitted variable bias: if the true model is $y = x\beta + z\gamma + u$, but we estimate $y = x\beta + u$, then $\hat{\beta} = \beta + (x'x)^{-1}(x'z\gamma)$. Because employees in the public sector work, on average, less hours than in the private sector (less than 35 as opposed to 39 hours per week in the main job), then $x'z < 0$, hence the estimated coefficient will be smaller than the true coefficient (with $\gamma > 0$).

In fact, sample selection appears to play a key role. On the one hand, under the assumption of random sector choice, quantile and mean regressions pointed to a pay differential in favour of the public sector workers averaging at about 14 per cent for women and 4 per cent for men. At the bottom quantiles the gap is slightly higher for women than for men; then it decreases substantially along the conditional wage distribution for men, whereas it remains relatively constant for women, so that the gender gap increases towards higher quantiles. The difference is generally higher for white collar workers (while managers in the public sector seem to be at an earning disadvantage with respect to those in the private sector), and in Southern Italy.

On the other hand, when we controlled for possible endogeneity of the sector choice we obtained substantially higher gaps, suggesting that additional motives, other than monetary ones, may induce people to seek employment in the public sector and that ignoring them may result in an underestimation of the overall advantage. Such additional motivations, which for men could be related to different propensities across sectors to take risks, appeared to be particularly significant above the median of the wage distribution, precisely where previous studies used to find lower premia or even penalties.

These results are robust to a different model specification (i.e., using labour experience instead of age as the proxy for human capital accumulation) and when considering monthly wages as opposed to hourly wages. Instead, when we restricted our comparison to large firms only, a pay differential in favour of public sector workers was confirmed for the sample of women, whereas for men no clear-cut evidence was found.

Using up-to-date techniques, we have documented the existence of a significant public sector pay gap in Italy. However, drawing from the analysis specific policy implications, such as differentiated pay by geographical area or by gender or horizontal wage cuts in the public sector, is far beyond our intention.

Moreover, we have said little or nothing about causes and consequences of such a differential. Nonetheless, understanding whether the public-private pay gap simply depends on institutional features, or rather reflects the true preferences of the policy-makers, and what its implications are for the quality of public services as well as for educational choices and labour market developments is of crucial importance. Thus, we take the analysis in this paper as a necessary intermediate step and leave the other questions for future research.

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Table 1: The evolution of public sector over time. Relative frequencies

Area	1998	2000	2002	2004	2006	2008
			Women			
NW	0.370	0.354	0.365	0.385	0.349	0.328
NE	0.340	0.314	0.302	0.333	0.320	0.311
C	0.456	0.397	0.422	0.313	0.398	0.368
S	0.572	0.537	0.489	0.491	0.472	0.403
IS	0.674	0.547	0.601	0.515	0.599	0.460
ITALY	0.432	0.394	0.397	0.379	0.388	0.354
			Men			
NW	0.207	0.190	0.182	0.207	0.186	0.157
NE	0.178	0.184	0.180	0.159	0.127	0.136
C	0.357	0.272	0.318	0.239	0.250	0.231
S	0.397	0.359	0.331	0.324	0.352	0.271
IS	0.313	0.287	0.316	0.321	0.351	0.320
ITALY	0.283	0.251	0.255	0.239	0.237	0.209
			Total			
ITALY	0.344	0.309	0.314	0.297	0.302	0.273

Table 2: Descriptive statistics for continuous variables, by Area, Gender and Sector.

Variable	Area	Mean	SD	p25	p50	p75	Mean	SD	p25	p50	p75
Women											
		Private					Public				
Ln(Wage)	NW	1.969	0.404	1.757	1.945	2.153	2.245	0.415	1.980	2.194	2.456
	NE	1.946	0.378	1.731	1.924	2.120	2.240	0.412	1.987	2.181	2.444
	C	1.912	0.427	1.683	1.889	2.120	2.269	0.405	2.011	2.227	2.504
	S	1.691	0.543	1.395	1.689	1.970	2.238	0.499	1.984	2.252	2.519
	IS	1.818	0.450	1.581	1.833	2.092	2.228	0.470	1.958	2.181	2.510
	ITALY	1.910	0.432	1.683	1.913	2.120	2.246	0.435	1.987	2.200	2.477
Age	NW	37.7	9.9	30.0	38.0	45.0	42.0	8.5	36.0	42.0	48.0
	NE	36.7	9.9	29.0	36.0	44.0	40.6	8.7	34.0	41.0	47.0
	C	37.9	9.8	30.0	38.0	44.0	43.4	9.4	37.0	42.0	51.0
	S	37.2	10.6	29.0	37.0	45.0	44.2	9.2	38.0	44.0	51.0
	IS	37.5	10.4	29.0	37.0	45.0	43.2	9.2	36.0	43.0	50.0
	ITALY	37.4	10.0	29.0	37.0	45.0	42.5	9.0	36.0	42.0	49.0
Men											
		Private					Public				
Ln(Wage)	NW	2.116	0.406	1.873	2.063	2.343	2.345	0.430	2.067	2.288	2.558
	NE	2.064	0.375	1.864	2.046	2.238	2.356	0.439	2.075	2.280	2.560
	C	2.071	0.436	1.833	2.056	2.270	2.265	0.438	2.005	2.200	2.506
	S	1.900	0.475	1.650	1.913	2.161	2.214	0.434	1.970	2.189	2.397
	IS	1.852	0.505	1.581	1.865	2.120	2.285	0.389	2.081	2.238	2.449
	ITALY	2.031	0.438	1.808	2.019	2.238	2.282	0.432	2.019	2.226	2.498
Age	NW	38.8	10.2	31.0	39.0	46.0	43.8	9.2	37.0	45.0	51.0
	NE	37.0	10.1	29.0	37.0	44.0	43.0	8.7	37.0	43.0	50.0
	C	39.6	10.8	31.0	39.0	48.0	43.9	10.0	37.0	44.0	51.0
	S	39.1	11.2	30.0	39.0	48.0	45.6	9.8	40.0	46.0	53.0
	IS	38.2	11.4	29.0	38.0	47.0	45.8	9.4	39.0	46.0	52.0
	ITALY	38.5	10.6	30.0	38.0	46.0	44.5	9.6	38.0	45.0	52.0

Table 3: Descriptive statistics for dichotomous variables, by Area, Gender and Sector.

Area	Sector	Married	Education			Positions			Partime	Obs.
			Low	Int.	High	Blue	White	Man.		
						Women				
NW	Private	0.583	0.331	0.511	0.090	0.455	0.539	0.006	0.198	2515
	Public	0.714	0.139	0.529	0.310	0.170	0.790	0.041	0.085	1516
NE	Private	0.560	0.332	0.541	0.064	0.521	0.476	0.003	0.263	2549
	Public	0.669	0.118	0.580	0.285	0.143	0.833	0.024	0.161	1286
C	Private	0.544	0.374	0.488	0.080	0.538	0.456	0.006	0.235	1988
	Public	0.629	0.112	0.525	0.347	0.128	0.833	0.039	0.061	1221
S	Private	0.501	0.333	0.382	0.086	0.654	0.343	0.003	0.279	1112
	Public	0.668	0.089	0.533	0.338	0.062	0.918	0.020	0.061	1215
IS	Private	0.476	0.363	0.451	0.051	0.599	0.399	0.003	0.328	573
	Public	0.637	0.149	0.549	0.284	0.100	0.883	0.017	0.119	633
ITALY	Private	0.553	0.342	0.495	0.078	0.523	0.472	0.005	0.241	8737
	Public	0.670	0.120	0.542	0.315	0.128	0.842	0.031	0.095	5871
						Men				
NW	Private	0.599	0.377	0.442	0.100	0.608	0.365	0.027	0.016	3864
	Public	0.758	0.240	0.426	0.314	0.140	0.751	0.109	0.019	951
NE	Private	0.581	0.408	0.477	0.057	0.689	0.294	0.017	0.017	3644
	Public	0.707	0.194	0.514	0.274	0.142	0.771	0.087	0.036	778
C	Private	0.595	0.402	0.443	0.082	0.620	0.345	0.035	0.027	3034
	Public	0.740	0.282	0.473	0.219	0.172	0.764	0.064	0.019	1113
S	Private	0.664	0.443	0.351	0.032	0.775	0.217	0.008	0.059	3072
	Public	0.806	0.267	0.505	0.175	0.144	0.812	0.044	0.033	1539
IS	Private	0.653	0.503	0.257	0.027	0.770	0.220	0.010	0.076	1554
	Public	0.865	0.299	0.473	0.197	0.143	0.808	0.049	0.046	748
ITALY	Private	0.611	0.413	0.416	0.067	0.676	0.303	0.021	0.032	15168
	Public	0.776	0.259	0.478	0.231	0.149	0.782	0.069	0.029	5129

Table 4: Log of hourly wage by Area, Gender, Sector and Position.

Area	Sector	Blue	White	Man.	Blue	White	Man.
			Women			Men	
NW	Private	1.838	2.039	2.391	1.987	2.190	2.647
	Public	1.957	2.277	2.560	2.053	2.305	2.675
NE	Private	1.836	2.045	2.298	1.969	2.195	2.549
	Public	2.002	2.267	2.405	2.116	2.331	2.629
C	Private	1.799	2.024	2.211	1.944	2.166	2.622
	Public	1.993	2.301	2.365	2.076	2.211	2.561
S	Private	1.598	1.821	2.178	1.831	2.058	2.503
	Public	1.805	2.265	2.298	1.976	2.222	2.444
IS	Private	1.770	1.872	2.148	1.772	2.001	2.591
	Public	1.939	2.253	2.381	2.061	2.277	2.616
ITALY	Private	1.790	2.011	2.305	1.920	2.154	2.605
	Public	1.960	2.275	2.419	2.047	2.260	2.580

Table 5: Conditional Quantile Regression for log of hourly wage. Coefficient of "Public" by Gender, Area and Position.

Women (Obs.: 14,378)

	10th	20th	30th	40th	50th	60th	70th	80th	90th	Mean
All Sample										
Public	0.167 ***	0.132 ***	0.117 ***	0.125 ***	0.122 ***	0.129 ***	0.134 ***	0.121 ***	0.121 ***	0.141 ***
R2	0.218	0.220	0.220	0.226	0.231	0.236	0.241	0.241	0.226	0.331
By Area										
NW	0.135 ***	0.088 ***	0.082 ***	0.085 ***	0.085 ***	0.077 ***	0.080 ***	0.081 ***	0.075 ***	0.100 ***
NE	0.144 ***	0.077 ***	0.071 ***	0.075 ***	0.074 ***	0.074 ***	0.086 ***	0.106 ***	0.136 ***	0.102 ***
C	0.188 ***	0.157 ***	0.147 ***	0.143 ***	0.149 ***	0.164 ***	0.167 ***	0.115 ***	0.100 ***	0.147 ***
S	0.473 ***	0.366 ***	0.343 ***	0.315 ***	0.294 ***	0.264 ***	0.251 ***	0.267 ***	0.215 ***	0.286 ***
IS	0.213 ***	0.209 ***	0.200 ***	0.179 ***	0.164 ***	0.172 ***	0.181 ***	0.124 ***	0.186 ***	0.182 ***
R2	0.223	0.228	0.228	0.232	0.237	0.240	0.245	0.243	0.228	0.336
By Position										
Blue	0.174 ***	0.099 ***	0.085 ***	0.078 ***	0.074 ***	0.071 ***	0.064 ***	0.052 **	0.067 **	0.092 ***
White	0.165 ***	0.144 ***	0.127 ***	0.145 ***	0.143 ***	0.152 ***	0.163 ***	0.147 ***	0.153 ***	0.161 ***
R2	0.218	0.220	0.221	0.227	0.233	0.238	0.243	0.243	0.228	0.333

Men (Obs.: 20,040)

	10th	20th	30th	40th	50th	60th	70th	80th	90th	Mean
All Sample										
Public	0.113 ***	0.070 ***	0.058 ***	0.045 ***	0.042 ***	0.025 ***	0.013	0.007	0.005	0.042 ***
R2	0.242	0.248	0.245	0.249	0.251	0.257	0.266	0.276	0.281	0.389
By Area										
NW	0.086 ***	0.029 *	0.031 *	0.023 *	0.007	0.001	-0.018	-0.048 ***	-0.052 **	-0.007
NE	0.045	0.006	0.017	0.017	0.036 **	0.028	0.027	0.071 ***	0.061 *	0.043 ***
C	-0.004	0.031 *	0.016	0.019	0.020	0.026	0.032 **	0.029	0.034	0.021
S	0.230 ***	0.147 ***	0.113 ***	0.075 ***	0.055 ***	0.015	0.000	-0.005	-0.019	0.068 ***
IS	0.473 ***	0.260 ***	0.182 ***	0.135 ***	0.114 ***	0.098 ***	0.059 ***	0.020	0.025	0.129 ***
R2	0.252	0.253	0.248	0.250	0.252	0.257	0.267	0.277	0.282	0.390
By Position										
Blue	0.156 ***	0.093 ***	0.087 ***	0.073 ***	0.063 ***	0.054 ***	0.045 **	0.021	0.034	0.077 ***
White	0.095 ***	0.068 ***	0.056 ***	0.041 ***	0.043 ***	0.022 **	0.011	0.002	0.011	0.044 ***
R2	0.242	0.249	0.246	0.250	0.252	0.258	0.267	0.277	0.282	0.390

Note: *** is 1% significance level; ** is 5% significance level; * is 10% significance level.

Table 6: Statistics for relevance and exogeneity of instruments from first step of IV regression by Gender. All years. Only coefficients attached to exclusion restrictions.

	Women			Men		
	Father	Home	Father & Home	Father	Home	Father & Home
Father pub.	0.100 ***		0.099 ***	0.085 ***		0.084 ***
Non-owner		0.002	0.006		0.010	0.010
Children		0.041 ***	0.039 ***		0.020 ***	0.021 ***
Obs.	11429	14378	11429	16668	20040	16668
R2	0.305	0.290	0.308	0.272	0.271	0.273
Joint sig.	39.2 ***	15.9 ***	21.0 ***	41.9 ***	8.6 ***	19.4 ***
J test		7.0 ***	10.8 ***		20.2 ***	16.5 ***
Redundancy	34.3 ***	23.5 ***		40.9 ***	14.5 ***	

Note: *** is 1% significance level; ** is 5% significance level; * is 10% significance level.

Table 7: Statistics for relevance and exogeneity of instruments from first step of IV regression by Gender. Year 2004. Only coefficients attached to exclusion restrictions.

Instrument	Father	Home	Father & Home	Discount	Risk	Social	FHDS[a]	FHDRS[a]
				2004 – Women				
Father pub.	0.111 ***		0.108 ***				0.058	0.003
Mother pub.	-0.019		-0.019				-0.061	-0.169
Non-owner		0.002	-0.015				-0.118 **	-0.005
Children		0.016	0.034 **				0.058	0.120 *
High disc.				0.018			-0.093	-0.343 **
Lower disc.				0.040			-0.104	-0.200 *
Risk av.					-0.002			0.074
Pro-social						0.137 **	0.121	0.178
Obs.	1396	2417	1396	808	291	380	319	134
RMSE	0.422	0.420	0.422	0.426	0.438	0.437	0.448	0.441
R2	0.229	0.252	0.230	0.245	0.232	0.210	0.187	0.226
J test	0.079	0.054	0.932	0.000	.	.	4.382	14.402 **
Joint sig.	5.287 ***	0.784	3.677 ***	0.590	0.001	3.906 **	1.456	2.000 *
Redundancy	0.748	5.859 *	6.266	2.943	0.835	2.839 *		
				2004 – Men				
Father pub.	0.132 ***		0.132 ***				0.149 ***	0.253 ***
Mother pub.	0.049		0.049				0.146 **	0.178 *
Non-owner		-0.001	-0.003				-0.014	-0.015
Children		0.014	-0.003				-0.014	0.047
High disc.				0.038			-0.017	0.132 *
Lower disc.				-0.025			-0.036	-0.047
Risk av.					0.069 **			0.099 *
Pro-social						-0.016	-0.047	-0.114 *
Obs.	2355	3366	2355	1856	653	851	712	260
RMSE	0.365	0.366	0.365	0.392	0.396	0.373	0.376	0.378
R2	0.280	0.265	0.280	0.266	0.269	0.341	0.341	0.373
J test	1.272	1.322	2.153	0.725	.	.	14.157 ***	5.875
Joint sig.	19.682 ***	1.187	9.857 ***	3.815 **	4.408 **	0.178	3.296 ***	3.950 ***
Redundancy	20.131 ***	0.707	21.555 ***	1.081	3.987 **	1.372		

Note: *** is 1% significance level; ** is 5% significance level; * is 10% significance level.
[a] FHDS: Family & Home & Discount & Social; FHDRS: Family & Home & Discount & Risk & Social.

Table 8: Instrumental Variable Quantile Regression for log of hourly wage. Coefficient of "Public" by Gender.

	10th	20th	30th	40th	50th	60th	70th	80th	90th	Mean
Women (Obs.: 11,429)										
Father	0.43 **	0.17 *	0.32 **	0.31 ***	0.30 ***	0.41 *	0.50 ***	0.57 ***	0.24	0.39 ***
Father & Home	0.76 *	0.42 ***	0.45 ***	0.44 ***	0.46 ***	0.62 ***	0.64 ***	0.79 ***	0.78	0.61 ***
Men (Obs.: 16,668)										
Father	0.11	0.24 ***	0.21 **	0.18 *	0.25 ***	0.31 ***	0.33 ***	0.45 **	0.95	0.34 ***
Father & Home	0.10	0.25 ***	0.25 ***	0.33 ***	0.39 ***	0.43 ***	0.49 ***	0.82 *	0.86	0.39 ***

Note: *** is 1% significance level; ** is 5% significance level; * is 10% significance level.

Table 9: Robustness checks, using various definitions and/or various model specifications to explain the pay gap. Coefficients of “Public”, “Small - Private” and “Large - Private” by Gender.

	Women				Men			
	Exp.	Big Ent.	Monthly	Monthly No Hours	Exp.	Big Ent.	Monthly	Monthly No Hours
	Exogenous							
Public	0.136 ***		0.072 ***	0.018 **	0.039 ***		-0.033 ***	-0.061 ***
Small		-0.172 ***				-0.079 ***		
Large		-0.041 ***				0.033 ***		
Obs.	14378	14327	14256	14363	20040	19947	19865	19972
	Endogenous							
Public	0.380 ***		0.253 **	0.181 *	0.309 ***		0.221 **	0.193 **
Small		-0.203 ***				-0.100 ***		
Large		-0.074 **				0.044 **		
Obs.	11429	11394	11331	11416	16668	16596	16528	16611

Note: *** is 1% significance level; ** is 5% significance level; * is 10% significance level.

Figure 1: The wage gap in National Account data

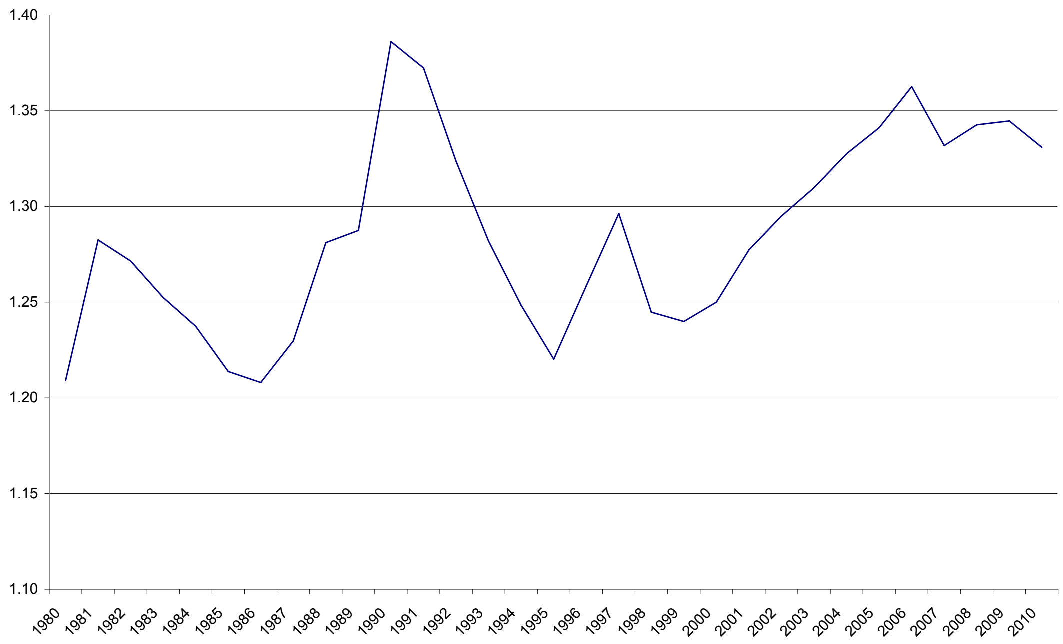
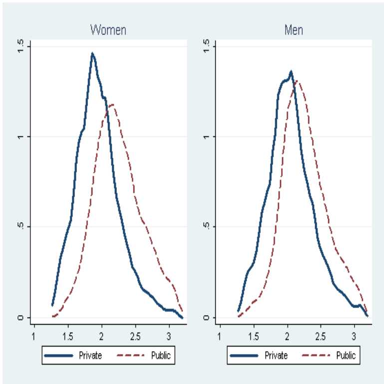
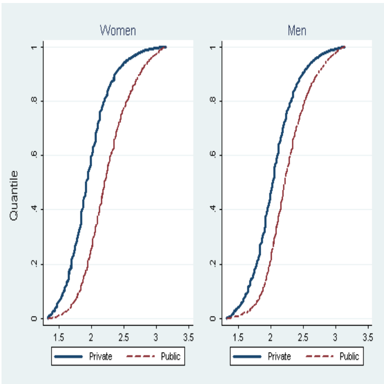


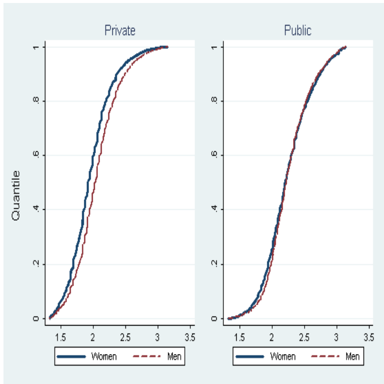
Figure 2: Distribution function for (log of) hourly wage, by Gender and Sector. Kernel density (a) and Cumulative distribution functions (b and c).



a)

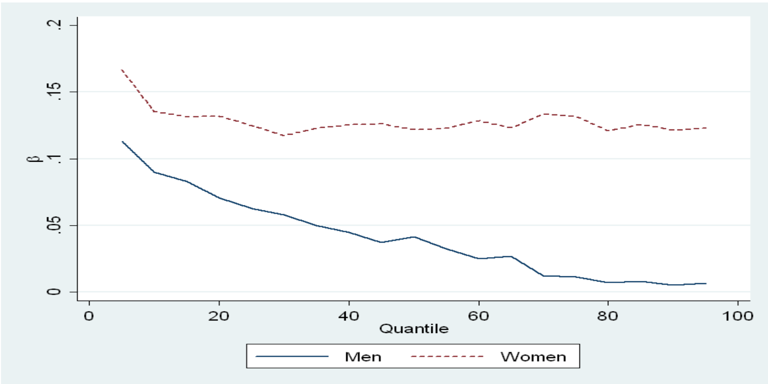


b)

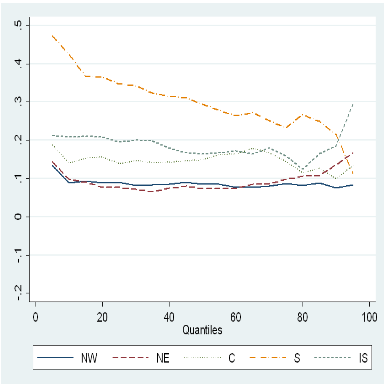


c)

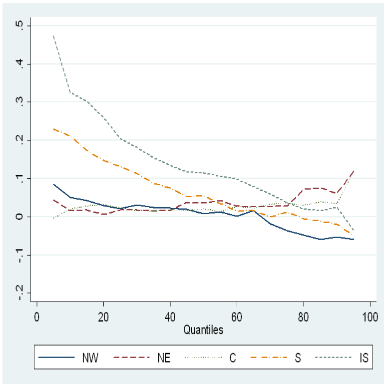
Figure 3: Coefficient of PUBLIC across quantiles. Overall estimates (a), estimates by geographical area (b and c) and by job position (d and e).



a)



b)



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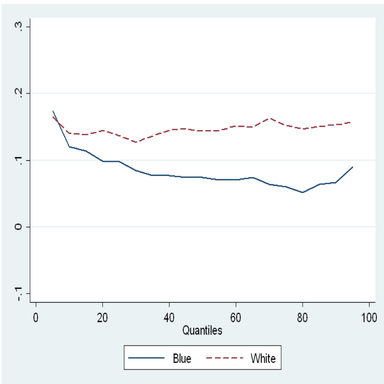
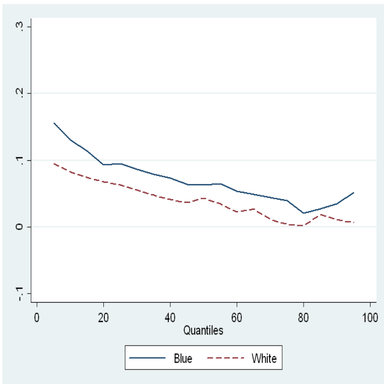
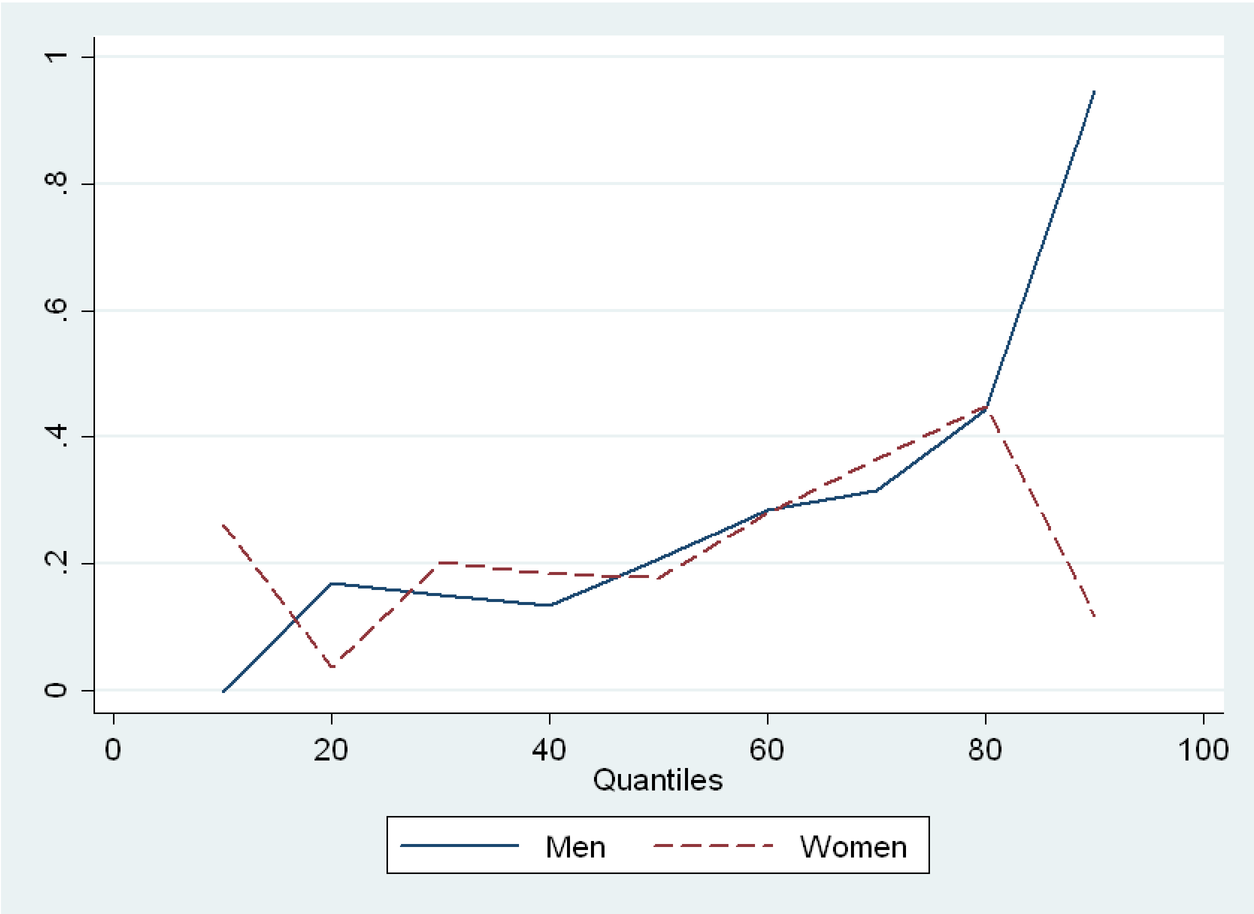


d)



e)

Figure 4: The bias averaged by year, by quantiles and gender.



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