THE EVOLUTION OF WAGE INEQUALITY IN THE PUBLIC SECTOR: EVIDENCE FROM ITALY

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Abstract

Empirical literature shows that wage inequality in Italy has been quite stable over the last decade. In this paper we argue that this is mainly due to the fact that public and private sector have always been considered jointly. Actually, carrying out the analysis only for the public sector the pattern of wage inequality (the 90/10 ratio) increases over the period between 1993 and 2006. Further, applying a quantile decomposition procedure we find out that the increase of wage inequality in the public sector is driven by an increase of the residual component in the upper tail of the wage distribution, which we relate to the reforms in the wage setting system introduced in 1993 and 1998.

Keywords: Inequality, Quantile Decompositon, Public Sector, Italy.

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1. **Introduction**

The analysis of changes in the distribution of wages has been an active research area in labour economics over the last two decades, especially because of the steep increase of wage inequality and schooling premia in United States and other Anglo-Saxon countries since the early 1980s (Bound and Johnson, 1992; Katz and Murphy 1992). To a less extent, increases of the educational wage premia (EWP, henceforth) and wage inequality are also documented for many other OECD countries (Gottschalk and Smeeding, 1997).[1] As for the Italian case, the empirical literature shows that wage inequality has not changed much. Using the SHIW Bank of Italy data Brandolini, Cipollone, Sestito (2002) point out that wage inequality slightly narrowed from 1977 to the end of the 1980s, then slightly increased at the beginning of the nineties, around 1992-1993, due to an important economic crisis in that years, and from 1993 to 2004 it remained quite stable (see also Lilla, 2005).

However, this literature for Italy considers private and public sector together. This paper revisits the changes in the wage structure in Italy showing that inequality trends are no longer stable when we restrict the analysis to the public sector. Using the *Survey of the Household Income and Wealth* (SHIW) of the Bank of Italy for the period 1993-2006, we point out that in the public sector the Gini index increased by 15.6%, from 17 to 20.6. The ratio between the 90 and the 10 percentile increased until 2002 (+15%), while it starts decreasing in 2004. Nonetheless, the 90/10 ratio in the 2006 is still 5.5% greater than the 1993 value

To identify which forces have played a role in explaining these observed patterns we implement a quantile decomposition methodology, developed by Machado and Mata (2005), Melly (2005), and Autor, Katz, Kerney (2005), which decomposes the changes of the wage distribution into changes in covariates, coefficients (between groups), and residual (within-groups) components.

Applying a standard mincerian wage equation, where education, experience and gender are included as covariates, we point out that the increase in wage inequality in the public sector is mainly driven by the residual-within component and, to a less extent, by the coefficients component, while the effect of covariates is negligible.

One possible explanation for these results claims for changes of the institutional setting introduced in the public sector in 1993 and in 1998. This reform process mainly concerned the so-called “privatization” of employment relations, i.e. making job conditions and wage setting more similar to those prevailing in the private sector, and as consequence this may have favoured the increases of residual component of the wage inequality. In particular, a new agency was set up (ARAN) in order to represent public employees in collective bargaining at the national level. A two-level bargaining setting was also introduced, similarly to the private

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♦ This paper has been carried out inside the research project IEIILM, funded by the Italian Ministry of University and Research (MIUR, PRIN research project)), and inside the the research partnership between DSE and ISFOL. We are also grateful to Andrea Brandolini, Lorenzo Cappellari, Piero Cipollone, Bart Cockx, Muriel Dejemeppe, Andrea Ichino, Claudio Lucifora, Blaise Melly, Franco Peracchi, Paolo Piacentini, Henri Sneessens, Stefano Staffolani, Bruno Van der Linden, Daniela Vuri, Robert Waldmann. We thank also the partecipants to the seminars at Brucchi Luchino 2006, LUISS, University of La Sapienza, University of Bologna, University of Bergamo, UCL-London, UCL-Louvain, UCD-Dublin.

[1] Peracchi (2006) distinguishes between returns to education, which is a measure of the causal effect of an extra level of schooling on the worker’s earnings, and educational wage premia, which is a measure of statistical association between levels of schooling and wages. We make use of this terminology in the paper.

sector. The first level of bargaining is at the National level, and determines not only the basic pay for all the different occupations in a given sub-sector, but also the margin for wage increases that have to be respected by local units in the second level of bargaining (the so called "wage drift"). In the new setting local administrative units are free to choose, in a given interval fixed by the first bargaining level, their own pay policy (Dell'Aringa et al. 2007). This means that after the reform the same kind of public employees (same age, occupation, tenure) might be paid differently in different administrative units, or even in the same units if assigned to different tasks.

The paper is organized as follows. In section 2 the characteristics of the sample are analysed and descriptive statistics presented. In section 3 we introduce the quantile decomposition analysis and in section 4 the results are presented and discussed. Section 5 concludes.

2. Data, descriptive statistics, and inequality trends.

The empirical analysis is based on the *Survey of the Household Income and Wealth* (SHIW) of the Bank of Italy, from 1993 to 2006. This database represents the main source used in the literature to investigate inequality issues in Italy. We consider this time period for different reasons. First, because former periods have been widely covered in the literature. Second, because it is a homogenous institutional period, since the reforms in 1992 and 1993 represented a breaking point for the Italian labour market. Nonetheless, as robustness check we replicated our decomposition analysis considering the periods: 1995-2006, 1993-2002, 1993-2004.

The sample consists of employees aged 18-64 in the public sector both men and women. We focus on employees with permanent jobs or with fixed term contracts ("contratti a tempo determinato"), while we do not consider atypical contracts introduced after the reforms in 1998 and 2003 ("collaborazioni coordinate e continuative", "contratti a progetto"), since we cannot identify these kind of contracts in 1993. We refer to the real monthly net wage, obtained by dividing yearly wage from employment (including overtime, bonuses, and fringe benefits), net of taxes and social security contributions, by the number of months worked in that year, and deflating by the consumer price index (base year 2004). We consider full-time equivalent individuals, controlling for differences in working time by taking into account the worked hours of part time workers, as in the section of the private sector.

Table 6 provides the descriptive statistics of the main variables in 1993, 2004, and 2006 for the public, and for comparison, for the private sector.[2] It is worth noting that the share of public sector decreased over time, from 33% in 1993 to 21% in 2006. This strong fall in the public employment share is mainly related to the trend of privatization of Public Companies and to the slowdown in the hiring process.[3] Most of the other figures shown in Table 6 are in

[2] We define a public employee using two variables in the database, APSETT and DIMAZ. APSETT provides us with self-declaration of the sector in which the individual works, including the public sector, while DIMAZ refers to the firm size, and it is specified when the employee declares that he/she is employed in the public sector. We consider as public employees those workers who declare for both questions that they are employed in the public sector. Results do not change much when we consider definitions of public employee based on APSETT and DIMAZ separately.

[3] One might argue that the privatization process might have affected the inequality trends in the private and public sectors. In particular, a composition effect might be at work whether the privatized parts of the public

line with the literature on the differences between public and private sector (Gregory and Borland, 1999, Dell'Aringa, Lucifora, Origo, 1999). First, wages are higher in the public sector, and are increasing over time, while in the private sector wages remain quite stable. In the public sector wages increased across the whole wage distribution: from 1009 to 1043 euros at the 10th percentile, from 1375 to 1411 at the median, and from 1962 to 2141 euros at the 90th percentile. Third, the shares of the labor force with upper secondary education and university degree are considerably higher in the public sector than in the private sector. In the public sector the share of upper secondary and graduated workers increase from 44% to 52% and from 24 to 31%, respectively. Fourth, the average level of labor market experience is higher in the public sector.[4] The average number of years of experience in 2006 is 22.3 while in the private sector is 19.2. Further, in the public sector there has been an impressive increase of the classes 26-30 and 31-35 years of experience and a decrease of individuals with less than 25 years of experience.

As for inequality trends, Brandolini, Cipollone, Sestito (2002), Lilla (2005), Boeri and Brandolini (2004) show that wage inequality in Italy has been quite stable from 1993 to 2002. However, these papers do not separately investigate the patterns of public and private sector, and do not make use of the SHIW last waves for 2004 and 2006. In this section we point out that separating these two sectors actually matters in understanding changes in the wage structure and inequality trends. We then focus on public sector trends.

Figure 3, which reports the time dynamics of the Gini index computed on real term wages, confirms that when we consider the whole set of employees in the Italian labour market inequality remains quite stable over time until 2004, while it starts decreasing in 2006. However, when public and private sectors are split trends change. In particular, wage inequality increased by 15.6% in the public sector, from 17 to 20.6 from 1993 and 2006, while it was pretty stable considering public and private sector together, at least until 2004, and declining in the private sector, as shown in the previous section.

Figure 4 shows the same patterns using another standard index of inequality, the ratio between the 90th and the 10th percentile. As for the public sector, inequality increased until 2002 (+15%), while it start decreasing in 2004. Nonetheless, the 90/10 ratio in 2006 is still 5.5% greater than the 1993 value.

Useful insights can be also derived from the analysis of the changes in the different parts of the wage distribution. In particular, we can investigate the lower (upper) tail of the wage distribution analyzing the 50/10 (90/50) ratio. As for the public sector, Figure 5 clearly shows

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sector had been particular compressed in 1993 (and then up to 2004). Unfortunately, we cannot address this issue directly, since we cannot identify in the SHIW data the different components in the public sector that have been privatized, nor the workers that have been privatized over time. However, we can say something about it. It is well known that the major privatizations took place in the transportation and communication industry, and in the energy sector. We cannot identify the energy sector among the private employees, since in the SHIW data it is included in the manufacture, while we can identify transportation and communication industries. In this sector, the 90/10 ratio in 1993 was equal to 2.7 (remaining stable until 2004), higher than the 90/10 ratio for the public sector (1.96) and slightly lower than the 90/10 in the private sector (2.74). Under the assumption that the parts of the transportation and communication sectors that have been privatized had inequality trends similar to those observed in their counterparts in the private sector, the composition effect should have played in the opposite direction, i.e. increasing inequality in the private and decreasing it in the public.

[4] Labour market experience is defined as the difference between the current age of the worker and the age at which that worker begun her/his labour market career.

that the increase in the 90/10 ratio in completely driven by the upper tail of the wage distribution, i.e. the 90/50 index, while the 50/10 ratio remains basically constant.

3. **Quantile Regression Decomposition**

In this section we disentangle the contribution of labour force characteristics and labour market prices in the dynamics of the Italian wage structure. This literature goes back to the seminal contributions in 1973 by Oaxaca and by Blinder, and has seen great developments over the last three decades. The most recent contribution in this literature is to consider a quantile regression setting, which explores the dynamics of the whole wage distribution. We make use of a methodology that has been recently developed by Machado & Mata (2005), Melly (2005) and Autor et al. (2005), papers that use the same general idea and slightly different techniques in the implementation.

This methodology takes as starting point the two quantile estimations in cross-section, for 1993 and 2006, using a Mincerian standard specification:

$$(1) \qquad \ln w_i^t = X_i^t \beta^t(\theta) + u_{i,\theta}^t$$

where *i=1,…N* is the number of observations in each year t, θ is the quantile being analyzed, u_i is an idiosyncratic error term, and X represents our set of explanatory variable that, according to the Mincerian specification, includes education, experience and gender.[5] As standard in this literature (Koenker and Basset, 1978), $\beta(\theta)$ can be estimated minimizing the following expression:

$$(2) \qquad \min_{\beta}\left[n^{-1}\left(\sum_{i=1}^{n} \rho_\theta (\ln w_i^t - X_i^t \beta) \right) \right]$$

where $\begin{cases} \rho_\theta(u) = \theta u \ \text{ if } u > 0 \\ \rho_\theta(u) = (\theta - 1)u \ \text{ if } u < 0. \end{cases}$

Once having derived the quantile parameters $\beta(\theta)$, this methodology allows to estimate the marginal distribution of wages as function of both X and $\beta(\theta)$, and to derive counterfactual distributions of wages.

3.1 Estimation of the marginal distribution of wages

This methodology is essentially developed in two main parts. In the first part, the conditional quantile distribution is estimated, $Q_\theta(w \mid X_i)$, for all θ given the set of covariates X. More specifically, quantile regression theory has shown that, using a linear specification, the conditional quantile distribution of wages can be defined as:

$$(3) \qquad Q_\theta(w \mid X_i) = X_i \beta(\theta) \quad \text{for all } \theta \in (0,1)$$

[5] Note that choosing this set of covariates leads to a human capital interpretation, which we argue have played a role in the last decades. Obviously, choosing different sets of covariates would shed light on other possible explanations. For instance, for an analysis focusing on a generational approach (age classes, and cohorts) see Rosolia and Torrini (2007).

where X_i is a vector for the set of covariates. For instance, X_i might stand for male graduates with less than 5 years of experience.[6] Basset and Koenker (1982) showed that, under some regularity conditions, the estimated conditional quantile function is a consistent estimator of the population conditional quantile function, uniformly in θ.[7]

It is then possible to use the estimated parameters to simulate the conditional distribution of w given X, using an application of the probability integral transformation theorem: if V is a uniform random variable on [0,1], then $F^{-1}(V)$ has distribution F. In our case, if $\theta_1, \theta_2, .., \theta_j .., \theta_J$ are drawn from a uniform (0,1), the corresponding j estimates of the conditional quantile at X_i, $\hat{w}_i \equiv \left\{X_i \hat{\beta}(\theta_j)\right\}_{j=1}^{J}$, constitute a random sample from the estimated conditional distribution of wages given X_i. Using this procedure, we can estimate the conditional distribution of wages for all the different combination of X.

The second part of the procedure consists in deriving an estimation of the marginal distribution of wages. Following Machado & Mata (2005) and Autor, Katz, Kerney (2005), the marginal density of wages depends upon both the conditional quantile function, $Q_\theta(w \mid X_i) = X_i \hat{\beta}(\theta_j)$ for given X_i and θ_j, and the distribution of the covariates, $g(X)$. In order to derive a random sample from the marginal density of wages, it is possible to multiply the matrix containing random observations (or all the observations) from $g(X)$ times the matrix of $\beta(\theta_j)$, with $j=1,\ldots,J$, in which the different θ_j are randomly chosen from the uniform (0,1) distribution. In this setting, each observation of the resulting matrix, $\hat{w}_i \equiv \left\{X_i \hat{\beta}(\theta_j)\right\}_{\substack{j=1,\ldots,J \\ i=1,\ldots,N}}$, can be considered as drawn from the estimated marginal distribution of wages.

By applying this procedure, it is possible to draw an arbitrarily large random sample from the marginal distribution of wages. Autor, Katz, Kearney (2005) claim that this procedure can be considered as equivalent to numerically integrating the estimated conditional quantile function $\hat{Q}_\theta(w \mid X)$ over the distribution of X and θ, i.e. $\int_X \int_\theta \hat{Q}_\theta(w \mid X) g(X) \partial\theta \, \partial X$, integral that produces a consistent estimator of the marginal distribution of wage, $f(w)$, which can be written as (Melly, 2005, 2006):[8]

$$f(w) = \int_x f(w,x)dx = \int_x \underbrace{f(w \mid x)}_{= \int_\theta Q_\theta(w|X)d\theta} g(x)dx = \int_x \int_\theta Q_\theta(w \mid X) g(x) d\theta dx$$

The insights behind the comparison between the Machado & Mata approach and the integral procedure are quite intuitive. More specifically, any given random observation of X_i is multiplied by all the possible $\beta(\theta)$, with θ ranging from 0 to 1, and this can be considered as the internal integration over the support of θ. Then, X is repeatedly drawn from the whole support $g(X)$, and this can be seen as the external integral in X. Melly (2006) shows that the Machado

[6] In this framework the quantile regression coefficients can be interpreted as rates of return to the different characteristics at the specified quantile of the conditional distribution.

[7] To validate the heteroscedasticity hypothesis, i.e. the fact that 'slope coefficients' are different for the same covariate across quantiles, we successfully test that the estimates of the coefficient vectors at different quantiles are statistically different from one another (Buchinsky, 1995, Koenker and Basset, 1978).

[8] Note that θ is uniformly distributed on the [0,1] interval, implying that the relative density $f(\theta)$ is equal to 1.

and Mata (2005) estimator and the integration procedure produce the same results when both the sample size and the number of quantiles chosen in (0,1) are sufficiently large.

We implement this methodology using the SHIW Bank of Italy data, in 1993 and 2006. We consider the monthly real wages (in log) as dependent variable, and as covariates education, experience and gender. Further, we implement 200 weighted quantile estimations on a regular grid, from 0 to 1 (0.005, 01, 015,…, 0.99, 0.995), deriving the coefficients $\beta(\theta)$ along all the θ distribution. We then derive the unconditional wage distribution multiplying the full matrix of X by the matrix containing all quantile regression coefficients, as in Autor et al. (2005). Each element of the resulting matrix can be considered as drawn from the unconditional wage distribution.[9] As for the fit of the estimation methodology, it is very accurate, as shown in Figure A1 in appendix.

3.2. Counterfactual wage distributions: covariates, coefficients and residual components

This methodology derives the marginal distribution of wages as function of covariates and coefficients, which implies the possibility to generate counterfactual densities, using different sets of $g(X)$ and $\beta(\theta)$. For instance, it would be possible to compute a counterfactual distribution keeping the covariates at the 1993 level and coefficients at the 2006 level.[10]

Furthermore, since the Machado and Mata (2005) methodology did not explicitly build up a direct measure for a within-residual component, i.e. between observationally equivalent individuals, our reference now moves on to Autor et al. (2005) and Melly (2005), who extend the Machado-Mata approach to identify three separate components in the computation of counterfactual distribution: covariates, coefficients and residuals.

Autor et al. (2005) and Melly (2005) define the coefficients component as a measure of between group inequality. In particular, following the notation of Melly (2005) and taking the median as a measure of the central tendency of the data, it is possible to derive the following wage equation for each year (1993 and 2006):

(4) $$\ln w_i^t = X_i^t \beta^t(0.5) + u_i^t, \quad t = 93,06$$

where $\beta^t(0.5)$ is the coefficients vector of the median regression in year t, which can be considered as a measure of between group inequality. To disentangle the effect of coefficients (between groups inequality) from the effect of residuals (within group inequality) it is important to note from (4) that the θth quantile of the residual distribution of u_i^t conditionally on X is consistently estimated by $X(\hat{\beta}^t(\theta) - \hat{\beta}^t(0.5))$.[11] Accordingly, Melly (2005) defines the following vector of coefficients as a measure for the within component:

[9] Note that instead of drawing observations from g(X) we consider the whole X matrix, as in Autor et al. (2005). Further, we do not consider, as in Machado and Mata (2005), only the elements on the diagonal of the resulting matrix generated by $\hat{w}_i \equiv \left\{ X_i \hat{\beta}(\theta_j) \right\}_{\substack{i=1,..,N \\ j=1,..,J}}$, but all the matrix. This means that we produce a much larger set of simulated values for the unconditional wage distribution (200 times larger), as in Autor et al. (2005).

[10] Note that all this literature, for instance Autor et al. (2005), Melly (2005, 2006), Machado and Mata (2005), make use of the partial equilibrium assumption that aggregate quantities of covariates do not affect labour market prices and vice versa. This assumption represents the major drawback of this research field. As shown by Di Nardo, Fortin, Lemieux (2006), supply and demand adjustments can be quantitatively important, accounting for 21% to 33% of the growth of male 90/10 log hourly wage inequality between 1979 and 1988.

[11] Note that it is possible to apply the conditional quantile process to (4), deriving: $Q_\theta(u \mid X) = Q_\theta(w \mid X) - X\beta(0.5) = X\beta(\theta) - X\beta(0.5)$.

$\hat{\beta}^{m06,r93}(\theta_j) = (\hat{\beta}^{06}(0.5) + \hat{\beta}^{93}(\theta) - \hat{\beta}^{93}(0.5))$, where the consistent estimate of the residual component given X, $(\hat{\beta}^{93}(\theta) - \hat{\beta}^{93}(0.5))$, is added to the between component, $\hat{\beta}^{06}(0.5)$, in 2006. In other words, we estimate the distribution that would have prevailed if the median return to characteristics had been the same as in 2006 but the residuals had been distributed as in 1993.

Using counterfactual distributions generated by applying different sets of covariates and coefficients, Melly (2005) computes how the variations over time of some quantile *q* of the wage distribution is attributable to covariates, coefficients and residuals. In particular, Melly (2005) estimates the residual component as the difference, at the quantile *q*, of the two following distributions, $\hat{q}(\hat{\beta}^{06}, X^{06})$ and $\hat{q}(\hat{\beta}^{m06,r93}, X^{04})$,[12] where the *X* and the $\beta^t(0.5)$ are constant at the 2006 level while the residual component is the only one that changes over time.[13]

Similarly, the difference between $\hat{q}(\hat{\beta}^{m06,r93}, X^{06})$ and $\hat{q}(\hat{\beta}^{93}, X^{06})$ is due to changes in coefficients since characteristics and residual are kept at the 2006 level.[14] Finally, the difference between $\hat{q}(\hat{\beta}^{93}, X^{06})$ and $\hat{q}(\hat{\beta}^{93}, X^{93})$ is due to changes of covariates.

To sum up, adding and subtracting $q(\hat{\beta}^{93}, X^{06})$ and $q(\hat{\beta}^{m06,r93}, X^{06})$ it is possible to decompose the variation over time of an estimated quantile of the wage distribution into the three components (residuals, coefficients, covariates), as follow:[15]

$$\hat{q}(\hat{\beta}^{06}, X^{06}) - \hat{q}(\hat{\beta}^{93}, X^{93}) = \overbrace{\left\{ \hat{q}(\hat{\beta}^{06}, X^{06}) - \hat{q}(\hat{\beta}^{m06,r93}, X^{06}) \right\}}^{\Delta residuals\ (within)} + \overbrace{\left\{ \hat{q}(\hat{\beta}^{m06,r93}, X^{06}) - q(\hat{\beta}^{93}, X^{06}) \right\}}^{\Delta Coefficients\ (between)}$$
$$+ \underbrace{\left\{ \hat{q}(\hat{\beta}^{93}, X^{06}) - q(\hat{\beta}^{93}, X^{93}) \right\}}_{\Delta \text{Covariates}}$$

Similarly, it is also possible to decompose the variations of any inequality index we are interested in, such as the ratios 90/10, 90/50, 50/10.

Other methodologies that compute the residual component have to assume independent error terms, as in the case of Juhn, Murphy and Pearce (1993). Methods based on quantile regressions can instead account for heteroscedasticity. This is actually crucial when the variance of the residuals expands as a function of education and experience (Lemieux, 2002) and when the population gets more educated and experienced, as in the Italian case.

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[12] The notation $\hat{q}(\hat{\beta}^{06}, X^{06})$ stands for the *q* quantile of the estimated distribution generated using the vector of coefficients $\hat{\beta}^{06}(0.5)$ and the set of covariates X^{06}. Note that it is different from $Q_\theta(w \mid X_i)$, which represents the conditional quantile distribution of wages.

[13] It is worth noting that the difference for each quantile *q* between the two distributions $\hat{q}(\hat{\beta}^{06}, X^{06})$ and $\hat{q}(\hat{\beta}^{m06,r93}, X^{06})$ can be rewritten as $\left| \hat{q}(\hat{\beta}^{06}(0.5) + \hat{\beta}^{06}(\theta) - \hat{\beta}^{06}(0.5), X^{06})) - \hat{q}(\hat{\beta}^{06}0.5) + \hat{\beta}^{93}(\theta) - \hat{\beta}^{93}(0.5), X^{06})) \right|$, from which it comes out clearly that the only component that changes over time is the residual one.

[14] A in the previous note: $\hat{q}(\hat{\beta}^{m06,r93}, X^{06})$ - $\hat{q}(\hat{\beta}^{93}, X^{06}) = \hat{q}(\hat{\beta}^{06}(0.5) + \hat{\beta}^{93}(\theta) - \hat{\beta}^{93}(0.5), X^{06}))$ $- \hat{q}(\hat{\beta}^{93}0.5) + \hat{\beta}^{93}(\theta) - \hat{\beta}^{93}(0.5), X^{06}))$

[15] Note that the sum of the three components exactly amounts to the estimated variation over time of that given quantile. This property is not shared with other methodology previously adopted.

4. **Decomposition results**

In order to identify what are the forces that are driving inequality trends in the public sector we carry out the decomposition analysis presented in the previous section. The set of covariate is the same: gender, education (4 dummies), experience (8 dummies).

As displayed in the panel A of Table 7, and confirming the descriptive statistics evidence, estimated wages increase along the whole wage distribution, i.e. 3.7% at the 10^{th}, 3.6% at the median and 10.4% at the at 90^{th} percentiles. As for the decomposition components, the negative coefficients component, which represented the driving force of inequality in the private sector, is quite negligible along the wage distribution. As expected, also in the public sector the covariates component is positive because of the increase in educational and experience levels, and it is quite stable along the wage distribution. The only component that strongly increase along the wage distribution is the residual-within, which ranges from -0.9 at the 10^{th} percentile to 6.1 at the 90^{th} percentile.

Panel B and C of Table 7 suggest that the whole distribution of wages is increasing over time. In particular, in the period 1993-2002 only the 90^{th} increased with respect to 1993, while the rest of the distribution decreased. In the period 1993-2004, wages remains pretty stable from the 10^{th} to the 75^{th} percentile while they increase at the 90^{th} percentile. In the period 1993-2006, all the wage distribution is shifted to the right. Furthermore, the behavior of the three decomposition components is similar in the three panels.

Table 8 shows the estimated changes of wage inequality for the public sector, for the period 1993-2004 and 1993-2006. As already stressed in the descriptive statistics, the increase in wage inequality is greater for the period 1993-2004 than for the period 1993-2006. However, for both periods the increases in the 90-10 ratio is associated to the increase in the 90-50, being the 50-10 quite stable over time. Furthermore, both the rises of the 90-50 and the 90-10 inequality indexes are driven by a positive residual contribution, being negligible the effect of covariates and coefficients components.

4.1. Explanations for the public sector: the role of institutions *vs* increasing experience levels

The increase in wage inequality observed in the public sector is concentrated in the upper tail of the distribution, and it is entirely related to the increase in residual inequality. In this section, we investigate the main changes occurred in the public sector in the period 1993-2004, namely the striking increase in the level of experience and the reforms of the wage setting system enacted in 1993 and 1998, as possible candidate explanations.[16]

As shown in Table 6, the share of individual with more the 20 years of experience increased from 44% to 59%, while the corresponding share in the private sector increased only from 41% to 43%. This is mainly due to the slowdown of hiring rates, related to binding public budget constraints, and to the very low rate of separations observed in this sector. Thus, since the main increase of public employment took place in the 70s and the 80s, these cohorts represent now the modal class in this sector. According to Lemieux (2002, 2006), the increase in experience levels might mechanically generate an increase in residual inequality, since

[16] Similar results are derived for the period 1993-2006.

experienced workers display usually higher wage dispersion.[17] Hence, the increase in the experience levels might be a suitable candidate to explain the rise in the upper tail of residual inequality observed in the public sector.

As for the wage setting system in the public sector, it has been substantially reformed in 1993 and in 1998. After the reforms, the same kind of public employees (same age, occupation, tenure) might be paid differently in different administrative units, or even in the same units if assigned to different tasks. For these reasons, also these reforms may be considered as suitable candidates for the increase in the residual inequality in the public sector.

Let us now investigate the pros and cons of these two competing explanations.

To the extent that the increase in the residual inequality were mainly driven only by a mechanical composition effect related to the increase in the share of experienced workers, we should expect similar wage dispersion by experience classes in 1993 and in 2004: since experienced workers are supposed to display higher wage dispersion, the increase in their shares would mechanically cause the increase in residual inequality (Lemieux, 2002, 2006).

In the left panel of Figure 6 we report standard deviations of residuals derived in a mincerian quantile regression computed at the median, using the same set of covariates as in the decomposition analysis.[18] It is straightforward to note that standard deviations computed by experience classes show patterns completely different between 1993 and 2004. This suggests that the composition explanation proposed by Lemieux cannot be considered as the sole driving force of the increasing residual inequality in the public sector. Nevertheless, the standard deviations for those experience classes that increase their employment shares (more than 20 years of experience) have exploded over time, suggesting that experience patterns matter for the explanations of the inequality trends, although not through the pure composition effect mechanism. At the same time, average wages by experience classes, reported in the right panel of figure 6, do not change much between 1993 and 2004, revealing that the changes occurred in the public sector have mainly affected dispersion but not levels.

In a similar way, standard deviations of residuals increase for all the educational levels. The left panel of Figure 7 shows also that graduate standard deviations, which displayed the highest dispersion already in 1993, strongly increases over time. Further, also for educational levels average wages are quite stable over time, with a slight premia only for individuals with primary education and for graduates.

As for occupational levels, figure 8 shows that standard deviation of residuals increased for all categories, and especially for white collar and teachers, while average wages are quite stable over time, with a slight increase for managers and teachers.

To sum up, residual wage dispersion increased for all experience, educational, and occupational levels, while average wages have not changed much over time. In this framework, the new system of employment reforms introduced in the public sector can be considered as the most convincing explanation. In this reformed institutional setting, as

[17] Further, also educational levels increased in the public sector, although less in magnitude. Note that, according to Lemieux (2002, 2006), our predictions regarding experience levels would be reinforced by the increase in educational levels occurred in the public sector.

[18] Actually we regress log wages on gender dummy, education (4 dummies), and experience (4 dummies, while in the decomposition analysis we had 8 dummies). As in the decomposition analysis, we drop the 0.25% at the tails of the distribution.

already stressed, the same kind of public employees (same age, occupation, tenure) might be paid differently in different administrative units, or even in the same unit if assigned to different tasks. Hence, while average wages do not change much, wage dispersion has exploded.

This explanation is also supported by the timing of changes in wage inequality. Actually, Dell'Aringa and Della Rocca (2007) claim that the first reform introduced in 1993 was basically ineffective, and for this reason the second reform in 1998 was enacted. Dell'Aringa and Della Rocca (2007) argue that the main effects of this second reform took place especially from 1998 to 2002, since local administrative units abused from the possibility to integrate the first national level of bargaining. After 2002 a stricter monitoring was introduced in order to keep under control the increasing public spending. Table 3 and 4 confirm these trends, with a clear increase between 1998 and 2002 and a fall afterwards.

However, we had to explain why residual inequality derived in the decomposition analysis increased only in the upper tail of the wage distribution, i.e. at the 90th percentile. This is probably because 64% of individuals at the last decile is graduated, 45% are managers and 35% are white collars, and around 70% have more than 20 years of experience. Since residual wage dispersion strongly increased for all these categories, it can explain the increase in residual inequality observed in the decomposition analysis at the 90th percentile.

5. Conclusion

Empirical literature shows that wage inequality in Italy has been quite stable over the last decade. We claim that this is no longer true when public and private sectors are analyzed separately as inequality increased from 1993 to 2006 in the public sector.

This paper investigates which are the forces that have generated this observed pattern of inequality in the private sector. Then we carry out a quantile decomposition analysis (Machado and Mata, 2005, Melly, 2005, Autor et al., 2005) to identify three main components, related to coefficients, covariates and residuals.

The main result of the analysis it that the increasing wage inequality in the public sector, which is almost entirely driven by an increase in residual inequality in the upper tail of the distribution, may be explained by the reforms of the wage setting system implemented in 1993 an 1998.

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Tables

Table 6: SHIW Sample descriptives. Public and Private Sector						
	Public			**Private**		
	1993	2004	2006	1993	2004	2006
Share on total employment	0.33	0.23	0.21	0.67	0.77	0.79
Female	0.46	0.52	0.54	0.33	0.38	0.39
Education						
Primary - no school	0.06	0.02	0.02	0.18	0.09	0.05
Lower secondary	0.26	0.20	0.15	0.45	0.38	0.38
Upper secondary	0.44	0.52	0.52	0.33	0.45	0.47
Univ. Degree or higher	0.24	0.26	0.31	0.04	0.08	0.09
	1.00	1.00	1.00	1.00	1.00	1.00
Experience (year)						
eps1 - 0-5	0.09	0.08	0.07	0.18	0.14	0.13
eps2 - 6-10	0.12	0.09	0.09	0.13	0.13	0.13
eps3 - 11-15	0.16	0.11	0.10	0.14	0.14	0.13
eps4 - 16-20	0.19	0.14	0.15	0.13	0.17	0.17
eps5 - 21-25	0.17	0.19	0.17	0.13	0.14	0.15
eps6 - 26-30	0.11	0.19	0.16	0.11	0.11	0.12
eps7 - 31-35	0.08	0.13	0.15	0.09	0.09	0.09
eps8 - >36	0.08	0.07	0.09	0.08	0.09	0.09
	1.00	1.00	1.00	1.00	1.00	1.00
Net Wages (Monthly)						
Mean	1466	1514	1592	1318	1304	1356
10th percentile	1009	1009	1043	723	775	822
50th percentile	1375	1356	1411	1177	1195	1224
90th percentile	1963	2179	2141	2006	1933	1994
Observations	1988	1312	1167	4052	4341	4386

Note: 0.025% of the observation in the right and left tails dropped. Weight used:pesofl

Table 7. Total observed variations at selected quantiles and quantile decomposition into the coefficients, covariates and residuals components in the public sector (in percentage points).

	Δ p10	**Δ p25**	**Δ p50**	**Δ p75**	**Δ p90**
Panel A	**Variations at selected quantiles - 1993-2006**				
Total estimated variation	3.7	4.8	3.6	4.7	10.4
Coefficients (between)	-1.0	0.3	-0.3	-0.4	-0.3
Covariates	5.5	4.5	3.9	3.4	4.5
Residual (within)	-0.9	0.1	0.1	1.7	6.1
Panel B	**Variations at selected quantiles - 1993-2004**				
Total estimated variation	-0.6	-0.5	0.2	2.7	10.9
Coefficients (between)	-1.2	-0.9	-0.8	-0.2	0.3
Covariates	3.8	3.1	2.8	2.3	2.8
Residual (within)	-3.2	-2.7	-1.9	0.6	7.8
Panel C	**Variations at selected quantiles - 1993-2002**				
Total estimated variation	-6.2	-4.0	-3.3	-1.3	5.6
Coefficients (between)	-6.2	-5.6	-6.1	-6.3	-6.3
Covariates	2.9	2.8	2.9	2.5	2.8
Residual (within)	-2.9	-1.3	-0.2	2.5	9.1

Source: SHIW data. 0.25% of observations are dropped at the two tails of the wage distribution.

Table 8. Decomposition of the inequality indexes into the between, within and covatiates component in the public sector (in percentage points), 1993-2006

	Period - 1993-2006		
	90/10	**50/10***	**90/50**
Total estimated variation	6.7	-0.1	6.8
Coefficients contribution (between)	0.76	0.68	0.08
	11%	-	*1%*
Covariates contribution	-1.07	-1.68	0.61
	-16%	-	*9%*
Residual contribution (within)	7.01	0.94	6.08
	105%	-	*90%*
	Period - 1993-2004		
	90/10	**50/10**	**90/50**
Total estimated variation	11.6	0.8	10.7
Coefficients contribution (between)	1.54	0.47	1.07
	13%	*58%*	*10%*
Covariates contribution	-0.98	-1.02	0.04
	-9%	*-126%*	*0%*
Residual contribution (within)	11.00	1.36	9.64
	95%	*168%*	*90%*

*Since the 50/10 variation is almost equal to zero, we cannot compute the percentages (they tend to infinity)

Figures

Figure 3. Gini Index - period 1993-2006

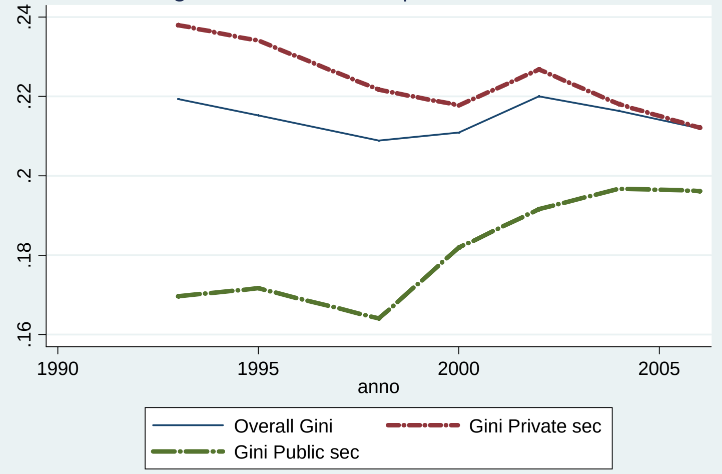


Figure 4. 90/10 Index - period 1993-2006

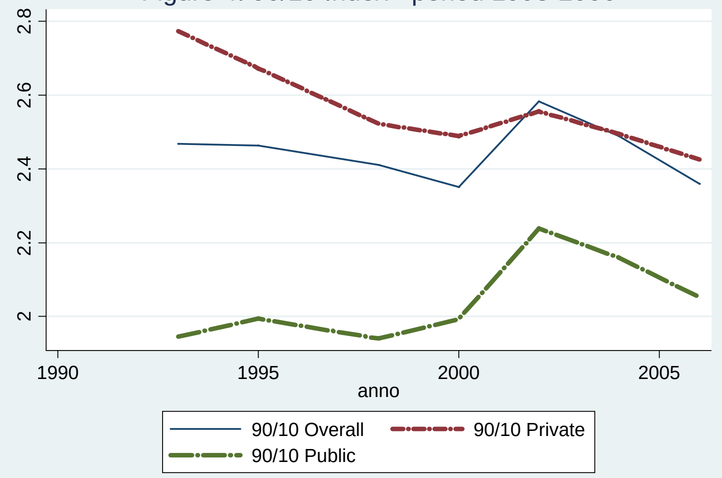
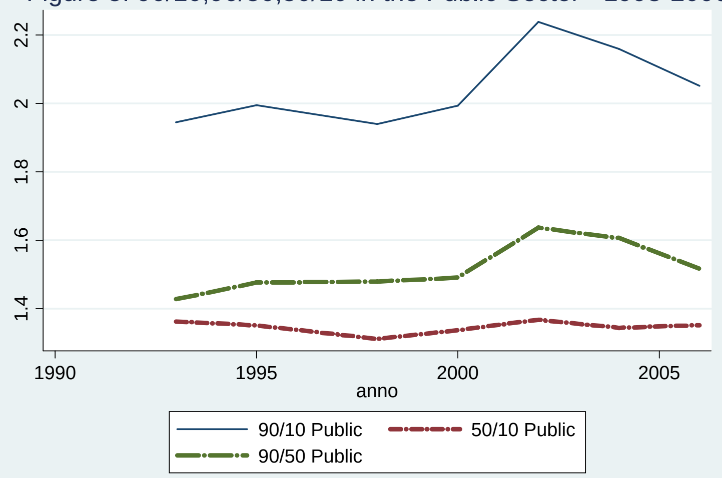


Figure 5. 90/10,90/50,50/10 in the Public Sector - 1993-2006



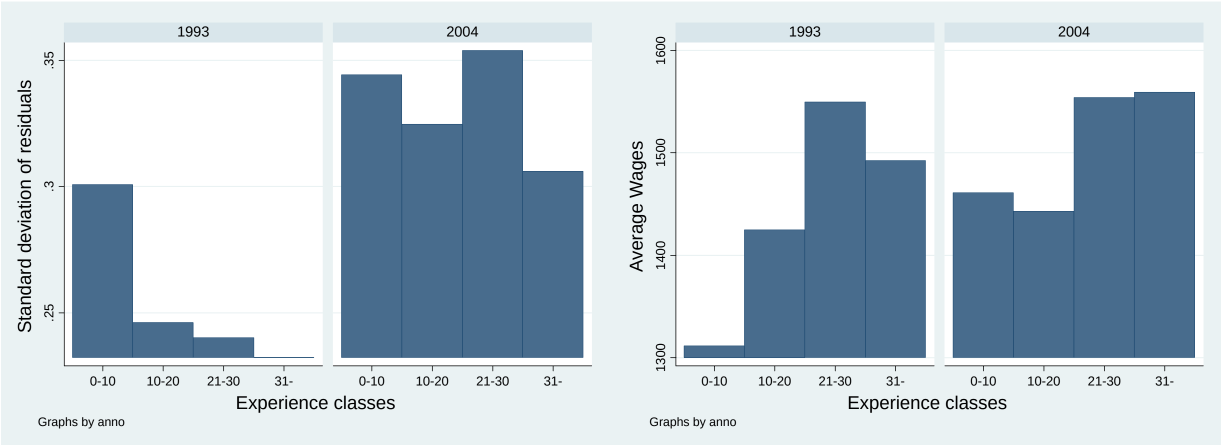


Figure 6: Residual dispersion and average wages by experience classes, 1993-2004

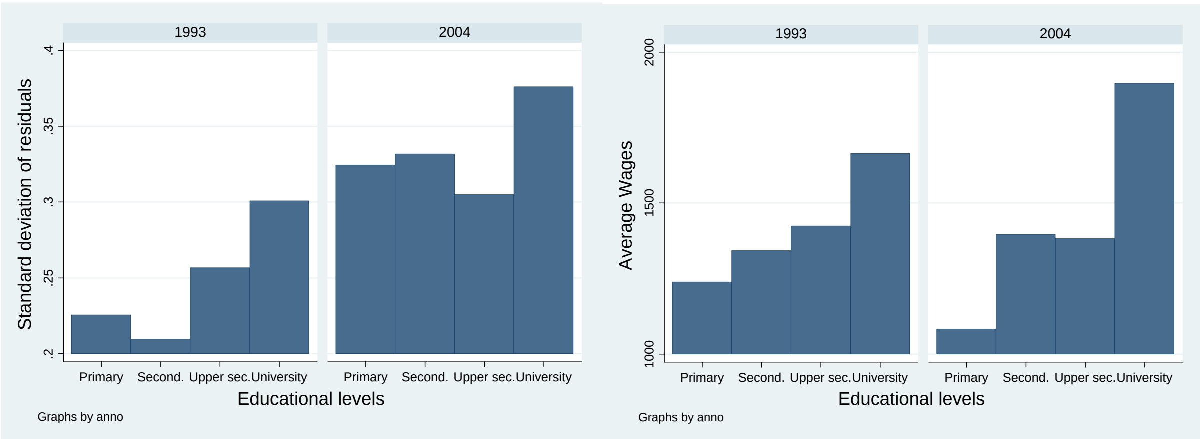


Figure 7: Residual dispersion and average wages by educational levels, 1993-2004

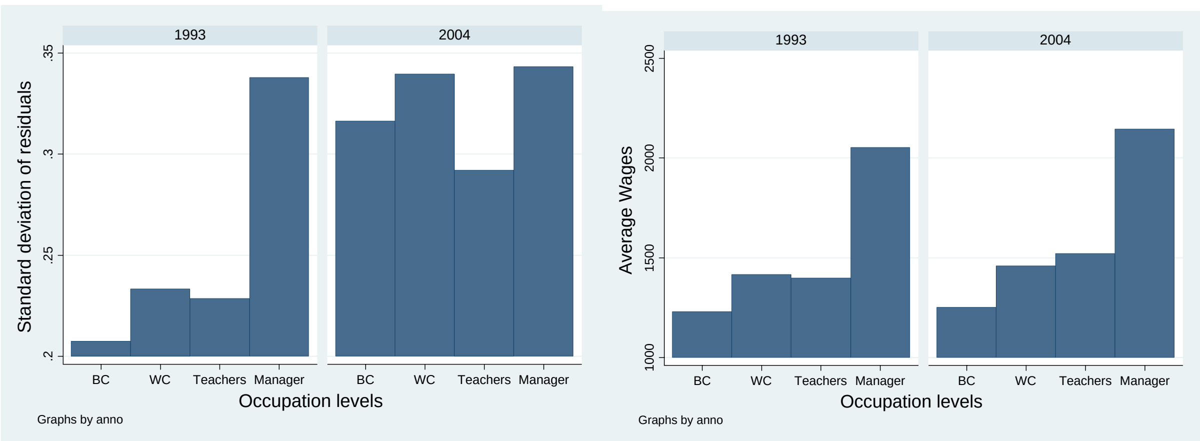


Figure 8: Residual dispersion and average wages by occupational levels, 1993-2004