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Gender wage differentials among young workers: methodological aspects and empirical results

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1. Introduction

The research on gender discrimination during the last decade has pointed out some stylised facts. Although recent figures do not always agree on the value of the gender wage gap and on the temporal evolution of gender wage differentials, studies on discrimination confirm the existence of a widening trend in the unexplained part of the wage gap. Returns to male individual characteristics are higher than returns to female workers and differences are increasing.

With regard to the literature on the Italian labour markets, Addis and Waldmann (1996)[1] estimate a level of unexplained wage differential equal to 13% in 1989. Higher discrimination –about 20%– is found by ITER (2001) and Flabbi (2001) using data from the Bank of Italy and the Ministry of Finance; moreover, these works confirm the increase in the unexplained gap between 1977 and 1995.

In the last decade, particular attention has been devoted to the analysis of wage differentials among young workers. In a work by Wood et al. (1993), the gender wage gap between graduates of the University of Michigan Law School classes of 1972-1975 is estimated to be relatively small at the outset of their careers and equal to 40 percent of the male wage after 15 years. This result is partially due to different choices in the hours of work and in worker qualifications by gender; but even controlling for these factors, women are found to gain a 20 percent less than men.

Many other studies on young workers estimate low levels of wage differentials; however, these results are mainly due to data problems. Much of these works cannot control for actual experience because of left-censored data and results are therefore biased by this lack. Loprest (1992), using a sample of 18-25 year old women and men drawn from the U.S. National Labour Survey, estimates an entry wage gap of 11 percentage points. Dolton and Makepeace (1987) reveal a 7 percent wage gap for a sample of U.K. graduates in 1970. Research on gender pay gap in the U.S. generally explains the increasing wage differential throughout the 1970s and 1980s in terms of a lower level of actual work experience of women compared to men and relatively higher returns to work experience for men [Corcoran & Duncan (1979), Polachek and Robst (2001)]. Even job changes appear to favour men compared to women in the first years of their careers [Loprest (1992)].

In our study we evaluate the gender wage gap for 15-29 year old female and male workers drawn from the administrative data set of the National Social Security System for the two Italian provinces of Treviso and Vicenza (North-Eastern Italy) in two different years (1990 and 1997). The data set contains the universe of workers employed in the private sector in the two provinces for the period 1976-1997. Time-series availability and the structure of the data set enable us to exactly measure work experience, to measure firm-specific human capital (tenure in the actual firm) and to control for individual and occupational characteristics, economic sectors and local labour markets.

From a methodological point of view, the approach generally adopted in the analysis of discrimination relies on the Oaxaca (1973) and Blinder (1973) proposal. Although

[1] They use data from the Bank of Italy survey on income and wealth.

recent extensions of this approach[2] control for some distributional aspects, the methodology is fundamentally based on the analysis of the average discrimination. This traditional analysis, as demonstrated by Jenkins (1994), could be fully informative of discrimination if the individual difference between the estimated wage and the wage of reference was constant along the entire distribution of female earnings; i.e., if the difference between estimated values and counterfactual values was distributed as a uniform probability density. If this condition does not hold, the analysis *à la Oaxaca-Blinder* is not a proper instrument to evaluate the discriminatory phenomenon.

The aim of this work is therefore to evaluate gender discrimination for young workers in two Italian provinces of North-Eastern Italy (Treviso and Vicenza) using the complete distribution of discrimination estimated at the individual level. Our data drawn from the archives of the National Social Security System allow us to match information on individual characteristics to information on the size and economic sector of firms. Moreover, we are also able to control for firm's location within local labour market systems.

The next Section presents some descriptive statistics of the dataset and quantifies the wage gap between male and female workers. Section 3 provides a description of the traditional approach to the analysis of wage discrimination and highlights some critical remarks. Earnings function estimates by gender are discussed in Section 4, while Section 5 introduces the distributional approach, which is then illustrated and implemented in Section 6. Section 7 concludes.

2. Data and descriptive statistics

We use administrative data made available from the National Social Security System (*Istituto Nazionale di Previdenza Sociale,* INPS*)* for the two North-Eastern Italian provinces of Treviso and Vicenza. The INPS database is made of different archives containing information on the universe of workers who had at least one employment spell eligible for the social security insurance scheme in any year between 1976 and 1997[3] and on the universe of firms which paid any social security contribution during the same period. The database allows the match of each worker to each different firm where he used to work since the beginning of his working experience; the matching enables to merge personal data with information on gross earnings[4], type of job, sector and firm-size. Unfortunately, the personal archive does not contain information on the household the worker belongs to and therefore we are not able to control for family characteristics. We are fully aware of the relevance of household information in

[2] This methodology has been extended in two different directions. Juhn, Murphy and Pierce (1991) have extended the Oaxaca decomposition by taking into account the residual distribution, useful when analysing the gender gap in a comparative settle or throughout a long temporal interval. Brown, Moon and Zoloth (1980) have extended the approach in a way to endogenize the distribution of women across occupations.

[3] For detailed information about the INPS archives (for the provinces of Treviso and Vicenza) see the work by Occari et al. (1996).

[4] Gross earnings include social security contributions. We use the Consumer Price Index (base=1995) to deflate earnings.

explaining both female decisions of work participation and gender differences in working career; however we also believe that analysing young workers aged less than 30 notably reduces the importance of these variables. The personal archives do not contain any information on education but we are able to circumvent this problem with a suitable proxy.

We generate two samples of young workers –aged between 15 and 29– for the years 1990 and 1997. The construction of the sample allows both to exactly determine the total effective experience of each worker, since his entrance in the private sector of the labour market as a regular employee, and to solve for the problem of left-censoring. We also limit the analysis to individuals working the whole year and full time, in order to have a homogeneous sample and to exclude seasonal working spells; moreover we exclude managers due to female under-representation in those types of occupations. The analysis is carried out for workers in the economic sectors included in manufacturing, commerce and credit-insurance[5]. Finally, following the debate on the risks in using INPS data for the analysis of wage gaps, we include in our sample only workers with a number of worked days equal or higher than 280[6]. The final sample is then made of 74759 workers in 1990 and 68882 workers in 1997 (Table 1)[7].

A first interesting figure emerges from the sample composition with respect to the educational level[8]; the most of workers, independently of gender, has a lower-secondary educational level in both years. In 1990, 69% of women and 71% of men have a compulsory diploma while in 1997 the percentage decreases to 52% for women and 59% for men. Although this trend could be partially influenced by a delay in the entrance in the labour market[9], we believe that most of the effect is due to the increasing investment in post-compulsory education. This is confirmed by the percentages of workers with a higher education degree which increase from 2.2% in 1990 to 4.4% in 1997 for men and from 1.4% in 1990 to 2.1% in 1997 for women. These figures suggest a less unequal distribution of education among women; in 1997 the percentage of females with an upper-secondary education diploma is 45% against a male percentage of 36.

Looking at other human capital variables (table 2), further gender differences appear, especially in 1990. In this year, gender is particularly relevant in determining the type of experience that workers accumulate; women tend to accumulate more specific human capital than men while the latter ones tend to invest more in general experience. Differences are quite relevant. In few years, however, women have almost caught up with men, reaching a level of general experience on average almost equal to that of men.

Before going through the econometric analysis aimed at isolating the unexplained part of the wage gap, we briefly present some features of the male and female distributions of earnings.

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[5] The economic sectors considered in the analysis are those included in the 3, 4, 6 and 8 one-digit ATECO 1981 ISTAT classification (manufacturing, commerce, credit, insurance).

[6] INPS data on wages seem to be overestimated for people working only few days in a year. For that reason, it is suggested to exclude individuals working less than 280 days a year. For the debate, see Ginzburg et al. (1998, 1999), Gavosto and Rossi (1999).

[7] The percentage of women is 47.6 in 1990 and 46.2 in 1997.

[8] See Section 4 for the definition of the education variable.

[9] By construction of the variable.

Recent Italian studies [ITER (2001), Flabbi (2001)] have shown different earnings trends depending on the data used. Data from the Ministry of Finance show an increase in the wage gap from 29% in 1982 to 36% in 1996; data from the same source, but for the sub-sample of year-round employees, show a lower gender wage difference, however steadily increasing from 19% in 1982 to 21% in 1994. Data from the Bank of Italy for the whole distribution of earnings show an increase in the first years of the 90s with a much stronger worsening for manual female workers than for white-collars. Conversely, considering only year-round employees, data from the Bank of Italy show an increasing relative wage.

Table 3 reports, for the whole distribution of year-round full-time workers, wage gap levels much lower than those calculated using data from the Bank of Italy and from the Ministry of Finance. However, strong differences appear between the distributions of white-collars and blue-collars. In 1990, the gap for low-paid blue-collars is higher than the gap for low-paid white-collars, while the gap for white-collar women in the right-hand side of the distribution is much higher than that of blue-collars in the same side of the distribution. In 1997, the gap for blue-collars becomes sharply higher than the gap for white-collars.

A temporal comparison shows an increase in the wage difference along the entire distribution, except for the last quartile. That means that young workers with high earnings experience lower wage gaps than those that earn low wages. However, once more, there exists a sharp difference conditional on the occupation: white-collars experience quite a sharp fall in the wage gap with respect to each single quartile, while blue-collars show an increasing trend at all levels of the distribution.

3. Methodological aspects

In the previous section we have shown that high earnings differences exist between genders and that these differences have tended to widen during the 1990s. That analysis, however, is not indicative of the existence of discriminatory behaviour towards female workers. Actually, high wage gaps could be explained by gender differences in the characteristics of individuals.

In order to explain how much of the gap could be justified by gender differences in human capital accumulation and how much could be the result of discriminatory employer behaviour, we depart our methodological approach from the traditionally used discrimination literature[10].

The literature on discrimination analysis generally adopts the approach of Oaxaca-Blinder (1973)[11]; gender discrimination is defined as the difference between the wage women earn and the wage they would earn if they were men, given the same characteristics. The methodology consists in estimating gender-separate 'earnings functions' explaining worker income as a function of the characteristics of the individual. Then, the total estimated difference is decomposed in two terms: one

[10] Antonji and Blank (1999) and Kunze (2002) present two interesting surveys on discrimination analysis.

[11] As already said in previous paragraphs, the traditional model has been developed but all versions essentially rely on the average analysis.

represents productivity differences explained by individual characteristics and a second explains earnings gaps in terms of differences in the remuneration of the characteristics. This second term represents the discriminatory component of the wage gap. Traditionally, the methodology then suggests the calculation of an index of discrimination defined as the average of the difference between the reference estimated wage –e.g. the wage that women may earn if their characteristics were paid as much as men– and the estimated female earnings[12].

The main problem in implementing this approach is the lack of data on workers productivity. In order to exactly estimate the discriminatory component of the wage gap, in fact, we should be able to evaluate individual productivity and to isolate the part of the gap which is explained by these differences. Unfortunately, few databases at the moment contain these information and economists have to rely on data that can explain only part of the differences in productivity. That implies the possibility that the decomposition procedure overestimates the discriminatory part of the wage gap. Moreover, very often, databases do not contain some relevant variables for explaining earnings functions and these variables are correlated with those included. As a consequence, the estimates can be affected by endogeneity problems [Brunello and Miniaci (1999)].

In our analysis, unfortunately, we can not correct our estimates to take account of all these concerns; in fact our database does not contain several pieces of information which would allow us to better evaluate the discriminatory part of the gap and to partially correct for the problem of endogeneity. Being aware of this issue, our aim is not to give an evaluation of discrimination but to decompose the wage gap conditional on the information we have at our disposal. Moreover, our research primarily aims at proposing a different methodological approach for the analysis of discrimination, starting from a critical analysis of the Oaxaca-Blinder approach.

An important drawback of the Oaxaca-Blinder methodology, apart from the econometric concerns we already discussed, is its low power in explaining the discriminatory phenomenon. This approach can provide little or, in some cases, no information on the phenomenon under study. First critical remarks to this approach go back to Dolton and Makepeace (1985) and Munroe (1988) who realise the shortcomings of the Oaxaca-Blinder methodology and the need to use the entire set of information contained along the distribution of earnings. These two contributions, however, have strong limits themselves. In Munroe (1988), the definition of 'absence of discrimination' is not precise and the meaning of the reference distribution for the female earnings is not satisfactory. Dolton and Makepeace (1985) correctly define the reference distribution and try to evaluate discrimination making use not only of the mean but also of other moments of the distribution, but fail to define a link between the different shapes of the distributions and discrimination.

In order to understand the limits of the Oaxaca-Blinder approach and the criticisms to it, let us consider a situation in which the reference distribution and the estimated distribution have the same average but a different shape. As an example, let us suppose that the estimated earnings is equal to 5 for 90% of working women and 55 for the other 10%, with an estimated average equal 10. Let us also suppose that the reference

[12] This is equivalent to the difference between the averages of the two distributions.

earnings are equal to 10 for all women. Given that information, the traditionally used approach would deny the existence of discrimination. This method, in fact, does not take into consideration the distribution of earnings and the fact that 90% of women earn half of the wage a men would receive with the same characteristics while the remaining 10% gets 5.5 times the wage the men would receive. As a consequence, only if the probability density functions were identical would it be possible to exclude situations as the one just described.

However, even in the case of identical probability density functions, it is still possible that focusing on the average of the two distributions hides relevant information about the distribution of discrimination. Let us consider the opposite situation in which all the moments of the two distributions are identical, i.e. the shapes of the two distributions are identical. However, also in this case the traditional approach could be misleading as it would be unable to reveal whether the earnings gap is the same for all women or whether each individual positive gap is exactly compensated by a negative gap of some other woman. In other words, the fact that the shapes of the distributions of the estimates and of the reference wages are identical is not a sufficient condition to ensure the absence of discrimination, but just a necessary condition.

Summarising, the Oaxaca-Blinder approach appears inadequate when the shapes of the two distributions are different (independently from the incidental coincidence of average values) and, in general, very little informative on the discriminatory phenomenon.

Jenkins (1994) proposes an analysis of discrimination using the complete information contained in the distributions of the estimated and of the reference female earnings. His proposal, however, is not incoherent with respect to the criticisms he himself raises against the traditional methodology, but also unable to provide a complete evaluation of the discriminatory phenomenon. The author, in fact, after criticising the traditional approach because of the impossibility to infer significant conclusions about discrimination from looking at the statistics for the two distinct distributions, still focuses on the two separate probability density functions. Its innovative contribution consists in comparing not summary statistics but the relative position of the Generalised Lorenz Curve (GLC) and the Generalised Concentrations Curve (GCC)[13], suggesting that discrimination exists when the GCC is above the GLC. However, the relative position of the two curves is not indicative of the existence of discrimination; in fact, by construction, the distance between the two curves depends on cumulated differences between estimated and reference earnings, so that it can be positive even though the marginal contribution of one more female worker is negative. Consequently, the fact that the GCC lies above the GLC does not allow concluding that there exists discrimination against women along the whole distribution. Therefore, the methodology proposed by Jenkins fails in implementing a truly distributional analysis

[13] In order to construct the GLC, after ordering female workers in ascending observed wage, plot cumulative estimated wage per capita $\sum_{i=1}^{k} \hat{y}_i / n_w$, against cumulative sample share $p \equiv k/n_w$, for each k = 1, …, n_w (n_w is th etotal number of femal workers). Similarly, to summarise the no-discrimnation distribution using a GCC, you plot cumulative reference wage per capita $\sum_{i=1}^{k} \hat{r}_i / n_w$ against p, for each k = 1, …, n_w, insuring women are order exactly as for the GLC.

of discrimination and can actually be interpreted as a simple extension of the traditional methodology. Moreover, a further critical remark on the methodology put forward by Jenkin concerns the discrimination index he suggests. The index actually does not evaluate discrimination but, rather, the diversity between the two distributions. The index, in fact, is defined as the weighted sum of the absolute values of the individual differences between estimated wages and reference wages[14]. In this way, discrimination against women and pro women would show up with the same sign.

Following these considerations, we proceed by adopting a different methodological approach in order to directly evaluate discrimination –both positive and negative– and the incidence of individual characteristics, without losing any information contained along the entire distribution[15].

As in the traditional analysis, we first estimate log wage equations for men and women as a function of their characteristics – X_m for men and X_f for women– and of an error term:

$$log(W_{m_i}) = X_{m_i}\beta_m + \varepsilon_i \quad (1)$$

$$log(W_{f_i}) = X_{f_i}\beta_f + \varepsilon_i \quad (2)$$

Once the characteristics coefficients, b_m for men and b_f for women have been estimated, discrimination emerges if there are differences between these coefficients. The reference wage for each woman is equal to the retribution she would receive if she were paid as a man for each characteristic.

As previously pointed out, the traditional approach would evaluate discrimination as the average of the difference between estimated and reference earnings, expressed in log transformation: $\overline{X}_f(b_m - b_f)$. Differently, we make proper predictions of reference and estimated earnings for each woman (Jenkins, 1994) in order to derive and study the shape of the distribution of the individual differences and to evaluate the incidence of each of individual characteristics.

Following Jenkins' notation, we call $\hat{r}_f$ the series of the reference earnings and $\hat{y}_f$ the series of the estimated earnings:

$$\hat{r}_i^f = \exp(X_i^f b^m), \quad (3)$$

$$\hat{y}_i^f = \exp(X_i^f b^f) \quad (4)$$

each $i \in F$

[14] Weights are given by the relative earnings of each woman; therefore those earning high wages have higher weights. The sum of the gaps is then weighted a second time with a parameter of discrimination aversion.

[15] See paragraph 6.

Finally, as previously pointed out, we are fully aware of the limitations in our dataset and conscious of the impossibility to exactly estimate discrimination as well. Actually, we aim at providing an evaluation of the unexplained wage gap, given the characteristics we observe. However, in order to ease the reading of the article we will call the unexplained gap simply discrimination.

4. Empirical analysis

The empirical investigation of the determinants of the earnings gap is based on the dataset described in Section 2[16]. Our dependent variable is the log transformation of daily earnings. The INPS archives provide information on yearly gross earnings (gross of social contributions and income taxes) and worked days for each year of worker-spell. We therefore derive our dependent variable making daily earnings at constant prices 1995[17]. As previously explained, we focus on 1990 and 1997; although the time interval is not extended, important transformations occurred in the Italian labour market during that period, such as the complete abolition of the wage indexation system (the '*scala mobile*' was definitely suppressed in 1992).

Among the explicative variables we include different individual characteristics aimed at evaluating the human capital endowment of each worker: effective experience in the labour market, tenure at the firm level and a proxy for education. Moreover, we control for the occupation (white-collar/blue-collar), the type of contract (full-time/part-time, apprenticeship and training contracts), the economic sector[18] and firm-size. We also include dummies for each local labour market system.

Experience is included as a measure of general experience accumulated since the first entrance in the labour market and is computed as the number of worked weeks net of the actual spell. Tenure is included to capture the return effect on wages of specific human capital accumulated inside the firm; it is measured as the number of worked weeks spent in the actual firm.

The INPS database lacks some information on the educational level of workers. We therefore generate a proxy of the educational level considering for each individual the age of the first non-seasonal entrance in the labour market. In order to identify this spell, we look at the first work experience either longer than 17 weeks or between 4 and 17 weeks but not registered between May and September; in that way we take into consideration only first work spells which are non-seasonal or non compatible with the regular attending at school.

Two dummies have been generated to isolate the effect of post-compulsory education. The dummy 'secondary education' take on a value of one when the age of the first-spell is between 19 and 25 (proxy for an upper-secondary diploma); the dummy

[16] We also estimated a model specification with fixed effects at the firm level. The coefficient estimates in this case did not differ from the coefficients we present in this section, confirming the robustness of our specification. We preferred to proceed using the OLS model because it allows us to separate firm-size effects from sector and local labour market effects.

[17] We use the ISTAT Consumer price index to deflate earnings.

[18] We use the ISTAT definition (ATECO 1981).

'university education' is equal to one when entrance age in the labour market is bigger than 25[19].

Finally, we include two different dummies to isolate the impact of both *Apprenticeship* contracts and *Contratti di Formazione Lavoro* (CFL; Working and training contracts)[20].

Regression results are reported in Table 4[21]. Estimates confirm the positive effect of the accumulation of human capital on the wage profile, both for women and for men. With regard to education, workers with a university degree gain higher wages in comparison to those with a diploma in both years; general and specific experience positively affect wages following a decreasing return function.

Relevant differences emerge between the two years. In 1990, women get a higher premium with respect to men if they stay longer in a firm while males have a greater return compared to females if they accumulate general experience. In 1997, the female advantage in the accumulation of firm-specific human capital reduces and becomes lower than that of males. Moreover the female disadvantage in the return to general experience increases throughout the years.

Gender differences in the returns to education are more pronounced in 1990 than in 1997. In 1990, women receive a 25-30% lower return to education compared to men, independently of the education level; the secondary education diploma guarantees a premium of 4.5% to women and of 6% to men, while the gain from a university degree is equal to 8.8% for women and 11.3% for men. In the considered period, the returns to female human capital have increased, reaching in 1997 levels much alike to those for men. While the secondary education diploma always gives a higher premium to men, the return to tertiary level education is equal for men and women. In general, our results are hence in line with those of previous studies; our human capital estimates are lower than in other researches but this is probably due to sample composition (young workers).

The other coefficients show some interesting figures. First of all, women in white-collar occupations receive a premium with respect to blue-collar women which is higher than the correspondent male premium; this difference is irrelevant in 1990 and becomes significant in 1997 (equal to 19.6% for women and 15.2% for men). Particularly interesting are the estimates of the dummies on workers with apprenticeship contracts and CFL; wages for these workers are sharply lower than for the other ones, but women are in some way less disadvantaged and experience a lower gap.

Firm-size dummies show a gain for both genders when working in medium and large-size firms. Men loose a maximum of 6.3% in 1990 and 7.5% in 1997 when working in firms with less than 20 employers, gaining up to 7 percentage points in 1990 and 6.3 percentage points in 1997 if employed in large-size firms. Women employed in

[19] Returns to education are generally underestimated when using OLS methods because of endogeneity problems (Brunello e Miniaci, 2001). In our case, because of the lack of data on the familiar situation of the worker, we do not have any instrumental variables to solve this problem.

[20] The CFL were introduced in 1985 in order to improve the youth chance (aged less than 30) to get a job. The employer is provided with a rebate on the labour cost and a full exemption from firing costs.

[21] We do not show the local labour market dummies (LLM); however we can provide the complete tables on request. The significant LLM dummies are different between the two genders and the estimated coefficient is often negative.

large firms gain a 10% in 1990 and an 11% in 1997, while loosing up to a maximum of about 3% in 1990 and 5% in 1997 if employed in a small-size firm.

These results are explicative of some relevant figures about the different impact of individual characteristics on wages, with respect to gender. In next section we proceed by analysing the distributions we generate from these estimates, the above-defined variables $\hat{r}_f$ and $\hat{y}_f$. The paragraph is a necessary step in order to define the methodology to use for evaluating discrimination.

5. Distribution Analysis

According to the traditional approach, we can detect the existence of gender discrimination in wages using an average index calculated as the difference between estimated and reference earnings for female workers. Following this approach, we have performed a *t*-test (for dependent samples)[22] to evaluate the difference in means between estimated and reference earnings. The results, shown in Table 5, confirm that the mean of reference earnings is significantly higher than the mean of estimated earnings in both 1990 and 1997 thus suggesting the existence of significant gender discrimination in wages.

However, as we pointed out in Section 3, this approach to the study of wage discrimination can lead to misleading results, particularly when the shape of the distributions of estimated and reference earnings is different. Indeed, an identical shape in the two distributions is just a necessary, but not a sufficient, condition for the validity of a test on wage discrimination based on sample means. In order to shed further light on the presence and extent of discrimination we can therefore move to the direct analysis of the shape of the distributions. Therefore, Figure 1 reports the Epanechnikov kernel estimates of the probability density functions for both estimated and reference earnings. The visual inspection of the density estimates shows that the differences between the two distributions are not confined to their means, as documented above, but extend to other distribution moments. A direct confirmation of this is given by the results of a Kolmogorov-Smirnov nonparametric test reported in Table 6. Moreover, the test also indicates that estimated earnings are significantly higher than corresponding reference earnings.

Two broad implications seem to emerge from these results. On the one hand, they suggest the existence of statistically significant differences between the two distributions which can be ascribed to wage discrimination not uniformly distributed among female workers. On the other hand, they emphasise the dangers of any inference based on the traditional approach. Consequently, an alternative approach to the analysis of wage discrimination will be presented in what follows. In particular, the approach developed here, making use of bivariate density estimates, studies directly the shape of

[22] In principle, this test requires the variables to be normally distributed but our estimates and reference wages fail to pass such test. However, in the presence of large sample sizes such as in our case, we know from asymptotic theory that sample means tend to be normally distributed. Consequently, we have chosen to employ a *t*-test on sample means (with heterogenous variances) rather than a less statistically powerful nonparametric test.

the distribution of wage discrimination, thus allowing to evaluate the incidence of discrimination not only in relation to the level of wage but also to the main workers' characteristics.

6. An analysis of wage discrimination

A first step in our distributional approach to the analysis of wage discrimination is represented by an estimate of the bivariate density function of estimated and reference earnings. Indeed, a contour plot of this density function allows us not only to understand whether there is discrimination but also to evaluate the relationship between the extent of discrimination and the size of the estimated wage. In particular, the 45-degree diagonal in the contour plot highlights absence of wage discrimination; if most of the graph were concentrated along this diagonal, then estimated and reference wages for female workers would coincide. On the contrary, a concentration of the probability mass away from the 45-degree diagonal and closer to the reference wage axis indicates that substantial wage discrimination against female workers exists.

Figure 2 shows the contour plots of the bivariate density function estimated via the Epanechnikov kernel both in 1990 and 1997. Three main features seem to emerge. First, in both years the probability mass is divided into two separate parts thus suggesting a different incidence of discrimination in relation to the wage level; while there appear little signs of discrimination for low wage levels, positive discrimination against female workers generally manifests for higher wage levels. Second, some features of the wage discrimination appear to have changed over time. To facilitate the comparison, Figure 3 thus combines the 0.001 contour lines for the two years on the same Cartesian plane. Starting from 1990, we can note that the upper part of the probability mass is not only entirely underneath the 45-degree diagonal but it also tends to get further away from the diagonal as the estimated wage (on the vertical axis) increases. In other words, it appears that in this year wage discrimination tends to be stronger for higher wage levels. Moving to 1997, the upper part of the probability mass becomes now less concentrated than in 1990, extending both upwards (beyond the diagonal) and downwards thus suggesting an increase in dispersion. Moreover, differently from 1990, the extent of wage discrimination does not appear to increase significantly as the estimated wage increases.

However, an important aspect of these results must be emphasised. The observed change in the shape of the probability mass over time may come from two different sources. On the one hand, it may be the outcome of a change in the differential reward to male and female workers for a given set of characteristics in the sample; on the other hand it may be brought in by a change in the composition of the sample in terms of worker characteristics, given the gender differences in reward.

In order to shed light on this aspect it is therefore necessary to study directly the effects of some of these characteristics on the shape and location of the bivariate distributions. However, while some of these characteristics, such as the tenure within the firm or the previous experience, are measured on a continuous space, others are entered into the analysis as dichotomic dummy variables as it is the case of the level of

education or the type of contract. In the latter case, we can easily isolate the role of a given characteristic by estimating the bivariate density function of estimated and reference earnings conditional on that characteristic. In other words, we simply estimate the bivariate density function for those workers who share the chosen characteristic and the interpretation of these conditional estimates is therefore analogous to the unconditional estimates described above.

However, when the characteristics are measured on a continuous space, we have to follow an alternative route. In these cases, we first calculate a wage discrimination variable, as the difference between reference and estimated earnings, and then produce a contour level plot of the bivariate density function of wage discrimination (on the vertical axis) and the chosen characteristic (on the horizontal axis). Obviously, in these types of plot, the absence of wage discrimination would be represented by a concentration of the probability masses along the horizontal line corresponding to a value of zero on the vertical axis.

A first example of this type of analysis is presented in Figure 4 that allows evaluating the role played by the overall experience (measured in years) gained in the labour market prior to present occupation. A first general feature common to both years is that most of the probability mass corresponds to positive values for the difference between reference and estimated earnings. Moreover the probability masses tend to concentrate around positive and increasing levels of discrimination as general experience level increases. However, other features should be emphasised. For very small levels of experience, the comparison between the 1990 and the 1997 plots suggests a noticeable downward shift in the probability mass. So, while in 1997 there still is significant wage discrimination against female workers with low experience, the extent of discrimination attributable to this characteristic appears to have declined. In contrast, the variability of wage discrimination appears to have increased for higher levels of experience. Finally, only for particularly high levels of experience there is a marked upward shift in the probability mass implying a significant increase in discrimination. These results allow us to highlight the clear advantage from the adoption of a distributional analysis approach. According to the traditional discrimination methodology, the increase in the gender premium differential together with a higher average experience would suggest an increased proportion of discrimination due to that characteristic. In contrast our results show a much more complex picture and, in particular, that a sharp increase in discrimination is confined to workers with the highest level of experience.

Rather similar results (Figure 5) emerge from the analysis of the role played by the tenure within the firm. As it was the case with experience, the location and the shape of the probability mass suggests the presence of significant wage discrimination against females in both years with an increase in the phenomenon throughout the period. However, a very interesting feature is the substantial increase in the share of the probability mass that lies underneath the horizontal no-discrimination line; female workers that in 1997 have a relatively small number of years of tenure appear to be less discriminated than both their colleagues with higher tenure and female workers with a similar level of tenure in 1990.

Figure 6 reports the bivariate density of estimated and reference earnings conditional to different levels of education. As we noted in Figure 2, the probability mass is divided into separate parts both in 1990 and in 1997, thus suggesting that the role played by the

education level on the degree of wage discrimination varies depending on the level of earnings. In general, in both years and for both education levels, there appears to be no sign of discrimination, or even discrimination in favour of female workers, for low wage levels. However, as this level increases, the extent of wage discrimination also appears to increase. Indeed, this feature manifests itself in both years in the case of primary education and is also evident for workers with secondary or university degree in 1990. Conversely, the situation appears to have significantly improved for female workers that achieved a higher level of education in 1997; the corresponding probability mass has not only extended upwards beyond the 45-degree diagonal, but its overall slope appears now steeper. To get a direct confirmation on this result, Figure 7 compares the probability masses in the two years for each level of education. While their shapes and locations do not appear to have changed substantially for primary education, a significant improvement manifests for workers with a higher level of education. Indeed, not only the upper part of the distribution has extended upwards as noticed previously, but there now appears to be no discrimination or even discrimination in favour of female workers, although for very low levels of earnings only.

A further interesting aspect is the relationship between discrimination and different types of contracts such as apprenticeship or “CFL”. Figure 8 thus plots the bivariate density of reference and estimated earnings conditional to being on an apprenticeship contract. It is immediate to note that the probability mass is concentrated on the 45-degree diagonal in both years, thus suggesting the absence of gender discrimination among workers on this type of contract. Moreover, these masses are located quite close to the origin of axes, thus corresponding to low levels of earnings. Therefore, a comparison between this plot and those in Figures 2 and 3 seems to suggest that the previously emphasised absence of discrimination among low pay workers can be attributed to this type of contract. Moving to the effect of the “CFL” contract (Figure 9) another very interesting result emerges: the noticeable degree of wage discrimination emerging from the 1990 estimates completely disappeares in 1997 even at relatively high levels of earnings.

7. Conclusions

This paper has addressed the issue of gender wage differentials. From a strictly methodological point of view, given the important limitations of the traditional approach, we have presented an alternative methodology. We shift our focus away from simply looking at the average level of discrimination and study directly the entire distribution of wage differentials making use of bivariate density function estimates.

Regression results confirm the positive effect of the accumulation of human capital on wage profiles, both for women and for men. With regard to education, workers with a university degree gain higher wages in comparison to those with a diploma, in both years; general and specific experience positively affect wages following a decreasing return function. Relevant differences have emerged between the two years. In 1997,

compared to 1990, the female advantage in the accumulation of firm-specific human capital has reduced and become lower than that of males. With regard to the returns to education, gender differences are more pronounced in 1990 than in 1997 in particular for tertiary level education. Our results are fundamentally in line with those of other works on gender wage differentials among young workers [Corcoran and Duncan (1979), Polachek and Robst (2001)].

The empirical illustration of the distributional approach to wage discrimination has provided some interesting results. Generally, the labour markets for young workers in the two North-Eastern Italian provinces appear to have been characterised by substantial wage discrimination against female workers during the 1990s. This is particularly so for workers with higher work experience, tenure and wages. However, some encouraging signs also emerge; a tendency towards a reduction in the extent of discrimination is associated with having higher educational qualifications or being on a vocational training contract. On the other hand, the positive association between wage discrimination and tenure and, even more significantly, between discrimination and work experience have become stronger in 1997.

Gender wage differentials among young workers: methodological aspects and empirical results

Summary

We evaluate the gender wage gap and the unexplained gender earnings differential for workers 15-29 year old in the labour markets of two Italian provinces during the period 1990-1997, using administrative data from the Italian Social Security System.
Our starting point is the critique of Jenkins to the traditional methodology employed in the analysis of discrimination, methodology which focuses on the comparison between the average of the estimated female wage distribution and the average of the reference wage distribution. After investigating the existence of differences between the two earnings distributions we evaluate gender discrimination by studying the entire distribution of the unexplained wage gap rather than focusing on mean wage differences only, as usually done in the more conventional Oaxaca approach. Finally we present the analysis of the incidence of different human capital characteristics on the dimension of the unexplained wage gap.

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Tables and figures

Table 1
Summary information on the sample

	1990		1997	
	Women	Men	Women	Men
	35583	39151	31819	37063
Blue-collar	23868	32357	18485	29869
White-collar	11715	6794	13334	7194
Full-time	34849	39093	30802	36971
Part time	734	58	1017	92
CFL	5862	6284	2214	2707
Apprenticeship	4821	4505	2482	2196
Metals, precision tools (sector 3)	5732	17225	7513	17742
Food, clothing, wood (sector 4)	21857	14664	16008	12866
Commerce (sector 6)	5351	6211	5309	5307
Credit (sector 8)	2643	1051	2989	1148
Firm size 1_5	9661	9257	7551	7115
Firm size 6_10	6056	5434	4116	4578
Firm size 11_20	8078	6303	6452	5850
Firm size 21_50	4553	6456	5069	6820
Firm size 51_100	2524	3976	3228	4924
Firm size > 101	4711	7725	5403	7776
Primary education	24558	27778	16501	21933
Secondary education	10530	10532	14438	13485
University education	495	841	880	1645

Table 2
Mean and standard deviation of experience and tenure

		1990		1997	
		Mean	S.D.	Mean	S.D.
Experience					
	All	99.66	135.61	112.21	141.18
	Women	91.79	134.20	109.62	144.77
	Men	106.11	136.50	114.42	137.99
Tenure					
	All	212.95	158.89	225.10	151.29
	Women	217.12	163.80	231.61	154.91
	Men	209.16	154.20	219.52	147.87

Table 3
Daily earnings (1000 lire) and wage gap by quartiles and occupation*

	1990			1997		
	Women	Men	Wage gap	Women	Men	Wage gap
Mean	77.24	89.43	13.6	82.30	95.73	14.03
S.D:	16.75	23.72		19.31	24.20	
10%	58.86	66.79	11.9	65.03	73.13	11.08
25%	70.51	76.65	8.00	72.19	81.12	11.00
Median	75.67	86.26	12.28	80.21	91.70	12.53
75%	83.52	98.58	15.28	90.11	105.64	14.70
90%	94.19	115.50	18.45	102.70	124.00	17.18
			White-collar			
Mean	88.98	109.90	19.03	94.12	112.57	16.39
S.D.	19.08	33.29		19.84	30.99	
10%	72.04	78.09	7.75	75.78	81.68	7.22
25%	77.18	86.20	10.46	81.60	89.89	9.22
Median	84.34	101.77	17.13	89.34	104.69	14.66
75%	95.54	126.52	24.49	100.80	128.35	21.46
90%	111.63	151.86	26.49	118.26	153.69	23.05
			Blue-collar			
Mean	71.84	85.17	15.65	74.43	91.73	18.86
S.D.	12.23	18.54		14.29	20.33	
10%	54.52	62.66	12.99	56.98	71.84	20.68
25%	68.22	75.34	9.45	69.02	79.48	13.16
Median	73.12	84.44	13.40	74.74	89.54	16.53
75%	78.46	94.93	17.35	82.42	101.70	18.96
90%	84.86	106.42	20.26	89.93	115.46	22.11

* Only full-time workers. Wage gap in percentage

Table 4
a) Earnings functions. Dependent variable logW*. Year 1990

	Women		Men	
	Coefficient	S.E.	Coefficient	S.E.
Constant	3.4843	.0129	3.6121	.0388
Experience**	.3380	.0208	.5248	.0192
Experience^2	-.4304	.0483	-.4945	.0404
Tenure**	.7997	.0279	.7790	.0268
Tenure^2	-.8382	.0384	-.7218	.0381
Secondary ed.	.0449	.0027	.0602	.0026
University ed.	.0882	.0120	.1134	.0085
White-collar	.1632	.0030	.1601	.0033
Full-time	.6129	.0101	.7788	.0375
CFL	-.0258	.0030	-.0816	.0028
Apprenticeship	-.1709	.0037	-.3049	.0039
Sector 3	.0594	.0059	-.0607	.0093
Sector 4	.0421	.0061	-.1140	.0093
Sector 6	.0796	.0054	-.0932	.0094
Firm-size 1-5	-.0336	.0027	-.0639	.0029
Firm-size 6-10	-.0115	.0025	-.0192	.0031
Firm-size 21-50	.0335	.0027	.0233	.0029
Firm-size 51-100	.0507	.0038	.0379	.0034
Firm-size 101+	1002	.0032	.0701	.0029
Observations	35583		Observations	39151
R^2	0.5001		R^2	0.5516

Table 4
b) Earnings functions. Dependent variable logW*. Year 1997

	Women		Men	
	Coefficient	S.E.	Coefficient	S.E.
Constant	3.4664	.0105	3.4678	.0302
Experience**	.3356	.0262	.5716	.0223
Experience ^{2}	-.3811	.0589	-.4046	.0506
Tenure**	.8281	.0304	.8503	.0269
Tenure^2	-.7741	.0441	-.6194	.0411
Secondary ed.	.0463	.0030	.0576	.0028
University ed.	.1096	.0109	.1097	.0058
White-collar	.1963	.0035	.1519	.0032
Full-time	.6287	.0087	.7862	.0297
CFL	-.0117	.0044	-.0654	.0036
Apprenticeship	-.1186	.0045	-.2379	.0046
Sector 3	.0383	.0026	.0633	.0021
Sector 6	.0454	.0039	.0278	.0032
Sector 8	.0357	.0069	.1323	.0085
Firm-size 1-5	-.0496	.0035	-.0756	.0034
Firm-size 6-10	-.0178	.0034	-.0193	.0037
Firm-size 21-50	.0468	.0033	.0289	.0033
Firm-size 51-100	.0652	.0037	.0301	.0036
Firm-size > 101	.1097	.0036	.0633	.0032
Observations	31819		Observations	37063
R^2	0.4595		R^2	0.4272

** Estimates corrected by the White var-cov matrix.*
*** 'Tenure' and' experience' are scaled by 1000*

Table 5

t-test on the equality of $\hat{y}^f$ and $\hat{r}^f$ averages

	$H_a : E(\hat{y}^f) - E(\hat{r}^f) < 0$	$H_a : E(\hat{y}^f) - E(\hat{r}^f) > 0$	$H_a : E(\hat{y}^f) - E(\hat{r}^f) \cong 0$
1990	$t = -310.6676$ $P < t = 0.0000$	$t = -310.6676$ $P > t = 1.0000$	$t = -310.6676$ $P > \|t\| = 0.0000$
1997	$t = -188.9448$ $P < t = 0.0000$	$t = -188.9448$ $P > t = 1.0000$	$t = -188.9448$ $P > \|t\| = 0.0000$

Table 6

Kolmogorov-Smirnov equality of distribution test (non-scaled $\hat{y}^f$ vs. non-scaled $\hat{r}^f$)*

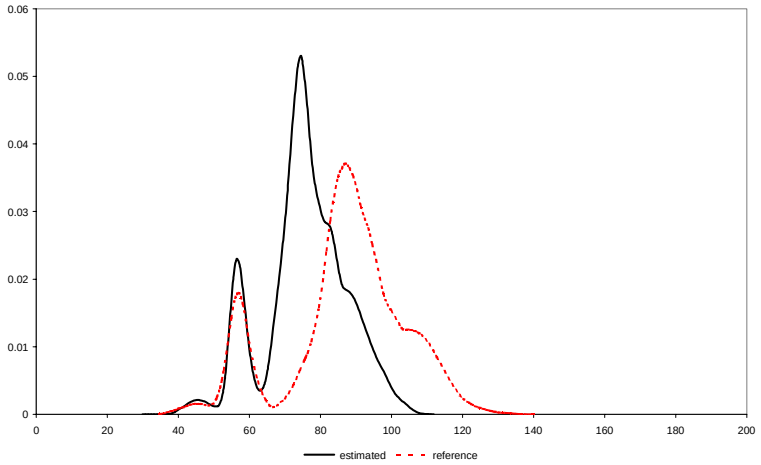
	Type	D	*p*
1990	1	0.4423	0.000
	2	-0.0072	0.159
	combined	0.4423	0.000
1997	1	0.3413	0.000
	2	-0.0135	0.003
	combined	0.3413	0.000

*Type 1 tests the hypothesis that $\hat{y}^f$ values are lower than $\hat{r}^f$ values. Type 2 *tests the hypothesis that* $\hat{y}^f$ *values are bigger than* $\hat{r}^f$ values.

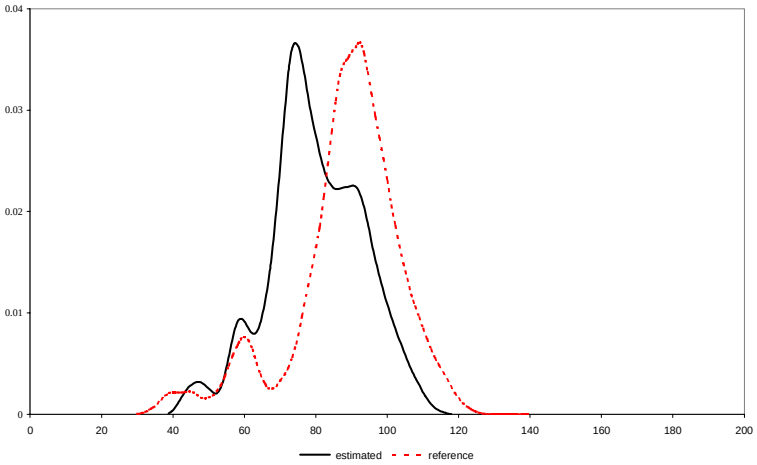
Figure 1 - Probability density functions (Epanechnikov kernel)

a) estimated and reference earnings (absolute)

1990

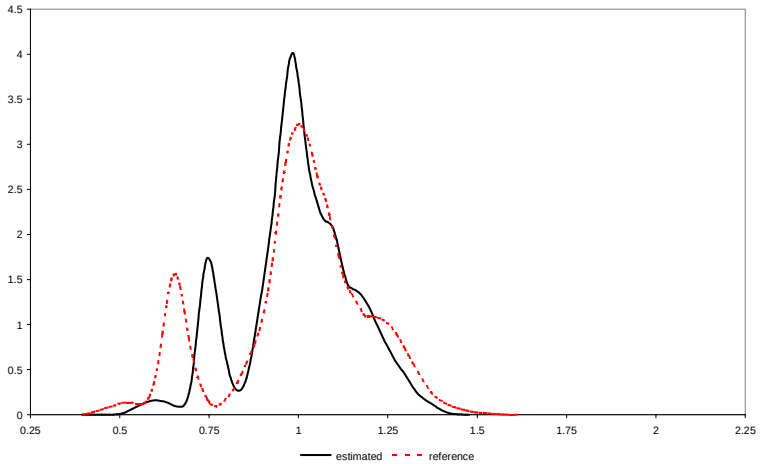


1997



b) estimated and reference earnings (relative to sample mean)

1990



1997

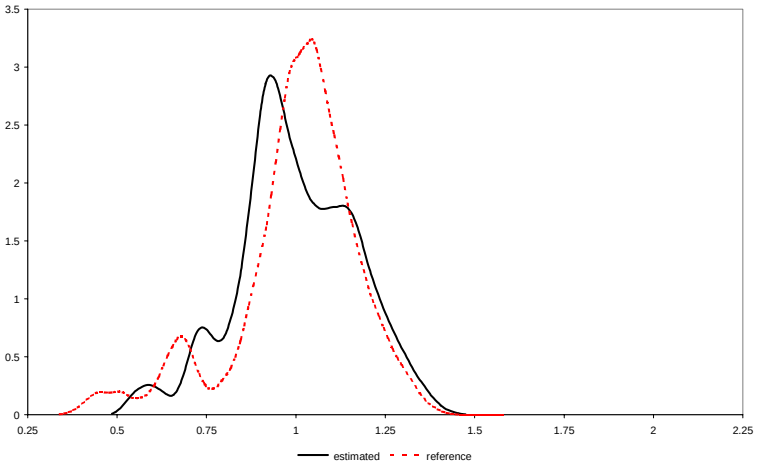
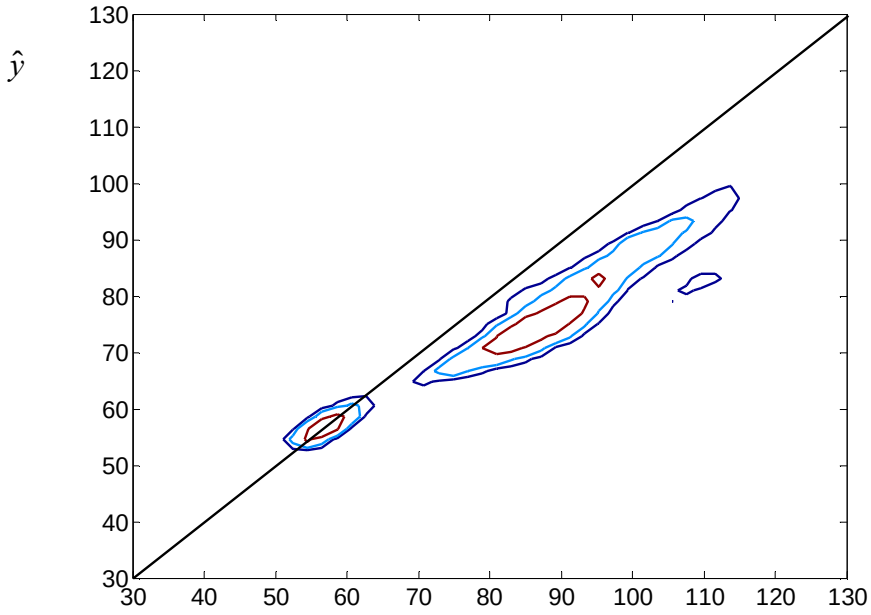


Figure 2
Bivariate Density of $\hat{y}$ and $\hat{r}$ - contour plot at levels 0.001, 0.002 and 0.005

1990



1997

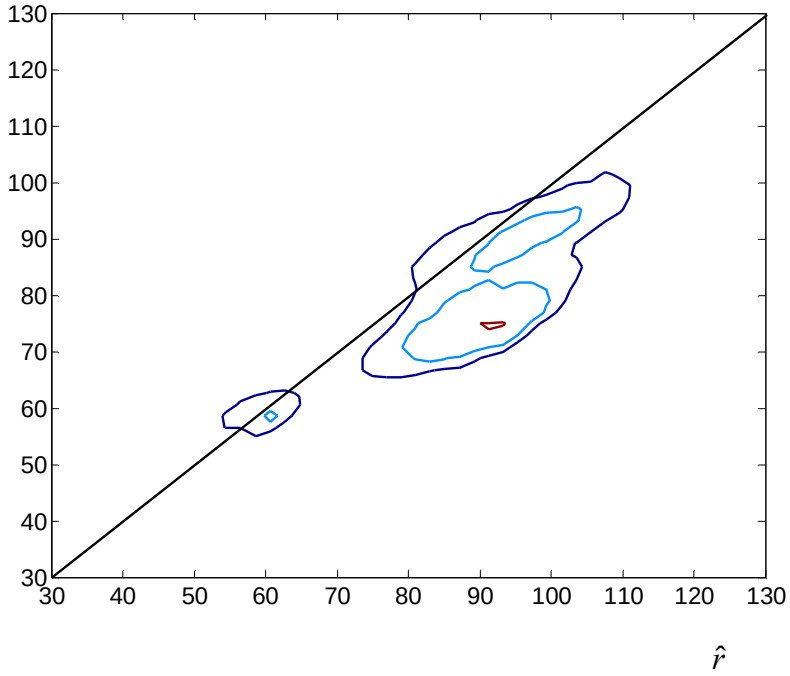
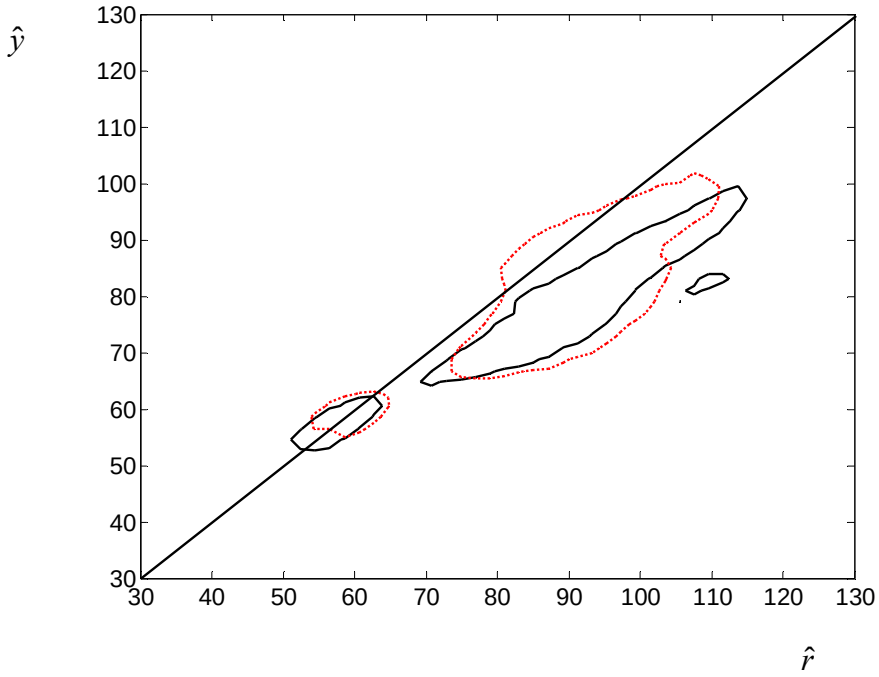


Figure 3
Bivariate Density of $\hat{y}$ and $\hat{r}$ - contour plot at level 0.001

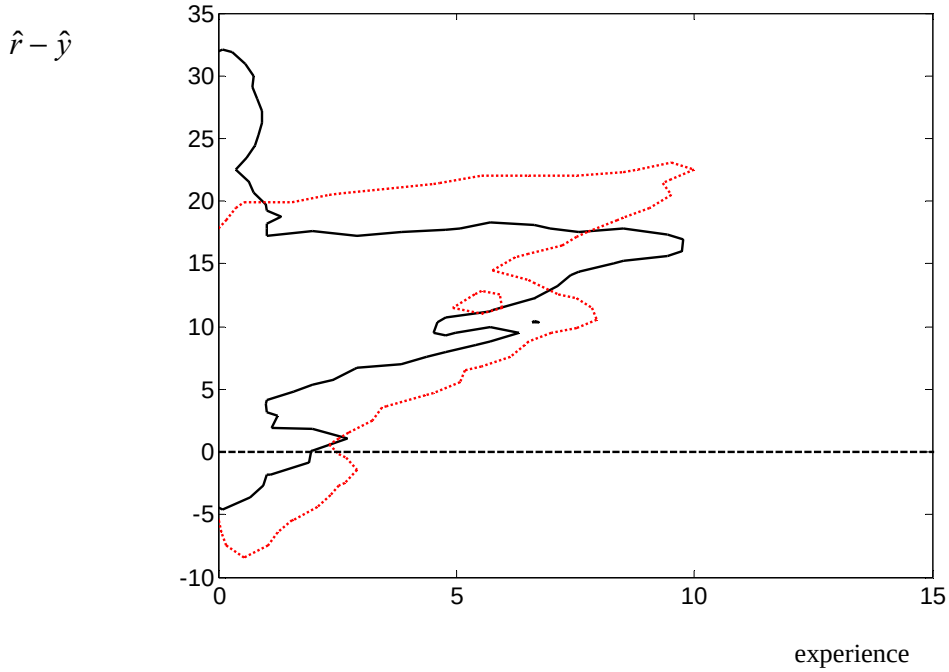
1990 - 1997



1990: solid (black) line
1997: dashed (red) line

Figure 4
Bivariate Density of Wage Discrimination and Experience - contour plot at level 0.001

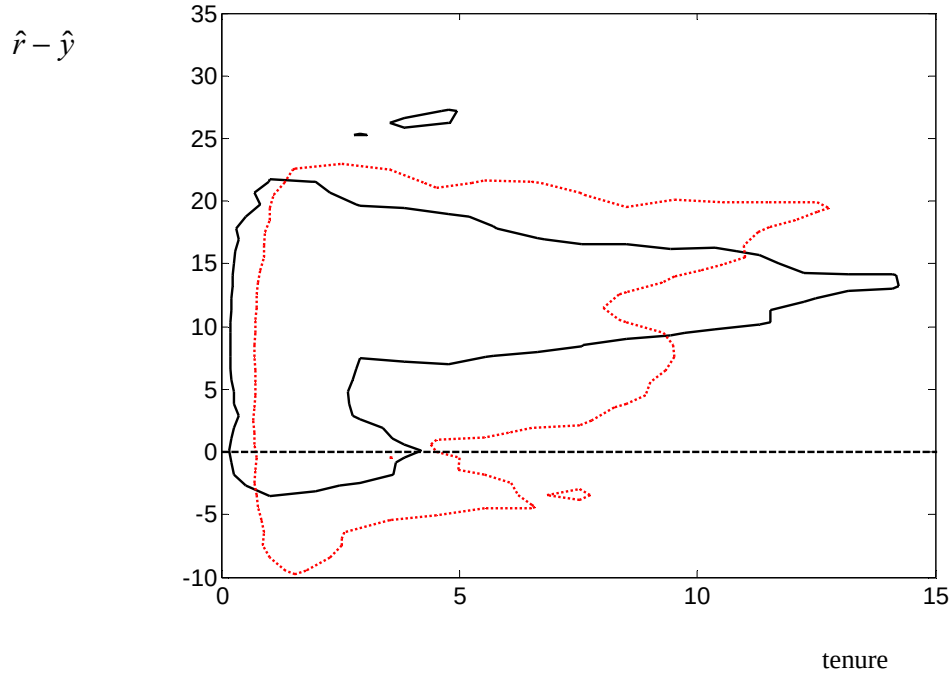
1990 - 1997



1990: solid (black) line
1997: dashed (red) line

Figure 5
Bivariate Density of Wage Discrimination and Tenure - contour plot at level 0.001

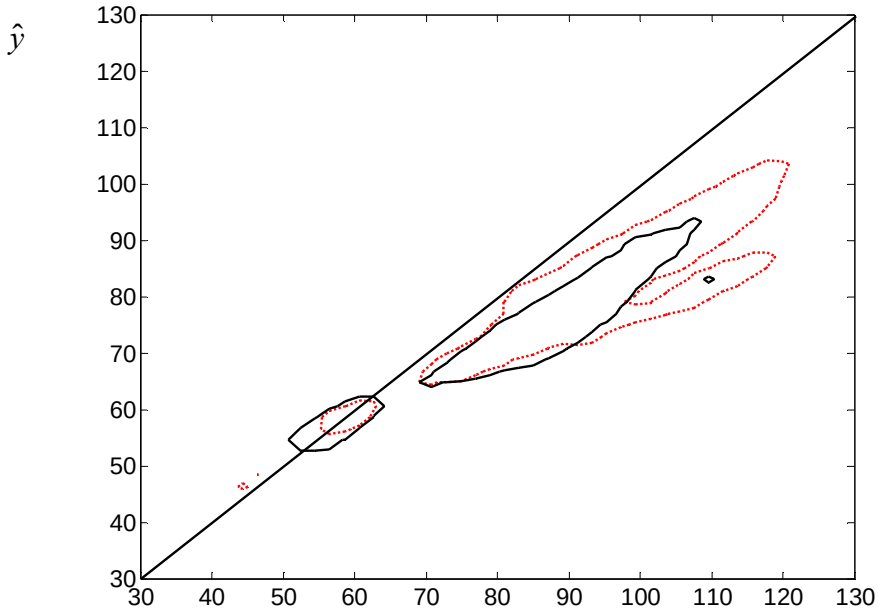
1990 - 1997



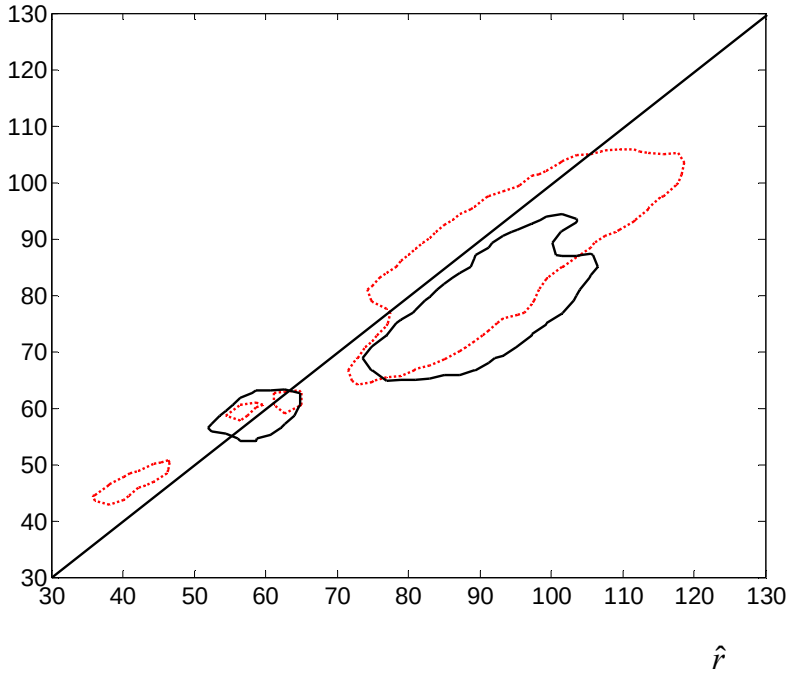
1990: solid (black) line
1997: dashed (red) line

Figure 6
Bivariate Density of $\hat{y}$ and $\hat{r}$ conditional to education level - contour plot at level 0.001

1990



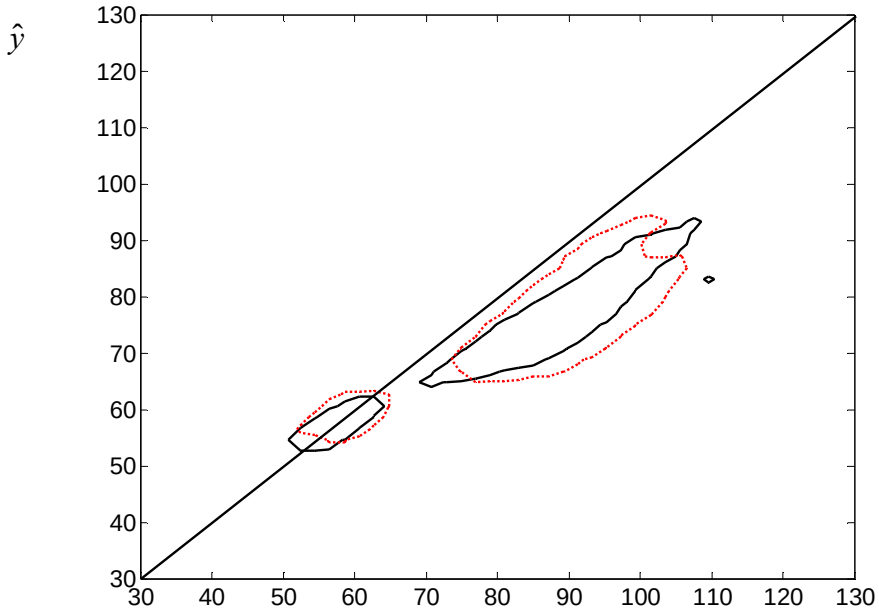
1997



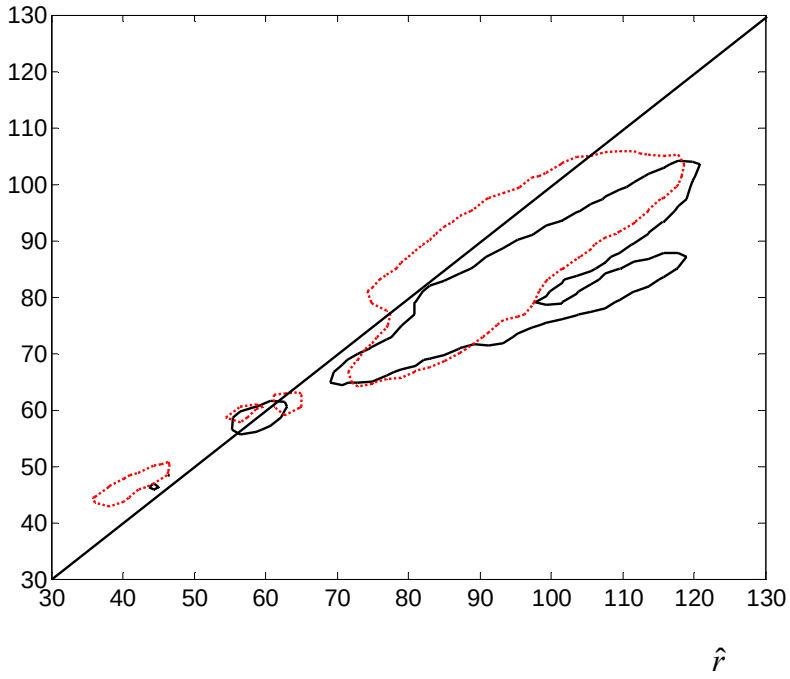
Primary: solid (black) line
Secondary or University: dashed (red) line

Figure 7
Bivariate Density of $\hat{y}$ and $\hat{r}$ conditional to education level - contour plot at level 0.001

Primary



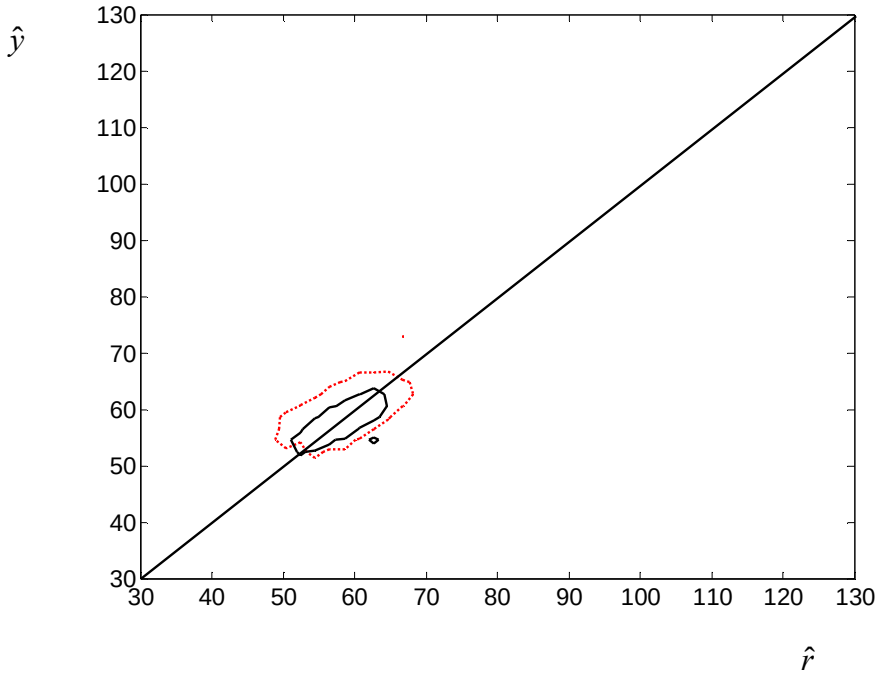
Secondary or University



1990: solid (black) line
1997: dashed (red) line

Figure 8
Bivariate Density of $\hat{y}$ and $\hat{r}$ conditional to apprenticeship contract - contour plot at level 0.001

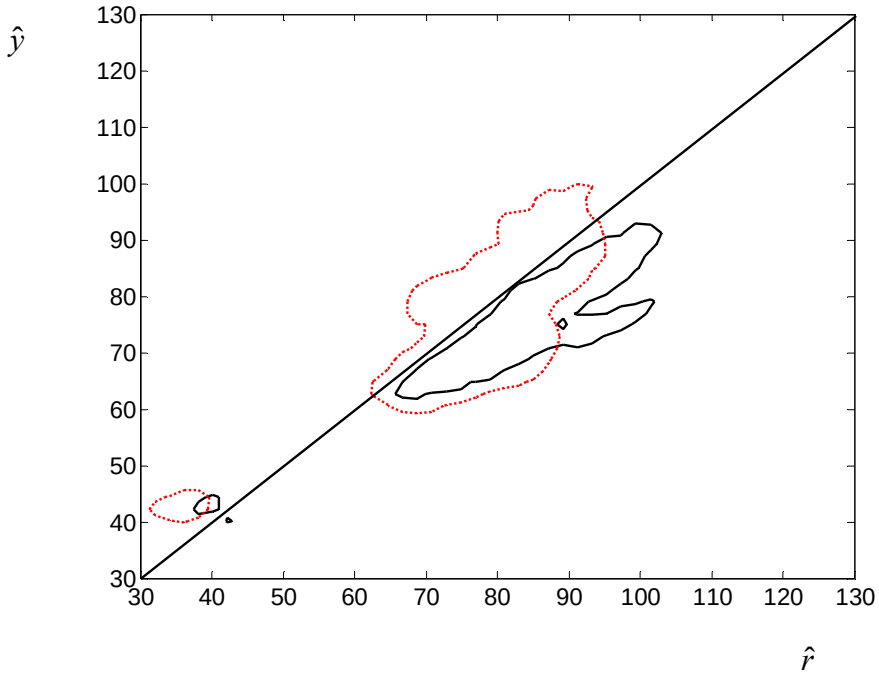
1990 - 1997



1990: solid (black) line
1997: dashed (red) line

Figure 9
Bivariate Density of $\hat{y}$ and $\hat{r}$ conditional to "CFL" - contour plot at level 0.001

1990 - 1997



1990: solid (black) line
1997: dashed (red) line

Working Papers già pubblicati

1. E. Battistin, A. Gavosto e E. Rettore, *Why do subsidized firms survive longer? An evaluation of a program promoting youth entrepreneurship in Italy*, Agosto 1998.
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