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**QUADERNI DELL'ISTITUTO DI
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The Wage Effect of Working in the Public Sector When Education and Sector Choices Are Endogenous: An Empirical Investigation for Italy

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Abstract: This paper uses a pooled cross-section from 1995, 1998, 2000 and 2002 household survey for Italy and estimate the public wage premium for male white collars. The empirical model consists of a wage equation with endogenous education and an endogenous dummy for sector affiliation with a random coefficient (the wage premium). I find that, when only the second source of endogeneity is taken into account by means of correction terms, selectivity issues concerning optimal schooling decisions affects the parameters' estimate: while the true value of returns to education is higher than OLS predictions, the public wage premium obtained is lower than the true one. For what concerns endogenous sector choices, results show that that public employees are negatively self selected and, accordingly, the wage premium they earn is still positive (9.6%), but lower than for an arbitrary individual (15%).

JEL codes: J31, J45, C3.

Keywords: Public Sector Wage Premium, Endogeneity, Italy.

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1. Introduction

In many industrialised countries the state accounts for a significant share of total employment and, both by producing good and services and by regulating the activity of the private sector, deeply influences the functioning of the entire economy. In this context, Italy is not an exception, and public intervention in the economy is substantial, and occurs at the central (government) and at the local level, as well as by means of public authorities.

By their nature, goods and services offered by the public sector are essential and (often) produced from a monopolistic position, and rules governing employment conditions, human resource management and pay determination are intrinsically different across public and private sectors. For example, in Italy large differences exist in recruitment, retention and incentive policies, as well as in careers and wage profiles. Moreover, the availability of most occupations is not the same across the two sectors.

In the public sector, once hired through a public examination, a civil servant enjoys a lifetime working contract and seniority plays a key role in wage progression. In addition, at least until mid '90s, incentives relating wages to productivity were often missing.

In the private sector, although the power of "insiders" is still substantial, the degrees of flexibility in wage determination are higher and the criteria to hire and promote workers are less strict than in the public sector. In addition, as a consequence of higher union power and because the State aims to be perceived as a "good employer" by offering (relatively) high wages to low skilled workers and (relatively) low wages to the high skilled, the wage structure in the public sector tends to be flat compared to the private sector.

In the light of the above considerations, this paper aims to estimate the wage premium earned by Italian public employees using a pooled cross section from the 1995, 1998, 2000 and 2002 Bank of Italy's Survey of Households Income and Wealth.

The wage effect of working in the public sector is estimated using the so-call treatment effect model, which, in my case, is a standard mincerian wage equation augmented by a binary indicator for public sector affiliation. Under this specification, which is commonly employed in the evaluation literature, the coefficient of the sector dummy measures the "treatment effect", which is the mean difference in outcomes (in this case wages) between the two alternative states of the world (working in the public or in the private sector) for public employees. The estimation of this equation, however, is complicated by a number of issues.

First of all, as a consequence of differences in job attributes, working condition and hiring requirement, some employees may display preferences for the public sector, self-selecting themselves according to unobservable characteristics. Since in this case the treatment dummy is endogenous, the OLS estimator of the wage premium is biased and inconsistent. A number of studies for Italy acknowledged the potential endogeneity of public employment by estimating models with controls for this selectivity source.

Second, as pointed out by the vast literature on returns to schooling and confirmed by recent studies for Italy – see, for example, Brunello and Miniaci (1999) and Colussi (1997) -, the educational attainment is correlated with unobservable wage determinants and, therefore, is likely to be endogenous in the wage equation. For the most part, the labour literature employs IV techniques to solve this problem. Quite surprisingly, a control for the potential endogeneity of education, which enters as a determinant of both the wage and the sector choice, has been never included in studies of public/private wage differentials for Italy[1]. However, since endogeneity of schooling decisions may affect the estimation of the whole vector of model's parameters – and not only the return to education -, existing evidence may report biased estimates of the wage premium. The intensity of the bias depends on the degree of correlation between education and sector decisions, which is likely to be present in Italy, where the schooling level plays a key role to be recruited by the public sector.

To avoid issues of endogenous female labour market participation, I focus on males only. Furthermore, I restrict the analysis to white collars, since blue collars occupations in the public sector are hardly comparable to their private sector counterpart[2].

This paper contributes to the existing literature on the public sector premium in many aspects.

First of all, from an econometric point of view it aims at bringing together the two strands of literature on endogenous education and endogenous sector affiliation in

[1] Brunello and Dustmann (1997) recognise the importance of this problem but they do not try to solve it.

an unified framework. Following the seminal paper of Dustmann and Van Soest (1998) I endogenise education in a model which contains yet an endogenous dummy for public sector affiliation. However, differently from the above authors, whose estimation procedure is fully parametric, I use a three stage procedure which combines IV techniques (for education) with control function methods (for sector affiliation). From an economic point of view, by modelling the relationship between wages and education and sector decisions, this paper may also contribute to the general debate about the labour market effects of the reforms recently introduced in the Italian schooling system and public sector[3].

Second, as an alternative to the standard treatment model, where the coefficient for the sector dummy is constant, and, therefore, the wage premium is constrained to be equal across different individuals, I also estimate a more flexible specification with random coefficients for the treatment dummy, i. e. where the effect of working in the public sector is heterogeneous across individuals. A general discussion about the estimation of random coefficient models in the context of treatment effects can be found in Heckman and Robb (1985). As explained in more detail below, the random component of the coefficient associated with the sector dummy has an interesting economic interpretation. In fact, it measures the difference between individual unobserved wage components (potential productivity) in the two sectors. If this difference is positive, public employees are efficiently allocated in the sector in which their productivity is higher. If the difference is negative, public employees are allocated into the sector where their potential productivity is lower. Since the empirical strategy provides an estimate of the mean of this difference for public employees, the efficiency of the State as an employer can be investigated.

I find that, first, the assumptions about the nature of education do have an effect on the estimation of the public wage premium. More specifically, under exogenous education the premium is positive but not statistically significant at the usual levels. Instrumenting years of schooling increases the return to education and, at the same time,

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[2] In principle, the occupation status may not be exogenouly determined. However, since I instrument the education level, the issue of endogenous occupation is attenuated to the extent that education maps into occupational levels.

[3] The so called "privatisation" of the public sector started at the beginning of the '90s with the aim to introduce higher degrees of flexibility and exposure to market forces in the public employment, in order to provide public sector employees to economic incentives comparable to those existing in the private sector (performance related pay schemes are an example).

improves the precision in the estimate of the wage premium, which benefits from the elimination of the simultaneous correlation between education and sector choice that contaminates the OLS results.

Second, as compared to the model with a constant premium, the specification with random coefficients is more informative since it clarifies the reasons why people selects into the public sector and the effect of their decision on the wage premium they earn.

Main results from the model with constant treatment are that sorting in schooling and in the public sector are both negative. By controlling only for selection in sector choices and not for endogenous education the public wage premium is positive but quite poorly estimated (11.4%). By assuming selection in both preocesses, the wage premium increases (16.8%) and becomes significant.

For what concerns the model with random treatment effects, the main additional findings are that the average wage premium in the population (average treatment) is higher than for public employees (the average treatment for the treated) because the latter are negatively self-selected. In other words, on average, the individual-specific part of the wage premium, which measures the difference between earnings potential in the public and private sector, is negative. Accordingly, on average, the sector choice of public employee is based on the presence of non-monetary gains that counterbalance potential wage losses.

The remaining part of the paper is organised as follows. Section 2 contains a brief review of the literature. In Section 3 the main features of the data are described. Section 4 introduces the econometric framework and discuss, under different assumptions, estimation techniques and the identification strategy. Main results are offered in section 5. Conclusions follow in section 6.

2. Literature Review

In the last thirty years, the evidence that both wages and working conditions greatly differ between the public and private sector has stimulated a large debate over the (competitive or non-competitive) explanations for the wage effect of working in the public sector (see Ehrenberg and Scwarz, 1986). Starting from Smith (1977), the empirical literature typically focussed on the estimation of (conditional) structural differences between public and private wages. For the most part, US studies report that the public sector has a less elastic labour demand curve and that rents are being earned by public sector employees with respect to private sector workers with comparable (observable) characteristics.

However, as a consequence of differences in job attributes, working condition and hiring requirement, the choice of the sector might not be random. In particular, some employees may exhibit a preference for the public sector, self-selecting themselves according to unobservable characteristics. Since in this context the OLS estimator is biased and inconsistent, a number of studies for different countries estimated wage regressions using maximum likelihood and two-step Heckman methods to compute public pay premium free of selectivity bias. Examples are Hartog and Oosterbeek (1993) for Netherlands, Belman and Heywood (1989) for the US, Disney and Gosling (1998) for UK and Adamchick and Bedi (2000) for Poland. As a consequence of differences in wage structures, institutional settings and workforce selection mechanisms across countries, results from these studies show a great deal of variation in the estimated premium.

Dustmann and Van Soest (1998), who add an equation for schooling decisions to a model with endogenous switching between two separate wage regimes (public and private), performed the analysis most close in the spirit to the one proposed here. Using data for German male workers they find that the assumption of schooling exogeneity affects the estimates of returns to education and to public employment. Albeit the statistical models are similar, the estimation procedures differ substantially. As it will become evident in Section 4, while they use maximum likelihood methods to estimate simultaneously all the equations under the assumption that the corresponding errors are jointly distributed as a multivariate normal, I will develop a consistent sequential procedure which requires less distributional requirements.

For what concerns Italy, OLS estimates of the public-private wage differential (Cannari et al, 1989: Brunello and Rizzi, 1993; Brunello and Dustmann, 1997; Lucifora, 1999; Comi and Ghinetti, 2002) varies in the range of 9-12% depending on the period considered, the sample used, the specification adopted and the definition of public sector employed.

Estimates obtained controlling for endogenous sector choices vary considerably more than those obtained from OLS: using the single equation model with an endogenous dummy for sector affiliation Cannari et al. (1989) and Brunello and Rizzi (1993) find that the wage differential is not significantly different from zero; using the more flexible specification with endogenous switching Brunello and Dustmann (1997) report that, at least for males, the premium is positive (21%) and it can be largely explained by observable workers' attributes, while Bardasi (1996), who uses a more sophisticated selection procedure – workers can choose to work in the public sector, in the private sector or to be self-employed -, finds that the observed differential is substantial for women (35%) and smaller for men (8.8 %), and that the larger contribution (40% for male, 50% for women) comes from different returns paid to similar characteristics while the effect of different (observed) characteristics is not significant[4].

Also the magnitude of selection effects varies considerably across studies: according to Cannari et al (1989) and Brunello and Rizzi (1993) they are weak, whereas Brunello and Dustmann (1997) find no evidence of such effects; Bardasi (1996) reports that significant and negative endogenous selection exists in the public-private occupational choice.

Overall, results from the above studies show a rather small and often negative or not statistically significant wage differential, and do not find sistematic evidence of selection effects in the choice of the sector.

On the one hand, differences in estimates across studies may reflect the high volatility of the wages in both sectors over the period considered (end of the 80s, beginning of the 90s), which is disturbing when using cross-sections from different years. On the other hand, these differences may depend on the sensitivity of results to model assumptions and identification strategies. A detailed discussion over these issues will be presented in Section 4.

[4] Cappellari (2002) takes an alternative route to the approach based on static differences in earnings between the two sectors and investigates the dynamic of earnings. He finds that life cycle considerations

3. Data and Variables

The data used in this paper are drawn from the 1995,1998, 2000 and 2002 waves of the Survey of Household Income and Wealth (SHIW). Each wave is based on a random sample of around 20,000 individuals. Although the sampling unit is the household, detailed information is available also at individual level, like maximum schooling degree obtained, gender, age, work experience, region of residence and of origin, occupation, (net) yearly earnings, average weekly hours of work and number of months of employment per year. The Survey provides also detailed family background information, like parents' level of education, occupation, sector of employment.

The sample used in the empirical investigation is drawn by the population of non-agricultural workers who are employed and aged from 25 to 60[5]. Younger workers have been excluded to mitigate the problem that positions in the public sector are rationed and the recruitment procedure is typically very long, and individuals with preferences for that sector may have to wait before being employed. Older workers have been excluded to avoid the problem of endogenous retirement.

Using this information, I have constructed a pooled cross-section for the years of interest. The pooling procedure has been employed to improve the asymptotic properties of the estimates by increasing the sample size. In addition, as compared to a single cross section, this structure might help to smooth the effect that the timing of contract renewals, which differs across occupational categories, have on the wage premium. Earlier waves have not been included because bargaining procedures until 1993 were significantly different from those adopted in the subsequent period and family background information was missing.

In order to avoid well-known sample selection problems associated with female labour market participation, I focus the attention on males only. In addition, as pointed out by previous studies on Italian public/private wage differentials (see Bardasi, 1996), the comparison of wages between the public and private sector should be restricted to those occupations that are equally available in both sectors. Since blue collars in the public sector are quite rare and are not directly comparable to manual occupations in the private sector, I restrict the attention to white collars. For similar reasons I exclude the

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matter in the formation of the differential; in the private sector careers are less stable and the growth rate of wages is more volatile that in the public sector, where wages are more homogeneous over the life.

highest occupational group which includes positions, such as magistrate and university professor, for which a private sector counterpart does not exist. These selection criteria reduce the sample to approximately 5,600 observations.

In the survey, one major difficulty is the definition of the public sector, which refers to the Italian "*Pubblica Amministrazione*", which excludes firms financed by the state but operating in the market. For this reason, public employees have been identified also through additional information from the variable "firm size", which classify public employees in a specific category. The second major shortcoming is that no information is available on the number of weeks worked on average in a month. According to all previous studies (see Bardasi, 1996, for a detailed discussion over this issue), hourly earnings are computed (at 1995 prices) assuming that an individual worked 52/12 weeks per month. Wages are inclusive of extra-time compensations and fringe benefits, and net of taxes and social security contributions. The third limitation is that years of schooling are not effective, but imputed on the basis of the higher degree obtained. Thus, education, measured as usual in terms of years of schooling, is a categorical (ordered) variable which can take only positive values. Nevertheless, to for computational purposes, it will be treated as continuous in the empirical analysis. Since the treatment of education as endogenous requires to estimate an equation where education is the dependent variable, due to its categorical nature and non negative values corresponding to corner solutions, this simplification may be thought to affect the results. But, in practice, this does not seem to be the case. A sensitivity check has revealed that results are unaffected when a continuous logaritmic transformation is used in place of the original variable.

For the three years of interest, summary statistics are given in Table 1[6].

< TABLE 1 AROUND HERE>

Public sector employment accounts for more than 40% of the sample. On average, public employees are older (44 years) than private employees (37 years). Among white collars, average years of schooling in the two sectors do not substantially differ (12 years in both sectors). As one may expect, public employees are more concentrated in the south of Italy, while private employees in the north. Differences

[5] Due to data limitations on family background information, the sample is further restricted to household heads, spouses and their children.

between the area of birth and of residence show that there is geographical mobility from the south to the north. Family background information show that older cohorts hold lower level of education and that being sons of parents public employees is more common event in the public sector. Hourly (unconditional) average white collars wages are higher in the public sector than in the private sector and they show also a lower dispersion. This is consistent with the idea that the public sector displays a wage structure more compressed even within non-manual occupations. The raw differential is about 6% and it is statistically significant[7].

4. The Econometric Methodology

In order to investigate whether individuals receive an equal remuneration in the two sectors or not, the simple comparison of wages between the public and the private sector does not provide enough information. The computation of the wage differential requires knowledge about the wage that an individual working in the public sector would receive in the private sector, maybe controlling for other determinants besides the sector. As the same person cannot be in two different labour market states at the same time and we observe the wage only for the sector in which a worker is actually employed, the counterfactual situation is not observable and can only be estimated using information on private sector's workers.

Let "working in the public sector" to be the treatment received by an individual, and her wage the outcome of that treatment. Private sector employees are the "control group". The parameter of interest is the "Average Treatment effect on the Treated", which is the mean difference between the wage actually earned by a public sector employee and the (potential and not observed) wage she would earn in the private sector (counterfactual situation). If assignment to sectors is not exogenous, the wage received by (comparable) workers in the private sector is not necessarily a good estimator of the wage earned by public employees had they worked in the other sector: persons who work in the public sector are different from persons who do not, in the sense that mean outcomes of participants in the non participation state would be different of those of non participants. Two main reasons are responsible for that.

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[6] The data were cleaned by excluding outliers and missing values for relevant variables. There are no reasons to belive that excluded individuals have any systematic relationship with the these variables.

[7] The p-value of a t-test for zero mean difference is 0.000.

The first one is the traditional endogeneity problem: the wage impact of the public sector is constant across individuals but public sector workers have *on average* a lower unobservable productivity than otherwise similar individuals working in the private sector. If productivity levels were correlated with the decision to work in the public sector and, at the same time, affected the wage received in both sectors, the public sector workers would earn less than otherwise similar workers, had they worked in the private sector. Failure to control for this difference would lead the lower wage of those with lower ability working in the public sector to be incorrectly attributed to their sector affiliation.

The second reason is self-selection: assume that the wage impact of working in the public sector is different across individuals and function of unobservable variables that, in turn, also affect the probability to be either a public or a private employee. Then, selection into the two sectors sector may not be random. If the sample used is selected, one might expect that the true average treatment effect for treated individuals is different than the average treatment effect estimated using information on self-selected private employees.

In both cases (endogeneity and self-selection) OLS estimates are biased and inconsistent. To solve for these problems, Heckman (1979) and Heckman and Robb (1985) developed a two-step procedure which, under specific distributional assumptions, consistently estimate endogenous treatment effects by modelling the stochastic dependence between the unobserved determinants of the outcome and the endogenous treatment. This dependence usually takes the form of control functions (correction terms) known up to some estimable parameters.

However, if additional covariates besides sector assignment are correlated with unobservable determinants of wages - and, maybe, with preferences for the public sector – standard two-step Heckman's model do not consistently estimate the parameter of interest. For example, suppose that unobserved wage determinants are correlated both with education and sector choices, and/or public sector workers are on average more likely to acquire above average education levels due to unobserved factors. In this case standard techniques that control only for selection in the choice of the sector are not able to consistently estimate the effect of working in the public sector and returns to education, and more sophisticated procedures are needed.

The next subsections illustrate the econometric model. The estimating version contains three equations: (i) a mincerian earning function augmented by a dummy for public sector affiliation, where both the education variable and the sector dummy are endogenous, and, in addition, the coefficient for the dummy may be random; (i) a reduced form for endogenous education; (iii) a reduced form for sector choices. Selection in sector choices is modelled *a la* Heckman, while selection in education is solved applying 2SLS techniques on the wage equation previously augmented by estimates of correction term(s).

4.1 The model and the parameter of interest

As suggested by Heckman (1990), let $\ln w_{iG}, \ln w_{iP}$ be, respectively, the (latent) wage that the *i-th* individual earns in the treatment status (G, standing for government or public sector) and in the control status (P, private sector):

$$\begin{aligned} \ln w_{Gi} &= \alpha_G + \psi_G S_i + X_i^{'}\beta_G + u_{Gi} \\ \ln w_{Pi} &= \alpha_P + \psi_P S_i + X_i^{'}\beta_P + u_{Pi} \end{aligned} \qquad (1)$$

S is the number of years of schooling and X is a vector of exogenous individual characteristics that influence earnings.

Let the binary variable D denote observed sector affiliation (D = 1 if the individual works in the public sector; D = 0 otherwise). Then, the individual wage may be written as:

$$\ln w_i = D_i \ln w_{Gi} + (1 - D_i) \ln w_{Pi}$$

Using (1) and imposing the restrictive assumption that the parameters are the same in each sub-sample except for the intercept, the wage equation may be expressed as:

$$\ln w_i = \alpha + \psi S_i + X_i^{'}\beta + \delta_i D_i + u_{Pi}$$

where $\alpha = \alpha_P, \psi = \psi_P, \beta = \beta_P, \delta = \alpha_G - \alpha_P$ and $\delta_i = \delta + (u_{Gi} - u_{Pi})$.

The estimating equation is:

$$\ln w_i = \alpha + \psi S_i + X_i^{'}\beta + \delta D_i + \varepsilon_i \qquad (2)$$

where $\varepsilon_i \equiv u_{Pi} + (u_{Gi} - u_{Pi}) D_i$

Note that (2) is a model with individual-specific effects (random coefficient) for D: due to the presence of u_{Gi}, u_{Pi} the effect of working in the public sector is individual-specific. Nested in this model there is also the specification with homogeneous effects (constant coefficient), for example when $u_{Gi} = u_{Pi}$.

I assume that the optimal individual investment in education is observable and can be expressed as a linear combination of variables including a random error term:

$$S_i = Z_i^{'}\vartheta + \tau_i \tag{3}$$

An individual works in the public sector (*D* = 1) or in the private sector (*D* = 0) as the outcome of an unobserved latent variable *D** which can be interpreted as the difference in expected utilities between public and private employment. Writing the reduced form[8] as:

$$D_i{}^* = Z_i^{'}\gamma + v_i$$

she chooses to work in the public sector only if this difference is positive (net benefit):

$$D_i = I(Z_i^{'}\gamma + v_i > 0) \tag{4}$$

where *I(A)* is an indicator function assuming value 1 whenever *A* is true.

The probability to be a public employee is influenced both by observable and unobservable factors like individual preferences, attitudes toward the risk, personal characteristics, family background and tastes for specific job attributes. Note that this is a model of "pure choice" since the decision to join the public sector is not constrained or rationed. In other words, once an individual chooses to work in the public sector (supply side), the (public) employer automatically is willing to hire him (demand side). Of course, this is an unrealistic simplification, especially for the public sector, where recruitment happens through public concourses where the number of applicants is traditionally much higher than the number of available positions. As explained in section 3, I try to mitigate the problem using a sample of individuals older than 25 years. As an alternative, the choice model outlined in (4) may be interpreted as a reduced form for both supply and demand decisions (for a discussion see also Bardasi, 1996)[9].

[8] As an alternative, the structural model of Lee (1978), where sector decisions also depends on the difference between public and private wages for the i-th individual may be employed. Of course, as compared to reduced form models, structural specifications are more flexible, since they allow for simultaneos effects. However, since the choice to work in the public sector is done once for all, it is typically not influenced by simultaneous wage differences (see also Dustmann and Van Soest, 1998).

[9] Limiting the entry by recruiting people through public examinations, the public sector generally produce a queue. In order to observe an individual in the public sector two events must happen: first, the individual must choose to join the queue; second, the individual needs to be selected out of the queue.

I assume that the disturbance terms contained in the choice equation and in the outcome equations are distributed as a trivariate normal[10]:

$$\begin{pmatrix} u_G \\ u_P \\ v \end{pmatrix} \sim MVN\left[\begin{pmatrix} 0 \\ 0 \\ 0 \end{pmatrix}\begin{pmatrix} \sigma_{GG} & . & \sigma_{GV} \\ & \sigma_{PP} & \sigma_{PV} \\ & & 1 \end{pmatrix}\right] \tag{5}$$

The covariance between the error terms of the two wage equations is not identified since the two wage regimes are not simultaneously observed. Thus, the sector of employment is endogenous to wages: some unobserved characteristics that influence the probability to choose a particular sector of employment could also influence the wage received by the individual once he is employed.

As a result, the error term in (2) is correlated with *D* and the OLS estimator is biased and inconsistent. Note that, if $u_{Gi} = u_{Pi}$ OLS is still biased due to the non zero correlation between u_P, u_G and v.

No restrictions are imposed on the way that the unobserved heterogeneity in education choices τ is correlated with u_P, u_G, v (education is potentially endogenous to sector choices and to wages). As usual, the reduced form for the two endogenous processes contain only the exogenous variables of the model. Consequently, *Z* includes all the regressors in *X* and a vector of instruments *H,* which consists of variables influencing decisions about sector of employment and investments in education but not the wage. *Z* is assumed to be independent of all the error terms. As it will be discussed in more details below, the adoption of a sequential estimation procedure allows to relax some of the assumptions concerning the composition of *Z*. In particular, the endogeneity of both *D* and *S* prevents the inclusion of the latter in the first step estimation of (4). In other words, the relationship between education and sector choices is completely captured by the (arbitrary) correlation between τ and v, which is a convolution of causal and simultaneous effects. Using first step results it is then possible to model the stochastic dependence between ε and v by means of correction term added to (2).

However, due to a lack of separate information on the relevant population for the two selection stages (which could be obtained if it was possible to identify those who applied to work for the public sector but were not chosen), the proposed selection process is a reduced form for supply and demand factors.

[10] Strictly speaking, this is an unnecessary assumption. In fact, it suffices that $v \sim N(0,1)$ and $E(u_j \mid v) = av$ (a linear function), with *j=P, G.* If the error terms are jointly normally distributed this condition follows automatically.

Once these control functions are included, D can be treated as exogenous to wages and, together with the correction term(s), included as additional exogenous regressor in the second step estimation of the education equation[11].

Using a first-order Taylor approximation, the wage differential (treatment effect, TE) between working in the public and the private sector for the *i-th* employee may be written as:

$$TE_i = \frac{w_{Gi} - w_{Pi}}{w_{Pi}} \approx \ln w_{Gi} - \ln w_{Pi} = \delta + (u_{Gi} - u_{Pi}) \qquad (6)$$

This term has two components: the first one is the coefficient associated to D in (2): $\delta = E(\ln w_{Gi} - \ln w_{Pi} | X_i, S_i) = ATE(X,S)$, where ATE is the (constant across individuals) average gain for a randomly chosen individual with given characteristics. The second component is the individual idiosyncratic effect. The "Average Treatment effect on the Treated (ATT)" may be expressed as:

$$ATT(X,S) \equiv \delta' = E(\ln w_{Gi} - \ln w_{Pi} | X_i, S_i, D_i = 1) = \delta + E(u_{Gi} - u_{Pi} | X_i, S_i, D_i = 1) \qquad (7)$$

which is the average gain in the population plus the average of individual-specific effects among public employees.

Let $u_{Gi} - u_{Pi} = \theta_i$, $u_{Pi} = u_i$, then we may rewrite (2) as follows:

$$\ln w_i = \alpha + \psi S_i + X_i^{'} \beta + \delta' D_i + \omega_i \qquad (8)$$

where $\omega_i = u_i + [\theta_i - E(\theta_i \mid X_i, S_i, D_i = 1)] D_i$

In the discussion of estimation techniques I treat separately the case of a constant effect from the case of individual-specific effects.

Constant public sector premium and "genuine" endogeneity

The composite error in equation (2) contains a term capturing the difference in unobserved factors that influence the wage of public sector workers, with and without working in the public sector. Under the assumption that these factors are the same for each worker ($u_{Gi} = u_{Pi} = u_i$) or mean independent of the decision to be a public

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[11] The inclusion of the correction term into the education equation suggests another reason for the exclusion of education from the first stage estimation of the sector choice equation: if included, since the correction term is a function first stage regressors, in the second stage we would have a situation with education being both the dependent variable and one of the arguments in a function used to explain education, which is, of course, not allowed.

employee ($E(u_{Gi}-u_{Pi}|X_i,D_i=1)=E(u_{Gi}-u_{Pi}|X_i)=0$)[12], the mean wage differential is equal to δ. In other words, conditional on X, the effect of working in the public sector is assumed to be the same for everyone and independent of sector status. In this case, ATE = ATT and, since there are no individual-specific gains from working in the public sector, δ = δ'. Under the assumption that $u_{Gi}=u_{Pi}=u_i$ (2) becomes:

$$\ln w_i = \alpha + \psi S_i + X_i^{'}\beta + \delta D_i + u_i \tag{9}$$

However, since D and S are not independent of u_i, the conditional expected value of the error term is different from zero and OLS do not consistently estimate the parameters of the model. Instrumental variable methods applied to a model with two endogenous variables offer a solution to this problem. This paper takes an alternative route. Endogeneity in education is eliminated using standard instrumental variables techniques, whereas correction terms control for endogeneity of sector affiliation by modelling the stochastic dependence between wages and sector choices. To explain the details of this approach, let assume for the moment that the level of education is exogenously assigned. Thus, the model contains one endogenous dummy (D). To derive an estimating equation, just write the conditional expected value of the log wage as[13]:

$$E(\ln w_i \mid D_i, Z_i) = \alpha + \psi S_i + X_i^{'}\beta + \delta D_i + D_i E(u_i \mid D_i = 1, Z_i) + (1-D_i)E(u_i \mid D_i = 0, Z_i) \tag{10}$$

Heckman (1978, 1979) showed that, under specific distributional assumptions, selectivity issues when the endogenous variable is a binary treatment can be solved by including the functional form of the conditional expectations for u_i in (10) as an additional variable. Therefore, the equation to be estimated becomes:

$$\ln w_i = \alpha + \psi S_i + X_i^{'}\beta + \delta' D_i + \mu\lambda 1_i + \xi_i \tag{11}$$

where $\mu = \sigma_{PV} = \sigma_{GV} = \sigma$, $\xi_i = u_i - E(u_i \mid D_i, Z_i)$. The conditional mean of the new error term is zero and:

$$\lambda 1_i = D_i \frac{\phi(-Z_i^{'}\gamma)}{\Phi(Z_i^{'}\gamma)} - (1-D_i)\frac{\phi(-Z_i^{'}\gamma)}{\Phi(-Z_i^{'}\gamma)} \tag{12}$$

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[12] In other words, u_G and u_P are mean independent of D given X, and the difference between the two error terms is unkown or ignored by individuals when they decide to work in the public sector. Thus, their best forecast for this difference is simply zero

[13] I'm leaving implicit the conditional dependence on *X* and *S*, since I'm assuming that *X* is included in *Z* and *S* is exogenous.

the inverse Mill's ratio for the entire sample. Clearly, by including (12) in the wage equation, D can be treated as exogenous. Still, (12) is not known but a consistent estimate for it - $\hat{\lambda 1}_i$ - corresponds to the generalised residual of the first-step probit for the probability to work in the public sector. The wage equation augmented by the correction term (11) can be consistently estimated by OLS in the second step.

If S is not exogenous to the wage, $E(u_i \mid S_i, D_i, Z_i) \neq E(u_i \mid D_i, Z_i)$, so the correction term is different from the generalised probit residual because the (joint) distribution of both the endogenous variables should be taken into account. Therefore, even if the error term has conditional mean zero,

$E(\xi_i \mid S_i, D_i, Z_i) = E[(u_i - E(u_i \mid D_i, Z_i)) \mid S_i, D_i, Z_i] = 0$,

it is still correlated with S[14] and, therefore, (12) does not model properly the stochastic dependence between the endogenous variables and the outcome. The model contains an endogenous variable and OLS do not deliver consistent estimates. It can be shown (see Wooldridge, 2002, pp 567-569) that standard IV techniques offer a straightforward solution to this problem. In fact, (11) can be consistently estimated with a three-stage procedure which combines Heckman methods and instrumental variables techniques: first, the computation of the sample generalized residual $\hat{\lambda 1}_i$ from the probit estimation of (4); second, the application of 2SLS to (11) using [Z, D, $\hat{\lambda 1}(Z,D)$] as instruments for S[15]. Of course, the implementation of this procedure requires an instrumental variable for S. As usual, the test of no selectivity bias (no correlation) is a t-test of $\mu=0$[16].

[14] In fact the subtraction of E(u | D, Z) from u "cleans" the original error term from the correlation between u and v. What is left in the new error term is the purely random component of u and the part correlated with τ.

[15] While education cannot be included in the first stage probit for sector choice since it is an endogenous variable, the sector dummy can be included in the reduced form for education becouse, once we include the correction term, it can be treated as exogenous.

[16] Similarly to the standard Heckman model, where OLS statistics are incorrect unless the null of no endogeneity is not rejected, 2SLS statistics should be corrected for the generated regressor bias unless $\mu = 0$. Since computing standard errors analytically can be very complex in this situation, coefficients estimated with this procedure are consistent but not efficient. However, the comparison of results obtained by estimating the original Heckman model (with education exogenous) both with the true variance covariance matrix and by just plugging the generalised residual in the second step reveals that standard errors in the two cases are very similar.

Individual-specific public sector premium and self selection

In a model with heterogeneous effect the common gain (ATE) is different from the effect of the average individual-specific gain (ATT≠ATE). By writing explicitly (8) and considering for the moment S as exogenous we obtain:

$$\ln w_i = \alpha + \psi S_i + X_i{}^{'}\beta + [\delta + E(u_{Gi} - u_{Pi} \mid Z_i, D_i = 1)]D_i + \{u_i + [\theta_i - E(\theta_i \mid Z_i, D_i = 1)]D_i\} \qquad (13)$$

Note that by construction the error term in (13) is zero mean since:

$E[\theta_i - E(\theta_i \mid Z_i, D_i = 1) \mid Z_i, D_i = 1] = 0$.

Under the assumption that Z, X and S are independent of u_{Pi}:

$$0 = E(u_{Pi} \mid Z_i) = E(u_{Pi} \mid D_i = 1, Z_i)prob(D_i = 1 \mid Z_i) + E(u_{Pi} \mid D_i = 0, Z_i)prob(D_i = 0 \mid Z_i)$$

which may be written as:

$$E(u_{Pi} \mid D_i = 1, Z_i) = -E(u_{Pi} \mid D_i = 0, Z_i)\frac{prob(D_i = 0 \mid Z_i)}{prob(D_i = 1 \mid Z_i)}$$

From (7)[17], the ATT may be then expressed as:

$$\delta^{'} = \delta + E(u_{Gi} \mid Z_i, D_i = 1) + E(u_{Pi} \mid D_i = 0, Z_i)\frac{prob(D_i = 0 \mid Z_i)}{prob(D_i = 1 \mid Z_i)}$$

Finally, using the fact that Z is independent of all the error terms and the distributional assumptions outlined in (6) the previous expression becomes:

$$\delta^{'} = \delta + (\sigma_{GV} - \sigma_{PV})\frac{\phi(-Z_i\gamma)}{\Phi(Z_i\gamma)} \qquad (14)$$

where $(\sigma_{GV}, \sigma_{PV})$ are the cross-equations correlation coefficients between disturbances[18]. The first term in (14) measures the "pure" public sector effect, explained by structural differences between the two sectors that are common to everyone. The second term in (14) capture, on average, the part of the wage differential based on unobservable (by the econometrician: observed by the individual) and individual-specific wage differences for public sector employees. A test of $(\sigma_{GV} = \sigma_{PV})$ is equivalent to test the hypothesis of no selection on unobservable gains and of ATT = ATE. Using (14), (13) may be written as:

[17] Here: $\delta' = \delta + E(u_{Gi} - u_{Pi} | Z_i, D_i = 1) = \delta + E(u_{Gi} \mid Z_i, D_i = 1) - E(u_{Pi} \mid Z_i, D_i = 1)$

$$\ln w_i = \alpha + \psi S_i + X_i^{'}\beta + \delta D_i + (\sigma_{GV} - \sigma_{PV})\frac{\phi(-Z_i^{'}\gamma)}{\Phi(Z_i^{'}\gamma)}D_i + \sigma_{PV}[D_i\frac{\phi(-Z_i^{'}\gamma)}{\Phi(Z_i^{'}\gamma)} - (1-D_i)\frac{\phi(-Z_i^{'}\gamma)}{\Phi(-Z_i^{'}\gamma)}] + \pi_i \qquad (15)$$

where π_i is a generic zero mean error term uncorrelated with the regressors.

Since the correction terms contains unknown coefficients, the empirical counterpart of (15) is:

$$\ln w_i = \alpha + \psi S_i + X_i^{'}\beta + \delta D_i + \theta \hat{\lambda 2}_i D_i + \mu \hat{\lambda 1}_i + \pi_i \qquad (16)$$

where hats denote probit estimates and $\theta \hat{\lambda} 2_i = (\sigma_{Gv} - \sigma_{Pv})\frac{\phi(-Z_i^{'}\hat{\gamma})}{\Phi(Z_i^{'}\hat{\gamma})}$

The correction term interacted with the dummy captures the individual specific component of the wage premium based on comparative advantages, while the coefficient associated to the generalised probit residual tests for the presence of self selection in the base state of the world (private sector).

Because in my case also education is considered endogenous, the estimation procedure consists of three steps, where in the first step $\hat{\lambda 1}_i$ and $\hat{\lambda 2}_i$ are computed from a probit for (4), and then used as instruments for S in conjunction with Z and D in a 2SLS which involves (16) and (3).

A sample estimate of ATT can be easily obtained as:

$$ATT = \hat{\delta} + \hat{\theta}\frac{\phi(-\bar{Z}_i^{'}\hat{\gamma})}{\Phi(\bar{Z}_i^{'}\hat{\gamma})} \qquad (17)$$

where $\bar{Z}_i^{'}$ is the sample mean of Z for public sector workers.

A test for θ = 0 is a test of self-selection on individual unobservable wage differences. Note that, in principle, θ may be positive or negative. The sign of the coefficient may also help to understand which theory – comparative wage advantages or compensating wage differences - is able to explain the behaviour of public sector workers. As discussed above, each individual chooses the sector in which he receives the highest utility (which includes both monetary and non monetary factors).

[18] In a context of switching regressions with endogenous switching these two covariances are the

Let us assume θ > 0: in this case, public employees are self-selected in the sector where they receive expected monetary gains from participation, and for them the wage premium is somehow larger than for an arbitrary person. In other words, - by assuming that unobserved factors are a proxy for motivation, good matches and other productivity-related wage determinants - they self-select themselves in the sector where they are more productive and where they benefit from a comparative wage advantage with respect to the entire population. Since the allocation of workers is based on productivity-based comparative advantages, it is also efficient.

Let instead θ < 0: the unobserved individual productivity of public sector workers is on average lower than their potential productivity in the private sector. Because sector choices are driven by both monetary and non monetary aspects, the amount of non monetary gains they receive from working in the public sector more than compensate wage losses due to low unobserved productivity and comparative disadvantages. In this case, higher qualitative job determinants – such as wage and employment stability, risk aversion, less stress and competition at the workplace, higher social protections – raise satisfaction at the workplace and more than compensate unobserved (potential) negative wage differences: public employees are willing to pay for these attributes though the wage advantage is lower than for a random individual. Still, since workers are employed in the sector in which they are less productive, the allocation of the workforce across sectors is inefficient.

4.2 The identification strategy

The base specification of the wage equation is parsimonious and includes (a) a set of controls for age and its square, and dummies for the geographic area of residence (to account for different labour market conditions between north, centre and south of Italy) (b) years of schooling, the dummy for the public sector (c) time dummies for 1998, 2000 and 2002, and additional family background variables – years of schooling of the parents.

Since education and the sector dummy are endogenous, identification usually requires valid sources of exogenous variation. Although identification in normally distributed models with self-selection is achieved through the non linearity of the

coefficient for the correction terms in the two equations.

correction term, in practice, since in certain regions of the support the Mill's ratio is linear, the inclusion of instruments for the sector choice may be important to guarantee identification. This means that at least two variables that affect the choice of the sector and the level of education but have no direct effect on wages are needed.

As usual, all the exogenous variables are included in the two equations for the endogenous processes. The structure of the model also requires the inclusion of D and the correction term(s) (which, in turn, is(are) also function of Z and D) into the education equation. Following the strategy of Brunello and Dustmann (1997), the first set of instruments is composed by three dummies for the father's occupation (blue collar, white collar and self-employed). Following Cappellari (2004), I use the geographic birth area as an additional instrument for schooling, since it is thought to influence human capital accumulation without residual effect on wages, once the impact of the area of residence has been controlled for.

To identify the choice of the sector I also use two binary variables capturing, respectively, employment of the mother and employment of the father in the public sector.

All of these variables should capture tastes and constraints influencing sector choices and schooling decisions. The use of parental background variables as instruments is common in the literature. Still, these variables are often suspected to be positively correlated with unobserved factors in the wage equation and their validity as instruments might be questionable, especially in models without distributional assumptions.

In principle, I assume that, once the impact of several personal or family characteristics, including parents' education, has been controlled for, their occupation and the fact that they worked in public sector it is unlikely to have any significant impact on wages.

In practice, the validity of my identification strategy will be evaluated in the empirical analysis by means of a number of exclusion restrictions' tests.

5. Main Results

For comparative purposes, the first column of Table 2 presents the estimates of (9) by OLS. Results show that (marginal) returns to education for male white collars are on average equal to 3.1% and those to age (capturing both general and specific on-the-job-formed human capital as well as life cycle wage effects) approximately 6%. The coefficient associated with the sector dummy (public) shows that public employees earn on average 3% more than comparable private. Also regional differences matter: being employed in the north guarantees a wage premium compared to the south equal to the 8%, while differences are less pronounced for workers employed in the centre of Italy. Time dummies show that during the last decade real wages of white collars were decreasing. In addition, the parents' level of education is statistically significant.

<TABLE 2 AROUND HERE>

The second column of Table 2 reports a parametric test on instruments validity. Overall, the data support the exclusion of the set of variables used as instruments (region of birth, dummies for the two parents employed in the public sector and for father's occupation) from the wage equation. If *S* (educ) and *D* (pub) were uncorrelated with the error term, then OLS consistently estimate ATT = ATE. However, this is no longer true when education and/or sector choices are not exogenous. The remaining part of the section is organised as follows: I first estimate the wage equation corrected for the endogeneity of sector decisions but not for endogenous education; then I endogenise both processes and estimate the model with the three step procedure discussed in section 3. Table 3 presents the results for the model where the public sector wage premium is constant across individuals. Table 4 presents the estimates of the model with random coefficients for the sector dummy.

Column 5 of Table 3 reports the first step probit results for the selection equation.

<TABLE 3 AROUND HERE>

The probability to work as a public employee increases when the father has worked in the same sector. Other significant determinants of sector choices are the age, the geographic area (of birth and residence) and the level of education of the parents.

Column 4 reports second step OLS estimates of a wage equation augmented by a correction term for endogenous sector choices. Results show that the return to education is equivalent to OLS estimates, while the public wage premium is higher (11.5%) but not significant at the usual levels. The negative coefficient for the correction term signals that public employees are negatively selected, but this effect is not statistically significant.

In order to account for both sources of endogeneity (education and sector decisions), columns 1 and 2 report respectively 3rd and 2nd step results from the application of 2SLS to (11), a wage equation augmented by the correction term estimated in the 1st step. As explained in Section 3, the estimation of the reduced form for education includes the full set of exogenous variables, including the sector dummy and the correction term. Exclusion of instruments is strongly rejected by the data, and, since some instruments which are significant in the selection equation are not relevant predictors of education, the model is identified without relying on non-linearities in the functional form of the correction term.

For what concerns the estimation of the earning equation (11) in the third step, the overidentification test supports the validity of the identification strategy. A version of Hausman-Wu test for endogeneity of education and sector choices[19] rejects their joint exogeneity with a p-value of 1%. As considered separately, both selection in education (see test 3) and in the public sector (captured by the coefficient for lambda1) are significant.

Interactions between the two endogenous processes affect the coefficients' estimates. In fact, as compared to results in columns 4, important differences emerge.

First, the return to education is approximately equal to 4.9%, almost two percentage points more than with standard OLS. This evidence is qualitatively similar to the results obtained by Brunello and Miniaci (1999), who found that returns are underestimated when education is erroneously treated as exogenous due to the sorting of less able individuals in the group with high educational endowment. This may happen because people with higher unobservable productivity drop out school earlier and start working. As an alternative, liquidity constraints may prevent able pupils with poor family background to make their optimal schooling choice.

This behaviour may be rational in Italy, where the (private) economy for the most part is populated by traditional and typically low-skill firms, where specialised positions

[19] It is implemented including the first stage residual of the education equation as an additional regressor in the wage equation. The residual for the sector choice is simply $\lambda 1$, which is already included. Testing whether they are jointly zero or not is equivalent to test for the presence of endogeneity in *S* (educ) and *D* (pub).

requiring high levels of cognitive skills acquired through general human capital are quite rare. Therefore, more productive or more able individuals may have no incentives to acquire additional years of (not requested by firms) education, because it may be relatively easy to find well-rewarded jobs for them, while being enrolled at school may have relatively high opportunity cost and low expected returns. In addition, the compressed structure of wage differentials by education levels in Italy may create a further disincentive for able workers to acquire high levels of education.

Second, the point estimate of the public wage premium is higher (16.8%) and more accurate than the previous one, and also negative selection in the choice of the sector, captured by the coefficient for the correction term, is now marginally significant. Accordingly, due to unobserved characteristics, workers employed in the public sector earn less in both sectors than the average worker.

Overall, the treatment of education as exogenous produces a less accurate estimate of the wage premium and of the relevance of endogeneity issues in the choice of the sector. Thus, the result obtained by many existing studies for Italy, namely that the wage premium is not statistically significant and that selection effects are negligible, may depend on the lack of control for endogenous education: if not adequately instrumented, the public premium captures both the true effects and a composition effect which confounds the estimates.

As discussed in section 3, the model with constant treatment effect has the disadvantage that, since unobservable wage determinants are assumed to be equal across the two sectors, the public wage premium is constrained to be the same for a random individual and for those actually working in the public sector, and the coefficient for the endogenous dummy estimate both the average premium (ATE) and the premium for public employees (ATT).

Table 4 reports the estimates of the more flexible specification (16), which allows for individual-specific returns to sector choices and self selection.

<TABLE 4 AROUND HERE>

Column 4 presents the results of the model estimated treating education as exogenous. The wage premium for an individual randomly selected from the population (ATE) is about 10%, but, similarly to what observed in Table 3, the statistical significance of the coefficient is quite weak.

The coefficient for lambda1 is not significantly different from zero. This means that workers in the base state of the word (private sector) are not self-selected on the basis of unobserved characteristic: their earning potential is not different from that of the average worker. On the opposite, there is significant negative self-selection in the public sector, as revealed by the coefficient for lambda2*pub (see Bardasi, 1996 for similar results). Since for public employee the difference between earnings potential in the two sectors is negative, their unobserved characteristics could allow them to earn more in the private sector. This implies that the allocation of skills is inefficient, and workers self-select themselves into the sector where their productivity is less rewarded.

In other words, the choice to work in the public sector seems to be based more on (unobservable) non monetary factors compensating monetary losses than on comparative wage advantages. As a consequence of negative self-selection, the coefficient associated with the dummy for sector affiliation (which represents the treatment effect in the population) overestimate the true average wage differential for public workers. In fact the Average Treatment for the Treated (ATT), obtained by adding the interaction term (lambda2*pub) evaluated at the average sample characteristics to the sector dummy coefficient as in (17), is still positive but equal only to 3.7%.

By comparing results from Table 3 and 4 it is possible to evaluate the advantages of the model with individual-specific premium and self-selection with respect to the model with constant premium and an endogenous dummy. First, treating selection in sector decisions not as an endogeneity problem (as in table 3), but instead as a self-selection problem (as in table 4) allows to separate two effects (selection in the private as different from selection into the public sector and based on different unobserved wage determinants) which are constrained to be the same when the premium is constant.

This is an important feature because these two effects work in opposite directions (see the coefficients for the correction terms in Table 4) and only one of them is significant. As a result, it is not surprising that the correction term in Table 3, being a sort of "average" of the two, is not significant. Second, since public employees are on average negatively selected and, accordingly, earn a premium which is very different from what is gained on average in the population, the assumption of constant return made in Table 3 it seems quite restrictive.

Next, columns 1, 2 and 3 present the results obtained when education and sector decisions are considered jointly endogenous. Results are consistent to those reported in table 3 and qualitatively similar: the return to education increases at the level of 4.5% and the public wage premium increases to 14.9% and becomes significant at the 10% level. The coefficient associated to the interaction term (lambda2*pub), which captures the individual-specific component of the wage premium, is again negative and marginally significant. As a consequence of negative self-selection into the public sector, average wage differential (ATT) for male white collars public employees is equal to 9.6%.

6. Conclusions

The aim of this paper was to investigate how endogenous education and sector decisions affect the estimate of the wage premium earned by workers employed in the public sector in Italy. To my knowledge, the existing literature on the public/private wage differential for Italy has tried to control only for the first source of selectivity. Still, a number of studies (also for Italy) have shown that education decisions are likely to be correlated with unobservable wage determinants.

The data set is obtained by pooling information from 1995, 1998, 2000 and 2002 Bank of Italy Households' Surveys. To avoid issues of endogenous female labour market participation, the study is limited to males. In addition, since manual occupations in the public sector do not have a close counterpart in the private sector, only white collars have been selected.

Results show that, if not adequately treated, the endogenous nature of education lowers returns to education - due to the (weak) sorting of less able individuals in the group with high schooling attainment -, and produces imprecise estimates of the public wage premium.

Main findings reveal that, under the more flexible specification with random coefficients for the sector dummy, the assumption of exogenous education leads to an estimate of the public wage premium for an arbitrary individual close to 11.5% but not significant. When also education is considered endogenous, the population average wage differential (ATE, Average Treatment Effect) for a male white collar raise at the level of 15% and gains significance.

For what concerns public employees, they are negatively self selected and display a lower unobservable productivity in the sector in which they actually work. As a consequence, the wage premium they earn (ATT) is still positive (9.6%) but lower than the population average (ATE). In other words, public workers are employed in the sector in which their earning potential is lower, and, as a results, they are willing to exchange the comparative wage advantage offered by the private sector with better non-monetary job attributes in the public sector, that more than compensate potential monetary losses.

These findings suggest a number of considerations about the relative efficiency of the education system and of the public sector's retaining, recruiting and pay policies. A well-known fact in Italy is that the wage structure is compressed, especially in the public sector. In addition, the structure of the Italian economy is for the most part traditional, and positions for people holding university degrees are scarce and rewarded proportionally less than jobs requiring intermediate education levels.

As a result, it seems that able individuals are more likely to drop out school and start working because it might be relatively easy for them to obtain well-rewarded white collar positions not requiring high levels of cognitive skills (especially in the private sector). On the contrary, less able individuals - with worst outside options - may have an incentive to acquire higher levels of education (public education in Italy is relatively cheap), partly to avoid unemployment, partly to increase productivity and maybe to buy a signal, which may be especially useful in the public sector, where holding high levels of education is crucial in the recruitment procedure.

For what concerns more specifically public employees, they would be more motivated, productive and better matched in the private sector. Still, they apply for a position in the public employment because it pays a ceteris paribus premium relative to the private sector and guarantees higher levels of valuable non-wage job attributes, like stability and flexibility.

From the public sector perspective, this situation creates inefficiencies: on the one hand, recruitment methodologies based on schooling performance as a signal of high ability are likely to select individuals with low ability; on the other hand, due to negative self-selection into the public sector, non-monetary job attributes are likely to attract individuals with comparable low productivity. As a result, the allocation of skills across sectors is inefficient. Albeit this situation may be optimal from the point of view of public employees, efficiency and equity considerations suggest that it might not be totally desirable from a social perspective.

Reforms of public sector wage determination aimed at increasing productivity and efficiency are desirable but probably do not suffice, and new rules and procedures to recruit and retain workers may be even more important. In this context, schooling reforms aimed at reducing the unbalance between the demand and the supply of education and creating incentives for a more efficient allocation of talent may be important *di per se* and for the positive spillover effects on the allocation of skills in the public sector.

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TABLES

Table 1. Summary statistics

		Whole sample		Private sector		Public sector	
Variable	Description	Mean	*Stdv*	Mean	*Stdv*	Mean	*Stdv*
Hwage	Hourly wage (€, 1995 prices)	8.09	*4.18*	7.99	*4.68*	8.21	*3.39*
Lnhwage	ln(hourly wage)	2.67	*0.39*	2.64	*0.41*	2.70	*0.35*
Educ	years of schooling	12.67	*3.33*	12.74	*3.11*	12.57	*3.60*
Pub	public sector = 1	0.42					
Age	Age (years)	42.61	*9.28*	41.28	*9.43*	44.42	*8.76*
resid_n	residence in the north = 1	0.44		0.54		0.31	
resid_c	residence in the centre = 1	0.20		0.19		0.22	
resid_s	residence in the south = 1	0.36		0.27		0.47	
1995	1995 = 1	0.27		0.22		0.33	
1998	1998 = 1	0.25		0.26		0.23	
2000	2000 = 1	0.26		0.28		0.23	
2002	2002 = 1	0.23		0.25		0.21	
Edufath	Father education	5.26	*4.65*	5.20	*4.88*	5.34	*4.31*
Edumoth	Mother education	4.59	*4.12*	4.59	*4.34*	4.60	*3.80*
Fathbc	Father blue collar = 1	0.37		0.38		0.37	
Fathwc	Father white collar = 1	0.32		0.32		0.31	
Fathse	Father self employed = 1	0.22		0.21		0.24	
Fathpub	father public employee = 1	0.19		0.16		0.23	
Mothpub	mother public employee = 1	0.06		0.06		0.07	
Birth_n	Birth in the north = 1	0.38		0.49		0.25	
Birth_c	Birth in the centre = 1	0.19		0.18		0.20	
Birth_s	Birth in the south = 1	0.43		0.33		0.55	
N. obs		*5678*		*3276*		*2402*	

Table 2. Wage equation with constant public sector premium: OLS results

	(1)		(2)	
Dep var: ln(hourly wages)	Coef.	*t-stat*	Coef.	*t-stat*
Intercept	0.722	*7.5*	0.726	*7.3*
Educ	0.031	*21.05*	0.031	*20.29*
Pub	0.030	*3.13*	0.030	*3.05*
Age	0.058	*12.87*	0.058	*12.72*
Age2	-0.001	*-10.14*	-0.001	*-10.05*
Resid_n	0.084	*7.63*	0.079	*4.16*
Resid_c	0.037	*2.81*	0.027	*1.18*
1998	-0.024	*-1.86*	-0.024	*-1.87*
2000	-0.025	*-2.04*	-0.024	*-1.96*
2002	-0.032	*-2.5*	-0.031	*-2.36*
Edufath	0.004	*2.57*	0.004	*1.83*
Edumoth	0.006	*3.16*	0.006	*3.24*
Exclusion Restrictions:				
Fathpub			0.020	*1.07*
Mothpub			-0.020	*-0.92*
Fathbc			-0.005	*-0.26*
Fathwc			0.006	*0.26*
Fathse			0.017	*0.83*
Birth_n			0.009	*0.46*
Birth_c			0.016	*0.78*
Test: Exclusion of instruments from the wage equation			*Pr>F = 0.533*	
R2		*0.199*		*0.200*
# observations		*5,678*		*5,678*

Note: Pooled data for 1995, 1998, 2000 and 2002. t-statistics (robust to heterosk. and autocorr.) in parenthesis. Excluded categories: year 1995, resid_s, birth_s. Test statistics are from Wald test for excluded instruments (H0: birth_n = birth_c = fathbc = fathwc = fathse = fathpu = mothpu = 0).

Table 3: Wage equation with constant public sector premium: two- and three-step procedures

	Education and sector endogenous						Only sector endogenous			
	3rd step: wage equation (1)		2nd step: education equation (2)		1st step: sector equation (3)		2nd step: wage equation (4)		1st step: sector equation (5)	
Variables	Depvar: ln(hourly wages)		Depvar: educ		Depvar: Pub		Depvar: ln(hourly wages)		Depvar: Pub	
	Coeff.	*z-stat*	Coeff.	*t-stat*	Coeff.	*z-stat*	Coeff.	*t-stat*	Coeff.	*z-stat*
Intercept	0.496	*3.1*	15.645	16.98	-2.163	*-5.96*	0.747	*7.62*	-2.323	*-6.22*
Educ	0.047	*5.77*					0.031	*20.91*	0.010	*1.83*
Pub (ATE = ATT)	0.168	*2.07*	-0.489	-0.25			0.114	*1.48*		
Age	0.057	*10.63*	-0.134	-1.82	0.098	*5.76*	0.055	*10.54*	0.100	*5.84*
Age2	-0.001	*-8.9*	0.001	1.49	-0.001	*-4.59*	-0.001	*-8.78*	-0.001	*-4.65*
Resid_n	0.114	*4.98*	-0.617	-2.32	-0.278	*-4*	0.105	*4.7*	-0.273	*-3.91*
Resid_c	0.047	*3.14*	-0.473	-2.26	0.006	*0.08*	0.044	*2.98*	0.011	*0.13*
1998	-0.005	*-0.29*	-0.063	-0.21	-0.384	*-7.88*	-0.012	*-0.7*	-0.384	*-7.88*
2000	-0.010	*-0.58*	0.257	0.83	-0.406	*-8.33*	-0.012	*-0.71*	-0.409	*-8.4*
2002	-0.018	*-1.09*	0.209	0.75	-0.355	*-7.07*	-0.021	*-1.23*	-0.358	*-7.13*
Edufath	0.002	*1*	0.101	6.19	-0.008	*-1.2*	0.004	*2.61*	-0.009	*-1.36*
Edumoth	0.005	*2.59*	0.044	2.56	0.003	*0.41*	0.006	*3.07*	0.002	*0.34*
Lambda1	-0.086	*-1.73*	0.398	0.33			-0.052	*-1.1*		
Instruments										
Fathwc			0.086	0.42	0.103	*1.28*			0.103	*1.28*
Fathbc			-0.032	-0.13	0.298	*5.58*			0.300	*5.6*
Fathse			-0.836	-4.74	0.092	*1.34*			0.101	*1.47*
Mothpub			0.501	2.73	0.003	*0.04*			-0.002	*-0.02*
Fathpub			0.106	0.49	0.176	*2.4*			0.176	*2.39*
Birth_n			1.040	2.89	-0.451	*-6.48*			-0.463	*-6.62*
Birth_c			0.801	2.67	-0.292	*-3.48*			-0.301	*-3.58*

Table 3. – continued -

Test 1: Exclusion of instruments from the sector equation			*0.000*		*0.000*
Test 2: exclusion of instruments from the educ equation		*0.000*			
Test3: exogeneity of educ	*0.05*				
Test 4: exogeneity of educ and pub	*0.06*				
Test 5. Overidentific. Restrictions	*0.655*				
Centred R2	*0.18*				
R2		*0.115*		*0.20*	
Log-likelihood					*-3,526.33*

Notes: Pooled data for 1995, 1998, 2000and 2002. n° of obs.: 5,678. Estimates: in columns (5) and (3) are obtained by probit; in column (4) by OLS; in columns (1) and (2) by 2SLS. t-(or z-)statistics robust to heterosk. and autocorr.. Excluded categories: 1995, South. Test 1-2 give statistics from Wald test of hypoteses, p-values reported in italics. Test 1: validity of instruments in the probit for the sector choice (H0 birth_n = birth_c = fathpub = mothpub = fathwc = fathbc = fathse = 0). Test 2: validity of instruments in the education equation (H0: birth_n=birth_c=fathpub=mothpub=fathwc=fathbc=fathse=0). Test 3: significativity of the residual from education equation when included in the wage equation, coeff. = -0.16, t-stat = 1.96 (H0: residual education equation=0). Test 4: joint significativity of residuals from the sector and education equation (H0: residual education equation = generalised residual of the sector equation (lambda1) = 0). Test 5: gives Hansen j-statistic for the overdentification of all instruments in a 2SLS prodedure.

Table 4. Wage equation with random public sector premium: two- and three-step procedures

	Education and sector endogenous						Only sector endogenous			
	3rd step: wage equation (1)		2nd step: education equation (2)		1st step: sector equation (3)		2nd step: wage equation (4)		1st step: sector equation (5)	
Variables	Depvar: ln(hourly wages)		Depvar: educ		Depvar: Pub		Depvar: ln(hourly wages)		Depvar: Pub	
	Coeff.	*z-stat*	Coeff.	*t-stat*	Coeff.	*z-stat*	Coeff.	*t-stat*	Coeff.	*z-stat*
Intercept	0.517	*3.23*	15.517	*16.22*	-2.163	*-5.96*	0.742	*7.55*	-2.323	*-6.22*
Educ	0.045	*5.56*					0.031	*20.92*	0.010	*1.83*
Pub (ATE)	0.149	*1.87*	-1.185	*-0.49*			0.099	*1.27*		
Age	0.058	*10.72*	-0.109	*-1.23*	0.098	*5.76*	0.057	*10.66*	0.100	*5.84*
Age2	-0.001	*-9.02*	0.001	*0.99*	-0.001	*-4.59*	-0.001	*-8.93*	-0.001	*-4.65*
Resid_n	0.107	*4.62*	-0.690	*-2.28*	-0.278	*-4*	0.098	*4.31*	-0.273	*-3.91*
Resid_c	0.046	*3.09*	-0.469	*-2.24*	0.006	*0.08*	0.044	*2.94*	0.011	*0.13*
1998	-0.008	*-0.46*	-0.161	*-0.46*	-0.384	*-7.88*	-0.015	*-0.86*	-0.384	*-7.88*
2000	-0.013	*-0.75*	0.153	*0.41*	-0.406	*-8.33*	-0.015	*-0.9*	-0.409	*-8.4*
2002	-0.021	*-1.24*	0.119	*0.36*	-0.355	*-7.07*	-0.023	*-1.39*	-0.358	*-7.13*
Edufath	0.002	*1.14*	0.100	*5.94*	-0.008	*-1.2*	0.004	*2.67*	-0.009	*-1.36*
Edumoth	0.005	*2.61*	0.045	*2.59*	0.003	*0.41*	0.006	*3.04*	0.002	*0.34*
Lambda1 (σ_{PV})	-0.036	*-0.62*	0.938	*0.58*			0.003	*0.05*		
Lambda2*pub (σ_{GV} -σ_{PV})	-0.068	*-1.92*	-0.205	*-0.5*			-0.080	*-2.23*		
Instruments										
Fathwc			0.113	*0.53*	0.103	*1.28*			0.103	*1.28*
Fathbc			0.047	*0.16*	0.298	*5.58*			0.300	*5.6*
Fathse			-0.812	*-4.45*	0.092	*1.34*			0.101	*1.47*
Mothpub			0.501	*2.73*	0.003	*0.04*			-0.002	*-0.02*
Fathpub			0.151	*0.64*	0.176	*2.4*			0.176	*2.39*
Birth_n			0.916	*2.1*	-0.451	*-6.48*			-0.463	*-6.62*
Birth_c			0.722	*2.13*	-0.292	*-3.48*			-0.301	*-3.58*
ATT	0.096						0.037			

Table 4. – continued -

Test 1: Exclusion of instruments from the sector equation			*0.000*		*0.000*
Test 2: exclusion of instruments from the educ equation		*0.000*			
Test3: exogeneity of educ	*0.07*				
Test 4. Overidentific. Restrictions	*0.825*				
Centred R2	*0.187*				
R2		*0.115*		*0.20*	
Log-likelihood			*-3,528.04*		*-3,526.33*

Notes: Pooled data for 1995, 1998, 2000and 2002. n° of obs.: 5,678. Estimates: in columns (5) and (3) are obtained by probit; in column (4) by OLS; in columns (1) and (2) by 2SLS. t-(or z-)statistics robust to heterosk. and autocorr.. Excluded categories: 1995, South. Test 1-2 give statistics from F-test of hypoteses, p-values reported in italics. Test 1: validity of instruments in the probit for the sector choice (H0: birth_n=birth_c=fathpub=mothpub=fathwc=fathbc=fathse=0). Test 2: validity of instruments in the education equation (H0: birth_n=birth_c=fathpub=mothpub=fathwc=fathbc=fathse=0). Test 3: significativity of the residual from education equation when included in the wage equation, coeff. = -0.14, t-stat = 1.80 (H0: residual education equation=0). Test 4: gives Hansen j-statistic for the overdentification of all instruments in a 2SLS prodedure.

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