# The economic returns to schooling for Italian men. An evaluation based on instrumental variables[1]

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Abstract

We study the relationship between (log) current earnings and educational levels in Italy. In line with other international evidence, we find that OLS under-estimate the marginal return to additional education. When the endogeneity of educational choice is taken into account, the marginal return from one additional year in junior high school increases from 3.2 to 5 percent. Similarly, the marginal return from one additional year in secondary school or in college increases respectively from 3.4 to 4.2 percent and from 6.4 to 7.2 percent. Using longitudinal data, we also find that individuals of the same age with higher education experience faster earnings growth. Hence, there is evidence that wage differentials by education widen as individuals grow older.

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# 1 Introduction

According to human capital theory, individuals can allocate their active time either to production or to the accumulation of human capital. One typical way of accumulating this capital is to spend time at school and acquire better (or higher) education. Human capital is useful because it affects individual productivity. Since earnings closely follow individual marginal contribution to output, individuals with more human capital earn, ceteris paribus, higher wages.

This prediction has led to substantial empirical work that investigates the relationship between education and earnings and, more specifically, the economic returns to education. As reviewed recently by Card (1994), the empirical assessment of this relationship is complicated by the fact that unobservable individual effects influence both earnings and education, making standard estimation techniques not adequate. Awareness of this problem has led to a new generation of research, starting perhaps with Griliches (1977), that focuses on instrumental variables methods.

The current paper is in this tradition and looks at the economic returns to education in Italy. Compared to the US or to the UK, there has been relatively little empirical research on this topic in Italy, partly because of the relative scarcity of suitable microeconomic data-sets. Our empirical work relies upon the most recent waves of the survey on the income and wealth of Italian households carried out by the Bank of Italy. Based upon these data, we construct both a sectional and a longitudinal data-set.

We use sectional data to study the relationship between (log) current earnings and educational levels, given other observed characteristics. Earnings at a given point of time, however, capture only part of the economic returns to education, that involve the stream of earnings over the entire working lifetime. Sectional data can also be used for inference about the relationship between earnings growth and educational outcomes, because they include information on individuals with different age and the same educational attainment. Results from these data, however, can be interpreted as informative of the shape of individual age-earnings profiles only when some restrictive assumptions hold. Because of this reason, we also use longitudinal data.

In line with other international evidence, we find that OLS underestimate the marginal return to additional education. When the endogeneity of educational choice is taken into account, the marginal return from one additional year in junior high school increases from 3.2 to 5 percent. Similarly, the marginal return from one additional year in secondary school or in college increases respectively from 3.4 to 4.2 percent and from 6.4 to

Table 1: Population with at least upper secondary education. 1992. Percentages. By age group.

| | 25-34 | 35-44 | 45-54 | 55-64 |
|---|---|---|---|---|
| Italy | 42 | 35 | 21 | 12 |
| OECD average | 65 | 58 | 50 | 38 |

Source: OECD (1995).

7.2 percent. We also find that family background variables significantly affect educational choice but do not affect the returns to education, contrary to the evidence presented by Altonji and Dunn (1996) for the US.

We apply both within groups and random effects estimators to our longitudinal data and find evidence that earnings growth increases with education but does not vary with individual age. This suggests that higher education not only yields higher earnings but also higher earnings growth. Hence, wage differentials by education widen as individuals grow older. With the partial exception of individuals with an upper secondary degree, these findings are confirmed by our cross section estimates, that are based on a larger sample of individuals.

The paper is organized as follows. Section 2 presents a brief description of the Italian educational system. Section 3 reviews previous work in the area. Section 4 introduces the data. Sections 5 and 6 focus respectively on sectional and longitudinal data. Conclusions follow.

# 2 Education in Italy

The current education system in Italy is composed of primary, secondary, upper secondary and tertiary education. Primary school is compulsory for children aged between 6 and 11 years. Lower secondary education is also compulsory, free of charge and lasts three years. Post-compulsory education is differentiated into the following categories: classical, scientific and pre-school teacher training; artistic education; technical schools and vocational education[1]. Upper secondary education lasts from three to five years, depending on the type of school. Since 1969, the selection of the type of school does not preclude access to tertiary education. Graduation from upper secondary schools requires a leaving certificate examination and access to tertiary education is only conditional on passing this exam.

As shown in Table 1, educational attainment, measured by the percent-

[1] See OECD (1995) for further details.

Table 2: Ratio of Upper Secondary and Tertiary Graduates to Population at Theoretical Age of Graduation. 1992.

| | Upper Secondary | Tertiary |
|---|---|---|
| Italy | 68.5 | 10.2 |
| OECD average | 84.8 | 20.8 |

Source: OECD (1995) and Checchi (1997).

Table 3: Net Enrolment Rates in Schools at 15 and at 17 Years of Age.

| | 16 Years | 17 Years |
|---|---|---|
| Italy | 65.0 | 55.0 |
| OECD average | 86.9 | 75.4 |

Source: OECD (1995) and Censis (1992).

age of individuals with at least upper secondary education, is much lower in Italy than in the OECD average, independently of the age group. An alternative measure of performance of the education system is the percentage of graduates in the population at theoretical age of graduation. As Table 2 suggests, this percentage is significantly lower in Italy than in the OECD average: less than 70% of individuals at theoretical age of graduation completed upper secondary education in Italy in 1992, compared to about 85% in the OECD average. This difference is partly explained by the high dropout rate in the Italian system. According to a study by CENSIS (1992), out of a cohort of 100 individuals entering the first year of lower secondary education, only 80 individuals enroll in upper secondary schools. Of these, 49 graduate after five years and 33 enroll in a university course. Among those entering tertiary education, only 10 individuals graduate [2].

The importance of dropouts in the Italian system is also highlighted by Table 3, that shows enrollment rates in schools by individuals aged 16 and 17. While in Italy only 6.5 teenagers out of 10 are still in school at 16, this proportion rises to close to 9 in the OECD average.[3]

Figure 1 to 4 present time series information for the four educational levels, primary school, junior high, upper secondary and college. Consider

[2] In a recent detailed study, Trivellato and Bernardi (1995) show that the Italian system is not only characterised by a high dropout rate but also by a relatively high number of irregular students, especially in the South of the country. According to a survey by ISTAT, the Italian national statistical o¢ ce, 21.3% of Italian individuals not in school, aged between 14 and 39 and without a college degree, had quitted school in 1989.

[3] A fairly recent discussion of the sociological aspects of the dropout problem in the Italian educational system can be found in Moscati (1989).

first primary school and notice that the average dropout rate over the five years was close to 30 percent of enrolled students after the war, declined to about 10 percent in the early sixties and converged to zero in the seventies[4].

This decline in the dropout rate has been accompanied both by an increase in the proportion of graduates in the population at theoretical age of graduation and by a substantial decline in the pupils to teacher ratio.

Figure 2 tells a qualitatively similar story for junior high school. It is perhaps worth noticing that, even after the 1962 reform made this school level compulsory, more than 15 percent of pupils enrolled in junior high schools quitted during the sixties and more than 5 percent did so during the eighties. Hence, implementation of the law has been rather poor[5].

Next, consider secondary school and Figure 3. The percentage of individuals enrolled in any secondary school has steadily increased from less than 20 percent of the relevant population cohort (aged 14) in the late fifties to slightly less than 80 percent in the early nineties. At the same time, the percentage of graduates in the population at the theoretical age of graduation (age 19) has reached about 60 percent in the late eighties, to increase further to slightly less than 70 percent in 1992. On the other hand, the percentage of individuals dropping out of school has significantly increased during the sixties and reached about 16 percent in 1992.

Finally, Figure 4 focuses on college education. While enrolment has increased over the years to about 20 percent of the relevant population, the percentage of graduates is close to 10 percent. The reason of this is clear from the third part of the figure, that plots the percentage of college graduates over individuals enrolled at college five years earlier. It turns out that this percentage has collapsed from close to 20 percent in the late sixties to about 10 percent in the early nineties. This can be partly explained with the fact that individuals tend to spend more than the required time at college to complete a degree.

In most developed countries, people with a lower level of educational attainment are more likely to be unemployed than those with a higher attainment[6]. This is also the case for Italy, with the exception of the young members of the labor force. As shown in Table 4, the youngest cohort of individuals in the labor force experiences high unemployment rates, quite independently of the level of education. More strikingly, the unemployment rate for this cohort is highest among individuals with primary and with college education.

[4]The dropout rate usually peaks after the first year at school.

[5]See Checchi (1997) for a detailed discussion.

[6]See for instance Layard, Nickell and Jackman (1991) and OECD (1994).

Table 4: Unemployment Rates by Education and Age Groups. Italy, 1993

| | 15-29 | 30-39 | 40-49 | 50-59 | 60-70 |
|---|---|---|---|---|---|
| Primary | 0.30 | 0.14 | 0.06 | 0.04 | 0.02 |
| Lower Sec. | 0.20 | 0.08 | 0.04 | 0.02 | 0.01 |
| Upper Sec. | 0.25 | 0.06 | 0.02 | 0.02 | 0.01 |
| Tertiary | 0.26 | 0.05 | 0.01 | 0.01 | 0.00 |
| Total | 0.23 | 0.07 | 0.04 | 0.03 | 0.02 |

Source: ISTAT (1993).

High unemployment among young individuals with relatively high education can be partly explained by regional and occupational mismatch between labor demand and labor supply. While labor demand concentrates in the Northern and Central areas of the country, unemployment is particularly high in the under-developed South[7]. Other important factors are both the lack of systematic links with private industry and the poor signalling role of education. On the one hand, private industry in Italy has traditionally been characterized by low intensity of education and by the heavy reliance on internal training rather than on formal education [8]. As remarked by Michael Porter (1989) in his well known study of national comparative advantage, the success of Italian industrial clusters has been based more on informal training, often provided by the extended family, that operates small artisan shops and small and medium firms, than on formal education.

On the other hand, the poor performance of the education system in Italy and the lack of emphasis on competition among students has limited

[7]See Brunello, Lupi and Ordine (1997) for a recent discussion of regional disparities in the Italian labor market.

[8]See Jannaccone Pazzi and Ribolzi (1990). The current Italian Prime Minister, an economist, has argued that ... Italy has experienced strong economic development with a strong entrepreneurship but without a su¢ cient stock of professional competencies.. (Prodi (1993)). According to David Marsden, ... it is fair to suppose that the main form of training for skilled labour in France and Italy consists of work experience and training organised by individual employers.. (quoted by Regalia and Regini (1995), p. 141, who add that ... more recent research largely con rm that this was indeed the principal way in which skills were formed and developed in much of Italian industry ). According to Blanch ower and Freeman, ... in Germany ... apprenticeships move youths from school to the industry in which they nd permanent work ... in Japan, rms tend to recruit from particular colleges and universities or from specific high schools ... in yet other countries, youths rarely work while in school and are often jobless for a long period after leaving school before they obtain their first job ... Italy and Spain are examples of this pattern. (p. 1).

Table 5: Distribution of employment by public\private sector and by education. Net of agriculture. Italy 1993.

| | Private | Public | Total |
|---|---|---|---|
| Primary | 78.9 | 21.1 | 100 |
| Lower Sec. | 70.5 | 29.5 | 100 |
| Upper Sec. | 56.5 | 43.5 | 100 |
| Tertiary | 25.3 | 74.7 | 100 |

Source: Bank of Italy (1993).

the signalling role of schooling[9]. This and the limited demand for highly educated workers by private industry imply that the main employer of high education workers in Italy is the public sector, as shown in Table 5 below.

# 3 The Returns to Education in Italy: Previous Work

Selected previous empirical work that estimates the returns to education in Italy is briefly summarized in Table 6[10]. While recent studies focus on the Bank of Italy nationally representative sample of Italian households and use IV techniques, the research carried out in the late eighties is based both on OLS and on more heterogenous and often less representative data. One of the first studies in this area, by Antonelli (1985), uses a regional data-set and estimates a standard Mincerian earnings function by ordinary least squares. He estimates that an additional year of schooling increases annual net earnings by 4.6 percent. A similar result is found by Cannari and Sestito (1989), who use a larger sample from the 1986 wave of the survey by the Bank of Italy (BI). On the other hand, Lucifora and Reilly (1990) estimate similar earnings functions using the ENI special survey on individual earnings and find that the marginal return to schooling is slightly higher for women than for men.

The new generation of papers on the issue starts with Cannari and D A-lessio (1995) s paper, based upon the 1993 wave of the BI survey. These

[9]See Spence (1974) for a discussion of the signalling role of education. Porter (1989) emphasises the poor quality of the Italian schooling system and argues that ... in order to sustain growth and to acquire professional competences, Italians need to improve their basic knowledge of mathematics, computers and other key disciplines ... (Porter (1989), p. 812.)

[10]A more detailed but older survey is in Lucifora and Sestito (1993).

authors use family background variables as instruments of educational outcomes and find that the marginal return to education is close to 7 percent, much higher than in previous research. A similar result is obtained by Colussi (1996), who uses the same wave and a similar set of instruments to find that the marginal return of an additional year of education is 7.6 percent. Finally, Flabbi (1996) uses the 1991 wave of the BI survey and estimates the returns to education separately for men and women with an IV approach based upon the identification of exogenous changes in the schooling system. He finds that the marginal effect of one year of additional education is respectively 6.2 percent for men and 5.6 percent for women.

Table 6: Selected previous empirical work on the returns to education in Italy.

| Author | Data | Method | Marginal return |
|---|---|---|---|
| Antonelli (1985) | ER | OLS | 4.6 |
| Lucifora-Reilly (1990) | ENI | OLS | 4.0 (males) 3.6 (fem) |
| Cannari-Sestito (1990) | BI | OLS | 4.6 |
| Cannari-D Alessio (1995) | BI | IV | 7.0 |
| Colussi (1996) | BI | IV | 7.6 |
| Flabbi (1996) | BI | IV | 6.2 (males) 5.6 (fem) |

Notes: ER: Emilia-Romagna regional data; ENI: Indagini retribuzioni di fatto; BI: survey on income and wealth of Italian households held by the Bank of Italy.

# 4 The Data

Our data are from the survey on income and wealth of Italian households carried out every two years by the Bank of Italy (BI) more or less continuously from 1977 to 1995. The frame of the survey has changed significantly over the years and a panel section has been added only recently [11]. These data have been extensively used by Italian labor economists, mainly because they are the only national data-set that includes information both on earnings and on individual characteristics.

In the current paper, we look at the relationship between (real hourly) earnings and educational outcomes using both cross section and longitudinal information. Notice that the available measure of earnings is net

[11] See Brandolini and Cannari (1994) for a detailed description of the survey.

of income taxes and pension contributions, therefore any comparison with alternative gross measures might result in a underestimation of return to schooling. In the next section, we use repeated cross sections from the 1993 and the 1995 waves of the BI survey. In the following section, we use longitudinal data and consider individuals who were interviewed and employed in 1991, 1993 and 1995.

# 5 The Returns to Education: Estimates based on Sectional Data

In this section, we estimate the returns to education for Italian male household heads using cross sectional data from the 1993 and the 1995 waves of the BI survey. We select male household heads to avoid issues of labor force participation and household formation and focus more closely on education. Since these data include individuals who were interviewed twice, both in 1993 and 1995, we randomly allocate half of the recipients of multiple interviews to 1993 and the other half to 1995.

A key feature of college education in Italy is that the average time required to complete a degree is significantly longer than the prescribed number of years. Because of this, one cannot rule out the possibility that young and working household heads are still enrolled in college in their late twenties. Moreover, there are very few household heads younger than 30 in our sample. In the paper, we focus on household heads who are at least 30 years old. We also exclude individuals who went to school before, during and in the immediate aftermath of the Second World War and consider only those who were born from 1942 onwards. Thus the younger and the older individuals in our sample are respectively 30 and 53 years old.

Our empirical strategy follows closely previous work by Vella and Gregory (1996) and Harmon and Walker (1996) and consists of estimating the following two equations

$$\ln w_i = X_i^{\prime}\beta + \sum_{h=1,3} \alpha_h E_{ih} + u_i \tag{1}$$

$$S_i^* = Z_i^{\prime}\gamma + \nu_i \tag{2}$$

where $w$ is the real hourly wage, $E$ are educational dummies, that correspond to the highest degree achieved by the individual, $X$ and $Z$ are vectors of observed attributes, $u$ and $\nu$ are normally distributed error terms with zero means and finite variances and $S^*$ is the latent level of education. As in Vella and Gregory (1996), we define $S$ as the observed level of education, that takes the following discrete values

$$\begin{aligned} S_i &= 1 \text{ if } S_i^* < \mu_o;\ S_i = 2 \text{ if } \mu_o \leq S_i^* < \mu_1; \\ S_i &= 3 \text{ if } \mu_1 \leq S_i^* < \mu_2;\ S_i = 4 \text{ if } S_i^* \geq \mu_2. \end{aligned} \tag{3}$$

and associate $S$ to the educational dummies by setting $E_{ih} = 1$ if $S_i = h$ and $E_{ih} = 0$ otherwise.

As discussed in a very large literature summarized by Card (1994), ordinary least squares estimates of the returns to education $\alpha_h$ are consistent only if the errors $u_i$ and $\nu_i$ are uncorrelated. In practice, a correlation emerges either because of common unobservable factors such as individual ability or because of measurement errors. One strategy for dealing with this problem is to find a set of variables that influence educational choice without affecting individual earnings (conditional on schooling)[12].

We use a two-step procedure. In the first step, we estimate an ordered probit equation for (2) and use the estimates to compute the relevant score $\sigma$[13]. In the second step, we augment equation (1) with the score $\sigma$, that captures the correlation between the error terms in equations (1) and (2) [14], and apply ordinary least squares. This method is closely related to instrumental variables estimation.

Our specification of the ordered probit equation includes individual age, a year dummy, the interaction of age with the dummy $D51$, equal to $1$ for individuals born from 1951 onwards and to $0$ otherwise, and a set of variables that measure family background, including both the highest completed educational level and the occupation held by the father and the mother of the interviewed household head.

The dummy $D51$ picks up an important exogenous event, the Law 910 of December 1969, that extended the possibility of enrolment in college to individuals who have completed secondary education, independently of the curriculum chosen in secondary school. Since expected age of completion of secondary school is in general 19 years, this opportunity was mainly open to the cohorts born from 1951 onwards. A rough indication of the impact of the reform can be obtained by comparing the percentage of 19 years old

[12] Given the non-linearity of (3), the treatment e⁄ ect of $Z$ on $S^*$ is identi ed even when vectors $X$ and $Z$ coincide. See Card and Vella (1997).

[13] See Idson and Feaster (1990) for details on the computation of the score.

[14] The assumption that educatonal choice can be modeled with an ordered probit is common in the literature but not without problems, as discussed in detail by Altonji (1993). In particular, the ordered probit model ignores that the choice of a given level of education is sequential to the completion of the immediately lower level.

Table 7: Distribution of educational attainment by cohort of birth. 2943 observations.

| | 1942-50 | 1951-65 |
|---|---|---|
| Primary | 24.22 | 11.17 |
| Junior High | 34.37 | 38.92 |
| Secondary | 31.18 | 38.57 |
| Tertiary | 10.23 | 11.35 |
| Total | 100 | 100 |

individuals enrolling in college shortly before and shortly after the reform. It turns out that enrolment rates were 16.3% of the relevant population for individuals born in 1949 and 27.3% for those born in 1952. [15] On the other hand, the percentage of high school graduates enrolling in college was 54% for the 1949 cohort and 66% for the 1952 cohort.

Higher enrolment in college after the reform, however, had a rather limited impact on the percentage of college graduates in the population at theoretical age of graduation, partly because the percentage of irregular students (fuori corso), who are enrolled at college longer than the number of years required to complete the curriculum, increased sharply for the cohorts enrolling in college from the early 70s (see Figure 5). Hence, the increase in the number of college students was accompanied by a reduction in the efficiency of the college system and by an increase in the average time required to complete the degree.

These combined effects are partially reflected in Table 7, that compares educational achievement for individuals in our sample born before and after 1951. It turns out that the percentage of college graduates is only 1% higher among the younger cohorts. At the same time, there is a consistent increase in the percentage of individuals with junior and upper secondary education.

The table suggests that the dummy $D51$ picks up both the exogenous event of December 1969 and the general increase in the level of schooling achieved by the population who went to school during the economic boom of the late 50s and later.

The selection of family background variables as additional instruments in the ordered probit has two potential problems. First, the individual

[15] We choose 1949 and 1952 to minimize the risk of including individuals born before 1951 who completed their secondary school later than at the expected age. By taking close years, we also try to reduce the impact of aggregate factors, such as the increase in real income per-capita and the general trend towards more education.

is asked to recall both the highest educational level and the occupation held by his parents when they had his current age. Beside the obvious measurement issues, it is not clear whether information based on the same age as the respondent is always the most relevant. This is especially the case for the profession of the parents, that could have changed with respect to the profession held during the schooling period of the interviewed individual.

Second, and perhaps more important, family characteristics could affect the returns to education[16], thus failing to satisfy the necessary condition for instruments validity. We deal with this problem by using the test of over-identifying restrictions proposed by Card and Vella (1997). The test can be run in two steps. In the first step, we obtain the OLS residuals from the regression of equation (1) augmented by the score $\sigma$. In the second step, these residuals are regressed both on the vector $X$ and on the vector of additional instruments $Z$. The resulting LM test has an asymptotic $\chi^2$ distribution with degrees of freedom equal to the number of over-identifying restrictions [17].

We start describing our results with Table 8, that shows the means and standard deviations of the main variables used in the paper.

Table 9 presents the estimates of the ordered probit, that includes also among the regressors individual age minus 40 ($AGE$), area, year and town size dummies. As expected, the higher the educational attainment of both parents, the higher the level of education. The occupation of the father also significantly affects educational choice. In particular, educational attainment is higher if the father was either a professional or an entrepreneur. Whether the mother was working also matters, and individuals with the mother not working reach lower educational levels. Moreover, individuals belonging to the cohorts born after 1950 have higher education. This effect is stronger for older individuals in the cohorts, a result consistent with Figures 4 and 5, that shows how the e¢ ciency of Italian colleges has declined over the years.

Table 10 presents both the OLS and the IV estimates, obtained by including in the original specification individual age minus 40, its square, educational dummies, the interactions of education both with individual age and its square, town, area and year dummies. It turns out that we cannot reject the exclusion of interactions between age and education. Hence, we drop this group of interactions from the regressions.

The test of the over-identifying restrictions is equal to 3.361, with a p-value of 0.849. Thus we cannot reject the validity of the over-identifying

[16]See Altonji and Dunn (1996) for a detailed discussion in a similar set-up.

[17]See also Main and Reilly (1993).

Table 8: Means and standard deviations of the variables used in the cross sectional analysis.

| Variable | Mean | S.D. |
|---|---|---|
| Real Wage | 0.110 | 0.05 |
| Age | 41.53 | 6.29 |
| Primary School | 0.166 | |
| Junior High | 0.371 | |
| Upper Secondary | 0.355 | |
| Tertiary | 0.109 | |
| Area: North-West | 0.207 | |
| Area: North-East | 0.202 | |
| Area: Centre | 0.234 | |
| Area: South | 0.306 | |
| Area: Islands | 0.050 | |
| Town size $<$ 20,000 | 0.237 | |
| 20,000$<$Town$<$400,000 | 0.221 | |
| Town$>$400,000 | 0.542 | |
| Mother:No degree | 0.290 | |
| Mother: Primary | 0.559 | |
| Mother: Secondary | 0.097 | |
| Mother: Tertiary | 0.053 | |
| Father: No degree | 0.244 | |
| Father: Primary | 0.543 | |
| Father: Secondary | 0.128 | |
| Father: Tertiary | 0.084 | |
| Mother not working | 0.718 | |
| Father blue collar, self-employed or unemployed | 0.783 | |

Table 9: Ordered Probit Estimates

| Variable | Coe¢ cient | p-value |
|---|---|---|
| age | -0.028 | 0.00 |
| D51*age | 0.029 | 0.00 |
| Year=1995 | 0.163 | 0.00 |
| Mother educ: primary | 0.498 | 0.00 |
| Mother educ: secondary or higher | 1.001 | 0.00 |
| Father educ: primary | 0.224 | 0.00 |
| Father educ: secondary or higher | 0.668 | 0.00 |
| Mother not working | -0.078 | 0.09 |
| Father professional or entrep. | 0.470 | 0.00 |
| $\mu_1$ | -0.439 | |
| $\mu_2$ | 0.821 | |
| $\mu_3$ | 2.219 | |
| NOBS | 2943 | |

restrictions[18].

As it happens in most of the international literature, we find that OLS results under-estimate the returns to additional education. In particular, the estimated coefficient of the junior high school dummy is 57.3 percent higher with IV estimates than with OLS estimates. This percentage is equal respectively to 36.3 and 25.4 for secondary and tertiary education. Figure 6 plots the estimated earnings-age pro les for an individual randomly drawn from the population living in a small town of the North-West and interviewed in 1993[19]. While OLS more or less coincide with IV in the case of primary education, they under-estimate the returns to junior, secondary and college education.

The use of educational dummies for each level of completed schooling, rather than the more standard years of schooling, implies that the estimated marginal return to schooling varies with the completed degree. Because of this, our results are not immediately comparable with previous literature summarized in Table 6. Assuming that the marginal return of an educational degree can be evenly distributed among the regular years of school required to complete the degree, we nd that the marginal return to an additional year of schooling is 5 percent in junior high school, 4.2 percent

[18]These results do not con rm the evidence for the US presented by Altonji and Dunn (1996).

[19]Expected log earnings for such an individual are computed by ignoring the score.

Table 10: OLS and IV Estimates. Dependent variable: ln wage

| Variable | OLS Coefficient | OLS p-value | IV Coefficient | IV p-value |
|---|---|---|---|---|
| Junior High | 0.096 | 0.00 | 0.151 | 0.00 |
| Secondary | 0.267 | 0.00 | 0.364 | 0.00 |
| Tertiary | 0.585 | 0.00 | 0.735 | 0.00 |
| age | 0.012 | 0.00 | 0.013 | 0.00 |
| age$^2$ | -0.001 | 0.00 | -0.001 | 0.00 |
| Junior High*age$^2$ | 0.001 | 0.02 | 0.001 | 0.02 |
| Secondary*age$^2$ | 0.002 | 0.00 | 0.002 | 0.00 |
| Tertiary*age$^2$ | 0.001 | 0.02 | 0.001 | 0.03 |
| Area: North-East | 0.038 | 0.04 | 0.039 | 0.04 |
| Area: Center | -0.002 | 0.92 | 0.002 | 0.93 |
| Area: South | -0.064 | 0.00 | -0.057 | 0.00 |
| Area: Islands | -0.031 | 0.22 | -0.027 | 0.29 |
| Small Town | 0.009 | 0.61 | 0.008 | 0.66 |
| Medium Town | 0.030 | 0.05 | 0.025 | 0.10 |
| Large Town | 0.069 | 0.01 | 0.057 | 0.03 |
| Year=1995 | -0.078 | 0.00 | -0.085 | 0.00 |
| SCORE | - | | -0.052 | 0.00 |
| NOBS | 2943 | | 2943 | |
| LM test $\chi^2(7)$ | | | 3.361 | 0.849 |

Note: heteroskedasticity-corrected standard errors. The standard errors in the IV regressions are corrected by using the procedure described in Newey (1984). The regression includes a constant term.

in upper secondary school and 7.2 percent in college[20].

Notice that the estimated coe¢ cient of the score attracts a negative and significant sign, implying that the covariance between unobservable shocks to earnings and to educational choice is negative. Recalling that the score is positive if an individual attains a higher educational level than predicted and negative otherwise, a possible interpretation of the negative sign attracted by $\sigma$ in the earnings regression is that abler individuals have a higher marginal cost of schooling in terms of foregone earnings, because they receive more attractive wage offers. Hence, these individuals tend to

[20] These are broad estimates. First, both secondary schools and college degrees could require a different number of years at school, depending on the specialization. Second, a substantial percentage of college graduates take a longer number of years to complete the curriculum.

acquire less than predicted education and to earn higher wages (See Vella and Gregory (1996)).

The estimates presented in Table 10 describe age-earnings profiles for a repeated cross section of individuals. Assuming that we could interpret these results as suggestive of individual age-earnings profiles[21], the jointly signi cant interactions between $AGE^2$ and educational dummies suggest that the slopes of these profiles vary with educational attainment.

It turns out that earnings growth is positive and constant for junior high school and for college graduates, positive but declining with age for primary school graduates, and positive and increasing in age for secondary school graduates. These results suggest that wage differentials by education do not converge over time, with the single exception of the earnings of secondary school graduates with respect to the earnings of college graduates.

It is important to stress, however, that interpreting cross section evidence in terms of individual earnings pro les is problematic whenever the standard steady state assumptions are unlikely to hold. For this reason, we turn in the next section to the analysis of longitudinal data.

# 6 The Returns to Education: Estimates based on Panel Data

An evaluation of the economic returns to schooling requires that we consider how lifetime earnings vary with educational outcomes. If age-earnings profiles share the same slope, so that earnings growth does not depend on education, cross sectional comparisons of the earnings of individuals with the same observed characteristics, who differ only in their educational level, provide sufficient information. When earnings growth varies across individuals with different educational levels, however, the comparison of individuals at a given point of time could be misleading.

A good example is when there is over-education in the labor market [22]. In this case, individuals with a college degree are found to work in occupations that require less schooling and earn lower earnings than individuals with similar levels of schooling who hold jobs that require the level of education they have obtained. Lower earnings, however, could be a temporary

[21] Age, time and cohort effects cannot be jointly identi ed due to their exact linear relation. In the paper we present results where time dummies and a second order polynomium in age are used, but results do not change substantially using cohort dummies and age polynomium.

[22] See Sicherman (1991) for a discussion.

phenomenon if these entry jobs allow individuals to access through job mobility more adequate occupations, that provide faster earnings growth.

With cross sectional data, differences in earnings growth by education can be captured by comparing individuals with the same schooling but with different age, that is, by interacting educational dummies with individual age or its square. Clearly, unless some restrictive assumptions such as a stationary environment hold, the age-earnings profiles estimated from cross sectional data need not re ect the shape of individual age-earnings profiles [23].

In this section, we use the longitudinal data included in the Bank of Italy survey to provide alternative estimates of the relationship between schooling and earnings growth. In particular, we select employed male household heads aged between 30 and 53 in 1991 who were continuously employed in the 1991, 1993 and 1995 waves of the survey. Our empirical specification is given by

$$\begin{aligned} \ln w_{it} &= \alpha_i + X_i'\beta + \sum_{h=1,3} \alpha_h E_{ih} + \zeta AGE_{it} + \delta AGE_{it}^2 + \\ & \sum_{h=1,3} \gamma_h E_{ih} AGE_{it} + \sum_{h=1,3} \psi_h E_{ih} AGE_{it}^2 + \xi T + \varepsilon_{it} \end{aligned} \tag{1}$$

where $\alpha_i$ is an unobservable time invariant individual effect, $X_i$ is a set of time invariant respondent characteristics, $T$ are time dummies and $\varepsilon$ is the error term, assumed to be orthogonal to the regressors. Since unobservable individual effects are likely to be correlated with educational outcomes we either use the within groups estimator, or the two steps IV technique presented in the previous section. Clearly, the within group estimator does not allow us to identify the coefficients that associate log earnings to educational outcomes or any other time invariant variable ($\alpha_h$ and $\beta$).

The results of the estimates based on a sample of 590 individuals always employed in 1991, 1993 and 1995 are presented in the first two columns of Table 11[24]. We find that both age and the interactions of age with educational dummies are jointly significant. On the other hand, both age squared and its interactions with educational dummies are not jointly significant.

Even though panel estimates are based on a smaller sample of stayers , the estimated effects of education obtained with IV random effects estimates

[23] See Jonsson and Klevmarken (1978) for a detailed discussion.

[24] The rst step order probit model has been re-estimated on the sub-sample used here with the same speci cation presented in the previous section. Results are similar to those presented in Table 9 and are avialable upon request. Allowing for autocorrelated errors in the wage equation does not change the main results.

Table 11: Panel estimates. Longitudinal data 1991-95. Number of individuals:590. Number of observations: 1770. Dependent variable: ln wage. Regressions include a constant term, town, area and year dummies.

| | Within groups | | IV random effects | |
|---|---|---|---|---|
| Variable | Coefficients | p-value | Coefficient | p-value |
| $AGE\psi$ | -0.0179 | 0.03 | -0.0026 | 0.57 |
| Junior High*$AGE\psi$ | 0.0225 | 0.02 | 0.0102 | 0.04 |
| Secondary*$AGE\psi$ | 0.0284 | 0.00 | 0.0195 | 0.00 |
| Tertiary*$AGE\psi$ | 0.0322 | 0.01 | 0.0275 | 0.00 |
| Junior High | - | - | 0.1817 | 0.00 |
| Secondary | - | - | 0.4001 | 0.00 |
| College | - | - | 0.6879 | 0.00 |
| $\sigma\psi$ | - | - | -0.0632 | 0.00 |
| $R^2$ within | 0.024 | | 0.022 | |
| $R^2$ between | 0.154 | | 0.395 | |
| $R^2$ overall | 0.116 | | 0.294 | |
| Tests for random effects: | | | | |
| Breusch-Pagan LM test $\chi(1)$ | | | 335.79 | 0.00 |
| Hausman test $\chi(5)$ | | | 9.88 | 0.08 |

are rather similar to those obtained with the IV method applied to sectional data. Moreover IV random e⁄ ects estimates confirm sectional data results : education cannot be considered exogenous and the score term attracts a negative sign.

A quick comparison of these results with those obtained from sectional data (IV estimates) reveal important differences. Both within group and random effects estimation procedures support the idea that earnings growth does not vary with age and is highest for college graduates. Estimated growth rates range between -1.8% to 2.4% for within group estimates and between 0% to 2.7% for IV estimates. In particular, earnings growth for college graduates based upon cross section data is much lower than growth estimated from longitudinal data (1.3% vs. 2.4%-2.7%). Even more strikingly different, earnings growth for secondary school graduates is increasing in age with sectional data and independent of age with longitudinal data. Finally, estimates based upon longitudinal data show that earnings growth increases monotonically with the level of education.

Hence, earnings growth varies with educational outcomes but not with age. In particular, we find that earnings growth is faster when educational attainment is higher. Notice that our estimates are based on a sample of household heads aged between 30 and 53. It is entirely possible that individuals with lower education, who have started working in their teens, experience steeper earnings growth before turning 30 and therefore are characterized by low or even negative growth when aged 30-53. On the other hand college graduates, who enter the labor market near the age 30, are more likely to be observed in the ascending portion of their age-earnings profile. Conditional on household status and age, however, our findings suggest that wage differentials by education widen as individuals grow older.

# 7 Conclusions

We have studied the economic returns to education in Italy using both cross section and longitudinal data. Our findings are briefly summarized as follows:

1. In line with other international evidence, OLS under-estimate the marginal return to additional education. The estimated coefficient of the education dummy is respectively 57.3, 36.3 and 25.4 percent higher with IV than with OLS estimates for junior high school, upper secondary and college education. Abler individuals, who received better wage offers, have lower education than predicted, because of the relative incentive to anticipate labor market entry.

2. Conditional on given characteristics, individuals with higher education have not only higher earnings but also higher earnings growth. Hence, wage differentials by education widen as individuals grow older.

We conclude with a remark on the research agenda. As already mentioned above, the assumption that educational choice can be modeled with an ordered probit ignores that the choice of a given level of education is sequential to the completion of the immediately lower level. In future work, we plan to address this issue by explicitly modeling educational choice as a sequential strategy.

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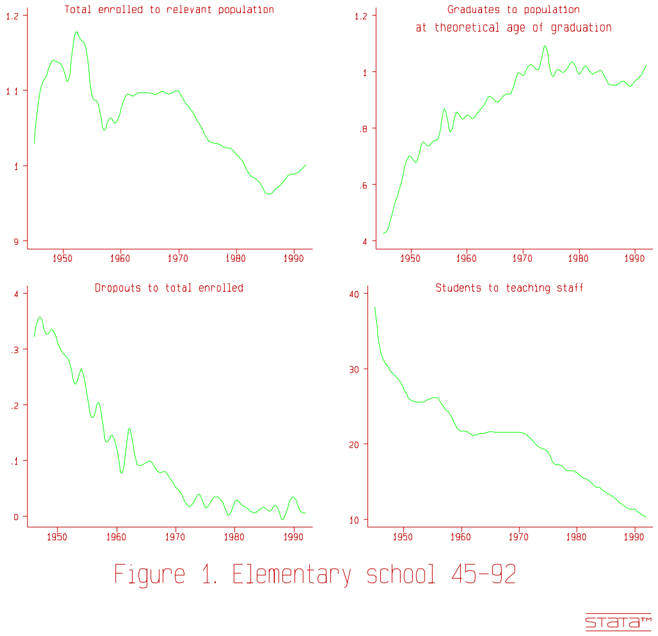

Total enrolled to relevant population
Graduates to population
at theoretical age of graduation
Dropouts to total enrolled
Students to teaching staff
Figure 1. Elementary school 45-92
STATA™

Figure 1:


Total enrolled to relevant population
Graduates to population
at theoretical age of graduation
Dropouts to total enrolled
Students to teaching staff
Figure 2. Junior high school 55-92
STATA™

Figure 2:

[1]

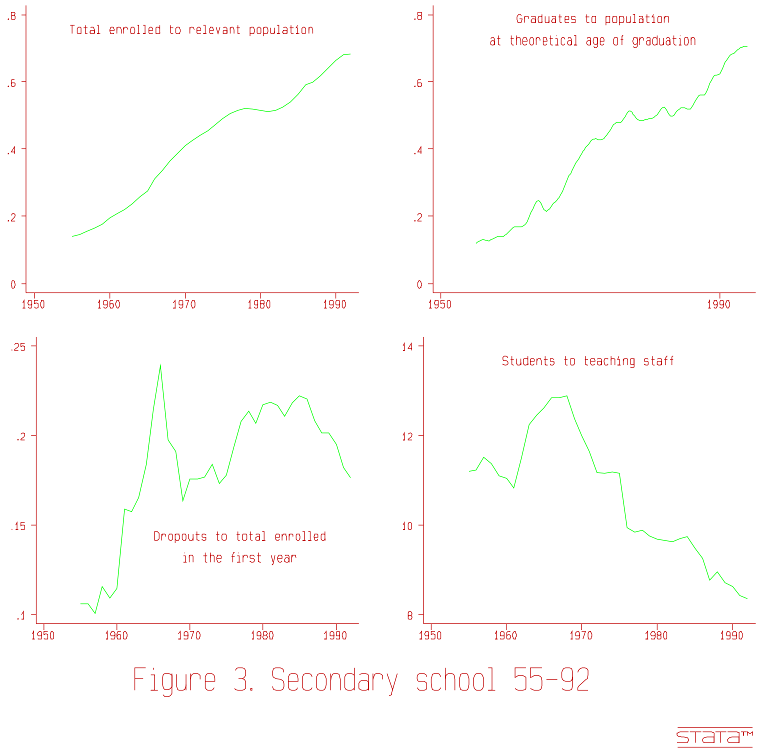
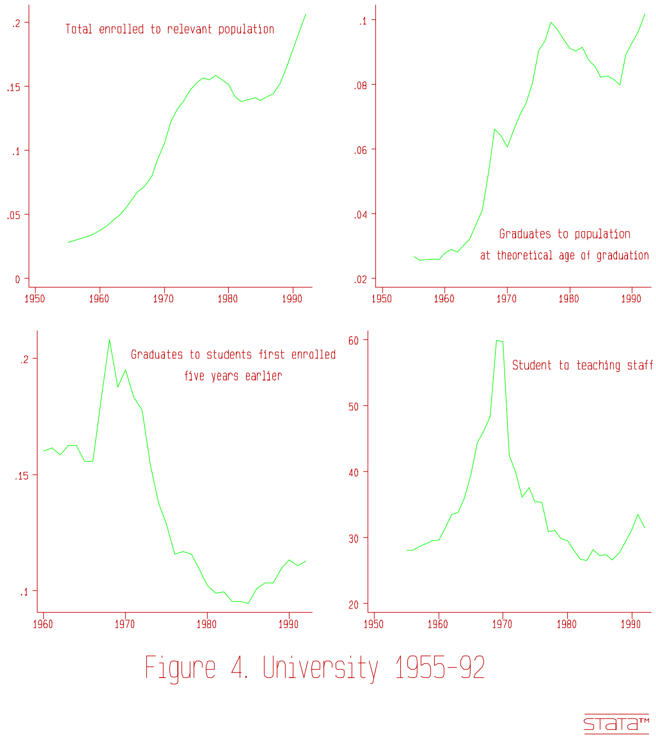

Total enrolled to relevant population
Graduates to population
at theoretical age of graduation
Dropouts to total enrolled
in the first year
Students to teaching staff
Figure 3. Secondary school 55-92
STATA™

Figure 3:

[1]



Total enrolled to relevant population
Graduates to population
at theoretical age of graduation
Graduates to students first enrolled
five years earlier
Student to teaching staff
Figure 4. University 1955-92


Figure 4:

[1]

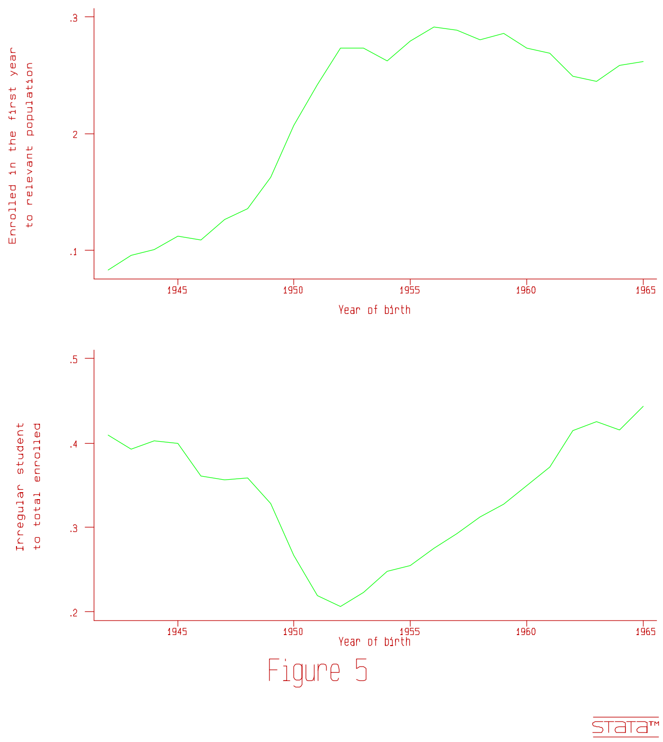

Enrolled in the first year
to relevant population
.3
.2
.1
1945
1950
1955
1960
1965
Year of birth
Irregular student
to total enrolled
.5
.4
.3
.2
1945
1950
1955
1960
1965
Year of birth
Figure 5
Stata™

Figure 5:

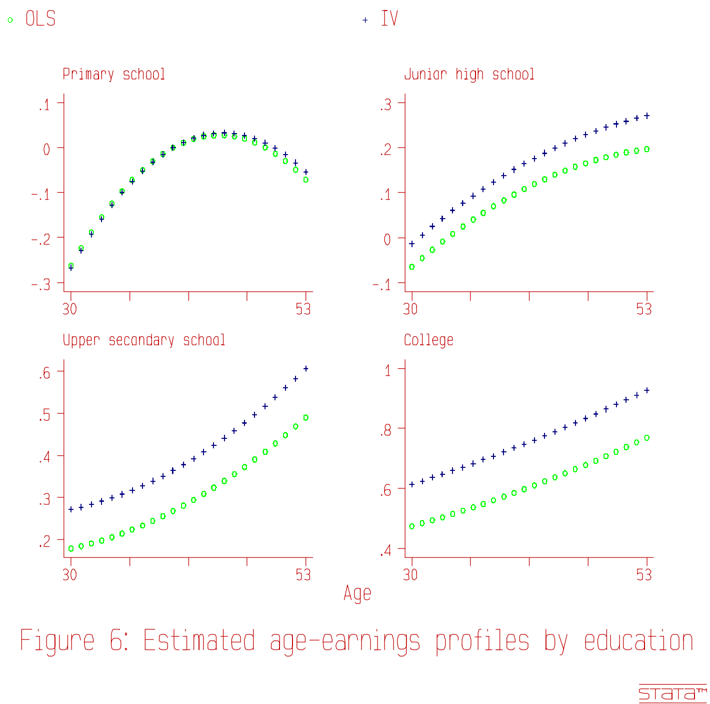

OLS
IV
Primary school
Junior high school
Upper secondary school
College
Age
Figure 6: Estimated age-earnings profiles by education

Figure 6: