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Economics of Education Review 24 (2005) 563–577
Economics of Education Review
www.elsevier.com/locate/econedurev

School quality and family background in Italy

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Received 22 January 2003; accepted 15 September 2004

Abstract

We study whether the combined significant reduction in the pupil–teacher ratio and increase in parental education observed in Italy between the end of the second World War and the end of the 1980s have had a significant impact on the educational attainment and the labor market returns of a representative sample of Italians born between 1941 and 1970. We find that the lower pupil–teacher ratio is positively correlated with higher educational attainment, but that the overall improvement of parental education has had an even stronger impact on attainment. We also find that the positive impact of better school quality on educational attainment and returns to education has been particularly significant for the individuals born in regions and cohorts with poorer family background. Parental education has had asymmetric effects, positive on attainment and negative on school returns. Better school quality has also had asymmetric effects on the returns to education, positive for individuals with poor family background and negative for individuals born in regions and cohorts with relatively high parental education. Our evidence suggests that better school quality, measured by a lower pupil–teacher ratio, is a technical substitute to parental education in the production of individual human capital. When school quality and family background are substitutes, an increase of public resources invested in education can be used to reduce the differences induced by parental education.

JEL classification: I21

Keywords: School quality; Family background; Italy

1. Introduction

There is a substantial literature that investigates the effects of indicators of school quality, such as class size and the pupil–teacher ratio, on test scores, educational attainment and the returns to education (see Hanushek (1986, 2002) and Card and Krueger (1996) for surveys of this literature). Most of this literature focuses on the US, but an increasing number of studies looks at different countries[1] and considers a comparative perspective.[2]

In spite of the large numbers of contributes in the area, there is no broad consensus on the economic effects of school quality. While there is agreement that better school quality improves educational attainment, the jury is still out on whether variations in class size

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[1]See for instance Dolton and Vignoles (1998), Wright (1999), Dearden, Ferri and Meghir (2000), Harmon and Walker (2000), Dustman, Rajah and van Soest (2002).

[2]See Wossman (2000) and Guldlach Wossman, and Gamelin (2001).

0272-7757/$ - see front matter
doi:10.1016/j.econedurev.2004.09.001

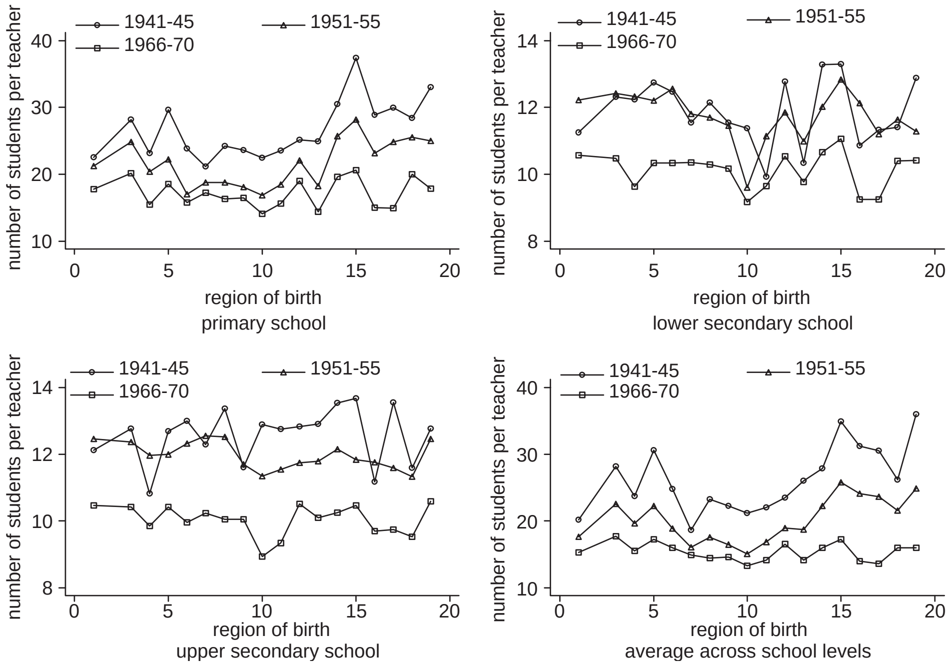

Fig. 1. Note to the figure: 1: Piemonte; 2: Lombardia; 3: Trentino; 4: Veneto; 5: Friuli; 6: Liguria; 7: Emilia; 8: Toscana; 9: Marche; 10: Umbria; 11: Lazio; 12: Abruzzo; 13: Molise; 14: Campania; 15: Puglia; 16: Calabria; 17: Basilicata; 18: Sicilia; 19: Sardegna. Pupil–teacher ratio by age cohort and region of birth—different school levels and average across school levels.

affect significantly performance tests and the returns to education.

Another factor with the potential of affecting both educational attainment and school returns is family background. Ermisch and Francesconi (2001) find that parents' educational attainments are very powerful predictors of their children's attainment. Sacerdote (2002) finds that being raised in a family with high socio-economic status greatly increases the probability that a child will attend college.[3] An important but somewhat overlooked question is how school quality and family background interact in the production of individual human capital. If they are technical substitutes, an improvement in school quality reduces the marginal contribution of family background to human capital, and therefore has the potential of reducing the differences induced by nurture. These differences widen when school quality and family background are technical complements. This paper considers this question, and investigates the relative contributions of family background and school quality to educational attainment and labor market performance in Italy.

The Italian institutional context is interesting but little explored. Primary and secondary education in this country is mostly public, virtually free, designed and organized centrally. Educational attainment is low by international standards (see OECD (2002)) but has increased significantly among the more recent cohorts. While the private returns to education are close to the European average (see Brunello, Comi, & Lucifora (2001)), the performance of Italian students in international cognitive tests is rather disappointing (see the PISA results in OECD (2002)).

Despite the centralization of school design, school quality in Italy has exhibited important variations both over time and across different areas of the country. One reason is that centralization is far from complete: while education standards and personnel are run by the Ministry, local authorities are responsible for school buildings. In particular, primary and lower secondary schools are built and managed by municipalities, and upper secondary schools are the responsibility of provinces. Another reason is that the hiring of new personnel occurs mainly if not exclusively at discrete intervals, since national competitions to fill vacancies take place every 4–5 years. These two factors, combined with demographic variations, produce significant variations in the pupil–teacher ratio across regions.

Fig. 1 shows the average pupil–teacher ratio in primary, lower secondary and upper secondary schools for the 19 Italian regions—tiny Val d'Aosta excluded—and three different age cohorts: compared to the cohort born between 1941 and 1945, the average pupil–teacher

[3]See also Betts (1996), Dustmann (2001) and Behrman, Foster, Rosenzweig, and Vashishtha (1999).

ratio experienced by the cohort born between 1966 and 1970 was almost half as high. If we consider the oldest cohort, the variation in the average pupil–teacher ratio across regions is in the range of 16 pupils in primary school (from 21 to 37), 3 pupils in lower secondary school (from 10 to 13) and 4 pupils in upper secondary schools (from 9 to 13).

Compared to the US, Italian society is less mobile, in terms both of educational attainment and of occupational outcomes. In Italy, less than 2% of the offspring of households where the father has not completed compulsory education has attained a college degree. In the US this percentage is close to 12%. Similarly, only 4% of Italians born in households where the father's income belongs to the lowest quartile of the income distribution receives college education, compared to 17% in the US (see Checchi, Ichino, & Rustichini (1999)). When family background plays an important role in the education and labor markets, it can generate persistence of social and economic stratification. In this environment, (public) school quality, decided and administered by the central government, can have a countervailing effect and increase the social and economic opportunities of the children of less fortunate households.

Gauging the relative importance of school quality and family background for educational attainment and labor market returns is important also because of the policy implications (see Hanushek (2002)). To illustrate, suppose that family background, measured by the educational attainment of parents, matters more than school quality in the production of individual human capital. Then policies that improve parental education, such as remedial adult schooling, continuous education and training, could be more effective than policies which reduce the number of pupils per class.[4]

The paper is organized as follows. Section 2 illustrates the empirical methodology, Section 3 presents the data, Section 4 shows the key results and Section 5 draws the main implications and concludes.

2. Methodology

Our investigation of whether changes in school quality, measured by the pupil–teacher ratio, and in family background, measured by the educational attainment of parents, affect educational attainment and the monetary returns to education follows the empirical approach taken by Card and Krueger (1992), Heckman, Layne, and Todd (1997) and Strayer (2002). An important part of the literature in this field (see Betts (1996) and Hanushek (2002) for reviews) uses school—specific measures of quality, which are not available in Italy. We use instead more aggregate measures, which vary with the region and the cohort of birth. On the one hand, the use of these measures leads to a systematic upward bias in the estimated effects of school quality. On the other hand, it reduces the measurement error bias associated to school—specific indicators (see Card and Krueger (1996)).

To estimate the impact of family background and school quality on the returns to education, we use a two - steps model. In the first step we perform the following regression

$$Y_{icrs} = \alpha_{crs} + \beta X_{icrs} + \gamma_{crs} E_{icrs} + \varepsilon_{icrs}, \tag{1}$$

where i is the individual, c the age cohort, r the region of birth and s the region of residence, Y is log annual earnings, α_{crs} are region of birth $\times$ age cohort $\times$ region of residence dummies, X is a vector of individual controls, including family background, E are years of education and γ measures the returns to education, which we allow to vary by cohort, region of birth and region of residence. In the second step we retrieve the estimated values of γ and estimate

$$\gamma_{crs} = \lambda_c + \lambda_r + \lambda_s + \lambda_{cs} + \lambda_{rs} + \phi Q_{cr} + \psi W_{cr} + \sigma Q_{cr} W_{cr} + \varepsilon_{cr}, \tag{2}$$

where λ are cohort, region of birth, region of residence, cohort $\times$ region of residence and region of birth $\times$ region of residence dummies, Q is school quality, which varies by region of birth and age cohort and W is average family background, by region of birth and age cohort. Therefore, family background affects individual earnings both directly and via its interaction with educational attainment. We define W as the highest number of years of education attained by parents in the household. Therefore, if the mother has only primary education (5 years) and the father has completed junior high school (8 years) the value of W for this household is 8 years. The value of W for the region and the cohort of birth is obtained by averaging over households.[5]

The region of residence dummies in (2) captures the effect of local labor markets on the returns to education. These effects can vary with the age cohort. The interaction of the dummies for the region of birth and

[4]Changing parental education in the short run may be very costly. However, "..long run policy may.... reasonably relate to family factors. For example, arguments for improving women's education in developing countries may reflect the potential impact of children's achievement more than normal arguments about the return to the mother of human capital investment....(Hanushek (2002), p. 39). Blöndal, Field and Girouard (2002) measure the returns to adult education and discuss policies to stimulate its acquisition.

[5]We have experimented with different combinations of parental education (average number of years in the couple, minimum level of education among the two), without finding significant differences. Results are available from the authors.

the region of residence accounts for the effects of endogenous migration across regions (see Heckman et al. (1997)). Variations in school quality and in family background capture the variability in the returns to education associated to the region and the cohort of birth. Finally, the interaction between Q and W is informative of whether these two factors are technical complements or substitutes in the production of human capital.[6]

Since educational attainment is measured as years of completed education, we use the following ordered probit model:

$$E_{icrs} = \Phi(\delta Z_{icrs}), \tag{3}$$

where Z is a vector of individual characteristics, including individual family background. Since educational attainment in (1) is the result of individual choice, we control for the endogenous selectivity by including in (1) the predicted score from the ordered probit model (3) (see Vella & Gregory (1996)).[7] The identification of Eqs. (1) and (3) require that at least one variable included in Z be excluded from the vector of regressors in (1). An important exogenous event in the recent history of Italian education is Law 910 of December 1969, which extended the possibility of enrolment in college to individuals with completed upper secondary education, independently of the type of secondary school attended. The possibility of enlarged access to university may have contributed to the rise of the demand for education, registered during those years in the country. Since expected age of completion of secondary school in Italy is 19 years, this opportunity was open to those born from 1951 onwards (see Brunello & Miniaci, (1999)). We define the dummy D51 as equal to 1 if the individual was born in 1951 or later and to 0 otherwise, and include it in vector Z but not in X.

3. The data

The data on individual annual earnings and educational attainment are drawn from the survey on the income and wealth of Italian households (SHIW), waves 1993, 1995, 1998 and 2000. Previous waves cannot be used because they lack information on family background. We consider only employees born between 1941 and 1970 who have a positive labor income and use the labor tax code to compute for each individual gross earnings from the original data on net earnings. We use annual rather than hourly earnings. There are two reasons for this. First, school quality and family background could also affect working hours; second, the available information is on weekly hours, which can be transformed into annual hours only by introducing an additional measurement error.

In Italy there are no data on school quality at the school level that cover a nationally representative sample. Therefore, we follow Card and Krueger (1992) and use aggregate measures of quality based on the region and the cohort of birth. This choice is driven by the fact that, due to privacy restrictions, the SHIW survey makes available to researchers only the information on the region of birth. We collect regional data on the pupil–teacher ratio for different types of schools, ranging from kindergarten to upper secondary education, every two years from 1944 to 1989. In 1944 the oldest cohort born between 1941 and 1945 was eligible to start kindergarten and in 1989 the youngest cohort born between 1966 and 1970 could have completed upper secondary education.[8]

The key assumption is that most individuals complete their schooling, from less than primary to upper secondary, in their region of birth. The plausibility of assigning to each individual the school quality of the region of birth could obviously be questioned. In the absence of individual information on the age when migration between regions eventually took place, we present in Tables 4 and 5 information on the percentage of individuals who live in the same region of birth.[9] Internal migration is very low among the young, with less than 5% of individuals younger than 21 living in a region different from the region of birth. Turning to older individuals, who were born between 1941 and 1970, there is evidence of short-distance migration in the more developed northern regions and of long-distance migration from the poorest areas of south to the more developed north. Assuming that the observed pattern of migration experienced by the young has not changed significantly over time, the evidence in the two tables suggests a tendency for migration to take place mainly after completing up to upper secondary education in the region of birth.

We match school quality to individuals in the sample by attributing to each individual the pupil–teacher ratio in the region of birth during the period when she went to school. To illustrate, an individual born in 1945 who

[6] See the appendix at the end of the paper.

[7] With only two threshold levels, a_1 and a_2, the score is equal to $-\varphi(a_1 - xb)/\Phi(a_1 - xb)$ if $y < a_1$, $\varphi(a_1 - xb) - \varphi(a_2 - xb)/\Phi(a_1 - xb) - \Phi(a_2 - xb)$ if $a_1 < y < a_2$ and $\varphi(a_2 - xb)/\Phi(a_2 - xb)$ if $y > a_2$, where φ is the density function, Φ is the cumulative distribution function, xb is the linear prediction and y is the dependent variable in an ordered probit model.

[8] We collect data every other year because of the limited short-term variability in the data. We exclude college education because the assumption that this type of education is completed mainly in the region of birth is not tenable in Italy.

[9] 84.4% of the individuals in our sample resides in the region of birth. This percentage is 80.9% for the oldest cohort and 90.1% for the youngest cohort.

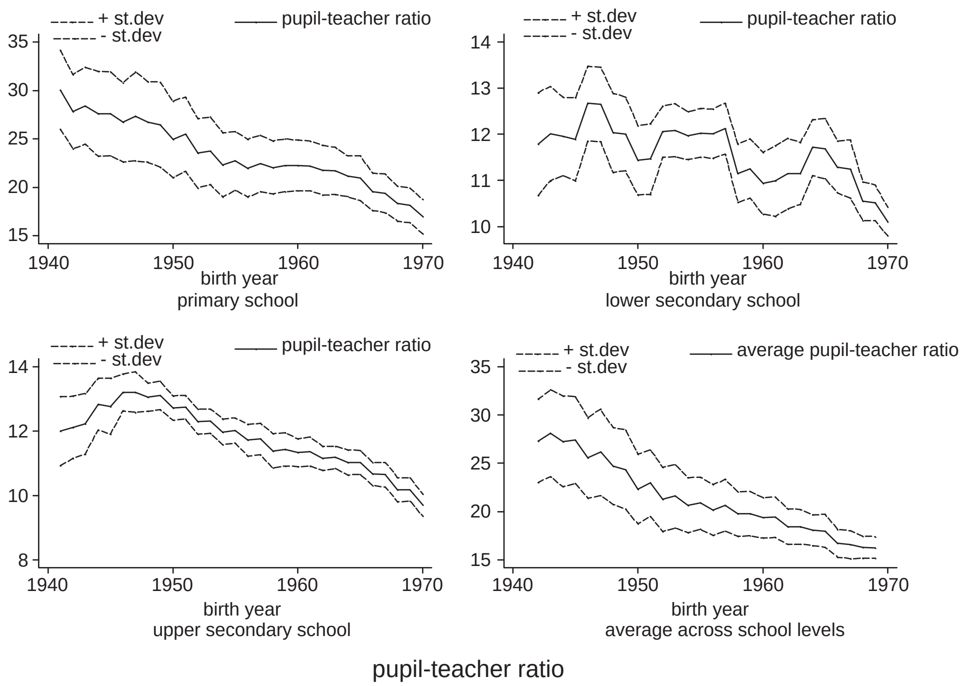

Fig. 2. Pupil–teacher ratio by year of birth—different school levels.

went to kindergarten between 1948 and 1951, to primary school between 1951 and 1956, to junior high school between 1956 and 1958 and to upper secondary school between 1958 and 1963 is assigned the pupil–teacher ratio associated to each type of school in the same sub-period. This matching can be performed because we know the highest educational level attained by each individual.

It is questionable, however, to associate to each individual only the school quality of the schools he/she graduated from. Consider for instance the choice of continuing education after junior high school. This choice is likely to be affected by the expected quality of upper secondary education, despite the fact that the individual could end up not enrolling. Moreover, we only have information on attained degrees and cannot rule out the possibility that an individual enrolls in a school level and is exposed to the associated school quality without completing the degree. These two arguments suggest that we should match each individual with the average school quality over the entire spectrum of school types, from kindergarten to upper secondary education. We compute average school quality by using enrolment rates during the relevant periods as weights. Table 6 in the data appendix provides further details on the matching of data.

While we prefer to match each individual with the average pupil–teacher ratio for all levels of education, we also present in the Appendix (Table 8) the results of the estimates of educational attainment when the matching is based on the average quality of the school—by cohort and region of birth—experienced by the individual up to her maximum education al attainment.

Fig. 2 shows for the individuals born between 1941 and 1970 the dynamics of the average pupil–teacher ratio and of its regional dispersion by school type, where dispersion is measured by one standard deviation around the mean. The average pupil–teacher has declined sharply, from 26.23 for the cohort born between 1941 and 1945 to 15.35 for the cohort born between 1966 and 1970. The regional dispersion has also declined, and the coefficient of variation has fallen from 0.183 to 0.086. The reduction in the pupil–teacher ratio has involved all types of schools but has been sharpest for primary education.[10]

[10]Further declines in the pupil–teacher ratio in primary schools took place after 1985, when a sweeping reform was introduced with the main purpose of maintaining employment levels among teachers in an environment characterized by a steady decline in the number of pupils. Brunello, Checchi, and Comi, 2002, test alternative measures of educational resources, such as class size, school size, the share of private schools, teacher status (tenured versus untenured) and failure rates, with no significant result.

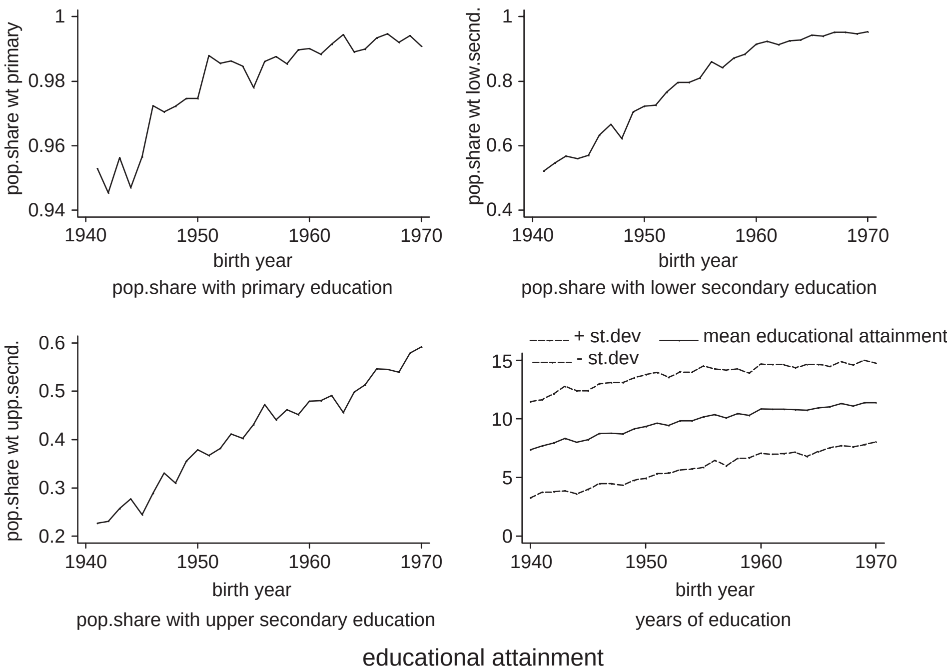

Fig. 3. Educational attainment in the population by year of birth and school type.

This drastic decline contrasts with the significant increase in the educational attainment of individuals born during the same period, shown in Fig. 3. Average educational attainment, measured as the number of years of attained education,[11] was equal to 8.23 years for the oldest cohort and to 11.26 years for the youngest cohort. The regional dispersion in educational attainment has also declined over time, and the coefficient of variation has fallen from 0.515 to 0.297.

We conclude the description of the data with our selected measure of family background W. As described above, this measure is the maximum number of years of education attained by parents within the household.[12] The average value of W in the sample has increased from the 5.29 years (standard deviation: 2.286) of the oldest age cohort to the 7.38 (2.304) of the youngest cohort. As in the case of the educational attainment of the offspring, the regional dispersion in the attainment of parents has declined over the years, from 0.432 to 0.312. The summary statistics of the main variables used in the paper are reported in Table 7 in the data appendix.

[11]We assign 5 years to primary school, 8 years for junior high school, 11 or 13 years for secondary school, depending on the type of school, 15 or 18 years for college, depending on the type of college.

[12]In our empirical estimates we also use separately the educational attainment of the mother and of the father.

4. The results

We start the description of our empirical results with educational attainment. We measure educational attainment as an ordered variable equal to 1 if the individual has attained at most primary education, to 2 if she has attained at most junior high school, to 3 if she has attained at most upper secondary education and to 4 for higher attainment (college and above), and regress this variable on individual family background, cohort, year, region of birth and region of residence dummies, a gender dummy, a dummy for individuals born from 1951 and the pupil–teacher ratio. The first two columns in Table 1 report the estimates based on an ordered probit specification and the remaining columns show the results based on ordinary least squares. The latter estimates are included because the marginal effects in the case of an ordered probit cannot be signed in an unambiguous way (see Greene (1990)).[13]

Each regression includes the interaction between family background W and school quality Q. In the first of each pair of columns we measure W as the maximum level of education attained by both parents, and in the second we report the estimates when W is measured separately for the father (W_{f}) and the mother (W_{m}). Table 8 in the appendix reports the estimates of the same

[13]The estimates are weighted with the population weights.

Table 1
Educational attainment. Ordered probit and OLS estimates. Dependent variable: educational attainment

Method of estimation	Oprobit (1)	Oprobit (2)	OLS (3)	OLS (4)
Pupil/teacher ratio Q	−0.0451**	−0.0440**	−0.0232**	−0.0220**
	(0.0063)	(0.0063)	(0.0041)	(0.0041)
Family educational background W	0.0240		0.0329**	
	(0.0139)		(0.0085)	
$Q \times W$	0.0066**		0.0035**	
	(0.0007)		(0.0004)	
Father years of education W_{f}		0.0380*		0.0334**
		(0.0154)		(0.0092)
$Q \times W_{\mathrm{f}}$		0.0033**		0.0018**
		(0.0008)		(0.0004)
Mother years of education W_{m}		−0.0233		−0.0030
		(0.0183)		(0.0111)
$Q \times W_{\mathrm{m}}$		0.0052**		0.0027**
		(0.0009)		(0.0005)
Dummies				
Gender	Yes	Yes	Yes	Yes
Age cohort	Yes	Yes	Yes	Yes
Region of birth (19)	Yes	Yes	Yes	Yes
Region of residence (19)	Yes	Yes	Yes	Yes
Educational reform in 1969 (born 1951)	Yes	Yes	Yes	Yes
Number of obs	31594	31594	31594	31594
R^2	0.163	0.171	0.341	0.353

Note: Robust cluster adjusted standard errors in parentheses with $p<0.05 = *$, $p<0.01 = **$.

model when school quality is conditional on the attained level of education by each individual.

Focusing on the ordered probit estimates, we find that educational attainment is higher when the pupil–teacher ratio is lower (so that school quality increases) and family background is better. We also find that the interaction between school quality and family background attracts a positive and significant coefficient, and interpret this as evidence that the impact of the pupil–teacher ratio on individual attainment varies with family background and is stronger—in absolute value—when parental education in the family is relatively low. These findings are confirmed when we estimate the same model by ordinary least squares. As shown in the second column, it is the educational attainment of the father that matters most.[14] The interaction with school quality, however, is stronger when the educational background of the mother is considered.

Table 2 replicates the ordered probit estimates separately by gender. We find that educational background W affects the educational attainment of females only via its interaction with school quality. Moreover, the estimated coefficient of school quality is lower for males than for females. Table 8 in the appendix reports the probit and OLS estimates when school quality is tailored to the highest attained degree of the individual. While the qualitative results are similar, the size of the estimated coefficients is higher than in Table 1.

The average maximum number of years of education attained by parents in the sample is 5.7 years, slightly above primary education. Suppose that this average number is increased to 8 years, corresponding to junior high school, equivalent to a 40.13 percent increase. The OLS results in column (3) of Table 1 suggest that, when the interaction effect is evaluated at the sample mean of the pupil–teacher ratio, this increase would raise the (expected) educational attainment of children by 10.40 percent, from an average value equal to 2.333–2.576. If we assign to the values 2 and 3 of the dependent variable 8 and 13 years of education respectively, this increase corresponds to moving from 9.665 to 10.880 years of education, more than one additional year of school, which implies strong intergenerational persistence in educational attainment.[15]

[14] The opposite occurs in Table 8, where the educational attainment of the mother has a statistically significant and stronger effect on attainment than the education of the father.

[15] Computed from the third column of Table 1 as $\partial E/\partial W\ W/E = (0.0329 + 0.0035^{*}20.887)5.709/2.333 = 0.259$ and $\bar{Q} = 20.887$.

Table 2
Educational attainment. Ordered probit by gender. Dependent variable: educational attainment

Method of estimation	Oprobit (1) men	Oprobit (2) men	Oprobit (3) women	Oprobit (4) women
Pupil/teacher ratio Q	−0.0266**	−0.0290**	−0.0652**	−0.0609**
	(0.0087)	(0.0087)	(0.0087)	(0.0083)
Family educational background W	0.0392**		0.0046	
	(0.0148)		(0.0183)	
$Q \times W$	0.0057**		0.0079**	
	(0.0008)		(0.0009)	
Father years of education W_{f}		0.0421*		0.0261
		(0.020)		(0.0295)
$Q \times W_{\mathrm{f}}$		0.0028**		0.0044**
		(0.0010)		(0.0015)
Mother years of education W_{m}		−0.0247		−0.0179
		(0.0242)		(0.0318)
$Q \times W_{\mathrm{m}}$		0.0055**		0.0047**
		(0.0012)		(0.0017)
Dummies				
Gender	Yes	Yes	Yes	Yes
Age cohort	Yes	Yes	Yes	Yes
Region of birth (19)	Yes	Yes	Yes	Yes
Region of residence (19)	Yes	Yes	Yes	Yes
Educational reform in 1969 (born 1951)	Yes	Yes	Yes	Yes
Number of obs	15476	15476	16118	16118
R^2	0.147	0.157	0.187	0.194

Note: Robust cluster adjusted standard errors in parentheses with $p<0.05$ = *, $p<0.01$ = **.

We also use the estimates in the third column of the Table 1 to compute the variations in educational attainment induced over the sample period by variations in school quality and family background. Given the important and persistent differences between the developed north and the less developed south, it is instructive to perform these computations not only for the full sample but also for the two sub-samples of northern/central and southern regions. We compute variations by comparing the oldest to the youngest cohort and by taking into account the interaction between school quality and family background. The data show that average educational attainment has increased by 41.45% in the full sample, by 42.49 in the northern/central regions and by 37.97 in the south.

The observed change in school quality Q between the oldest and the youngest cohort accounts for 4.7%,[16] 1.7% and 10.1% of the total variation of attainment in the full sample and in the two sub-samples, respectively. At the same time, the observed change in family background W explains 41.54%, 44.5% and 37.9% of the total variation. We conclude that family background has had a significantly higher impact on educational attainment than the pupil–teacher ratio, especially in the most developed areas of the country.[17] We also notice that about half of the total variation in educational attainment is not accounted by these two variables.

Next, we turn to the two–step model (1) and (2) and examine the impact of school quality and family background on the estimated returns to education. In order to have a sufficient number of observations in each cell identified by the region of birth, the region of residence and the cohort of birth for the estimation of the second stage implied by Eq. (2), we now organize our sample into 6 age cohorts, each comprising five years, the 19 regions of birth and aggregate the 19 regions of

[16]Computed from the third column of Table 1 as $\Delta E = (-0.0232 + .0035\,\bar{W}) \cdot \Delta Q = (-0.0232^{*}0.0035^{*}5.709) \cdot (-11.65) = 0.037$, which corresponds to 4.7% of the total variation in E (2.654−1.875 = 0.779).

[17]Our methodology does not allow us to assess whether this effect is the outcome of cultural resources associated to better educated parents, or the consequence of a reduction of liquidity constraints, which takes place more likely among poorly educated families. See the discussion in Cameron and Heckman (1998).

Table 3
Second stage estimate. Dependent variable: estimated returns to education

	(1)	(2)	(3)	(4)	(5)
Pupil/teacher ratio Q	−0.002	−0.002	0.008*	−0.002	0.006^
	(0.002)	(0.002)	(0.004)	(0.002)	(0.003)
Family educational background W		−0.008	−0.023**		
		(0.011)	(0.006)		
$W \times Q$		0.000			
		(0.001)			
Percentage of households with parents having at most primary education $P \times Q$			−0.012**		
			(0.003)		
Father years of education W_{f}				−0.007	−0.008
				(0.041)	(0.008)
W_{m}				−0.008	−0.018*
				(0.046)	(0.009)
$W_{\mathrm{f}} \times Q$				0.001	
				(0.002)	
$W_{\mathrm{m}} \times Q$				−0.001	
				(0.002)	
Percentage of households with fathers having at most primary education $P_{\mathrm{f}} \times Q$					−0.009*
					(0.004)
Percentage of households with mothers having at most primary education $P_{\mathrm{m}} \times Q$					0.000
					(0.004)
#obs	161	161	161	161	161
R^2	0.717	0.720	0.749	0.733	0.751

Note: Standard errors in parentheses with $p<0.10 =$ ^, $p<0.05 =$ *, $p<0.01 =$ **. Each regression includes cohort, region of birth, region of residence, cohort × birth, cohort × residence dummies.

residence into 3 macro areas (north, center and south).[18] In the estimate of Eq. (1) we use the following individual characteristics: gender, marital status, labor market experience and its square, dummies for the dimension of the town of residence, for part time and temporary jobs, and the number of months worked in the year. Since individual data belong to different years, we capture aggregate effects with time dummies. Finally, the endogenous selection of years of education is controlled for by adding to the regression the score computed from the ordered probit model of educational attainment. It turns out that the coefficient attracted by the score is positive but not statistically significant.[19] The estimated values of γ are retrieved from the first step estimate, together with the standard errors. In the second step we estimate (2) by weighted least squares, using as weights the standard errors (see Betts (1995)). Since the data in (2) are cell averages (by region of birth, region of residence and cohort), we only retain in the estimate the cells with more than 20 observations.[20]

The estimates of (2) are presented in Table 3. The first column in the table adds to the set of cohort, region of birth, region of residence and other dummies the pupil–teacher ratio Q. This specification is augmented in column 2 with family background W and its interaction with Q. In the next column we replace this interaction with the interaction of Q with the percentage P of households in the cell with at most primary education. The remaining three columns repeat the exercise by separating family background variables into the background of the mother and the father.

We find that the interaction between family background W and school quality Q is never statistically significant. However, if we replace in the interaction the variable W with P, which measures family background with the percentage of households having primary education as the highest attainment of parents, there is evidence that the relationship between the pupil–teacher ratio and the estimated returns to education varies with

[18]Since the region of current residence reflects local labor market conditions, one may wonder about the appropriateness of such macro regions. The large literature on the Italian Mezzogiorno, however, clearly suggests that the regional divide between the industrialized northern and central part of the country and the underdeveloped south accounts for the bulk of regional labor market disparities (see Brunello, Lupi and Ordine (2001) and the references therein).

[19]The estimated coefficient is equal to 0.0167 (standard error: 0.0166).

[20]By so doing we end up with 161 usable coefficients out of a theoretical number of 342 estimated coefficients (6 cohorts × 19 regions of birth × 3 macro regions of residence).

the average family background in the region and cohort of birth. These results suggest that the social environment where one was raised (whether she lived in a community with many or few educated parents/households) matters for the returns to education.

To illustrate, we compute the elasticity of returns with respect to the pupil–teacher ratio for different values of the distribution of P, using the estimates in the third column of Table 3. When P is low and equal to the 10 percentile value of its distribution ($P = 0.45$ for regions/cohorts with limited illiteracy), the estimated elasticity is 1.211,[21] which implies that a 1% reduction in the pupil–teacher ratio (i.e., an improvement in school quality) *reduces* the returns to education by close to 1 percent. This elasticity falls to 0.343 when P is equal to the 25 percentile value ($P = 0.60$) and turns negative (-0.267) at the median value of P ($P = 0.71$). The negative elasticity increases further in absolute value when P is at the 75 percentile value (the elasticity is -0.716 for $P = 0.79$) and at the 90 percentile (the elasticity is -1.193 for $P = 0.87$).

The average elasticity, evaluated at the sample averages of Q, P and γ , is negative and equal to -0.108. During the sample period spanning the 6 age cohorts, the pupil–teacher ratio has declined by 42.6%. Conditional on average family background, this corresponds to a 4.60% ($[-42.6 \times (-0.108)]$) increase in the returns to education, from 0.058 to 0.06.

Turning to the impact of family background W on school returns, we find that, conditional on school quality, the returns to education are higher in the regions and cohorts of birth with lower W. Again using the estimates in the third column of Table 3, the average elasticity of returns to changes in W is equal to -0.263.[22] With decreasing marginal returns to education, this negative elasticity can be explained with the fact that better parental education increases the educational attainment of children. Ceteris paribus, individuals born in regions and cohorts with better family background have higher education and lower returns to education. On average, family background has improved dramatically over time and the maximum number of years of education attained by parents has increased by 64.2% between the oldest and the youngest cohort. Using the estimated elasticity, we find that, following this increase and conditional on school quality, the returns to education should have declined across cohorts by 24.2% ($[+64.2 \times (-0.263)]$), from 0.058 to 0.048.

The estimates in the last two columns of the table decompose family background in the relevant cell into the educational attainment of the father and the mother and in the percentage of fathers and mothers with at most primary education. The last column in the table suggests that the negative impact of parental education on the returns to education of the offspring is mainly due to the mother's education. Conversely, the interaction between school quality and the percentage of parents with at most primary education is negative and statistically significant in the case of the father.

To summarize, the observed improvement in family background has had two contrasting effects. On the one hand, it has stimulated the increase in educational attainment of the younger generations. On the other hand, it has reduced the estimated returns to a year of education. Our results also point out that changes in family background have been quantitatively more important than changes in the pupil–teacher ratio for the evolution of educational attainment and of the returns to education.

Following Card (1999), we can interpret this evolution as the result of the interplay between the costs and the benefits of education. Assume that marginal benefits decline and that marginal costs increase with educational attainment. For any level of educational attainment, better school quality and improved family background are expected to improve marginal benefits and reduce marginal costs. The combination of the outward shift of marginal benefits and of the downward shift of marginal costs should have increased educational attainment, as we find, with uncertain results for the returns to education.

Based on our estimates, we find that the combined effect of changes in W and Q should have generated a 12.28% reduction in the returns to education from 0.058 for the oldest cohort to 0.051 for the youngest cohort. In practice, however, returns have declined only slightly, from 0.058 to 0.056. The failure of returns to significantly decline can be explained if other intervening factors have partially or totally compensated the estimated effects on returns of improved school quality and family background. There is a large recent empirical literature which shows that skill biased technical change is one of these factors, which has shifted relative demand in favor of better educated workers (see Katz and Autor (1999)), thereby sustaining returns.

Our results are consistent with the hypothesis that family background and school quality are technical substitutes in the production of human capital. As shown in the appendix, we find that a reduction in the pupil–teacher ratio by one unit reduces the contribution of a 1% increase in parental education to individual human capital by 0.032%. These findings support the view that an improvement of public school quality, decided and administered in Italy by the central government, can have a countervailing effect with respect to the differences induced by parental

[21] $\frac{\partial \ln \gamma}{\partial \ln Q} = (0.008 - 0.012P)\frac{\bar{Q}}{\bar{\gamma}}$, where P is evaluated at the percentile value and the other variables are at their sample means.

[22] $\frac{\partial \ln \gamma}{\partial \ln W} = \left(-0.023 - 0.012\bar{Q}\frac{\partial P}{\partial W}\right)\frac{\bar{W}}{\bar{\gamma}}$, where $\partial P/\partial W = -0.083$ and all other variables are evaluated at their sample means.

background on educational attainment and the returns to education.

5. Conclusions

In this paper we have investigated whether the combined significant reduction in the pupil–teacher ratio and increase in parental education observed in Italy between the end of the II World War and the end of the 1980s have had a significant impact on the educational attainment and the labor market returns of a representative sample of Italians born between 1941 and 1970. We have found that a lower pupil–teacher ratio is positively correlated with higher educational attainment, but that the overall improvement of parental education has had an even stronger impact on attainment. Our empirical evidence also suggests that the positive impact of better school quality on educational attainment and returns to education has been particularly significant for the individuals born in regions and cohorts with poorer family background. Parental education has had asymmetric effects, positive on attainment and negative on school returns. Better school quality has also had asymmetric effects on the returns to education, positive for individuals born in regions and cohorts with relatively poor average family background, and negative for individuals born in regions and cohorts where average parental education was relatively high.

We have also shown that a better school quality, measured by a lower pupil–teacher ratio, has been a technical substitute to parental education in the production of individual human capital. When school quality and family background are substitutes, an increase of public resources invested in education can be used to reduce the differences induced by parental education.

The relationship between parental education and the educational attainment of the younger generations suggest that a positive shock to attainment, triggered for instance by a school reform which extends compulsory education, can have multiplicative and long lasting effects in the medium and long run, as better educated generations will give birth over time to new generations, which will enjoy a higher parental education and therefore invest in even higher educational attainment. This self-sustained mechanism is empirically stronger in Italy than a reduction in the pupil–teacher ratio.

According to our results, the significant reduction in the pupil–teacher ratio has helped most the individuals born in the less developed regions of the country, endowed with poorer parental education, and has only partially compensated the decline in the returns to education associated to higher educational attainment. A question for future research is whether the benefits associated to the observed increase in school quality have been sufficient to compensate the costs borne by the Italian taxpayer.

Acknowledgements

Previous versions of this paper were presented at seminars in Bologna, Modena, Rome (Bank of Italy) and Salerno (AIEL). We are grateful to Simona Comi for efficient research assistance. The usual disclaimer applies.

Appendix

Consistently with our empirical specification and with the assumption that earnings equal productivity, assume the following production function of individual human capital H:

$$H = e^{\gamma E}, \tag{4}$$

where E is educational attainment and γ is the return to education. The second derivative of (log) H with respect to the pupil–teacher ratio Q and (log) family background W is

$$\frac{\partial^2 \ln H}{\partial Q \partial \ln W} = \left[\frac{\partial^2 \gamma}{\partial Q \partial W} E + \frac{\partial \gamma}{\partial Q}\frac{\partial E}{\partial W} + \frac{\partial \gamma}{\partial W}\frac{\partial E}{\partial Q} + \frac{\partial^2 E}{\partial Q \partial W}\gamma\right] W. \tag{5}$$

Based on our estimates, define

$$\gamma = 0.008Q - 0.0232W - 0.012QP$$
$$E = -0.0232Q + 0.0329W + 0.0035QW$$

and recall that $\frac{\partial P}{\partial W} = -0.083$.

Then the first and the last term on the right-hand side of (5) are positive and the remaining terms are negative. Evaluating the left-hand side of (5) at sample averages, we obtain

$$\frac{\partial^2 \ln H}{\partial Q \partial \ln W} = 0.032.$$

Since a higher pupil–teacher ratio Q implies lower school quality, our evidence points to technical substitutability between family background and school quality.

Data appendix

The data appendix is summarized in Tables 4–8.

Table 4
Population younger than 21 by region of birth (column) and region of residence (row)—bank of Italy surveys 1993–1995–1998–2000 (weighed)—percentages

Residence→ Birth↓	Piem	Aosta	Lomb	Tren	Vene	Friu	Ligu	Emil	Tosc	Umbr	Marc	Lazi	Abru	Moli	Camp	Pugl	Basi	Cala	Sici	Sard
Piemonte	**94.63**		1.29		0.15	0.08	0.30		0.08	0.15	0.38	0.15	0.38	0.23	0.38	0.45	0.08	0.53	0.53	0.23
Valled'aosta		**95.00**										5.00								
Lombardia	1.34		**90.33**	0.17	0.56	0.06	1.45	1.79	0.17	0.11	0.50	0.50	0.17	0.34	0.78	0.61	0.34	0.11	0.56	0.11
Trentino				**99.11**		0.22						0.67								
Veneto	0.21		0.10	2.06	**94.14**	1.75	0.10	0.51	0.21		0.21	0.62				0.10				
Friuli			0.67		0.45	**98.21**							0.22			0.45				
Liguria	2.79		0.82				**94.25**	0.16	0.66			0.16	0.16			0.16		0.66		0.16
Emiliaromagna	0.08		0.16	0.16	0.23		0.08	**97.97**	0.16	0.16	0.31		0.23		0.08	0.08	0.08	0.08	0.16	
Toscana			0.79		0.18		0.61	0.26	**96.40**	1.05		0.44			0.18		0.09			
Umbria							0.15		0.31	**98.78**	0.31	0.15	0.31							
Marche			0.36					0.61		0.73	**97.09**	0.12	0.36						0.73	
Lazio			0.63	0.08	0.08	0.08		0.16	0.16	0.16	0.32	**96.99**	0.48		0.40			0.16	0.16	0.16
Abruzzi			0.27		0.27				0.14		0.82	0.55	**97.26**	0.55			0.14			
Molise														**97.21**		2.79				
Campania	0.18		0.67	0.07	0.21			0.56	0.39	0.18	0.35	0.92	0.07	0.07	**95.62**	0.07	0.32	0.14	0.14	0.04
Puglia	0.73		0.73	0.21			0.05	0.21	0.05		0.42	0.37	0.21	0.16		**96.18**	0.63			0.05
Basilicata				0.88									0.44		0.44	0.88	**97.37**			
Calabria	0.44		1.42	0.11		0.44	0.22	0.44	0.66					0.11				**96.07**	0.11	
Sicilia	0.54		0.89		0.05	0.05	0.20	0.45	0.15		0.15					0.15		0.15	**97.08**	0.15
Sardegna	0.33						0.11	0.22								0.11		0.11	0.11	**99.01**

Table 5
Population born between 1940 and 1970 by region of birth (column) and region of residence (row)—Bank of Italy surveys 1993–1995–1998–2000 (weighed)—percentages

Residence→ Birth↓	Piem	Aosta	Lomb	Tren	Vene	Friu	Ligu	Emil	Tosc	Umbr	Marc	Lazi	Abru	Moli	Camp	Pugl	Basi	Cala	Sici	Sard
Piemonte	**88.44**	0.11	4.13	0.11	0.66	0.16	2.57	0.50	0.22	0.07	0.21	0.87		0.01		0.44	0.08	0.94	0.36	0.09
Valled'aosta	41.64	**47.81**	1.33				9.22													
Lombardia	1.10	0.03	**91.34**	0.21	1.19	0.24	1.03	2.44	0.61	0.08	0.15	0.29	0.13	0.01	0.52	0.18	0.01	0.25	0.16	0.02
Trentino	2.52		3.20	**84.76**	6.58		0.46	1.31	0.30		0.05	0.77								0.06
Veneto	2.99		4.63	0.63	**86.18**	1.57	0.39	1.53	0.22	0.04	0.08	1.15	0.07		0.01	0.29		0.01	0.16	0.05
Friuli	0.98		4.30	0.50	1.66	**89.83**	1.12	0.76	0.32	0.06	0.03		0.13		0.22				0.09	
Liguria	2.80		5.09	0.14	0.05	0.41	**88.23**	0.16	0.97		0.57	0.35	0.13		0.05	0.25		0.13	0.29	0.38
Emiliaromagna	0.81		3.03	0.22	0.13	0.05	0.73	**91.94**	0.81	0.26	0.44	0.83	0.07			0.15	0.03	0.07	0.25	0.18
Toscana	0.35		1.99		0.31	0.04	1.26	1.16	**92.20**	0.42	0.13	1.57			0.19				0.20	0.19
Umbria	3.19		1.33	0.07	0.17		0.20	0.67	2.81	**85.11**	0.63	5.15	0.67							
Marche	0.08		3.08	0.59	0.12	0.53	0.28	3.48	1.52	0.94	**82.19**	5.66	0.23	0.05		0.26			0.99	
Lazio	0.49		0.98	0.09	0.24	0.06	0.29	0.72	0.73	0.49	0.41	**93.89**	0.36	0.02	0.53	0.20		0.34	0.13	0.04
Abruzzi	1.23		1.73	0.28		0.03	1.80	1.14	0.60	0.35	1.90	4.59	**85.56**	0.14		0.10	0.46	0.10		
Molise	2.74		0.80	1.08				3.13	0.72		0.76	18.12	3.57	**66.63**	0.97	1.24		0.17		0.09
Campania	2.96	0.01	4.51	0.31	0.45	0.43	0.27	2.55	1.95	0.16	0.29	4.72	0.21	0.29	**78.81**	0.64	0.38	0.47	0.47	0.12
Puglia	4.95	0.02	6.16	0.18	0.84	0.34	0.26	1.44	0.48	0.05	0.49	2.52	0.34	0.21	0.20	**80.42**	0.27	0.11	0.62	0.11
Basilicata	7.41		8.85	0.14			1.07	1.05	2.45	0.23	0.09	2.95	0.15		1.72	2.25	**71.41**		0.23	
Calabria	11.31	0.21	10.67	0.16	0.86	0.14	2.49	2.07	1.34	0.16	0.20	4.62	0.01	0.05	0.61	0.43		**64.31**	0.25	0.11
Sicilia	4.61	0.08	6.26	0.12	0.46	0.31	1.14	0.77	1.02	0.10	0.19	2.01	0.01	0.05	0.27	0.23		0.09	**82.18**	0.13
Sardegna	6.25	0.04	4.14	0.45	0.83	0.04	0.90	1.11	0.97	0.16	0.20	5.31	0.04		0.18	0.19			0.40	**78.78**

Table 6
Pupil–teacher ratios by birth cohort

Cohort	Year of birth	Relevant years for school quality: kindergarten	Relevant years for school quality: primary school	Relevant years for school quality: junior high school	Relevant years for school quality: upper secondary school
1	1941–45	1946–48–50	1948–50–52–54–56	1952–54–56–58	1956–58–60–62–64
2	1946–50	1950–52–54–56	1952–54–56–58–60	1958–60–62–64	1960–62–64–66–68
3	1951–55	1954–56–58–60	1958–60–62–64–66	1962–64–66–68	1966–68–70–72–74
4	1956–60	1960–62–64–66	1962–64–66–68–70	1968–70–72–74	1970–72–74–76–78
5	1961–65	1964–66–68–70	1968–70–72–74–76	1972–74–76–78	1976–78–80–82–84
6	1966–70	1970–72	1972–74–76–78–80	1978–80–82–84	1980–82–84–86–88

Note: Weights are enrolment rates in each type of school during the period indicated in the table.

Table 7
Descriptive statistics for the relevant variables

Variable	Mean	Standard deviation
Sample used for the ordered probit model (31594 observations)		
Female	50.7	
Cohort:		
Born 1941–45	15.5	
Born 1946–50	17.8	
Born 1951–55	16.2	
Born 1956–60	16.2	
Born 1961–65	16.8	
Born 1966–70	17.4	
Residence:		
North-west	22.3	
North-east	19.5	
Center	21.1	
South	11.4	
Islands	25.7	
Age	40.64	9.10
Years of education	10.06	4.17
Years of education of father	5.53	4.17
Years of education of mother	4.79	3.70
Student/teacher ratio—average	20.44	4.74
Sample used for the two-step estimate of the returns to education (16.471 observations)		
Female	49.0	
Cohort:		
Born 1941–45	10.7	
Born 1946–50	18.2	
Born 1951–55	18.5	
Born 1956–60	18.3	
Born 1961–65	17.8	
Born 1966–70	16.5	
Residence:		
North-west	25.3	
North-east	21.8	
Center	21.8	
South	10.6	
Islands	20.5	
Age	40.05	8.33
Years of education	10.98	4.01
Years of education of father	5.86	4.16
Years of education of mother	5.04	3.62
Student/teacher ratio—average	20.05	4.44
Gross annual wage:		
1993 survey (thousands lire)	26893.5	14910.6
1995 survey (thousands lire)	28644.5	16313.8
1998 survey (thousands lire)	31621.0	19296.7
2000 survey (thousands lire)	34920.5	22586.5
Months employed per year	11.39	2.01
Years of potential experience	23.07	9.58

Table 8
Educational attainment. Ordered probit model. Dependent variable: educational attainment

Method of estimation	Oprobit (1)	Oprobit (2)	OLS (3)	OLS (4)
Pupil/teacher ratio Q	−0.4619**	−0.4581**	−0.1834**	−0.1810**
	(0.0436)	(0.0437)	(0.0115)	(0.0114)
Family educational background W	0.0791**		0.1031**	
	(0.0140)		(0.0083)	
$Q \times W$	0.0034**		−0.0016**	
	(0.0008)		(0.0005)	
Father years of education W_{p}		0.0203		0.0533**
		(0.0219)		(0.0099)
$Q \times W_{\mathrm{p}}$		0.0043**		−0.0001
		(0.0012)		(0.0006)
Mother years of education W_{m}		0.0939**		0.0760**
		(0.0281)		(0.0126)
$Q \times W_{\mathrm{m}}$		−0.0015		−0.0023**
		(0.0016)		(0.0007)
Dummies				
Gender	Yes	Yes	Yes	Yes
Age cohort	Yes	Yes	Yes	Yes
Region of birth (19)	Yes	Yes	Yes	Yes
Region of residence (19)	Yes	Yes	Yes	Yes
Educational reform in 1969 (born 1951)	Yes	Yes	Yes	Yes
Number of obs	30482	30482	30482	30482
R^2	0.353	0.357	0.573	0.578

Note: Robust cluster adjusted standard errors in parentheses with $p<0.05 = *$, $p<0.01 = **$.
School quality measured with reference to the highest educational attainment of the individual.

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