The Evolution of Primary and Secondary effects in Italy in the '90s

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1. Introduction

It is often held that educational expansion narrows social inequalities within nations by promoting a meritocratic basis for status attainment, yet substantial research indicates that the relative advantages of elite children over children with less privileged background have changed little in the last decades (Shavit and Blossfeld, 1993; Breen and Jonsson, 2000; Hannum and Buchmann, 2003); on average higher status children perform better in school and attain higher educational levels. In this light, equality of opportunity (EOP) in education is still a highly relevant issue in the international educational policy agenda.

Class differentials in educational attainment are related in the sociological literature to *primary* and *secondary* effects (Boudon, 1974). The former refer to the influence of social origin on ability early in children's educational careers: high status parents are more likely to sustain and motivate the school work and provide a stimulating environment to their offspring. The latter operate through the choices that families make within the educational system *given* the level of ability. The rational action approach (Goldthorpe, 1996; Breen and Goldthorpe, 1997), assuming that families wish to avoid intergenerational downward mobility, provides a theoretical explanation for the evidence that, at given levels of ability, school choices vary across social background. Ability is intended as an observed measure of school performance (typically grade point average) as opposed to unobserved measures of cognitive abilities, since it is held that it is the former that affects the decision process through the perceived probability of schooling success.

The evaluation of primary and secondary effects is particularly relevant at the end of compulsory schooling (lower secondary in Italy), where in many countries students face the decision whether to enrol into the academic track[1] (which gives access to tertiary education), to enrol into a vocational track, or to enter the labour market. EOP is obviously affected by institutional features. Interventions aimed at containing primary effects enhance the performance of children of less advantaged background, especially at the primary school level. Secondary effects can be reduced by endorsing the enrolment of lower status children into the academic track or, possibly, by regulating access through ability assessments.

The evaluation of the relative importance of primary and secondary effects is the aim

[1] The term *tracks* is often used in the literature to indicate the different secondary school educational paths available to students in a certain educational system. The *academic* track is the one conceived to prepare for university studies (even if in some countries it is not strictly necessary to access tertiary education).

of a growing body of literature (Erikson *et al,* 2005; Jackson *et al.,* 2007; Stocké, 2007; Kloosterman *et al.,* 2007). This research – based on surveys carried out at a national level – provides empirical evidence of the relevance of secondary effects in the creation of class differential in educational attainment. The methodology, briefly sketched in Section 2, combines the estimates of the distribution of school performance and of the probability of choosing a specific track given school performance, at each level of social background. A counterfactual argument is carried out: the probability of entering the academic track that individuals would face if they had the ability distribution of class *j*, but the transition probability given ability of class *k*, is evaluated. Observed and counterfactual odds-ratio are compared, and a decomposition of log(OR) based on counterfactuals provides an estimate of the relative importance of primary and secondary effects.

Aim of this paper is to provide an assessment of primary and secondary effects in secondary school choices in Italy. Other countries studies (UK, Sweden, Germany, Netherlands) are based on panel surveys recording data on children's schooling careers, but prospective longitudinal data is not available for Italy. For this reason the analysis is based on the data of the cross-sectional repeated survey *Percorsi di studio e di lavoro dei diplomati* (ISTAT, 2007), collecting detailed information of individual educational histories up to three years after the secondary school degree. A major issue to deal with is self-selection (see Section 4), as only secondary school graduates are interviewed[2]. By integrating the survey data with administrative and census information we derive estimates of the relevant distributions, correcting for selection-bias.

As lower secondary school final marks are assigned on a 4-level scale (*satisfactory, good, very good, excellent*[3]), a semi-parametric version of the standard approach is adopted. Results are described in Section 5. The main conclusion is that secondary effects are more important in shaping social origin differentials in secondary schools decisions than primary effects, and that this sharp prevalence has remained substantially unchanged over time. By comparing our estimates with those reported in the recent literature for UK, Sweden, Germany and the Netherlands, it turns out that the relative contribution of primary effects is substantially weaker in Italy than in the other countries.

2. The methodology

Let A be a continuous measure of students' school performance before track choice and S a discrete variable representing social status. Then $f(A|S)$ is the distribution of the performance scores for each social group; assuming a normal distribution, the relevant parameters can be estimated by group sample mean and variance.

Define Y as a binary variable taking value 1 if the academic track is chosen and 0 otherwise (i.e. if the student chooses a different track or if he does not continue to secondary education). Note that Y refers to the first choice after the end of compulsory schooling and not to possible subsequent changes. The transition probability *given*

[2] Employing data from PISA (*Programme for International Student Assessment*; OECD, 2005) in this context would greatly weaken the sample selection problem, since students are interviewed at 15, i.e. at the beginning of secondary school. However this option proves impossible since PISA does not include information on students' performance before track choice. PISA may however be appropriate to evaluate the *total effect* of social background on track choice (see for example Contini et al., 2007).

[3] As translated from *Sufficiente*, *Buono*, *Distinto*, *Ottimo* in Italian.

performance $P(Y=1|A,S)$ can be estimated with binary logistic regression for each class separately. The integrals:

$$P_{jj} = \int_{-\infty}^{+\infty} f(A|S=j)\, P(Y=1|A, S=j)\, dA \tag{1}$$

evaluated for each S by numerical integration, represent the predicted probability $P(Y=1|S=j)$ which can be compared with its observed counterpart, the percentage of those belonging to social class j enrolling into academic schools.

On the other hand, the integral:

$$P_{jk} = \int_{-\infty}^{+\infty} f(A|S=j)\, P(Y=1|A, S=k)\, dA \tag{2}$$

is a "counterfactual" probability. Expression (2) is the probability that an individual would experience if he had the performance distribution of social class j and the transition probability of class k. With K social classes, there are $K(K\text{-}1)$ counterfactual probabilities.

The total effect of class j over class k on the propensity to continue to the academic track is represented by the odds ratio:

$$Q_{jj.kk} = \frac{P_{jj}/(1-P_{jj})}{P_{kk}/(1-P_{kk})} \tag{3}$$

Define also:

$$Q_{jj.kj} = \frac{P_{jj}/(1-P_{jj})}{P_{kj}/(1-P_{kj})}$$

The numerator represents the odds of continuing to academic education for an individual exposed to the performance distribution and the transition probability of class j, while the denominator represents the odds for an individual with performance distribution of class k and transition probability of class j. Since the difference here lies only in the performance distributions, this quantity is informative on primary effects. Similarly:

$$Q_{kj.kk} = \frac{P_{kj}/(1-P_{kj})}{P_{kk}/(1-P_{kk})}$$

provides information on secondary effects, as what varies here is the transition probability while the performance distribution remains fixed.

Total effect (3) can be factorized in two distinct ways:

$$Q_{jj.kk} = Q_{jj.kj}\, Q_{kj.kk}$$

$$Q_{jj.kk} = Q_{jk.kk}\, Q_{jj.jk}$$

By taking the logarithms, we obtain:

$$L_{jj.kk} = L_{jj.kj} + L_{kj.kk} \tag{4a}$$

$$L_{jj.kk} = L_{jk.kk} + L_{jj.jk} \tag{4b}$$

where in each case the first term on the right hand side refers to situations with different performance distribution but the same transition probability, and the second term to situations with the same performance distributions and different transition probability. The relative importance of secondary effects can be evaluated by $L_{kj.kk} / L_{jj.kk}$ or $L_{jj.jk} / L_{jj.kk}$. Estimates based on (4a) and (4b) generally differ, although in practice not to a great extent (see Erikson *et al.* (2005) for details).

Assuming that there are only two social levels to ease the understanding: *H* (high) and *L* (low), we obtain the following expressions:

$$L_{HH.LL} = L_{HH.LH} + L_{LH.LL}$$

where the log total effect is given by the primary effect evaluated with the transition probability of the *high* class and the secondary effect with the performance distribution of the *low* class;

$$L_{HH.LL} = L_{HL.LL} + L_{HH.HL}$$

where the first term is the primary effect evaluated with the transition probability of the *low* class and the second term is the secondary effect with the performance distribution of the *high* class.

It is worthwhile to note that under the linear probability model: $P(Y=1|A,S)=\mu+\lambda S+\theta A$, with $A=\alpha+\beta S+\varepsilon$, the following would hold:

$$P(Y=1|S=j+1)-P(Y=1|S=j)=\beta\theta+\lambda$$

In this case primary effects are represented by $\beta\theta$ and secondary effects by λ. Instead, it can be shown that under the logistic model (even in the absence of interaction effects between *A* ad *S*):

$$\ln\frac{P(Y=1|A,S)}{1-P(Y=1|A,S)}=\mu+\lambda S+\theta A$$

primary and secondary effects – as measured by means of (4a) and (4b) – are functions of the parameters of both the model for *A* and the model for *S*, although the component related to primary effects is much more sensitive to β and θ, and the component related to secondary effects is much more sensitive to λ.

3. The analysis for Italy

3.1 Institutional features

Although compulsory education starts at age 6 and ends at age 16, the last two years are a very recent formal requirement; for the cohorts considered in this work (students born from 1976 through 1985) the end was still set at 14. There are five years of primary school and three years of comprehensive lower secondary education, after which students choose their upper secondary school among many different programmes.

As in most European countries, in spite of the wide range of different secondary school types, a broad distinction between an academic and a technical/professional track can be made. The academic track lasts five years and includes different types of *lyceum*: *classical*, *scientific*, *linguistic, artistic*[4]. The technical and vocational tracks (lasting respectively five and three years) lead directly to a professional qualification.

There are no special admission requirements, such as ability tests or marks, to enter the different tracks. After five years of schooling (with two integrative years for vocational schools), all tracks give access to university (Eurydice, 2006). In practice, only few students from the vocational track enter tertiary education: in the ISTAT sample little more than 20% did, while the proportion for lyceums was higher than 90% and about 50% for technical schools.

3.2 The data

Differently from other countries, no adequate panel survey recording schooling careers is available for Italy. Given this limitation, we employ cross-sectional data from the survey *Percorsi di studio e di lavoro dei diplomati* carried out by ISTAT on higher secondary school *graduates*, recording the relevant longitudinal information retrospectively. The survey takes place every three years since 1998, and graduates are interviewed three years after the degree attainment, with the aim to investigate the transition from secondary school to tertiary education or work. As we will point out in Section 4, the nature of the survey implies the existence of significant sample selection, which will have to be dealt with.

All four available surveys have been employed in this paper, starting with 1995 graduates (i.e. students choosing school track in 1990) with the last survey focusing on 2004 graduates (ISTAT, 2007), who chose their track in 1999. The survey has not changed much over time, the data being collected with a two stage sampling scheme and a sample size of about 20000 units: the most recent one involved 1868 schools and 20408 individuals[5]. Essential to our analysis is the recording of the final student's mark at the end of lower secondary school, the subsequent track choice, and a set of variables describing parental occupational and educational status[6].

3.3 Final marks in lower secondary school

According to the rational choice theory (Breen and Goldthorpe, 1997), families make their educational choices with the aim to avoid downward mobility, according to future employment prospects and the probability of schooling success of their children relative

[4] At the time of track choice for the ISTAT samples, the *Istituto magistrale* prepared for the primary school teaching career. Although this type of school has been later redefined as *socio-pedagogic lyceum,* and further university education is now required to enter the teaching profession, until a few years ago this school gave direct access to it. For this reason we will not consider this school type as belonging to the academic track.

[5] Interviews were carried out with CATI.

[6] Although we do not employ this classification in the present paper, the data allow to define individuals with respect to three social classes as in the simplified British *National Statistics Socio-Economic Classification,* used for example in Jackson *et al.* (2007).

to each option. This assessment is made by taking into account children's ability, conceived as an observed measure of school performance as opposed to unobserved measures of cognitive abilities.

In Italy, the final lower secondary school mark[7] is the main observed information on children's ability before track choice. We highlight three possible sources of measurement error:

(i) Final lower secondary grades encompass in Italy only four distinct proficiency levels (*excellent, very good, good, satisfactory*). This highly discrete grading system appears to be quite a rough measurement of students' ability when compared with other countries marks, based on finer scales (e.g. ten levels in the British case).

(ii) Exams are set up by the school teachers, and are not based on standardised national tests[8]. An indirect evidence of the existence of a bias is that, although international assessments such as PISA (OECD, 2005) show a significantly lower average level in Southern Italy with respect to the North, in the South the percentage of *ottimo* is higher than in the rest of the country.

(iii) Related to point (ii), if marks were given with some reference to the average ability within the school, higher performing schools could evaluate their students somewhat more severely. The issue is particularly relevant in highly socially segregated environments, since on average high status children perform better.

The problem of measurement error is not explicitly addressed here. The reasons are twofold. First, we think that the main source of bias in the Italian case is likely to be given by sample selection, due to employing data on secondary school graduates.

Perhaps more importantly, the second reason has to do with the rationale of the analysis. If it is true that people make their educational choices on the basis of observed school performance[9], the "correct" measure of ability for secondary effects is given by marks, even if they are affected by measurement error. On the other hand, the "correct" measure for identifying class differentials in the performance distribution should be latent ability.

Nevertheless, it is important to note that the decomposition method described above involves investigating the role of manifest ability in *shaping school choices.* In fact*,* social class transition probabilities - see formula (1) - are a weighted average of the class transition probabilities given ability (marks), where weights are given by the relative proportion of individuals with each level of ability (again, marks) within the class. In this light it is not relevant whether the school mark is a measurement error version of true ability (measurement error is instead very relevant for the assessment of inequality of opportunity in the true ability distribution across social classes). Thus, when we come to interpret primary effects in this context, we should acknowledge that what is here called "primary effects" has to do with the distribution of latent ability *and* the way this ability is actually translated in marks[10].

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7 The mark is attributed after a national exam, detached from normal school activity, held at the and of lower secondary school (*Esami di Stato conclusivi del I ciclo*).

8 This issue is likely to become less relevant in the future (from2007 onward), since final exams will include two standardized tests with common evaluation guidelines.

9 Stocké (2007) addresses this issue for Germany and finds out that educational choices are driven mainly by school marks, although a minor additional effect can be ascribed to parents' perception on their children ability.

10 It is nevertheless obvious that in the extreme case where marks are hardly related to ability, the

Yet, this caveats would not hold if people were aware of their true level of ability and shaped their decisions accordingly: transitions rates would have to be estimated given true ability, and weights defined consequently. It can be shown that if marks were employed, the relative contribution of primary effects would be underestimated[11].

4. Sample selection

As we have pointed out, no adequate panel survey recording school careers is available for Italy, and for this reason we employ the ISTAT cross-sectional survey on secondary school graduates in 1995-2004, recording the relevant information retrospectively. Since the survey target population does not include those who have enrolled into a secondary school and exited the educational system before graduation[12], the sample is affected by selection bias.

We now deal with the consequences of sample selection on the estimates of the relevant distributions for primary and secondary effects. We will show that without corrections, we would *underestimate* both the differences in the ability distribution across social background levels, and the effect of social background on school choices. Note that traditional methods for dealing with sample selection (e.g. Heckman's method) cannot be employed in this context because micro-data on dropouts is not available.

Primary effects

As before let A be the school performance before track choice and S a measure of families social status. Define G as a binary variable taking value 1 if the individual has attained a secondary school degree and 0 if he has dropped-out of the educational system. The distribution of interest is $P(A|S)$, while the observable distribution is $P(A|S, G=1)$. The two distributions are related by:

$$P(A|S,G=1)=P(A|S)\frac{P(G=1|A,S)}{P(G=1|S)}=P(A|S)\frac{P(G=1|A,S)}{\int_A P(G=1|S,A)P(A|S)dA}$$

The observable distribution and the distribution of interest coincide if the second factor in the right hand side is equal to 1, i.e. if performance A does not affect the graduation probability given social status. Since this is obviously very unlikely, the survey estimate of the performance distribution given social status is biased.

Let us recall that in the Italian system ability is measured on a 4-level ordinal scale, which we will code as: *satisfactory* (1), *good* (2), *very good* (3), *excellent* (4). We will make the assumption that school drop-outs come exclusively from the population of low performers (see next section for empirical evidence on this):

decomposition itself looses much if its meaning, in that secondary effects would become the only source of class differentials.

[11] With the aim to investigate this issue we have developed a simulation study (not presented here). The bias appear to be little for measurement error of type (i) and (ii), somewhat bigger but not dramatic for type (iii).

[12] Children who have chosen a vocational program and attained a *qualifica professionale* (after three years) but not a *diploma* (after five years) are also excluded from the survey. To simplify the exposition, we will refer to the term "dropouts" to indicate this people too.

$$P(G=0|S,A=j)=\begin{cases} f(S)>0 & if\ \ j=1 \\ 0 & if\ \ j=2,3,4 \end{cases} \tag{5}$$

For j=2,..4 this implies that:

$$P(A=j\,|\,S,G=0)=\frac{P(G=0\,|\,A=j,S)P(A=j\,|\,S)}{P(G=0\,|\,S)}=0 \tag{6}$$

Since:

$$P(A\,|\,S)=P(A\,|\,S,G=1)P(G=1\,|\,S)+P(A\,|\,S,G=0)P(G=0\,|\,S) \tag{7}$$

by combining (5) and (6) we obtain:

$$P(A=j|S)=\begin{cases} P(A=j|S,G=1)P(G=1|S) & if\ \ j=2,3,4 \\ 1-\sum_{j=2}^{4}P(A=j|S) & if\ \ j=1 \end{cases} \tag{8}$$

In order to estimate $P(A|S)$, we employ the ISTAT graduates' survey to assess $P(A|S, G=1)$, but we also need to estimate the graduation probability given social status $P(G=1|\mathrm{S})$. Since:

$$P(G=1\,|\,S)=\frac{P(S\,|\,G=1)P(G=1)}{P(S)}$$

we will estimate $P(S|G=1)$ from the graduates survey and exploit the official statistics derived from administrative data sources for the overall graduation probability $P(G)$ and the social status distribution $P(S)$ (see Section 5.2).

Secondary effects

Let Y represent again secondary school choice: Y=1 for the academic track and 0 otherwise. We are interested in $P(Y=1|A,S)$, but we can only estimate $P(Y=1|A,S,G=1)$. Since:

$$P(Y=1\,|\,G=1,A,S)=P(Y=1\,|\,A,S)\frac{P(G=1\,|\,Y=1,A,S)}{P(G=1\,|\,A,S)} \tag{9}$$

the survey estimate is unbiased if, given ability and social status, the graduation probability does not depend on the chosen track. Note that Y refers to the *first* choice undertaken at the end of compulsory school, while graduation can be achieved in *any* track. Students may change track if they fail or if they are not satisfied with their initial choice, and then graduate. In this light, the likelihood to attain a secondary school degree will not depend on how difficult or selective the specific track is. The enrolment into the academic track will be considered instead as a signal of higher *aspirations*.

The consequences of employing directly the graduates' survey to estimate $P(Y=1\,|\,A,S)$ can be easily grasped by assuming the simple linear probability models:

$$P(G=1|A,S,Y)=\alpha+\beta A+\gamma S+\delta Y$$

$$P(Y=1|A,S)=\lambda+\xi A+\theta S$$

We obtain:

$$P(G=1\,|\,A,S) = P(G=1\,|\,A,S,Y=1)P(Y=1\,|\,A,S) + P(G=1\,|\,A,S,Y=0)P(Y=0\,|\,A,S)$$
$$= (\alpha+\beta A+\gamma S+\delta)P(Y=1|A,S)+(\alpha+\beta A+\gamma S)P(Y=0|A,S)$$
$$= \alpha+\beta A+\gamma S+\delta P(Y=1|A,S)$$

The second factor in the right hand side of (8) is:

$$\frac{P(G=1|Y=1,A,S)}{P(G=1|A,S)} = \frac{(\alpha+\beta A+\gamma S)+\delta}{(\alpha+\beta A+\gamma S)+\delta(\lambda+\xi A+\theta S)}$$

Representing a probability, $(\lambda+\xi A+\theta S)<1$ and is an increasing function of *A* and *S*. Thus, the above ratio is always greater than 1 and a decreasing function of *A* and *S*. Unless $\delta=0$, the difference in the probability to choose the academic track across social backgrounds is going to be *underestimated* with the graduates survey data, leading to underestimation of secondary effects. However, as we will show in the next section by employing a different data source, empirical evidence suggests that δ should be nearly 0, meaning that aspirations are entirely captured by school performance and social status. No corrections are needed in this case, implying that $P(Y=1|A,S)$ can be estimated directly from the survey data.

4.1 Supporting the assumptions

We now turn to data analyses carried out in order to provide empirical support to the assumptions described in the previous section.

Primary effects

Let us recall the relevant assumption, described by equation (4), stating that, for each social background, only low performers eventually drop-out from school. The marginal distribution of performance can be written as:

$$P(A=j) = P(A=j\,|\,G=1)P(G=1) + P(A=j\,|\,G=0)P(G=0)$$

from which we obtain the performance distribution for school-drop-outs:

$$P(A=j\,|\,G=0) = \frac{P(A=j) - P(A=j\,|\,G=1)P(G=1)}{P(G=0)} \tag{10}$$

This distribution can be roughly estimated by employing the graduates survey data – providing information on $P(A\,|\,G=1)$ – and aggregate administrative data from ISTAT – which records the overall distribution of lower secondary final examination marks $P(A)$ for the year 1996, as well as an estimate of the overall national percentage of school dropouts $P(G=0)$ for the same year. We obtain from (10):

$\hat{P}(A=1\,|\,G=0) = 0{,}96$ $\hat{P}(A=2\,|\,G=0) = 0{,}05$

$\hat{P}(A=3\,|\,G=0) = 0{,}005$ $\hat{P}(A=4\,|\,G=0) = -0{,}02$ [13].

[13] Small inconsistencies among the combined data sources produce a negative probability, which is

strongly supporting the assumption.

Secondary effects

We now evaluate the assumption:

$$\frac{P(G=1\,|\,Y=1,A,S)}{P(G=1\,|\,A,S)}=1 \tag{11}$$

i.e., that the effect of aspirations is entirely captured by that of school performance and social status. As we have pointed out before, no longitudinal micro-data on schooling careers is available for the estimation of the conditional distribution of *G*.

The survey carried out jointly by CISEM and IARD[14] in 2006 on 3600 upper secondary school students in the area of Milan is employed for this purpose. The sample includes students in each of the five grades of the upper secondary schools; information on school careers as well as family characteristics, including parental educational and occupational status are recorded. The survey is cross-sectional and does not include dropouts; nevertheless, by comparing 1° grade students (including all future dropouts) with 5° grade students (assuming that nobody exits the school system thereafter), we can roughly assess the dropouts profile.

By a simple application of Bayes' theorem[15]:

$$\frac{P(G=1|Y=1,A,S)}{P(G=1|A,S)}=\frac{P(Y=1|G=1,A,S)}{P(Y=1|A,S)}$$

The right hand-side can be estimated by the ratio of the proportion of the academic track students in 5° grade and that for 1° graders. Considering *S* as the highest parental education and modelling both $P(Y=1|G=1,A,S)$ and $P(Y=1|A,S)$ with binary logit regressions, these ratios are all close to 1[16], supporting the validity of (11).

5. The empirical analysis

5.1 Semi-parametric approach

In Section 2 school performance *A* was taken as a continuous variable – as in most countries marks follow a fine scale and in some cases the grade point average is employed – which can be approximated quite well with a normal distribution; although not strictly necessary, this is also useful for the numerical evaluation of integral (1).

Because of the highly discrete scale (see Section 3.3), the normal distribution is clearly not appropriate for Italy. In this context, let *A* be the discrete variable taking values 1 to 4, corresponding to the four proficiency levels from lowest to highest.

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however so close to 0 to be reasonably considered negligible.

[14] CISEM stands for *Centro per l'Innovazione e Sperimentazione Educativa Milano* and is a research centre on educational problems of *Provincia di Milano*. IARD - Istituto Franco Brambilla is a research centre focusing on life problems and opportunities of young people. The authors would like to thank both CISEM and IARD for the collaboration and availability of data.

[15] Since $P(G=1|Y=1,A,S)=\frac{P(G=1,Y=1|A,S)}{P(Y=1|A,S)}=\frac{P(Y=1|G=1,A,S)P(G=1|A,S)}{P(Y=1|A,S)}$.

[16] This ratios vary from 0.93 for high status-high ability students to 1.32 for low-status-low ability students.

Expression (1) becomes:

$$P_{jj} = \sum_{A=1}^{4} P(A|S=j) P(Y=1|A, S=j) \tag{12}$$

and counterfactual probability (2a):

$$P_{jk} = \sum_{A=1}^{4} P(A|S=j) P(Y=1|A, S=k) \tag{13}$$

The performance distribution $P(A|S)$ is estimated non-parametrically, given gender and geographical area (*North West*, *North East*, *Center*, *South and Isles*). The transition probability $P(Y=1|A, S)$ could be estimated non-parametrically as well, or with binary logit models as in the original approach, to privilege parsimony and keep results simple; the approaches lead to very similar final results.

Note that although in the relevant literature social class - derived from parental occupation (Erikson and Goldthorpe, 1992) - is generally considered, for the moment we operationalize S with reference to the highest parental educational attainment. The main reason is that, having to correct for sample selection by employing national aggregate statistics on official reports, there seems to be stronger coherence between the two data sources[17]. In what follows, the terms "family status" or "social status" will always refer to this concept.

5.2 Sample selection correction factors

As we have seen in Section 4, in order to correct for sample selection, for the evaluation of $P(A|S)$ we need to estimate $P(G=1|S) = P(S|G=1) P(G=1) / P(S)$. The three factors have been obtained separately by gender[18] as follows:

- $P(S|G=1)$ has been estimated directly from the graduates survey data;
- $P(G=1)$ is the marginal graduation probability; it has been computed as the ratio of the number of graduates in the relevant years (data directly obtained from the Education Ministry Statistical Office) to the number of students who passed the lower secondary final exam five years before (data from ISTAT *Annuario di Statistiche Demografiche*, official publications).
- $P(S)$ is the national distribution of the highest parental educational level for the 1976, 1979, 1982 and 1986 birth cohorts (i.e. the 19 years old at the time of graduation), derived from ISTAT *Censimento Nazionale della popolazione* (the 1991 Census was used for the first two cohortsyears,moving to the 2001 Census for 1982 and 1986)[19].

[17] Note also that the odds ratio between Y and S when status is measured by social class is much lower than that relative to the highest parental educational level (around 5.6 vs. 11). Moreover, some recent works seem to be going in the same direction (see e.g. Kloosterman *et al.,* 2007).

[18] They were also computed separately for a set of four macro-regions (North-West, North-East, Centre, South and Isles). However this could introduce further distortion issues due to internal migration waves, and in any case only limited differences emerged. Thus the national values were employed in the end.

[19] In principle we should derive the highest parental educational level distributions for the children who have obtained the lower secondary school degree in a given year. This information is not available. As a proxy we calculate $P(S)$ for the all children born 14 years before according to Census data. The populations do not overlap for two reasons: first, some students may graduate earlier or later, due to

The first estimates of $P(G=1|S)$ were not always strictly smaller than one[20]. This could be due to the fact that we employ different data sources, which are likely to be affected by non-sampling error in various ways[21]. Another potential small source of bias is that the data employed for the estimation of $P(S)$ refers to parental educational level at a time that can differ from that consistent with the definition of status in the survey (i.e. when children are 14). For example, for the cohort born in 1982 parents educational status should be referred to 1996, while it is here estimated for 2001.

Since these first estimates varied little among geographical areas and inconsistencies were found to be weaker on aggregate data, the distributions were evaluated at the national level. Moreover, when raw values were slightly larger than one they were set to unity with no further adjustment[22]. As an example, final $P(G=1|S)$ estimates for the 1982 birth cohort are reported in Table 1 below.

***Table 1**. Estimated probabilities $P(G=1|S)$ of attaining the upper secondary school diploma by parental educational level and gender – 1982 birth cohort*

Parental education	tertiary	upper secondary	lower sec./primary
males	0.97	0.92	0.50
females	1.00	1.00	0.59

5.3 Results

Following the approach outlined in Section 4.1, primary and secondary contributions to $P(Y=1|S=j)$ are evaluated for each year. The procedure is outlined in Figure 1.

A full set of results regarding the lower secondary school final mark distribution and the transition rate to the academic track by highest parental educational level are presented for the 1982 birth cohort[23]. Counterfactual probabilities and the primary vs. secondary effect decomposition are instead reported for all years.

Estimates of $P(A|S)$ – conditional on parental highest educational level, but also on gender and geographical area – derived from the graduates survey and corrected for sample selection (as outlined in Section 4) are reported in Table 2. The distribution varies greatly across levels of S, being more favourable for children from well educated families; in line with the international evidence, females are better performers than males; more positive marks are observed in the South and Isles[24].

repeatitions; however, if grade failure is roughly stationary, the difference should be negligible. Second, the information derived from the Census includes children who have not obtained the lower secondary school degree. We assume that the students failing to pass the lower secondary school examination belong to lower educated families; the assumption is highly reasonable, since, as we have shown above, the great majority of the students passing the exam with the lowest mark *sufficiente* come from the lowest social strata. The number of students of the lowest social strata has been adjusted by subtracting from it the number of students who have not passed the final exam (data provided by the Ministry of Education); the distributions were then evaluated accordingly.

[20] In some cases, students from upper level families appeared to be slightly more likely to to drop out than those from middle level families. No adjustment was made on this.

[21] Sampling variability should enter here only via $P(S|G=1)$, but the standard errors of the estimates from the graduates survey are very small, and cannot alone explain the inconsistencies.

[22] Different sets of $P(G=1|S)$ were applied to check robustness of results: decompositions (4a) and (4b) appeared to be only slightly affected by mild changes in these percentages.

[23] Ad discussed before, this is here considered the same as those choosing secondary school track in 1996.

[24] As noted in Section 3, there is evidence of some measurement error across the country, as this result is not in line with the standardized resulted by the international assessment PISA. their child

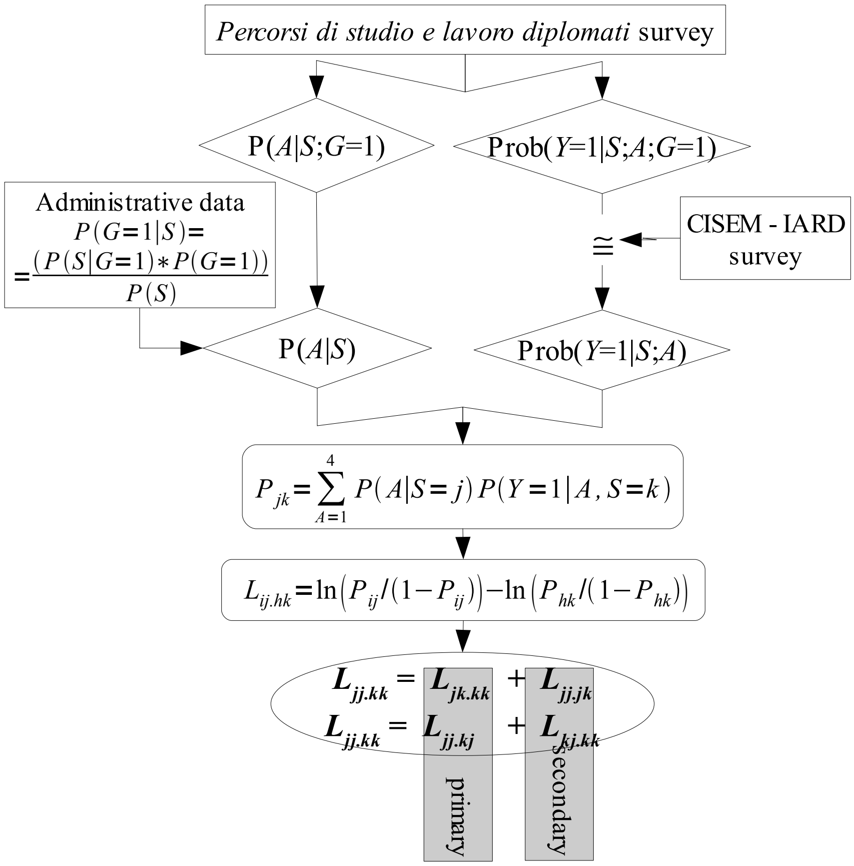

Figure 1. *Scheme of the applied decomposition technique*

Table 2. *Lower secondary school final mark distribution P(A|S) after sample selection correction, by highest parental educational level, gender and area-1982 birth cohort*

	Male					Female				
	Parental education	Lower sec. school final mark				Parental education	Lower sec. school final mark			
		satisfactory	good	very good	excellent		satisfactory	good	very good	excellent
North-West	tertiary	0.26	0.23	0.25	0.26	tertiary	0.12	0.23	0.22	0.43
	upper sec.	0.38	0.29	0.17	0.15	upper sec.	0.17	0.29	0.29	0.25
	lower sec./prim.	0.68	0.17	0.10	0.04	lower sec./prim.	0.56	0.23	0.13	0.08
North-East	tertiary	0.27	0.30	0.25	0.18	tertiary	0.06	0.25	0.29	0.40
	upper sec.	0.42	0.28	0.19	0.11	upper sec.	0.17	0.34	0.27	0.23
	lower sec./prim.	0.71	0.16	0.09	0.05	lower sec./prim.	0.58	0.22	0.13	0.08
Centre	tertiary	0.28	0.26	0.18	0.29	tertiary	0.10	0.17	0.28	0.45
	upper sec.	0.40	0.28	0.18	0.14	upper sec.	0.21	0.27	0.24	0.28
	lower sec./prim.	0.72	0.16	0.07	0.06	lower sec./prim.	0.58	0.21	0.12	0.09
South Isles	tertiary	0.25	0.17	0.24	0.35	tertiary	0.06	0.18	0.23	0.54
	upper sec.	0.39	0.26	0.18	0.17	upper sec.	0.18	0.23	0.22	0.37
	lower sec./prim.	0.68	0.18	0.08	0.07	lower sec./prim.	0.54	0.20	0.13	0.14

Table 3. *Raw transition rates to the academic track P(Y=1|A,S) by highest parental educational level, lower secondary school final marks, gender and area-1982 birth cohort*

	Male					Female				
	Parental education	Lower sec. school final mark				Parental education	Lower sec. school final mark			
		satisfactory	good	very good	excellent		satisfactory	good	very good	excellent
North-West	tertiary	0.37	0.61	0.80	0.90	tertiary	0.29	0.61	0.82	0.84
	upper sec.	0.13	0.30	0.47	0.77	upper sec.	0.07	0.25	0.50	0.67
	lower sec./prim.	0.04	0.10	0.14	0.46	lower sec./prim.	0.03	0.12	0.28	0.48
North-East	tertiary	0.31	0.58	0.78	0.96	tertiary	0.21	0.61	0.72	0.91
	upper sec.	0.08	0.20	0.43	0.73	upper sec.	0.08	0.24	0.43	0.64
	lower sec./prim.	0.03	0.07	0.18	0.59	lower sec./prim.	0.03	0.10	0.21	0.56
Centre	tertiary	0.51	0.68	0.84	0.92	tertiary	0.32	0.55	0.83	0.93
	upper sec.	0.10	0.26	0.51	0.72	upper sec.	0.15	0.26	0.53	0.74
	lower sec./prim.	0.03	0.09	0.20	0.45	lower sec./prim.	0.06	0.21	0.34	0.50
South Isles	tertiary	0.36	0.68	0.69	0.86	tertiary	0.23	0.42	0.68	0.85
	upper sec.	0.10	0.21	0.39	0.63	upper sec.	0.13	0.22	0.50	0.68
	lower sec./prim.	0.02	0.08	0.28	0.47	lower sec./prim.	0.09	0.13	0.28	0.51

Table 3 shows the raw observed transition rates to the academic track for all sub-groups. As anticipated in Section 4.1, in order to compute counterfactual probabilities and the ensuing decomposition into primary and secondary effects, $P(Y=1|A,S)$ was both estimated nonparametrically and modelled with binary logit regression on *A*, gender and area - separately for each value of *S*. For the latter approach, since preliminary log-linear analysis showed no significant interactions among these regressors, only main effects of the three covariates were included in the model. Since the parameter estimates for the distinct *A* levels are remarkably close to a progression with unit steps for all three educational levels, models where *A* was taken as a quantitative covariate (taking values 1-4 from *satisfactory* through *excellent*) were preferred. Results for the logit approach are shown in Table 4.

As expected, the propensity to enrol into a *liceo* is much higher among better performing students, although this effect seems to be slightly weaker for parents with tertiary education. At the same level of demonstrated ability, however, the transition probability is much higher among high status children[25]. Gender differences are less marked. Gender is significant for lower and upper *S*: females with low educated parents are more likely than males to enter the academic track given ability, while transition probabilities are higher for males from families with tertiary education. Geographical effects are not very clear. For these reason – and since the probabilities *Pij* (factual and counterfactual) resulting from the simplified models without the dummies for area do not change much with respect to the ones coming from the extended model – the more parsimonious specification was employed in the end.

[25] This can be seen from the raw probabilities in Table 3 and is reflected in the values of the constant in the logit models in Table 4.

Table 4. *Logit models for the transition probabilities to academic track – 1982 birth cohort*

	***S* = tertiary**			***S* = upper secondary**			***S* = lower sec./primary**		
Full models	β	Sig.	Exp(β)	β	Sig.	Exp(β)	β	Sig.	Exp(β)
ind_North-West	0.286	0.014	1.33	0.208	0.001	1.23	-0.147	0.101	0.86
ind_North-East	0.252	0.053	1.29	-0.034	0.646	0.97	-0.196	0.055	0.82
ind_Center	0.594	0	1.81	0.256	0.000	1.29	0.126	0.162	1.14
gender (female)	-0.290	0	0.75	0.069	0.168	1.07	0.385	0	1.47
ind_buono	1.030	0	2.8	0.955	0	2.6	1.050	0	2.86
ind_distinto	1.827	0	6.22	1.976	0	7.22	2.032	0	7.63
ind_ottimo	2.765	0	15.88	2.895	0	18.09	3.088	0	21.94
constant	-0.778	0		-2.242	0		-3.314	0	
Simplified models	β	Sig.	Exp(β)	β	Sig.	Exp(β)	β	Sig.	Exp(β)
ind_North-West	0.289	0.013	1.34	0.210	0	1.23	-0.149	0.097	0.86
ind_North-East	0.259	0.046	1.3	-0.031	0.669	0.97	-0.197	0.053	0.82
ind_Center	0.595	0	1.81	0.257	0	1.29	0.125	0.164	1.13
gender (female)	-0.287	0	0.75	0.069	0.169	1.07	0.385	0	1.47
A	0.899	0	2.46	0.971	0	2.64	1.023	0	2.78
constant	-1.614	0		-3.213	0		-4.327	0	
No geographic area	β	Sig.	Exp(β)	β	Sig.	Exp(β)	β	Sig.	Exp(β)
gender (female)	-0.288	0	0.75	0.072	0.15	1.07	0.382	0	1.47
A	0.871	0	2.39	0.965	0	2.62	1.026	0	2.79
constant	-1.315	0		-3.098	0		-4.365	0	

In Table 5, rows refer to school mark distributions according to parental education, while columns indicate which level of S is used to model the transition probability to the academic track. The square matrix is shown for each of the four surveys. In each, the numbers located on the diagonals are the actual estimated transition probabilities P_{jj}= $P(Y=1|S=j)$ for each family status. These values are always higher for females, and the stronger gender differences are observed for low S: females with low parental education are almost twice as likely to enrol into a *liceo* than males.

Off diagonal elements P_{jk} are instead counterfactuals, combining lower secondary school marks distribution and conditional transition probabilities for different parental educational levels.

For example, P_{11}=0.660 is the transition rate of a male whose parents have tertiary education for those entering secondary school in 1999. For the same year, the transition probability of an hypothetical child with the ability distribution of the upper class but the conditional propensity to choose *liceo* of the lower class is given by P_{13}= 0.191; similarly, the transition probability when the ability distribution is that of the lowest class and the conditional propensity is that of the upper class is P_{31}= 0.489.

There is a noticeable tendency - somewhat stronger for males - to decline faster along rows than along columns, indicating that the differences in family preferences for Y=1 due to S *given* children's marks are more relevant in determining the track choice with respect to school performance differences due to S. Only modest changes are observed over time, and no clear trend emerges.

Table 5. *Estimates of P_{ij} by cohort*

Choosing track in...	**Male**			**Female**		
1990	***P(Y=1|S;A) referring to...***			***P(Y=1|S;A) referring to...***		
P(A|S) referring to...	tertiary	upper sec.	lower sec. /primary	tertiary	upper sec.	lower sec. /primary
tertiary	0.730	0.439	0.238	0.767	0.564	0.322
upper secondary	0.686	0.351	0.172	0.682	0.454	0.244
lower sec. /primary	0.622	0.221	0.082	0.538	0.296	0.139
1993	***P(Y=1|S;A) referring to...***			***P(Y=1|S;A) referring to...***		
P(A|S) referring to...	tertiary	upper sec.	lower sec. /primary	tertiary	upper sec.	lower sec. /primary
tertiary	0.767	0.453	0.276	0.820	0.536	0.329
upper secondary	0.679	0.343	0.181	0.753	0.439	0.252
lower sec. /primary	0.575	0.214	0.085	0.636	0.301	0.147
1996	***P(Y=1|S;A) referring to...***			***P(Y=1|S;A) referring to...***		
P(A|S) referring to...	tertiary	upper sec.	lower sec. /primary	tertiary	upper sec.	lower sec. /primary
tertiary	0.709	0.415	0.236	0.726	0.504	0.344
upper secondary	0.619	0.304	0.153	0.642	0.415	0.271
lower sec. /primary	0.491	0.185	0.076	0.448	0.248	0.152
1999	***P(Y=1|S;A) referring to...***			***P(Y=1|S;A) referring to...***		
P(A|S) referring to...	tertiary	upper sec.	lower sec. /primary	tertiary	upper sec.	lower sec. /primary
tertiary	0.660	0.356	0.191	0.720	0.459	0.285
upper secondary	0.590	0.280	0.143	0.683	0.407	0.242
lower sec. /primary	0.489	0.172	0.085	0.556	0.260	0.135

Table 6 presents the results of the decomposition into primary and secondary effects. Both formulas (4a) and (4b) are computed, and produce similar results; average contributions are reported here. The main finding is that *secondary effects* tend to prevail in all contexts, the sole exception being that of medium vs. low status females for the 1996 choice cohort.

It is important to recognize that this result does not imply that class differentials in children's *ability* are weak (see the discussion on measurement error in Section 3), nor that differentials due *S* in children's *school marks* are weak. Results imply instead that *differentials due to S in secondary school choices* are mainly driven by differences in the transition probabilities given previous school performance, while differences in the performance distributions play a weaker role. This may occur either because performance distributions vary little across social status, or because performance does not affect much school choices[26]. Distinguishing between these two alternatives is possible by looking directly at the estimates of $P(A|S)$ and $P(Y=1|A,S)$.

No clear time trend is present; while small-scale variability is widespread, a general overall stability of the prevalence of secondary effects can be safely inferred. The relative importance of secondary effects was somewhat stronger for males than for females in the first survey; however, much weaker differences are found in the following years.

A clearer picture can be obtained looking across social origins: secondary effects are consistently stronger across time and gender when comparing upper and middle status

[26] In principle, there could be wide family status differences in the observed level of ability, but if school choices are affected little by performance, depending mainly on social status, these differences would not exert a relevant role.

with respect to middle and low: highly educated parents appear to attach a stronger value to the academic track, regardless of the proficiency level demonstrated by their offsprings.

Table 6. *Primary and secondary effects decomposition for each of the four cohort years*

		Male			Female		
Choosing track in...		tertiary > upper sec.	tertiary > lower sec./primary	upper sec. > lower sec./primary	tertiary > upper sec.	tertiary > lower sec./primary	upper sec. > lower sec./primary
1990	average % primary	0.185	0.257	0.413	0.316	0.351	0.420
	average % secondary	0.815	0.743	0.587	0.684	0.649	0.580
1993	average % primary	0.246	0.323	0.439	0.224	0.305	0.418
	average % secondary	0.754	0.677	0.561	0.776	0.695	0.582
1996	average % primary	0.259	0.332	0.432	0.285	0.419	0.543
	average % secondary	0.741	0.668	0.568	0.715	0.581	0.457
1999	average % primary	0.202	0.270	0.423	0.147	0.295	0.469
	average % secondary	0.798	0.730	0.577	0.853	0.705	0.531

6. Conclusions

The results described in Section 5.3 are particularly interesting when considered within the international context. The most striking finding is that the relative contribution of *primary effects* is *much lower in Italy* than in the other countries for which the analysis has been carried out. Let us review the main results. Primary effects[27] account for about 76% of the total social background effect in UK (Jackson *et al.*, 2007, for year 2001), 58% in Stockholm, Sweden (Erikson, 2007, for 1990), 47% in the German lander Rhineland (Stocké, 2007, for 2003), 58% in the Netherlands (Koosterman *et al.*, 2007, for 1999). The corresponding estimates for Italy are much lower: 27% for males and 30% for females for the most recent data analyzed, and there is no evidence of a move towards a more balanced situation in the decade under study. Although these values are not fully comparable, because of cross-country institutional differences, definitions of social status[28] and because ability assessments are not always standardized, differences are however large, and it would be of great interest to understand the reasons laying behind them.

We can think of different topics for further research:

(i) In order to interpret the results from a comparative point of view, the *absolute* contributions of primary and secondary effects should be evaluated together with the *relative* ones. This implies recovering comparable estimates of the total effect,

[27] The percentage with respect to the high-low status comparison is reported here.

[28] In UK and Sweden father's social class, in Germany mother's social class, in the Netherlands and Italy the highest parental educational level.

given by (3)[29]. Note however that cross-country comparisons are even more problematic in this case: employing parental education or social class can give rise to big differences within countries (we can see this from PISA, for which common alternative definitions are possible)[30].

(ii) The low importance of primary effects in Italy with respect to other countries can have two alternative interpretations: (i) social background differentials in the school performance distributions are relatively weak; (ii) the role of ability in educational decisions is weak. Evidence from the international assessment carried out on 4th graders, PIRLS (*Progress in International Reading and Literacy Study;* Mullis *et al.*, 2003) can help to shed light on this issue. Simple regression analysis indicate for example that Italy is one of the countries with the *lower* inequality of opportunity with respect to performance scores near the end of primary school.

(iii) The assessment of how specific institutional features – in particular, early tracking – affect equality of opportunity in education is the focus of an interesting body of work (Hanushek and Woessman, 2006; Woessman, 2007; Brunello and Checchi, 2007): by employing international surveys like PISA, the school design effect is identified by exploiting the cross-country variability. To our knowledge no attempt has been done yet to deepen the understanding of how institutional features promote or discourage primary and secondary effects[31]. In order to put forward educational policies with the aim to reduce educational inequality it would be very useful to separate the effects on school performance from those on choices given performance. At the moment this aim is difficult to accomplish, as on one hand it is difficult to harmonise national data to allow for adequate cross-national comparisons, on the other hand international data such as PISA cannot be employed for this purpose, because no measure of ability before school choice is available. This could be an interesting challenge for future research.

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[29] Sticking to the highest parental educational level as a measure of status, we obtain the following raw odds-ratio between high and low social class for a selection of countries: 4.69 for the Netherlands, 6.87 for Italy, 12.97 for Germany (no estimates can be derived for UK and Sweden, as tracking occurs at age 16, after the assessment has been carried out). A similar picture is obtained by employing the syntethic measure of social background ESCS, provided by PISA analysts.

[30] Moreover, as discussed in Contini and Scagni (2008), estimates of logit regression coefficients are biased with unobserved heterogeneity .

[31] For example, why is it that in Italy primary effects are so low? Could it be due to the fact that the compulsory school system is quite highly standardised in Italy? (standardization refers to the degree to which the quality of education meets the same standards nationwide; Allmendinger, 1989). On the other hand, secondary effects are strong. Is this related to the absence of performed-based restrictions to the academic track, at work in other countries (Netherlands for example)?

References

Allmendinger J. (1989) *Career Mobility Dynamics: A Comparative Analysis of the United States, Norway, and West Germany*, Berlin: Max-Planck-Institute fur Bildungsforschung.

Boudon R. (1974) *Education, Opportunity and Social Inequality.* New York, Wiley

Breen R., Goldthorpe J.H. (1997) Explaining educational differentials. Towards a formal rational action theory, *Rationality and Society*, Vol 9(3), 275-305

Breen R., Jonsson J.O. (2000) A Multinomial Transition Model for Analyzing Educational Careers, *American Sociological Review* 65:754-772.

Brunello G., Checchi D. (2007) Does school tracking affect equality of opportunity? New international evidence, *Economic Policy*, 52

Contini D., Scagni A.(2009) Equality of Opportunity in Secondary School Enrollment. Comparing Italy, Germany and the Netherlands, forthcoming on *Quality and Quantity*

Erikson R. (2007) Social Selection in Stockholm Schools: Primary and Secondary Effects on the Transition to Upper Secondary Education, pp. 61-81 in: *From Origin to Destination. Trends and Mechanisms in Social Stratification Research*, edited by Stefani Scherer, Reinhard Pollak, Gunnar Otte, and Markus Gangl, Campus Verlag, Frankfurt

Erikson R., Goldthorpe J.H. (1992) *The constant flux: a study of class mobility in industrial societies*, Oxford, Clarendon Press

Erikson R., Goldthorpe J.H., Jackson M., Yaish M., Cox D.R. (2005) On class differentials in educational attainment, *Proceedings of the National Academy of Sciences,* vol. 102, n. 27, 9730–9733

Eurydice, Italian Unit (2006) *Eurybase the Information Database on Education Systems in Europe,The Education System in Italy 2005/2006*, http://www.eurydice.org

Goldthorpe J.H. (1996) Class analysis and the reorientation of class thoery: the case of persisting differentials in educational attainment, *British Journal of Sociology,* 45(3): 481-506

Hanushek E.A., Woessman L. (2006) Does educational tracking affect performance and inequality? Difference-in-differences evidence across countries, *Economic journal, vol* 116

Hannum E., Buchman C. (2003) The consequences od global educational expansion, *Occasional paper of the American Academy of Arts and Sciences*, Cambridge, MA

ISTAT (2004) *Percorsi di studio e di lavoro dei diplomati*, Roma

Jackson M., Erikson R., Goldthorpe J.H., Yaish M. (2007) Primary and Secondary Effects in Class Differentials in Educational Attainment: the Transition to A-Level Courses in England and Wales, *Acta Sociologica,* 50(3), 211-229

Kloosterman R., de Graaf P., Ruiter S., Kraaykamp G. (2007) Parental education and the transition to higher secondary education. A comparison of primary and secondary effects for five Dutch cohorts (1965-1999), *Proceedings of the RC28 Spring Meeting in Brno*, Czech Republic

Mullis I.V.S., Martin, M.O., Gonzalez, E.J., Kennedy, A.M. (2003) *PIRLS* 2001 *International Report: IEA's Study of reading literacy achievement in primary schools*, Boston College

OECD (2005), *PISA 2003 Technical report*

Shavit Y., Blossfeld, H.P. (1993) (eds.) *Persistent inequality: changing Educational attainment in thirteen countries*, Westview, Boulder, Colorado

Stocké V. (2007) Strength, Sources, and Temporal Development of Primary Effects of Families´ Social Status on Secondary School Choice, Sonderforschungsbereich 504, WP series, n° 07-60

Woessman L. (2007) Fundamental Determinants of School Efficiency and Equity: German States as a Microcosm for OECD Countries, *CESifo Working Paper* 1981