Comparable Estimates of Intergenerational Income Mobility in Italy

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ABSTRACT: This paper examines the degree of intergenerational economic mobility in Italy. It adds to the growing number of international studies of the extent to which economic status is passed on across generations. On the basis of recent econometric innovations used in the literature (Bjorklund and Jantti, 1997), we are able to overcome some of the data limitations for Italy. We use the Historical Database of the Bank of Italy households survey, which contains information from 1977 to 2002. Retrospective information in the repeated cross-sections may be exploited by applying a two-sample two stage least squares estimation. We estimate the intergenerational income elasticity for Italy and find that mobility is limited. From an overall comparison, the evidence provided in this paper hints at Italy in the low-mobility group among advanced societies in the range of values found in the US and the UK. The analysis of the results allows a characterization of the main patterns in the transmission of economic status in Italy.

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1. Introduction

A large body of social science research has long questioned whether all individuals have the same opportunity to achieve social and economic success irrespective of their parents' status. A rapidly expanding international literature has been investigating the degree to which socio-economic status is passed on between generations. Recent reviews of the latest efforts by both economists (see Solon, 2002; Corak, 2004*a*) and sociologists (Breen, 2004; Breen and Jonsson, 2005) find significant differences in the degree of inequality persistence across countries.

Economists have tended to choose income or earnings as the preferred dimensions along which to characterize one's position. The increasing literature has shed light on a number of problems with correctly defining and measuring intergenerational mobility when focusing on this particular individual dimension of social status. One criterion has been the estimation of single-number expressions for levels of mobility, in the form of degrees of association of the economic outcome of an individual with her family background. This can be done, for instance, by studying the relationship between the economic outcomes of members of the same family in different generations.

While most of the literature regards the United States, a number of studies have been devoted to the analysis of the transmission of economic status in other countries (among others: Britain, Germany, Canada, Sweden, South Africa). International studies are important not only for the natural interest in characterizing an important facet of a country's income inequality but also because comparisons between countries can contribute to an understanding of the mechanisms underlying generational income mobility. How are countries with different institutional settings in the labour market, different educational systems and different levels of cross-sectional inequality doing in terms of intergenerational mobility?

In this perspective, Italy certainly represents an interesting case for comparison: its labour market is considered to be heavily regulated, with fairly centralized wage-setting institutions and a high proportion of its workforce covered by collective bargaining; the school system is extremely centralized and egalitarian; the level of cross-sectional income inequality is lower than in the United States but higher than most Western European countries (Brandolini and Smeeding, 2005).

Being notably characterized by limited availability of data, Italy has received relatively little attention in the economic literature. Lack of data has constrained previous studies to focus on measures of "socio-economic" condition such as occupational class or educational attainments. In a widely quoted study, Checchi *et al.* (1999) find that intergenerational mobility between occupations and between education levels in Italy is lower than in the United States. International comparisons by sociologists indicate that Italy displays low levels of intergenerational mobility in terms of social fluidity (Breen, 2004).

Taking advantage of recently developed empirical methods used to overcome similar data limitations in a number of other countries (Bjorklund and Jannti, 1997; Dunn, 2004; Ferreira and Veloso, 2004; Lefranc and Trannoy, 2005), we are able to produce new internationally comparable estimates of the degree of intergenerational mobility in Italy. The evidence seems to confirm the existence of considerable intergenerational economic persistence. The magnitudes of the estimates are within the range of values found in the "least" mobile advanced countries (Britain and United States), and noticeably higher than those estimated in the Scandinavian countries.

The rest of the paper proceeds as follows. Section 2 presents the basic econometric model of the intergenerational transmission of economic status and outlines the empirical methods used for the estimation. Section 3 describes the data and the process of selection of the observations into the final samples. Section 4 presents and discusses the estimates. Section 5 concludes.

2. Measurement Issues and Methodology

2.1 *Econometric Model and Previous Literature*

When attempting an empirical analysis of the degree of intergenerational economic mobility, a large number of economic researches have looked at measures that summarize in a single number the joint distribution of income at two points in time. If Y_i^s is a measure of the long-run economic status of sons and Y_i^f the corresponding value for fathers, then the intergenerational relationship can be specified as:

$$Y_i^s = \alpha + \beta Y_i^f + \varepsilon_i \qquad (1)$$

From which it is evident that β can be interpreted as a summary measure of the degree of intergenerational persistence. Conversely, $1 - \beta$ can be thought of as a summary measure of intergenerational mobility.[2]

In order to estimate the relationship, a measure of the long-run economic status as well as a combination of representatives of two different generations is needed. Most authors use income or earnings as preferred measures of economic status, while fathers and sons are predominantly chosen to represent the two succeeding generations. The standard linear equation (1) is the base of most empirical works in the economic literature. Typically, a regression of a logarithmic measure of sons' income on a logarithmic measure of fathers' income is performed. Then, the coefficient β – termed the intergenerational income elasticity – is estimated by applying ordinary least squares (OLS). By construction, the elasticity β will indicate the percent difference in sons' income observed for each 1 percent difference across the incomes of the fathers. Conceivably, any real number could be obtained from the estimation of equation (1); a negative value would indicate a situation where children of parents who were high in their distribution of income tend to be low in their own generation distribution. On the contrary, a positive value would indicate intergenerational persistence of incomes where higher parental income is associated with higher child income. In fact, all empirical studies in the rich countries have found β to lie between zero and one. Within this range, regression to the mean occurs, but at a rate inversely proportional to β.

The existing evidence can be broadly split into two major waves. The earlier studies on the issue resulted in estimates of β at around 0.2, leading to the conclusion that

[2] See Bowles and Gintis (2002) for a possible derivation of the above formulation. Note that the elasticity β differs from the intergenerational correlation coefficient r. The correlation is the regression coefficient multiplied by the ratio of the standard deviations of income in the two generations: $r = \beta(\sigma_f / \sigma_s)$.

advanced countries were characterized by strong mobility.[3] It is well recognized that the earlier lower estimates from these studies were biased downwards by two major problems (Solon, 1999). First, they relied on single-year measures of fathers' income which, because of both response errors and transitory fluctuations, represent an erroneous proxy for permanent status.[4] The second major problem of some earlier studies depended on the use of particularly homogeneous samples which naturally tended to underestimate the degree of intergenerational persistence. Aware of these two problems, empirical studies in the 1990's based their analysis on *longitudinal* samples and used *representative* data. By averaging fathers' earnings over more than only one year (generally five years) these studies reported estimates of the elasticity consistently above the previous ones (Table 1).

An alternative strategy for dealing with measurement errors in incomes is to resort to instrumental variable (IV) estimation. This approach consists of using one or more variables to instrument for father's income or earnings (the most common variables are occupational status and education). The idea is that the instruments will possibly suffer less from transitory variation than the single-year measures of income, thus representing a better proxy for long-run economic status. The greater their ability to capture the variance in permanent income the better job IV estimates will do. A problem of this method concerns the possibility of instruments being correlated to son's economic status independently of fathers' income (e.g. fathers' education). This problem will generally cause an upward bias in the IV estimator, for the instruments have a separate *positive* impact on the dependent variable.[5] However, recent results by Mazumder (2005) lead to question whether the IV estimates over-balance the downward bias induced by noisy measures of permanent status. By using a larger panel of US Social Security data, he is able to substantially increase the time span over which earnings are averaged. He estimates an intergenerational elasticity of earnings in the order of 0.6 or higher when averaging earnings over the longest period (16 years).

[3] Belief that is best synthesized by the often-quoted presidential address to the American Economic Association by Gary Becker in 1988: "In all these countries, low earnings as well as high earnings are not strongly transmitted from fathers to sons."

[4] Bowles (1972) first pointed out some problems with proxying permanent income by single year measure of income. Errors in measuring *sons'* permanent income do not lead to biased coefficient in a regression framework, even though they may cause imprecision.

[5] See Bjorklund Jantti (1997) and Solon (1992) for a formal treatment of both the multiple-years average and IV correction to measurement errors in permanent income.

Estimates of the intergenerational elasticity also vary with the *age* at which economic status for both fathers and sons is observed (life-cycle bias). Grawe (2001) finds evidence of higher mobility estimates when basing his analysis on mature fathers rather than young fathers. Son's age also is a concern. In particular, there seems to be a downward bias for those estimates based on measures of son's earnings taken at early stages of their careers as showed by Reville (1995).

Table 1.

Selection of international studies using longitudinal data

Country	Study	Elasticity	Estimation Method	Sons ages (average or range)	Fathers ages (average or range)
United States	Solon (1992)	0.41	OLS	25-33	44
	Solon (1992)	0.53	IV	25-33	44
	Mazumder (2005)	0.61	OLS	30-35	27-69
UK	Dearden *et al.*	0.58	IV	33	47.5
Germany	Wiegand (1997)	0.34	OLS		
Canada	Corak and Heisz (1999)	0.23	OLS	29-32	42.5
Sweden	Osterberg (2000)	0.13	OLS	25-51	52
Finland	Osterbacka (2001)	0.13	OLS	34.9	46

Source: author's selection from the review in Corak (2004*b*).

2.2. *Methodology used for estimating β in Italy*

The procedures by which the results in Table 1 are obtained are not directly applicable to the case of Italy. Like most countries, Italy does not have a long enough intergenerational panel that permit the explicit observation of father-son pairs. However, repeated cross-sections from household surveys can represent a good alternative, in that one can exploit retrospective information on parental background by sons. Among the parental characteristics reported by sons it is very hard to find income or earnings, while occupational status, level of education and other demographic characteristics are more common. These latter variables can be used to infer income from a sample of older men (pseudo fathers) and thereby to estimate intergenerational correlations. This procedure is a special case of the "two sample instrumental variable" (TSIV) technique examined in Angrist and Krueger (1992) and Arellano and Meghir (1992). Bjorklund and Jantti (1997) first applied this methodology to intergenerational mobility estimation (with Swedish and US data). They performed a two stage regression based on two samples: a sample of sons who have reported their fathers' socio-economic characteristics and a sample of adult men (pseudo fathers) whose age is consistent with that of the actual fathers. Once the samples are selected, the steps required for this empirical strategy are the following: (i) estimate an income equation from the older sample; (ii) use the estimated coefficients to predict fathers' income on the basis of sons' recollections; (iii) regress sons' income on the predicted fathers' income.

Let the relation between father's current (Y_{it}^{f}) and permanent income (Y_i^{f}) be given by:

$$Y_{it}^{f} = Y_i^{f} + e_{it}^{f} \tag{2}$$

where e_{it}^{f} includes both measurement error and transitory fluctuations in current income. Let X_i^{f} be a vector of time-invariant characteristics of fathers as recalled by their sons and consider father's permanent income as determined by the following relation

$$Y_i^{f} = \lambda X_i^{f} + \upsilon_i^{f} \tag{3}$$

with υ_i^{f} and X_i^{f} independent. Substituting into equation (2) gives

$$Y_{it}^{f} = \lambda X_i^{f} + \upsilon_i^{f} + e_{it}^{f} \tag{4}$$

Since Y_{it}^{f} cannot be directly observed from the sample of sons, an estimate of λ is obtained by regressing equation (4) using the distinct sample of pseudo fathers. The coefficient $\hat{\lambda}$ thus obtained will permit a prediction of income for the actual fathers',

$\hat{Y}_i^f = \hat{\lambda} X_i^f$. The standard linear intergenerational regression is then performed on fathers' predicted income

$$Y_{it}^s = \alpha + \beta \hat{Y}_i^f + \omega_{it} \quad (5)$$

where $\omega_{it} = \varepsilon_i + e_{it}^s + \beta \upsilon_i^f + \beta X_i^f \left(\lambda - \hat{\lambda}\right)$.[6]

The studies of intergenerational mobility that have used this approach (see Table 2), provide estimates of β based on the estimation of equation (4) and (5) from separate samples. Inoue and Solon (2005) refer to this procedure as a computationally convenient two sample two stage least squares (TS2SLS) variant of Angrist and Krueger's estimator. By analyzing the properties of the two estimators, they show that the commonly used TS2SLS estimator is more asymptotically efficient than the TSIV estimator because it implicitly corrects for differences in the empirical distributions of the instrumental variables between the two samples.

Table 2.

Existing international studies using two-sample procedures

Country	Study	Elasticity	Sons ages (average or range)	Fathers ages (average or range)
United States	Bjorklund and Jantti (1997)	0.42-0.52*	28-36	45
Sweden	Bjorklund and Jantti (1997)	0.28	30-39	43.3
France	Lefranc and Trannoy (2005)	0.41	30-40	55-70
Brasil	Dunn (2004)	0.69	25-34	30-50
	Ferreira and Veloso (2004) (wages)	0.58	25-64	25-64

Source: study data from Corak (2004*b*) and individual papers.
* Lower estimate is obtained not controlling for age of both fathers and sons.

[6] Equation (5) results from equations (1), (2) and (3) assuming that current and permanent income are related in a similar way for sons and that e_{it}^f and e_{it}^s are uncorrelated.

Like the standard IV procedure, the method is most vulnerable if one of the predictors of father's income is itself a predictor of son's income. Bjorklund and Jantti (1997) compare their two sample two stage estimate for the US with the value found by averaging fathers' earnings over five years on a data set in which they also have real fathers. The US elasticities between sons' current earnings and the five-year average of fathers' earnings are very close to those of Solon (1992) and are lower than the TS2SLS estimates (0.33-0.39 compared to 0.42-0.52). However, we noted that Mazumder (2005) shows that even a five-year average of father's income still cause a serious downward bias in the estimated value (suggesting that the true value of the parameter is about 0.6). On the basis of his study we cannot conclude that the procedure used in this paper overestimates the true value of the intergenerational coefficient in Italy.

The estimated values in Table 2 confirm the difference in mobility levels between Sweden and the US shown in the previous table, and point at France as a sort of intermediate case. Also, the results for Brazil are consistent with the conjecture of stronger intergenerational persistence in less developed countries (see Solon, 2002).

An alternative estimator is obtained by predicting income for *both* fathers and sons. That is, one can run the regression

$$Y_{it}^{s} = \theta Z_{i}^{s} + \upsilon_{i}^{s} + e_{it}^{s} \tag{6}$$

and then use the predicted value $\hat{Y}_{i}^{s} = \hat{\theta} Z_{i}^{s}$ to calculate the intergenerational coefficient from the following equation

$$\hat{Y}_{i}^{s} = \alpha + \beta_{1} \hat{Y}_{i}^{f} + \psi_{i} \tag{7}$$

Note that Z_{i}^{s} represents the vector of observable socio-demographic characteristics used as predictors of income in the sample of sons. Typically the coefficient β_{1} from equation (7) will differ from β obtained by estimating equation (5). β_{1} will measure the degree to which the observed components in the permanent income of fathers (X_i^f) and sons (Z_i^s) are associated. We expect β_1 to be smaller than β in the presence of a positive association between fathers' observed characteristics and the unpredicted part of sons' income. Dearden *et al.* (1997) and Bjorklund and Jantti (1997) do indeed find lower values of the income elasticity when using the "prediction approach". [7]

Finally, Italian data also permit to perform a direct OLS estimation relying on contemporaneous income report by co-residing fathers and sons. Obviously, this sample will be smaller and less representative of all individuals, possibly leading to display a different intergenerational income association than would a more representative sample. Generally, the use of single-year measures of fathers' income, the observation of relatively younger sons and of older fathers will tend to underestimate the true intergenerational coefficient. For co-residing pairs, we regress the natural log of sons' income on the natural log of fathers' income with controls for age and age squared for both fathers and sons, as shown by equation (8).

$$Y_{it}^s = \alpha_0 + \beta_2 Y_{it}^f + \alpha_1 age_i^s + \alpha_2 agesq_i^s + \alpha_3 age_i^f + \alpha_4 agesq_i^f + \varepsilon_i \tag{8}$$

I will interpret the estimates obtained from this procedure as providing a lower bound to the true intergenerational income elasticity.

3. Data and Sample Selection

In this paper, we use data from the Bank of Italy Survey on Household Income and Wealth (SHIW), a nationally representative household survey based on a random sample of approximately 8,000 households per year that is available form 1977 annually and at odd years after 1987. The SHIW is the only easily accessible source of micro data on income that spans over this long period. Brandolini (1999) describes Italian data sources and concludes that the SHIW still represents the best source of income distribution in

[7] They also provide a formal discussion of the two estimators.

Italy.[8] In order to enhance comparisons over time, the Bank of Italy has constructed a Historical Database, elaborating data from the 1977 survey and addressing the numerous changes the SHIW questionnaire has undergone over the years (D'Alessio and Gallo, 1997; D'Alessio, 1997). In the recent waves of the survey, head of households are asked to recall some characteristics of their parents (among which there are year of birth, educational qualifications and employment status). The information for fathers is indicatively referred to the same current age of the respondent.

All income is recorded net of taxes and social security contributions, with separate records for each recipient along with basic individual characteristics such as age, sex, education, work status and employment sector. In what follows, we will use annual disposable personal income from the historical database. This includes earned income from wages, salaries, and self-employment and other cash income from property, but does not include income from financial assets. This income definition is narrower than total market income (defined as before tax income from all market sources), and broader than earnings.[9]

3.1. *Sample selection*

The sample of fathers is taken from the survey conducted in 1977, which is the oldest wave of the SHIW available. The selection of fathers into the final sample tries to follow the standard procedure adopted in most of the similar studies of economic mobility. We consider employed males who are head of households and father of at least a co-resident child. We include all fathers aged 30 to 50 (i.e. born between 1927 and 1947). Following the majority of previous studies, individuals who report a non-positive income are excluded, for a final first-stage sample of 953 individuals.

The sample of sons is taken from the 2002 SHIW. They are male heads of household aged 30 to 45, whose fathers' were born between 1927 and 1947. Consistently with studies of mobility in other countries, we consider employed individuals with positive income and a report of their fathers' socio-demographic characteristics for a final

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[8] As with most survey data, major problems of SHIW regard the pattern of non-responses, mis-reporting of earnings and a relatively small sample size.

[9] Previous studies have found evidence of higher intergenerational coefficients for broader income concepts. See Mulligan (1997), Corak and Heisz (1999) and Mazumder (2005).

sample of 612 sons. Table 3 reports descriptive statistics for both samples and the magnitude of sample exclusions in the population under consideration.

We note that once the standard sample exclusions are made, individuals seem to maintain similar characteristics compared to the reference group. Selected sons appear to have higher levels of income with respect to all males in the same age range. The results may thus be biased by the selection of unrepresentatively high-earnings groups. However, given the correspondence of the above procedure with the standard exclusions adopted in the literature, the extent of the selection biases should be consistent with that of the studies to which this paper aim to be compared.

Table 3.

Descriptive statistics for selected fathers and sons

	Pseudo-Fathers (1977)		Sons (2002)	
	All Males 30-50 in 1977	Selected sample	All Males 30-45 in 2002 (whose fathers were born b/w 1927-1947)	Selected sample
N	1133	953	733	612
mean age	41.41 (4.99)	41.39 (4.99)	38.02 (4.13)	38.09 (4.13)
mean log income	9.65 (0.53)	9.69 (0.50)	9.87 (0.56)	9.93 (0.47)

Notes: Standard deviations in parentheses. Income is in 2002 Euros, deflated by the consumer price index.

3.2. *First stage variables*

To perform the first two empirical estimations described in section 2, we need a set of observable variables from the recall information in 2002 that we can use in the first stage sample for predicting fathers' income. We use the following categorical variables: educational attainments; work status; employment sector and geographic area.[10] Information on education is in term of maximum educational achievements. There are six categories: no school (less than one year of schooling); elementary (five years); lower secondary school (eight years); high school (thirteen years); bachelor (seventeen-

[10] Most similar studies have used one or more of these variables.

eighteen), and post graduate studies (more than eighteen years of education). Contrary to most similar studies, the occupational categories are not recoded following a social class schema. As noted by Checchi and Dardanoni (2002) the SHIW data does not provide a detailed classification of occupational status, making it difficult to construct a ranking of occupation based on social prestige or any other social grading. Furthermore, the detail of information on work status and sector of activity in the fathers' sample does not perfectly match with the characteristics sons are asked to recall in 2002. In order to perform the analysis, we rearrange the more disaggregated information on work status recalled by the children to be comparable with the available classification for parents. A similar rearrangement is effectuated for the variable "employment sector" where sons' recollections are less detailed than the information directly observed from fathers. As a result, based on the available information, we obtain four work status categories (blue collar, office workers/teachers, managers/professionals/entrepreneurs, self-employed) and four sectors of employment (agriculture, industry, public administration, private services). The last variable is a geographic dummy, which indicates whether the father was residing in the South.[11]

The relationship between the observable characteristics and income in the pseudo fathers sample is assumed to be valid for the true fathers. Since this cannot be verified, we check if the distribution of the characteristics self-reported in the 1977 sample is consistent with the distribution of the characteristics recollected by the sons. Table 4 compares fathers' own reported characteristics with sons' recollections.

The distributions appear to be consistent. The existing differences are of the same nature found in previous studies that use this technique. In particular, pseudo-fathers have, on average, higher schooling and more skilled occupations than the actual fathers. Typically, these discrepancies are ascribed to differential childbearing according to occupation and educational attainment.[12] Solon and Inoue (2005) note that the TS2SLS

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[11] We include this variable in order to account for the well-known geographic disparities in Italy. Considering a larger set of geographical areas only has a minor impact on estimates. While we can observe directly the area of residence in the sample of pseudo-fathers, we do not have the same information for the actual fathers. We use sons' place of birth as a proxy for the geographic area where the sons were mainly living when they grew up.

[12] To the extent that fathers with many sons are over-represented in the sample of sons, the differences are to be expected when occupation and education are correlated with the number of sons.

estimator corrects for differences between the two samples in the distribution of the observable characteristics.

Table 4.

Descriptive statistics for fathers and pseudo-fathers.

	Sons' report of fathers characteristics	Fathers' own report of their characteristics
Mean Age	42.36 (5.18)	41.39 (4.98)
Education		
None	0.08	0.06
Elementary	0.53	0.50
lower secondary	0.25	0.25
high school	0.11	0.13
Bachelor	0.03	0.06
Work Status		
blue collar	0.47	0.47
office worker & teacher	0.17	0.19
manager/profess/entrepr.	0.11	0.06
self-employed	0.25	0.28
Work sector		
agriculture	0.16	0.08
industry	0.30	0.44
public administration	0.14	0.14
private services	0.40	0.34
Area		
north/centre	0.65	0.69
south	0.35	0.31

Notes: All frequencies are weighted using sampling weights

3.3. *Co-residing samples*

The direct OLS estimate for co-residing father-son pairs is obtained from a different sample. We construct a sample of 231 pairs from the 2002 survey among those individuals who were employed and reporting positive income.[13] Table 5 reports summary statistics for selected individuals.

Table 5.

Characteristics of co-residing sample

	2002	
	Fathers	Sons
N	231	231
mean age	53.36 (5.78)	25.10 (4.29)
mean log income	9.97 (0.55)	9.19 (0.58)

Notes: Standard deviations in parentheses.

The sample mean age for sons in 2002 is 25, while the sample mean age for fathers is around 53. Because sons are observed at an earlier stage of the life cycle their mean income is lower. Compared to the previous sample, these sons are younger and their fathers are older. On the basis of the existing literature, both age differences are likely to cause lower estimated values for β.

4. Empirical results

4.1. *Regression results*

Table 6 reports regression coefficients for the intergenerational income equation for Italy. The values shown are the results of the estimation of equation (1) under the three measurement procedures outlined in section 2. First-stage estimates of fathers' income are shown in table A1 in the appendix.

In the light of the coefficients reported in Table 6, intergenerational persistence of economic status in Italy appears to be high and significant.[14] The TS2SLS estimate of β is

[13] We exclude observations from households where more than one co-residing child is working and earning a positive income.

0.479 or 0.509 depending on whether or not we control for age of both fathers and sons. Broadly speaking, this indicates that about half of the economic advantage of Italian fathers is passed on to their children. As expected, the values obtained predicting incomes for both generations are lower: 0.333-0.339. Both pairs of values are of a similar magnitude of those found in the US by Bjorklund and Jantti (1997) using the same technique.[15]

Table 6.

Estimated Intergenerational Elasticity in Italy

Technique	uncorrected for age	corrected for age
1. TS2SLS	0.479 (0.076)	0.509 (0.071)
2. Predicted incomes	0.333 (0.059)	0.339 (0.059)
3. Co-residing	___	0.327 (0.082)

Notes: Bootstrapping standard errors are in parenthesis. Incomes are predicted by educational, occupational and geographical dummies.

I check these results against various sensitivity tests. Tables A2-A5 in the appendix show the results from a number of alternative regressions estimating the coefficients in rows 1-2. Choosing different years for fathers' income does not significantly alter its effect on sons' income (Table A2, upper panel). Coefficients do not drastically vary either when using sons from the 2000 survey wave (Table A2, lower panel). If anything, the estimates from the selected pair of years (2002-1977) are slightly lower than the

[14] To calculate standard errors, we use the bootstrap procedure. First, we draw a bootstrap first-stage sample of fathers, from which we estimate the parameters used to generate predicted incomes. Then a bootstrap sample of sons is drawn and used for running the second-stage regression on fathers predicted incomes. After repeating this process 1000 times, the bootstrap standard error is estimated by the standard deviation of the distribution of the bootstrap estimates.

[15] Their TS2SLS values are 0.417-0.516, while the estimates obtained by predicting both incomes are 0.294-0.327. Note that their corresponding values for Sweden (uncorrected for age) are 0.282 and 0.216 respectively.

averaged values across the alternative regressions. Results did not dramatically change for minor variations of the age at which we observe sons (Table A3) and were not sensitive to a rough correction for under-reporting of self-employed income (table A4).[16] Somewhat more variation is found when we test for different sets of predictors of income. However, even in a slightly larger range of values, the results still point to low level of overall mobility in Italy (Table A5). The fact that persistence is high even when education is not included in the predictors of father's income (row 2, in Table A5), excludes the possibility of a serious upward bias in the preferred estimate.

Returning to Table 6, the coefficient in row 3 is obtained form the estimation of equation (7) using the smaller sample of co-residing father-son pairs in 2002. We confirm our expectations of lower estimates when using single-year measures of fathers' income, relatively younger sons and older fathers. As a check, we construct a similar sample for the year 2000 and we find a very similar value (0.336). The estimates have to be treated with caution, however, given that the samples are of limited size and possibly unrepresentative of the reference population.

4.2. *Transition Matrix*

An alternative way to characterize intergenerational mobility is provided by transition matrices. This approach relies on discrete categorizations and investigates the conditional probabilities of transition among ordered income quantiles/groups. We construct four income classes for both fathers and sons: (i) "low-income", which includes individuals with income below two-thirds of the median; (ii) "lower-middle", from higher incomes up to the median; (iii) "higher-middle", for incomes from the median to 150% of the median, and (iv) "high-income", for the rest of the individuals. Table 7 gives the fraction of sons in each income class given the predicted class of their fathers.

The information in Table 7 enables to investigate about the direction and the pattern of mobility in a way that cannot be accomplished by mean regression measures.[17] We note the existence of a "wealth trap", with richer people being very likely to pass on their

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[16] Following Checchi and Dardanoni (2002), we revise income from self-employment upward by 40%. This is the discrepancy of self-employment figures with corresponding values based on national accounts averaged over the period 1980-93.

[17] Assuming a Markovian process, each cell in Table 7 can be interpreted as the probability for a son to be in class i^{th}, conditional on his father's being in class j^{th}.

economic status to their children. The probability to move up to the two highest classes from the bottom is 28.5% against a probability of 86.02% to be there from the top. Stated differently, transition from rags to riches is far from a credible possibility for most low-income Italians. Poorer individuals have higher chances to move upwardly in the lower part of the matrix so that the persistence in the first class is not particularly high. The overall picture emerging from the above matrix is consistent with the findings of Checchi *et al.* (1999) based on different data and a different analytical procedure.

Table 7.

Transition Matrix by Income Classes

Son / Father	Low-income	Lower middle	Higher middle	High-income
Low-income	20.14	51.37	19.87	8.63
Lower middle	11.59	47.38	29.32	11.71
Higher middle	11.75	35.12	26.98	26.14
High-income	2.83	11.15	38.07	47.95

Notes: Values expressed in percentages.

4.3. *Family background and educational attainments*

Considering the way higher education is financed in Italy, high levels of economic persistence may appear as a rather surprising result. In Italy, access to higher education by poor families is facilitated by a largely public system financed through taxation. We would then expect parental income not to be critical for children's educational attainments. Table 8 reports mean fathers' income (in logs) for five classes of sons' educational attainments. It shows the existence of a relationship between levels of education and fathers' income, with graduates having on average richer fathers. What can be said about the role of this relationship in the transmission of economic status in Italy? Is it possible to evaluate the income persistence through superior educational attainments of richer children?

Table 8.

Sons' educational attainments and fathers' income

Son's education	*Mean fathers' log income*
elementary school	9.30
lower secondary school	9.52
high school	9.76
bachelor	9.90

Notes: Author's own elaboration from survey data

For the US, Bowles and Gintis (2002) seek to uncover the channels through which parental incomes influence offspring incomes by decomposing the intergenerational correlation into additive components reflecting the contribution of various mechanisms. We follow their strategy to estimate the size of the "education channel" in the transmission of economic status. This will allow an assessment of how much of the intergenerational association is accounted for by the fact that richer parents have higher educated children. Suppose that fathers' income (Y_i^f) directly affects sons' income (Y_i^s), but sons' income is also affected by sons' education (E_i^s), which is correlated with fathers' income. It is a property of correlation coefficients to be decomposable into additive parts: the intergenerational correlation (or the intergenerational income elasticity) can be expressed as the sum of the standardized regression coefficients of fathers' income ($\beta_{Y_f Y_s}$) and children education ($\beta_{E_s Y_s}$) in a multiple regression predicting Y_i^s, each multiplied by the correlation between Y_i^f and the regressor (which for fathers' income itself is just 1).[18]

$$r_{Y_f Y_s} = \beta_{Y_f Y_s} + r_{Y_f E_s} \beta_{E_s Y_s}$$

[18] A standardized regression coefficient is the change in the dependent variable, in standard deviation units, associated with a one standard deviation change in the independent variable.

The education component of this decomposition ($r_{Y_f E_s} \beta_{E_s Y_s}$) of the intergenerational correlation is called an *indirect effect*, while the *direct effect* of fathers' income is the standardized regression coefficient of fathers' income from this regression ($\beta_{Y_f Y_s}$). Note that this decomposition should only be seen as a descriptive device along the lines suggested in Bowles and Gintis (2002) and not as an analysis of causal effects. We estimate the size of these direct and indirect effects by applying this decomposition to the estimate of the intergenerational regression coefficient (uncorrected for age).

0.48	=	0.345	+	0.135
(total)	=	(direct effect)	+	(indirect effect)

The above simple exercise shows that only a small fraction (less than one third) of the intergenerational coefficient is accounted for by the fact that the children of rich parents are also more educated. In other words, assuming that the only channel of intergenerational income correlation would work through the association of father's income and child's education, the income regression coefficient for our sample of Italian men would be equal to 0.135.

5. Concluding Remarks

Data limitations have restricted mobility studies on Italy: there is not a longitudinal survey that is long enough to provide information on actual incomes of both parents and children. Recent econometric innovations employed in the literature permit to use the Historical Database of the Bank of Italy Survey of Households' Income and Wealth and to overcome some of the data problems for Italy. Retrospective information in the repeated cross sections is exploited by applying a two sample two stage least squares estimation. The remaining limitations of the database are addressed by performing several auxiliary regressions.

When comparing the results of this study with those obtained from other countries, one has to be aware that cross-country comparisons of intergenerational mobility are very difficult. This is essentially due to the fact that international comparisons have much more severe data requirements than single-country studies. Comparability of results demands similar information on both fathers and sons' income, as well as special attention on sample selection rules applied to the datasets. If comparative studies do not attempt a parallel analysis for the countries involved then it is difficult to say whether differences in the intergenerational income persistence reflect true differences in mobility or are driven by different income measures, age ranges or other sample selection criteria. Looking at the results from existing studies that have used similar estimation procedures, the findings of this paper indicate that Italy displays levels of economic persistence that are most similar to United States and Britain. Italy appears to be markedly less mobile than Sweden.

More international evidence is available if we consider also the studies that used different data and methodologies. Corak (2004*b*) does a thorough review of the international literature regarding rich countries, with an explicit comparative perspective. He puts together a set of comparable estimates across a number of countries, taking into account the specifics of different studies' design. Although it is not possible to rank countries in a rigorous way according to their level of "overall" mobility, he concludes that among rich countries, Scandinavian countries and Canada are the most mobile societies. At the other extreme, the US and the UK stand out as being the least mobile societies, with 40% to 50% of fathers' income advantage being passed on to sons.

Taking account of the possible biases arising from the data and the procedure adopted, the evidence provided in this paper hints at Italy in the low-mobility group among advanced societies. Obviously, this evidence has to be considered suggestive, not conclusive, as richer data and different estimation methods have demonstrated to significantly improve the reliability of the estimates in other countries. However, my results are consistent with the image of Italy as a "rigid" society that had emerged from previous studies by both sociologists and economists.

Various tentative explanations of the low mobility in Italy can be suggested. We find that the inheritance process operating through superior educational attainments of those with well-off parents, while important, accounts for less than one third of the intergenerational transmission of economic status. This result gives credibility to the hypothesis that in Italy equally educated children have unequal chances depending on their family background. The fact that the Italian labour market is characterised by the presence of extensive barriers to entry into a wide range of occupations might play an important role.[19] The high standardization of the Italian higher education system can also be blamed, for it deprives poorer children of a tool to signal their ability when competing with children from richer families, who can benefit from parental connections. This result may appear puzzling to the extent that a public education system is expected to favour intergenerational mobility. As a matter of fact, countries can differ significantly in the impact that education spending may have on intergenerational mobility. It will depend on a larger set of "intangible" advantages richer parents are able to pass on to their children, which includes not only family connections but also beliefs and motivations.

[19] See Schizzerotto and Bison (1996).

APPENDIX

Table A1.

First-stage regression of pseudo-fathers income on four categorical variables (1977)

			N = 953
			R^2 = 0.3285
Variable	coefficient	robust st.err.	t
Education			
elementary school	0.181	0.078	2.34
lower secondary school	0.271	0.084	3.24
high school	0.474	0.094	5.06
bachelor	0.708	0.108	6.53
Work status			
office worker/teacher	0.135	0.044	3.05
manager/professional/entrepreneur	0.499	0.098	5.07
Self-employed	0.258	0.045	5.78
Sector			
industry	0.329	0.073	4.5
public administration	0.210	0.077	2.72
private services	0.360	0.078	4.6
Geographic dummy			
south	-0.199	0.036	-5.46
Const.	9.076	0.082	110.33

Notes: reference categories are: no education, blue collar, agriculture and north.

Table A2.

Two-sample estimates for different pairs of years for fathers and sons.

	Sons sample: 2002					
pseudo-fathers sample	1978		1979		1980	
	(a)	(b)	(a)	(b)	(a)	(b)
TS2SLS	0.505	0.531	0.457	0.477	0.502	0.525
	(0.066)	(0.058)	(0.061)	(0.067)	(0.076)	(0.071)
Predicted incomes	0.365	0.368	0.342	0.346	0.354	0.359
	(0.026)	(0.024)	(0.023)	(0.023)	(0.028)	(0.026)

	Sons sample: 2000					
pseudo-fathers sample	1977		1978		1979	
	(a)	(b)	(a)	(b)	(a)	(b)
TS2SLS	0.538	0.559	0.518	0.558	0.478	0.520
	(0.049)	(0.051)	(0.051)	(0.050)	(0.047)	(0.048)
Predicted incomes	0.338	0.345	0.324	0.337	0.302	0.317
	(0.021)	(0.022)	(0.023)	(0.023)	(0.019)	(0.020)

Notes: Incomes are predicted by educational, occupational and geographical dummies. (a) does not control for age; (b) includes control for age

Table A3.

Two-sample estimates for different age ranges for sons.

	fathers age: 30-50							
Sons age	27-42 (n=542)		30-40 (n=404)		33-48 (n=613)		35-50 (n=560)	
	(a)	(b)	(a)	(b)	(a)	(b)	(a)	(b)
TS2SLS	0.458 (0.067)	0.496 (0.063)	0.464 (0.079)	0.486 (0.077)	0.497 (0.058)	0.509 (0.061)	0.484 (0.066)	0.498 (0.061)
Predicted incomes	0.327 (0.024)	0.333 (0.025)	0.342 (0.031)	0.343 (0.029)	0.366 (0.024)	0.368 (0.023)	0.354 (0.027)	0.354 (0.027)

Notes: Incomes are predicted by educational, occupational and geographical dummies. (a) does not control for age; (b) includes control for age

Table A4.

Estimated Elasticities correcting for under-reporting of self-employed income (2002-1977).

Technique	uncorrected for age	corrected for age
1. TS2SLS (N = 612)	0.470 (0.063)	0.497 (0.062)
2. Predicted incomes (N = 612)	0.363 (0.027)	0.368 (0.027)
3. Co-residing (N=231)	___	0.349 (0.073)

Notes: Incomes are predicted by educational, occupational and geographical dummies.

Table A5.

Estimated Intergenerational Elasticities (2002-1977) for different sets of predictors of income.

Predicting variables	2S2SLS		Predicted Incomes	
	(a)	(b)	(a)	(b)
1. education, work status	0.510 (0.072)	0.546 (0.073)	0.264 (0.024)	0.271 (0.024)
2. work status, sector, area	0.444 (0.058)	0.468 (0.058)	0.210 (0.024)	0.213 (0.023)
3. education, work status, area	0.530 (0.066)	0.553 (0.066)	0.367 (0.029)	0.372 (0.029)
4. education, work status, sector	0.414 (0.062)	0.453 (0.061)	0.234 (0.021)	0.234 (0.021)
5. education, sector, area	0.525 (0.079)	0.556 (0.077)	0.393 (0.040)	0.391 (0.039)
6. education	0.594 (0.093)	0.642 (0.093)	0.301 (0.029)	0.305 (0.028)

Notes: (a) does not control for age; (b) includes control for age.

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