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INTERGENERATIONAL EARNINGS INEQUALITY: NEW EVIDENCE FROM ITALY

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Using an innovative dataset built by merging survey and administrative data, we provide new estimates of intergenerational earnings' inequality between fathers and sons in Italy, extending previous evidence in several directions. We rely on the TSTSLS method to predict fathers' earnings and compute intergenerational elasticities and imputed rank–rank slopes, trying to reduce estimation biases. Confirming previous evidence, we find that Italy is characterized by a high intergenerational inequality in cross-country comparison. Extending previous analyses, we show that the intergenerational association increases when sons at older ages and multi-annual averages of pseudo-fathers' and sons' earnings are considered. We also find that the intergenerational persistence differs across geographical macro-areas and is high also for daughters, especially when family earnings are considered. Furthermore, estimates where possible mediating factors of the parental influence are included among the covariates show that a high intergenerational association persists when sons' education and occupation are controlled for.

JEL Codes: J24, J31, J62

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1. INTRODUCTION

In the last few decades, a growing strand of the economic literature has investigated the transmission of socio-economic advantages from parents to offspring. Scholars have focused their attention mostly on measuring the degree of income persistence across two generations,[1] usually estimating the intergenerational income elasticity coefficient (henceforth IGE or β), which measures the extent to which income differences among parents persist among children.[2]

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[1]We use "income" and "earnings" as synonyms until Section 3, where the monetary indicator used in this article is defined.

[2]Due to data limitations and some methodological issues, most studies focus on father–son pairs and investigate the intergenerational association of labour earnings (Bjorklund and Jäntti, 2009; Blanden, 2013).

Even if reliable estimates of the IGE are available only for few countries, a generally accepted ranking has emerged for developed countries (Solon, 2002; Corak, 2013; Blanden, 2013): Nordic European countries are the most mobile, since they are characterized by relatively low values of the IGE, while the US, the UK, and Southern European countries are among the countries with the highest estimated values of the IGE ($\beta > 0.40$). These rankings are also confirmed by studies on EU countries that, instead of computing the intergenerational income association, have analyzed the correlation between parents' socio-economic characteristics and children's midlife earnings (Raitano and Vona, 2015a, 2015b).

Collecting data to estimate the income association between subsequent generations is a very demanding task, because panel datasets recording parents' and children's incomes over time are required. In particular, researchers should be able to observe both parents and children for more than one year, and not at early stages of children's working lives (when children are too young), in order to minimize the so-called attenuation and life-cycle biases (Haider and Solon, 2006; Bjorklund and Jäntti, 2009; Gregg *et al.*, 2017). However, long panel datasets covering subsequent generations in a proper way are available only in few countries (e.g. the UK and the US, where cohort panel surveys have been carried out since the 1950s, and Northern European countries, where detailed administrative longitudinal datasets are available).

Datasets collecting incomes for subsequent generations in Italy are not available. To estimate the IGE between fathers and sons in Italy, Mocetti (2007) and Piraino (2007) have relied on the two-sample two-stage least squares (henceforth, TSTSLS) method. This method allows researchers to estimate the association between parents' and children's incomes when parents' incomes are not recorded. This is achieved by exploiting repeated cross-sectional datasets where retrospective information on parents' characteristics are available. These characteristics are indeed used to predict parents' earnings through a sample of "pseudo-parents" observed when children were young (Bjorklund and Jäntti, 1997). Drawing on the TSTSLS method and using waves of the Survey of Household Income and Wealth (henceforth, SHIW) carried out by the Bank of Italy, Mocetti (2007) and Piraino (2007) computed point-in-time measures—observing pseudo-parents' and children's incomes in a single year—of the intergenerational elasticity of net earnings and found an IGE equal to 0.45–0.50. This IGE value places Italy among the least mobile developed countries.

Recent studies have suggested relying on a further measure of intergenerational persistence, the rank–rank slope, which captures the association between the relative positions (percentiles) of parents and children in their respective income distribution (Dahl and DeLaire, 2008, Chetty *et al.*, 2014). To the best of our knowledge, no studies have estimated rank–rank slopes for Italy so far.

Some studies have also pointed out that EU countries differ with regard to the role played by some factors—mostly education—that may mediate the association between parents' characteristics and children's incomes, (Raitano and Vona, 2015a, 2015b). Indeed, when children's education is controlled for, a significant association between parental background and children earnings persists in countries characterized by a high IGE—Italy, Spain, and the UK—whereas no significant residual association emerges for Northern European countries. This evidence suggests a potential role of further background related factors (e.g. unobservable abilities or social connections) in the intergenerational transmission process in the most unequal countries, while the relatively lower IGE of Northern European countries might be related mostly to differences in educational opportunities.

The economic literature has also argued that the intergenerational transmission process in a country might be heterogeneous by geographical areas (Chetty *et al.*, 2014, Chetty and Hendren, 2018a, 2018b) and children's gender (Chadwick and Solon, 2002, Cervini-Plà, 2015). However, no studies have investigated whether intergenerational inequality differs in Italy by regions and by gender.

As mentioned, only two major studies have estimated point-in-time IGE of net earnings between fathers and sons in Italy (Mocetti, 2007, Piraino, 2007). In this article, still applying the TSTSLS method to overcome the lack of information on parents' incomes, we aim at improving the knowledge about intergenerational earnings' transmission from fathers to sons in Italy, extending, along several directions, the evidence provided by these two seminal studies.

Foremost, we provide new estimates of the IGE of gross earnings by exploiting an innovative panel dataset, recently built merging survey and administrative data. The 2005 wave of the Italian component of the European Union Statistics on Income and Living Conditions (EU-SILC; henceforth, the Italian component is named IT-SILC), where respondents are asked retrospective questions about parents' characteristics, has been indeed combined with the social security longitudinal administrative records about each working spell since the entry in the labor market of all respondents interviewed in IT-SILC. Thanks to the panel dimension of our dataset, where the same individuals are observed for multiple years, we may compare point-in-time (yearly) estimates with estimates where sons and pseudo-fathers are followed for a 5-year period and sons are also observed in different age classes, thus reducing possible attenuation and life-cycle biases. Furthermore, using predicted fathers' incomes in the first stage, we obtain the parental income distribution and estimate rank–rank slopes. Differently from previous studies, we focus on fathers' and children's gross earnings (i.e. before tax redistribution) instead of net earnings, thus analyzing the intergenerational persistence due to labor market forces.

Moreover, we extend the analysis of the intergenerational inequality in Italy along three directions suggested by the aforementioned recent research. First, we estimate the association between sons' and fathers' incomes adding covariates representing some children's outcomes (i.e., education, occupation, experience), to observe whether the intergenerational persistence is mediated wholly by these intervening factors. Second, we compare the intergenerational association in the Italian geographical macro-areas, also controlling for sons' mobility across these areas. Finally, following the method proposed by Chadwick and Solon (2002) and Cervini-Plà (2015), we compare the intergenerational earnings' association between fathers and sons and fathers and daughters, considering both family and individual earnings as outcome variables, to take into account a possible employment selection bias by females.

In more detail, the paper is structured as follows. Section 2 reviews the empirical approaches proposed to estimate the association between parents' and children's incomes. Sections 3 and 4 present the dataset and the empirical strategy, respectively. Section 5 shows results obtained estimating the IGE and the rank–rank slope, using both point-in-time and multi-year averages of pseudo-fathers' and sons' earnings and observing sons at different ages. Section 6 extends the analysis, showing the role of possible mediating factors (Section 6.1), and the differences in the intergenerational association by macro-areas (Section 6.2) and children's gender (Section 6.3). Section 7 concludes, summarizing our main results.

2. How to Estimate the Intergenerational Income Association: Empirical Issues

Over the last two decades, economists have analyzed broadly the extent to which economic advantages are transmitted from one generation to the next, assessing these advantages by focusing on monetary indicators, usually either labor earnings (in most of the cases) or total incomes (when also information about other income sources is available). To evaluate the degree of intergenerational income elasticity, one should observe permanent incomes for both generations and estimate the following equation:

$$y_i^c = \alpha + \beta y_i^p + \varepsilon_i \tag{1}$$

where y_i^c and y_i^p are the logarithms of children's and parents' permanent incomes, respectively, and β measures how much of the income gap within parents persists within children (Bjorklund and Jäntti, 2009). Therefore, the higher the earnings' elasticity, the lower the degree of economic mobility across generations. However, scholars have to deal with major methodological issues to estimate equation (1).

First, even the few datasets covering two generations usually record incomes in a limited number of points in time (years), thus not providing a good proxy of permanent incomes. This implies that, under classical measurement errors' assumptions, estimated elasticities obtained using yearly, instead of permanent, parents' earnings are likely to be downward-biased due to the so-called attenuation bias (Solon, 1992; Zimmerman, 1992). To reduce this bias, parents' incomes should be averaged over as long a period as possible (Mazumder, 2005). Measurement errors in children's earnings, instead, do not bias the estimate of the IGE but lead to a loss of precision and larger standard errors.

Second, the lack of permanent earnings might also cause the so-called life-cycle bias if children's earnings are observed when they are too young. More specifically, estimated elasticities are influenced by the amount of income dispersion, which tends to rise as children get older, since earnings' profiles are steeper for those with higher permanent earnings. To minimize this bias, since permanent measures of earnings are not available, empirical studies suggest focusing on men in their median age (around 35–45; Haider and Solon, 2006), whereas a simple rule does not emerge for women, as they display more variety in their life-cycle income profiles (Bohlmark and Lindquist, 2006).

Finally, and most importantly, because of the lack of either long surveys or administrative panel data covering subsequent generations, information about parents' incomes is absent in most developed countries and in almost all developing countries.

The TSTSLS method—developed starting from the two-sample instrumental variables estimation (Angrist and Krueger, 1992)—allows researcher to overcome this limit (Bjorklund and Jäntti, 1997; see Jerrim *et al.*, 2016, for a review of the studies carried out by using the TSTSLS approach). This method exploits two separate samples observed in different time periods: i) a sample of adult children, where their incomes and the socio-economic characteristics of their parents (e.g. education and occupation) are recorded; ii) a sample of "pseudo-parents"—i.e. individuals who are not the real parents but are observed in their middle-age during the childhood of the children— where incomes and socio-economic characteristics are recorded. In the first stage of the TSTSLS approach, the pseudo-parents' sample is used to estimate an earnings' equation using socio-economic characteristics as explanatory variables; in the second stage, the prediction of parents' incomes obtained through this earnings' equation is added to the children sample, replacing missing parents' earnings with the best linear prediction and allowing researchers to estimate the IGE.

In formal terms, in the first stage, the sample of pseudo-parents — observed at time *t-m* — is used to run the following equation:

$$y_{i,t-m}^{pp} = \alpha + \theta_1 Z_i^{pp} + v_{i,t-m} \quad (2)$$

where $y_{i,t-m}^{pp}$ are log earnings of pseudo-parents' and Z_i^{pp} is a vector of their time-invariant socio-economic characteristics (e.g. education, occupation). The estimated coefficient $\hat{\theta}_1$ is then used to predict missing parents' earnings in the children sample by merging the two samples according to the characteristics of real parents (e.g. education, occupation) reported by the children in their sample. The IGE β is then estimated in the second stage according to the equation:

$$y_{i,t}^{c} = \alpha + \beta \hat{y}_i^{p} + \varepsilon_{i,t} \quad (3)$$

where $y_{i,t}^{c}$ is the log of children's earnings and $\hat{y}_i^p = \hat{\theta}_1 Z_i^{pp}$ is the prediction of the log parents' earnings obtained in the first stage.

Available retrospective variables may influence the estimated IGE (Olivetti and Paserman, 2015). Since $0 \leq R^2 \leq 1$, the variance of parent's predicted earnings is either lower than or equal to the variance of actual parent's earnings and $\hat{\beta}_{TSTSLS} = \hat{\beta}_{OLS}$ when $R^2 = 1$ in the first-stage regression. However, it is impossible to obtain a perfect prediction of fathers' earnings by using the set of socio-economic characteristics available in empirical analyses. Therefore, as clarified by Olivetti and Paserman (2015), the TSTSLS estimator might be affected by two sources of biases with respect to the $\hat{\beta}_{OLS}$ estimator obtained by using fathers' actual earnings (see Appendix B for a formalization): i) a downward bias, due to the unpredicted share of fathers' income positively correlated with sons' income; ii) an upward bias, since the variance of predicted fathers' earnings is lower than the actual variance. Both types of bias increase if the share of fathers' earnings that is not predicted in the first stage is positively correlated with unobservable variables correlated across generations (e.g. soft skills, social networks, preferences). On the contrary, in a given sample, when the number of variables associated with earnings in the first stage rises, R^2 increases and both biases reduce (Nicoletti and Ermisch, 2007), even if quantifying the amount of reduction in each bias is not possible.

Apart from these major issues, recent studies have suggested relying on a different measure of intergenerational persistence, the rank–rank slope, which captures the association between the percentiles of parents and children in their respective income distribution and, differently from the IGE, is not affected by extreme income values and by the change in inequality across the two generations (Dahl and DeLaire, 2008; Chetty *et al.*, 2014). Actually, the IGE captures both the degree to which the position of sons in the income distribution is related to that of their fathers (re-ranking across generations) and the degree of inequality in each generation, whereas the rank–rank coefficient captures only relative positions. Furthermore, Chetty *et al.* (2014) point out that the relationship between mean children's ranks and parents' ranks is almost perfectly linear and much more robust across specifications, and they suggest relying mostly on rank–rank slopes when comparing intergenerational inequality across geographical areas within a country.

3. Data and Sample Selection

We use the AD-SILC dataset (where the "AD" stands for "administrative), built merging (using fiscal codes as matching key) the 2005–2010 waves of the IT-SILC survey with information collected from administrative archives managed by the Italian Social Security Institute (INPS) that record employment and earnings histories of all individuals working in Italy from the moment they entered the labor market up to the end of 2013. In other terms, the AD-SILC dataset enriches the IT-SILC cross-sectional waves with the whole longitudinal working histories of the individuals sampled in IT-SILC.

For our purposes, this dataset is very rich, since it matches individuals' characteristics – concerning the pseudo-parents, the adult children, and their real parents (recorded through retrospective questions made to the children)—to detailed longitudinal information about the earnings of sons and pseudo-fathers (see Table 1

TABLE 1
Description of the Main Variables Available in the AD-SILC Dataset, According to their Source

Source of the Variable	Variable Description
IT-SILC cross-sections	In all 2005–2010 waves: Year of birth; Age; Gender; Education (ISCED level); Occupation (2-digit ISCO) In the 2005 wave: Father's Education (ISCED level); Father's Occupation (2-digit ISCO); Father's main activity status (e.g., employee, self-employed, retired)
INPS archives (tracking individuals every year since their entry in the labor market up to 2013)	Year of birth; Age; Gender; Region of birth; Region of residence; Pension Fund where the worker pays contributions (distinguishing employees in the private sector, in the public sector, and the various self-employed categories); Annual worked weeks; Experience in the labor market (computed by adding annual worked weeks since the entry); Annual gross earnings

for the list of the variables included in AD-SILC useful for our purposes, distinguished according to the source; i.e. IT-SILC survey or administrative archives).

In more detail, on the one hand, the administrative archives record individuals' demographic characteristics (gender, year of birth, region of birth, and residence) and provide detailed information on every job spell that individuals experience during a year; e.g., the duration (measured in weeks), gross earnings (including personal income taxes and pension contributions paid by the worker), and the fund into which workers' pay contributions (which allows us to distinguish public and private employees and the various groups of self-employed: farmers, dealers, craftsmen, professionals, and dependent self-employed, i.e. those working as self-employed in legal terms but "economically dependent" on a single client). Hence, the panel structure of our data allows us to compute the multi-year averages of individual earnings of both children and pseudo-parents and to measure exactly the time of entry to the labor market and the effective labor market experience since entry.

On the other hand, the 2005 wave of IT-SILC contains a specific module in which variables about parents' features when children were aged approximately 14 are recorded retrospectively through questions made to the children. These questions record parents' education (coded through the ISCED-97 classification), occupation (coded through the 2-digit ISCO-88 classification), and main activity status (which distinguishes employees and self-employed). The 2005–2010 IT-SILC waves also include other information absent from administrative archives, such as educational attainment and details of occupation.[3]

Using the AD-SILC dataset, we select two subsamples, one for the children and the other for the pseudo-parents (note that, even if both generations are extracted by the same panel dataset, actual parents are not included in the AD-SILC dataset, except for the very select sample of adult children who still co-reside with their parents). The sample selection rules are as follows.

First, except for the analyses shown in Section 6.3, we focus on father–son pairs, consistent with most studies on intergenerational mobility.[4]

Second, we define the sample of adult sons by extracting individuals born in the period 1970–1974 from the 2005 wave of IT-SILC. Exploiting information available in the administrative archives, we follow them for five years, from age 35 to age 39 (i.e. during the period 2005–2013, according to their year of birth). Earnings of sons in the age class 35–39 are also averaged over the 5-year period to get a proxy of multi-year earnings.[5]

[3]Administrative archives record occupations for private employees only, distinguishing apprentices, blue-collar workers, white-collar workers, and managers.

[4]Children and pseudo-fathers without Italian citizenship are excluded from the samples.

[5]In our baseline estimates we average only positive annual earnings. In additional specifications we also compute 5-year sons' average earnings, including observations of individuals absent in the administrative archives in a year; i.e. assuming a zero yearly earning for that year. Note that the absence in a year cannot be due to attrition (all workers are tracked by INPS archives), but it corresponds to a whole year spent without working in Italy. However, an absence depends on reasons—e.g. long-term unemployment, informal work, voluntary inactivity, and migration abroad—not distinguishable in our dataset. By contrast, periods spent receiving sickness, maternity, or temporary job suspension (*Cassa Integrazione*) allowances are recorded as working periods in our dataset.

Third, pseudo-fathers are selected among those individuals—interviewed in the 2005–2010 waves of IT-SILC — born in the period 1940–1944. Exploiting longitudinal information available in the administrative archives, pseudo-fathers are followed for five years, from age 40 to age 44, during the period 1980–1988, i.e. when the sons in the sample were in their youth. Earnings of pseudo-fathers in the age-class 40–44 are also averaged over the 5-year period.

Accordingly, we observe both generations in their middle-age to cope with the issue of life-cycle bias, as suggested by Haider and Solon (2006), and we use multi-year averages of yearly earnings to produce a better estimate of fathers' permanent earnings in the first stage. Note, instead, that the SHIW used by Mocetti (2007) and Piraino (2007)—being made of repeated cross-sections (every year from 1977 to 1984, and every 2 years afterwards), with a small and short panel component—constrains researchers to rely on point-in-time estimates of the intergenerational persistence.

Our outcome variable, for both fathers and sons, is gross earnings, from either employment or self-employment, converted into real terms using the Consumer Price Index (we do not consider the top and bottom 1 percent of the earnings' distribution for both pseudo-fathers and sons). In our baseline specifications, we rely on 5-year mean earnings for both generations, but we also compute measures of intergenerational persistence using annual earnings for pseudo-fathers and sons (observed in 1985 and 2009, respectively). Differently from the studies based on SHIW data that record net earnings, we thus estimate the intergenerational association of gross earnings, which provides a more accurate measure of the intergenerational earnings' transmission engendered by markets before tax redistribution. Note also that we make use of earnings recorded in administrative archives, which are, by their nature, less affected by measurement errors than are the self-reported earnings included in surveys.

Descriptive statistics presented in Table 2 show that the two final samples include 1,468 sons and 2,742 pseudo-fathers, respectively, and that gross annual real earnings are higher and more dispersed in the sample of sons than in the sample of pseudo-fathers.[6] Note that the analyses for Italy carried out by Mocetti (2007) and Piraino (2007) were based on less homogenous samples: in their baseline estimates, Mocetti (2007) uses a sample of 4,903 pseudo-fathers and 3,216 sons, both in the 30–50 age-band, and Piraino (2007) uses a sample of 3,015 pseudo-fathers born between 1927 and 1949 (observed when aged 30–50) and 1,956 sons aged 30–45. The larger size of the samples used by Mocetti (2007) and Piraino (2007) is then due to the much larger age classes of both pseudo-fathers and sons considered in their articles.

[6]The size of the pseudo-fathers' sample is larger than is that of the sons' sample since, as mentioned, pseudo-fathers are extracted from the 2005–2010 waves of IT-SILC, whereas we extract sons only from the 2005 wave.

TABLE 2
Descriptive Statistics of Sample of Sons' and Pseudo-Fathers' (Mean Values, Standard Deviations in Parenthesis)

	Sons	Pseudo-Fathers
Age (5-year Mean)	36.97	41.97
	(0.44)	(0.44)
Gross 5-year Earnings (Euros)	26,236	22,388
	(14,344)	(10,886)
Log Gross 5-year Earnings	10.00	9.90
	(0.66)	(0.52)
Observations	1,468	2,742

Note: Earnings deflated by using the consumer price index.
Source: elaboration on the AD-SILC dataset.

4. Empirical Strategy

We apply the TSTSLS method described in Section 2 to obtain measures of intergenerational earnings' association, using the set of retrospective information about fathers' characteristics available in our dataset to predict fathers' gross earnings by making use of the sample of pseudo-fathers observed approximately 25 years before the period when sons are observed.

In more detail, before estimating the first-stage regression in the sample of pseudo-fathers, we predict the residuals of a regression of pseudo-fathers' log gross earnings (annual or 5-year average, according to the earnings' measure considered) on their birth cohort dummies to get rid of the transitory component of earnings related to age. The estimated residuals are then used as a dependent variable in the first-stage regression on individual time-invariant characteristics available in the dataset.

The characteristics considered in the first stage to predict fathers' earnings are: education (coded through four categories: at most primary, lower secondary, upper secondary, and tertiary), occupation (identified through 27 occupational categories, according to the 2-digit ISCO-88 classification),[7] main activity status (a dummy distinguishing employees from self-employed), and 20 regions of residence.[8]

The IGE in the second stage is then obtained by regressing the log of sons' earnings (annual or 5-year average according to the earnings' measure considered) on the log of fathers' predicted earnings obtained by making use of the same

[7]We use fathers' occupation (coded through the 2-digit ISCO-88) as a time-invariant variable and associate it with the variable about the current (or past, in the case of retirees) occupation of pseudo-fathers recorded in IT-SILC. We prefer to rely on this information instead of using the time-varying occupational classification available in INPS archives that, as mentioned, has only four categories and is only available for employees in the private sector. Note that occupation is rather constant for most workers over their individual careers (except for the starting years), and our results do not change by much if we use the occupation recorded by INPS in the first-stage regression. Detailed results are available upon request.

[8]For comparison, note that we consider as covariates in the first-stage regression four educational categories, 27 occupations, two types of activity, and 20 regions of residence, whereas Mocetti (2007) includes five educational groups, six occupations, four sectors of activity, three geographical areas, and age, and Piraino (2007) includes five educational groups, four occupations, four sectors of activity, and two geographical areas.

covariates used in the first stage and controlling for sons' birth cohort dummies.[9] The distributions by education, occupation, type of activity and region of residence of samples of pseudo-fathers' and actual fathers' (according to retrospective answers made by sons) are rather similar (see Table A1 in the Appendix).

Our dataset allows us to take into account some limits of estimates of the intergenerational association highlighted in Section 2. Indeed, we can consider both fathers' and sons' 5-year average of annual earnings, thus reducing the influence of transitory shocks. Furthermore, longitudinal features of our dataset allow us to compare estimates for sons observed at different ages—at the ages of 25–29 and 35–39—thus providing an empirical measure of the life-cycle bias emerging if sons are observed when they are still too young.

As mentioned, we also estimate rank–rank slopes. Consistently with Olivetti *et al.* (2018), we estimate an "imputed rank–rank slope" by considering percentiles of the predicted distribution of fathers' (annual or 5-year average) earnings obtained by the first-stage regression (in other terms, we do not run a different first-stage regression to estimate percentiles as the dependent variable, but we build percentiles starting from the earnings prediction obtained in the first stage). We then regress percentiles of sons' earnings on percentiles of fathers' predicted earnings. Both fathers' and sons' percentiles are computed with respect to their birth cohort.

From a statistical point of view, it is not easy to understand to what extent our imputed rank–rank slope can be compared to rank–rank slopes obtained by percentile ranking actual fathers' earnings. Obviously, when we impute the percentile of the father through a predicted variable, we may incur some errors in placing all fathers in the right percentile of their earnings' distribution. Hence, our estimates are likely to be affected by attenuation bias. However, this kind of positional measurement error cannot be intended as "classical" (Nybom and Stuhler, 2017). That is why we should exercise caution when comparing our imputed rank–rank slope to estimates obtained in previous studies for other countries.

As mentioned in the introduction, we also extend our analysis in three directions.

First—even if we do not propose a strategy for identifying mechanisms of intergenerational persistence in a causal sense—we inquire whether the association between fathers' and sons' earnings is mediated wholly by a positive link between fathers' earnings and some sons' traits that their economic returns positively depend on. Therefore, we inquire whether a residual association between fathers' and sons' earnings still emerges when possible mediating factors are controlled for. To this end, we run our second-stage regression, including among the covariates sons' labor market outcomes that may drive the intergenerational earnings association; specifically, we run three additional specifications, adding, step by step, i) education; ii) occupation (summarized by the 2-digit ISCO categories and the specific pension fund of the worker, which allows us to distinguish public and private employees and the various self-employed groups); iii) weeks of experience (and the

[9]We link the retrospective father's sector of activity reported in IT-SILC to the pseudo-father's pension fund, that allows us to exactly distinguish employee and self-employed, and we link sons' region of birth to pseudo-fathers' region of residence.

square value) since the entry in the labor market, which allow us to control for a possible influence of parents' characteristics on sons' employability.

Second, following Chetty *et al.* (2014), who notice that the US is characterized by large heterogeneities in rank–rank slopes across areas, we focus on heterogeneities in intergenerational persistence by sons' geographical area of birth. To this end, we compute IGE and rank–rank slopes between fathers and sons within the five Italian geographical NUTS-1 macro-areas, also controlling for sons' mobility across areas.[10]

Third, we compare intergenerational persistence for sons and daughters. Except for a few studies,[11] estimates of intergenerational income elasticities have usually been carried out only for males to get rid of participation constraints, which are particularly cumbersome to address for females. Estimating intergenerational income persistence for daughters is not an easy task, since a non-casual selection of females in the labor force might bias estimates (e.g. if females coming from a better background are more likely to participate in the labor market). To deal with this potential bias, following suggestions by Chadwick and Solon (2002) and Cervini-Plà (2015), we use, for daughters and sons, both individual and couples' earnings (obtained by the adult child and by his/her co-residing partner) as the dependent variable of the second-stage regression.

5. Estimates of the Intergenerational Earnings' Association

5.1. *Estimates of the Intergenerational Earnings' Elasticity*

We run the TSTSLS procedure to compute the IGE on both annual and 5-year average earnings, for fathers and sons, to observe how the estimated β changes when multiple individual observations are considered instead of point—in-time observations (results of the first-stage regression of pseudo-fathers' 5-year average earnings are shown in Table A2 in the Appendix).[12]

Table 3 (top panel) presents the estimated values of the IGE in various specifications when sons are observed at ages 35–39. Our preferred estimate is obtained when both generations are observed for 5 years (Column 1). We estimate a β coefficient equal to 0.501, which is in line with estimates by Mocetti (2007) and Piraino (2007), who find, in their favored specifications, an IGE equal to 0.500 and 0.435, respectively.[13] However, both authors carried out point-in-time estimates on net earnings, and these two aspects—due to the attenuation bias and to tax

[10]Consistently with Chetty *et al.* (2014), fathers' and sons' are ranked according to their (cohort specific) national earnings' distribution. Note that we do not have information on actual fathers' area of birth in our dataset (fathers' area of residence is imputed from sons' area of birth).

[11]See Chadwick and Solon (2002), Hirvonen (2008) and Cervini-Plà (2015), referring to the US, Sweden, and Spain, respectively, and Jäntti et al. (2006) and Raaum *et al.* (2008), who compare Nordic countries, the UK, and the US.

[12]The R^2 of our first-stage regressions is more than 0.40 in all specifications, whereas it does not exceed 0.32 in the previous estimates for Italy by Mocetti (2007) and Piraino (2007). Note also that all second-stage regressions are run through 1,000 bootstrap replications.

[13]The IGE does not change—the estimated coefficient amounts to 0.500—when we select pseudo-fathers among those born in the 1945–1949 cohorts (observed when aged 35–39), to avoid a possible bias in the pseudo-fathers' sample due to the effects of the Second World War on birth and survivorship rates.

TABLE 3
IGE Between Sons' and Fathers' Gross Earnings

	Observation Span of Fathers' and Sons' Earnings				
					5 years-5 years
	5 years-5 years	5 years-1 year	1 year-5 years	1 year-1 year	Imputing Zeros[a]
Prime age sons (aged 35-39)					
Father's earnings	0.501***	0.438***	0.398***	0.379***	0.577***
S.E.	[0.055]	[0.055]	[0.055]	[0.055]	[0.076]
Obs	1468	1367	1468	1367	1468
R^2 first stage	0.410	0.410	0.403	0.403	0.410
Young sons (aged 25-29)					
Father's earnings	0.270***	0.225***	0.233***	0.237***	0.303***
S.E.	[0.059]	[0.080]	[0.056]	[0.077]	[0.086]
Obs	1395	1151	1395	1151	1395
R^2 first stage	0.410	0.410	0.403	0.403	0.410

[a]The 5-year average of sons' earnings is computed assigning a zero value to sons who do not report earnings in administrative archives in a certain year.
***$p < 0.01$.
Note: IGE stands for Intergenerational Elasticity.
Source: elaborations on AD-SILC dataset.

redistribution (tax progressivity reduces the net incomes of well-off workers relatively more; well-off workers are, on average, sons of richer fathers)—should decrease the estimated IGE. Indeed, when we reconstruct net earnings, the value of the estimated β reduces to 0.430, a value slightly lower than those found in previous estimates for Italy.[14]

Note, however, that our results are not perfectly comparable with those of Mocetti (2007) and Piraino (2007), since, as mentioned, they considered pseudo-fathers and sons belonging to a much wider range of age classes than those considered in this paper. Moreover, following Olivetti and Paserman (2015), the lower number of covariates used in the first stage in the two previous studies for Italy might increase the upward bias in the TSTSLS estimate.

Nevertheless, our estimates confirm that Italy is a relatively immobile country within the group of developed countries, with IGE values similar to those estimated for the UK and the US, while the IGE is lower than 0.20 in Norway, Denmark, and Finland (Corak, 2013). Similar values of the intergenerational association between fathers and sons are found by Cervini-Plà (2015) for Spain, with an estimated IGE between 0.39 and 0.46, according to the age of adult sons.

As expected, the IGE reduces when annual earnings are considered instead of multi-year averages (Columns 2–4). In particular, when we observe fathers and sons in a single year, the estimated β reduces substantially and becomes 0.379 (Column 4) — a value that is fairly lower than is that estimated by previous analyses for Italy (Mocetti, 2007, Piraino, 2007)—and further reduces to 0.340 when we reconstruct net earnings. Although the two generations are taken at middle-ages, as proposed by Haider and Solon (2006), our findings suggest that using a single year measure of earnings may cause a downward bias in the estimated IGE, due to both left-hand side and right-hand side measurement errors (Jerrim *et al.*, 2016; Gregg *et al.*, 2017).

A better parental background might be associated with sons' economic outcomes, since it may affect both the probability of achieving high-paying jobs or of unemployment risks. Annual earnings allow us to consider both influences when children earn a positive wage during that year. However, if parents' earnings were negatively associated with the probability of long-term unemployment—i.e. zero working weeks over a year—estimating the IGE on individuals with positive earnings only would produce a downward bias. In our dataset, we do not know the reason why an individual has a zero income in a given year (e.g. voluntary inactivity, informal work, or unemployment), even if zero annual earnings are associated, presumably, with involuntary unemployment, since we have excluded from the sample those sons without positive earnings over the 5-year period.[15] Interestingly, when zero-earnings' years are included in the computation of sons' 5-year mean

[14]Our data do not allow us to precisely compute net earnings since detailed information needed to compute tax deductions and exemptions is not available (e.g. household composition is recorded only in the years when IT-SILC waves are carried out and information about other income sources and deductible expenses is not available). Thus, we reconstructed net earnings approximately by applying tax rates and social contribution rates paid by the worker.

[15]In the sample of sons, 90.7 percent of individuals have positive annual earnings over the whole 5-year period, 5.2 percent have four positive earnings, and 4.1 percent have, at most, three positive annual earnings. Note that our results do not change if we exclude from the sample sons with, at most, three positive annual earnings.

earnings, the IGE increases considerably and becomes equal to 0.577, an extremely high value (Column 5). This increase in the estimated IGE suggests that sons of poorer fathers are likely to have more unstable working careers than are sons of richer fathers. However, note that, as pointed out by Chetty *et al.* (2014), IGE estimates are highly sensitive to the treatment of children with zero incomes.

As mentioned, our dataset also allows us to obtain a direct measure of changes of the estimated IGE when younger sons are observed (Table 3, bottom panel). Confirming the existence of a life-cycle bias, the β largely reduces in all specifications if the same sample of sons is observed when individuals are aged 25–29 instead of 35–39. In our preferred specification, when both fathers and sons are observed for 5 years (not including possible zero annual earnings), the IGE reduces from 0.501 to 0.270.

5.2. *Estimates of Rank–Rank Slopes*

After estimating predicted fathers' earnings through the first-stage regression, we group fathers and sons according to percentiles of their respective earnings' distribution (conditional on their birth cohort) and estimate the rank–rank slope by regressing sons' percentiles on fathers' percentiles using OLS. Estimates are carried out for the same specifications shown for the IGE, thus considering both multi-year averages and point–in-time earnings and including zero annual earnings to compute 5-year averages in an additional specification.

As expected, since we focus only on the association between ranks without considering the extent of earnings' dispersion (and also due to possible errors when imputing fathers' rank), the estimated values of rank–rank slopes are much lower than are the estimated values of the IGE (compare Tables 3 and 4). In our preferred specification, where fathers' and sons' earnings are averaged over the 5-year period, the value of the rank–rank slope is 0.260, meaning that, on average, the positions of sons of two fathers with a gap of 10 percentiles differ by 2.6 percentiles.[16]

Capturing only relative positions, rank–rank slopes should be less affected by attenuation and life-cycle biases than is the IGE (Chetty *et al.*, 2014; Gregg *et al.*, 2017). Accordingly, on the one hand, the decrease in estimated rank–rank slopes when either sons or parents are observed over a single year, rather than 5 years, is lower than is the reduction observed in the estimated IGE, since measurement errors and transitory shocks mainly cause scale mis-measurement instead of positional inaccuracy in the earnings' distribution (Table 4, top panel). On the other hand, estimated rank–rank slopes decrease considerably when sons aged 25–29, rather than 35–39, are considered (Table 4, bottom panel); however, the extent of the decrease related to sons' age is lower than is the estimated reduction in the IGE, suggesting that son's ranking is associated with father's ranking at the beginning of the working career, when the income dispersion between those coming from more and less advantaged backgrounds is, in any case, lower.

Finally, it must be pointed out that the estimated value of rank–rank slopes does not change when zero annual earnings are taken into account to compute

[16]Rank–rank slopes remain rather constant when net incomes are considered, since the procedure followed to compute net values produces a limited re-ranking across workers. The estimated rank–rank slope remains also constant (0.259) when pseudo-fathers are selected among those born in 1945–1949.

TABLE 4
RANK–RANK SLOPE BETWEEN SONS' AND FATHERS' GROSS EARNINGS

	Observation Span of Fathers' and Sons' Earnings				
					5 years-5 years
	5 years-5 years	5 years-1 year	1 year-5 years	1 year-1 year	Imputing Zeros[a]
Prime age sons (aged 35-39)					
Father's earnings	0.260***	0.236***	0.230***	0.216***	0.255***
S.E.	[0.024]	[0.027]	[0.026]	[0.028]	[0.025]
Obs	1468	1367	1468	1367	1468
R^2 first stage	0.410	0.410	0.403	0.403	0.410
Young sons (aged 25-29)					
Father's earnings	0.179***	0.184***	0.167***	0.178***	0.146***
S.E.	[0.026]	[0.030]	[0.027]	[0.030]	[0.025]
Obs	1395	1151	1395	1151	1395
R^2 first stage	0.410	0.410	0.403	0.403	0.409

[a]The 5-year average of sons' earnings is computed assigning a zero value to sons who do not report earnings in administrative archives in a certain year.
***$p < 0.01$.
Source: elaborations on AD-SILC dataset.

5-year averages (Table 4, Column 5). Consistent with findings of Chetty *et al.* (2014) for the US, this suggests that zero-earning years mostly characterize individuals who have low earnings when working, thus supporting the idea that zero earnings do not depend on temporary voluntary choices by well-off workers.

Differently from the IGE, comparing the estimated rank–rank slope with values found for other countries is not easy, since the use of this alternative index is recent and, to the best of our knowledge, no study has so far computed rank–rank slopes on fathers' percentiles by using the prediction of fathers' earnings obtained through the TSTSLS approach. Nevertheless, the value estimated in our preferred specification for rank–rank slope (0.260) is not very high in cross-country comparison. For instance, for the US, Chetty *et al.* (2014) and Mazumder (2015) estimated a rank–rank slope of 0.34 and 0.40, respectively, whereas Gregg *et al.* (2017) estimated a 0.34 rank–rank slope for the UK; Bratberg *et al.* (2017) estimated rank–rank slopes for different countries, obtaining values equal to 0.383 for the US, 0.257 for Germany, 0.233 for Norway, and 0.215 for Sweden, and rank–rank slopes that are much lower than is that estimated for Italy were found by Chetty *et al.* (2014) for Denmark (0.180) and Canada (0.174).

6. Further Analyses of the Intergenerational Earnings' Association

6.1. *The Role Played by Mediating Factors*

The mainstream economic view of intergenerational inequality focuses on the key role played by family background in the accumulation of human capital, usually proxied by education in empirical studies (Becker and Tomes, 1979, 1986; Solon, 2004). Differences in earnings and occupational attainments are usually explained by means of differences in human capital by individuals coming from different backgrounds. However, intergenerational inequality might also depend, especially in non-competitive markets, on rewards of background-related traits that are different from human capital, on social connections above all (Franzini *et al.*, 2016), especially in countries, such as Italy, where family ties seem to play a role in determining occupational achievements and earnings (Pellizzari, 2010).

In this section, we do not aim at identifying whether human capital or social connections drive the intergenerational transmission process, since we do not have exhaustive proxies of both factors at our disposal. However, we are interested in verifying whether statistically significant IGE or rank–rank slopes still emerge when we consider possible mediating factors (e.g. sons' education and occupation) of the association between fathers' and sons' incomes. In other terms, if the intergenerational association was due only to the role played by these factors, it would disappear when they are controlled for.

Interestingly, when we run our preferred specification on 5-year average earnings for both pseudo-fathers and sons, adding sons' education to the covariates in the second stage, estimated IGE and rank–rank slopes are reduced only slightly (Figure 1): the IGE decreases from 0.501 to 0.411, a high value in cross-country comparisons, and the rank–rank slope reduces from 0.260 to 0.213. Therefore, a large and significant residual correlation between fathers' and sons' earnings still emerges when the main mediating factor, i.e. education, is controlled for.

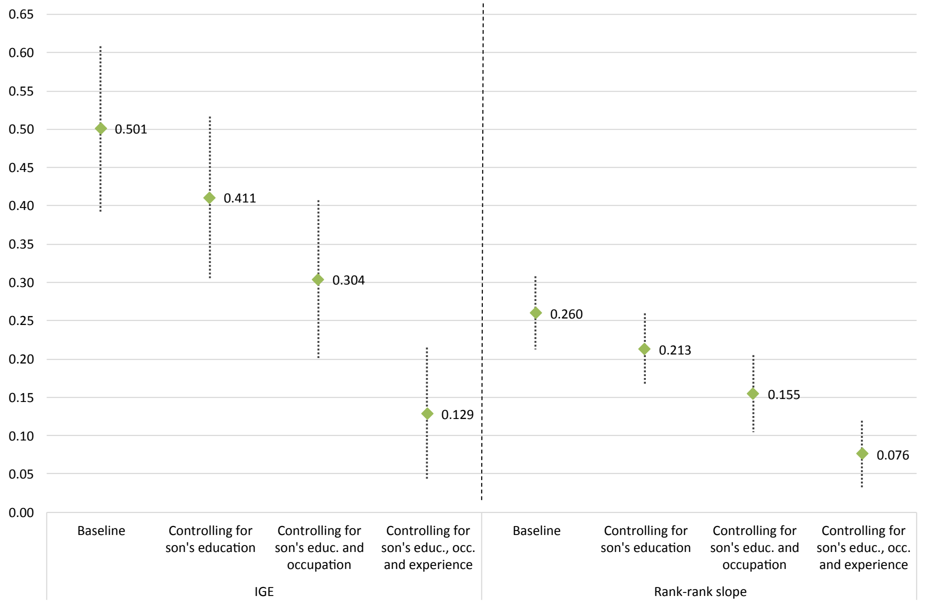

Figure 1. IGE and Rank-Rank Slopes Estimates Including Sons' Outcomes among the Covariates. Fathers and Sons Observed for 5 years. 95% confidence intervals.
Note: IGE stands for Intergenerational Elasticity.
Source: elaborations on AD-SILC dataset. [Colour figure can be viewed at wileyonlinelibrary.com]

Hence, one may ask what drives the existence of a "residual" intergenerational association within sons with the same education. A more advantaged background might be associated with higher cognitive and non-cognitive abilities that are not captured by the dummies on educational attainment. Actually, children with a better background often benefit from higher-quality education (Bratsberg *et al.*, 2007) and more extra-schooling activities (Duncan and Murnane, 2011), are advantaged in early-age skill formation (Cunha and Heckman, 2007), and have a more profitable endowment of soft skills, which are increasingly rewarded in the labor market (Bowles *et al.*, 2001).

Hence, education might be a poor proxy of individual skills. Educational attainment being equal, workers endowed with better background-related skills might get higher earnings, since these skills (unobservable in usual datasets, but observable by the employers) allow them to achieve better occupations (however, this result could also be attained by having better social connections). Indeed, if the intergenerational association were mostly led by skills observable by the employer, it should reduce significantly when sons' occupation is controlled for. By contrast, we find that, also adding detailed occupational dummies (on the 2-digit ISCO and the type of activity), the estimated IGE and rank–rank slopes remain large and statistically significant (0.304 and 0.155, respectively; Figure 1).

Both measures of intergenerational association decrease considerably, even if they remain statistically different from zero, when we add experience and its squared value to the covariates: the IGE and rank–rank slopes become equal to 0.129 and 0.076, respectively (Figure 1). On the one hand, such reduction may suggest that experience—a major driver of earnings' inequality in Italy, where many contractual arrangements link wages to tenure—is a proxy of unobservable abilities, since greater experience increases workers' productivity through learning by doing, and individuals with better abilities accumulate greater experience and have a lower risk of unemployment over their career. On the other hand, this evidence is not enough to argue that the intergenerational effect is related wholly to background-related abilities, since unemployment risks might be influenced by parental features through social connections.

6.2. *Intergenerational Association by Geographical Macro-Area of Birth*

Italy is a country characterized by large internal geographical differences. According to computations on IT-SILC data for 2005, for instance, annual labor incomes of males aged 35–39 living in the South or in the Islands were 22.1 percent and 16.8 percent lower than those living in the North or the Centre, respectively, and these gaps enlarge up to 31.5 percent and 24.1 percent when equivalized disposable incomes are considered. Consistent with the lower economic opportunities available in the Southern Italian regions, we observe a level of internal migration that is not negligible. Comparing NUTS-1 macro-areas of birth and residence of our adult sons' sample we notice, for instance, that the share of those born either in the South or in the Islands is 30.9 percent, whereas the share of those residing in these two areas amounts to 24.3 percent, and the share of individuals born in the South and in the Islands who moved towards the Northern and Central areas amounts to 27.1 percent and 23.1 percent, respectively (Table 5).

Following suggestions by Chetty *et al.* (2014) and Chetty and Hendren (2018a, 2018b), who stress the crucial role of within-country heterogeneities in intergenerational persistence, we have computed intergenerational association within each macro-area of birth, controlling or not, in different specifications, for individuals' mobility across areas (captured by interaction dummies between area of birth and residence). To obtain proper sample sizes within each area, we consider a larger sons' age span (35–44) and average their incomes over the period 2005–2013, and we consider 5-year predicted fathers' earnings at age 40–44 as the independent variable in the second stage. Note, however, that this larger sons' sample does not change IGE and rank-slopes' estimates at the national level when we do not control for sons' mobility (compare col. 1 of Tables 6 and 7 with col. 1 of Tables 3 and 4, top panels). Moreover, consistent with findings of Olivetti and Paserman (2015) for the US, both IGE and rank–rank slopes are considerably lower when controlling for sons' mobility, thus signaling that a fraction of the intergenerational persistence of earnings is related to the association between the area of birth and the residence of sons.

When we analyze the degree of intergenerational earnings' persistence within areas, we find not negligible heterogeneities when estimating either the IGE or the imputed rank–rank slope and either controlling or not for sons' mobility. As concerns the IGE, the Islands, followed by the North-West, are characterized by the highest intergenerational inequality, whereas the fathers'–sons' earnings' persistence in the Centre is very low (Table 6). When computing rank–rank slopes—the

TABLE 5
Sons' Mobility by Area of Birth and Area of Residence

		Area of Residence					
		North-West (%)	North-East (%)	Centre (%)	South (%)	Islands (%)	Total (%)
Area of birth	North-West	90.9	3.7	3.4	1.4	0.6	22.9
	North-East	2.1	96.4	1.2	0.0	0.3	25.4
	Centre	1.9	2.0	95.3	0.6	0.2	28.4
	South	10.6	6.7	9.8	72.5	0.4	15.6
	Islands	12.5	5.3	5.4	0.6	76.3	7.7
	Total	19.7	23.1	26.2	21.1	9.8	100.0

Source: elaborations on AD-SILC dataset.

TABLE 6
IGE Between Sons' and Fathers' Gross Earnings in Italian Macro-Areas

	Italy	North-West	North-East	Centre	South	Islands
Not controlling for mobility across areas						
Fathers' earnings	0.500***	0.464***	0.404***	0.171***	0.402***	0.518***
S.E.	[0.028]	[0.066]	[0.057]	[0.052]	[0.078]	[0.098]
Controlling for mobility across areas						
Fathers' earnings	0.339***	0.426***	0.399***	0.148***	0.347***	0.509***
S.E.	[0.028]	[0.074]	[0.055]	[0.055]	[0.074]	[0.095]
Obs	5354	1117	1240	1406	1089	502

Average earnings of sons aged 35–44 in 2005–2013 are considered. Fathers' earnings predicted as the 5-year average when aged 40–44.
***$p < 0.01$.
Note: IGE stands for Intergenerational Elasticity.
Source: elaborations on AD-SILC dataset.

TABLE 7
Rank–Rank Slope Between Sons' and Fathers' Gross Earnings in Italian Macro-Areas

	Italy	North-West	North-East	Centre	South	Islands
Not controlling for mobility across areas						
Fathers' earnings	0.260***	0.231***	0.210***	0.112***	0.231***	0.265***
S.E.	[0.013]	[0.033]	[0.028]	[0.027]	[0.035]	[0.048]
Controlling for mobility across areas						
Fathers' earnings	0.186***	0.212***	0.209***	0.107***	0.210***	0.269***
S.E.	[0.015]	[0.034]	[0.028]	[0.027]	[0.037]	[0.042]
Obs	5354	1117	1240	1406	1089	502

Average earnings of sons aged 35–44 in 2005–2013 are considered. Fathers' earnings predicted as the 5-year average when aged 40–44.
***$p < 0.01$.
Source: elaborations on AD-SILC dataset.

most appropriate measure to compare intergenerational persistence within regions, according to Chetty *et al.* (2014)—the Islands and the Centre still emerge as the least and the most mobile areas, respectively, whereas no differences characterize the North-West and the South (Table 7).

6.3. *Intergenerational Association by Children's Gender*

Table 8 shows estimates of IGE and rank–rank slopes by gender, considering three different earnings' measures for both daughters and sons: family annual gross earnings (obtained by adding the earnings of the daughter/son and her/his co-residing partner, also including single-person households; cols. 1–2); couple earnings (i.e., excluding single-person households; cols. 3–4); and individual earnings (cols. 5–6).

For comparison with estimates shown in previous sections, note that we take the 5-year average of fathers' earnings, whereas we observe children's earnings only in 2005 (we consider children aged 35–39, thus born in 1966–1970), since partners can be associated precisely in our dataset only in that year (the household composition, allowing us to match adult children and their partners, may change over time and is recorded only in the years when the cross-sections of IT-SILC are carried out).[17]

When we also consider single-person households (cols. 1–2), the IGE for daughters is higher than is that computed for sons, even if the difference is not statistically significant, whereas the rank–rank slopes are very similar. When we consider only couples (cols. 3–4), the IGE and the rank–rank slopes of sons become higher than are those computed from the daughters' sample, even if, also in this case, estimated confidence intervals cross at the 95 percent level. Large differences across gender emerge when looking at individual earnings of daughters and sons (cols. 5–6), since both IGE and rank–rank slope are much higher for males than for females. However, as remarked, focusing on daughters' individual earnings might be misleading due to the Italy's low rate of female employment, which might be affected by features of the parental background (according to Eurostat data, employment rates in the age class 35–39 in 2005 were 61.9 percent and 90.7 percent for females and males, respectively).

Note, however, that, when considering couple earnings, both measures of intergenerational association increase considerably with respect to measures computed using males only, thus suggesting that assortative mating—i.e., the tendency of males and females with similar socioeconomic and cultural characteristics to mate (Lam and Schoeni, 1994; Ermisch *et al.*, 2006)—strengthens the intergenerational transmission process.

7. Concluding Remarks

In this article, applying the TSTSLS method to overcome the lack of information on parents' incomes, and using a recently developed dataset that allows

[17]When comparing daughters' and sons' households, we select the subsample according to the age class of the daughter or the son, whereas we do not restrict the partner's age. Therefore, sample sizes in daughters' and sons' couples differ, since we select daughters or sons aged 35–39 in 2005 in columns 3 and 4, respectively.

TABLE 8
Association Between Daughters' or Sons' and Fathers' Gross Earnings

	Family Earnings		Couple Earnings		Individual Earnings	
	Daughters	Sons	Daughters	Sons	Daughters	Sons
IGE						
Fathers' earnings	0.474***	0.433***	0.534***	0.595***	0.349***	0.451***
S.E.	[0.054]	[0.053]	[0.055]	[0.065]	[0.069]	[0.052]
Rank–rank slope						
Fathers' earnings	0.216***	0.220***	0.265***	0.309***	0.164***	0.261***
S.E.	[0.026]	[0.025]	[0.029]	[0.030]	[0.030]	[0.026]
Obs	1523	1450	1159	944	1143	1407

Daughters' and sons' earnings observed in 2005. Daughters and sons aged 35–39 in 2005 are selected (no selection of partners' age is considered). When considering family and couple earnings, we take total earnings of daughters (sons) and their co-residing partner, including possible zero earnings of a partner; single-person households with zero earnings are not considered when considering family earnings. Fathers' earnings predicted as the 5-year average when aged 40–44.

***$p < 0.01$.

Source: elaborations on AD-SILC dataset.

researchers to improve the accuracy of the prediction of fathers' earnings, we have extended the knowledge about the transmission of intergenerational earnings between fathers and sons in Italy in several directions.

Different from previous estimates for Italy that were based on net earnings, we estimated the IGE of gross earnings, thus computing the level of intergenerational persistence ascribable to labor market forces, before tax redistribution. The longitudinal dimension of our dataset also allowed us to take into account the two sources of bias that may jeopardize the robustness of our estimates, i.e. the attenuation and the life-cycle biases. With this aim, we compared point-in-time estimates with 5-year averages of both fathers' and sons' earnings and considered sons at different ages, confirming that computing point-in-time earnings' associations and considering sons when they are still too young may result in large underestimation of the IGE. We also provided estimates of "imputed" rank–rank slopes for Italy for the first time.

Our preferred estimate is obtained considering sons aged 35–39 and averaging earnings over a 5-year period for both generations. We find an IGE equal to 0.501, which does not differ much from the values 0.500 and 0.435 that were obtained by Mocetti (2007) and Piraino (2007), respectively. However, these researchers relied on two factors—i.e. point-in-time estimates and net of taxes earnings—which tend to reduce the estimated IGE with respect to the case when averaged gross earnings are considered. Therefore, compared to previous studies, our analysis shows a slightly lower level of intergenerational persistence in Italy, which confirms, however, that Italy is in the group of developed countries with the highest levels of intergenerational inequality.

Furthermore, if we do not discard zero annual earnings' records when computing 5-year earnings' averages, the estimated IGE increases up to 0.577. Even if we are unaware of the very reasons for possible zero annual earnings in our dataset (e.g. either voluntary inactivity or long-term unemployment), the large increase in the estimated IGE emerging when zero earnings are considered suggests that sons coming from poorer backgrounds are more likely to experience unstable working careers than are sons born in well-off families. As a consequence, estimates on permanent earnings (or on observation periods longer than 5 years) might provide higher values of the intergenerational earnings' association.

When computing rank–rank slopes, we found, instead, the size of the intergenerational association to be relatively lower in a cross-country comparison perspective. However, a cross-country ranking based on this measure is not easy to obtain since rank–rank slope measures are available only for a few countries, and, to the best of our knowledge, existing studies did not use fathers' percentiles computed according to predicted earnings obtained through the TSTSLS approach.

We also computed measures of intergenerational earnings association by geographical macro-area and by gender, finding that both IGE and rank–rank slopes differ considerably across areas and that the intergenerational persistence increases when couple earnings are considered, consistent with the existence of an assortative mating mechanism that strengthens the process of intergenerational transmission of inequality.

Finally, when we ran additional estimates including sons' characteristics (education and occupation) that might mediate the association between fathers' and sons' earnings as covariates, we found that high and statistically significant IGE and rank–rank slopes still emerged. We are not able to infer if the residual association between fathers' and sons' earnings that emerged when we controlled for children's outcomes influenced by their parents is driven by unobservable abilities or if it is also related to the influence of less meritocratic factors, such as social connections, on labor market achievements and earnings. The identification of the mechanisms behind this residual association will be the aim of future research.

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