the review of income and wealth

Review of Income and Wealth
Series 67, Number 4, December 2021
DOI: 10.1111/roiw.12505

INTERGENERATIONAL MOBILITY IN LIFETIME INCOME

BY MOSHE JUSTMAN* and HADAS STIASSNIE

Ben Gurion University

Estimates of intergenerational mobility in lifetime income derived from incomplete income histories routinely incorporate in the estimation process, necessary life-cycle adjustments to annual income data. The two-stage method presented here first estimates proxies for fathers' and sons' lifetime family incomes from annual income observations, schooling and race; and then uses these income proxies to derive mobility measures. Applying this to United States PSID data for sons born between 1952 and 1981, we find a decline in intergenerational mobility in lifetime family income, as measured by the intergenerational elasticity of income, the rank-rank correlation, absolute upward mobility, and other indicators.

JEL Codes: D31, I32

Keywords: intergenerational income mobility, lifetime income, PSID, intergenerational income elasticity, rank mobility

1. Introduction

Measures of intergenerational income mobility quantify the extent to which the economic circumstances of a person's upbringing correlate with their lifetime income or earnings. The intergenerational elasticity (IGE) of income, estimated by regressing the logarithm of sons' or daughters' incomes (or earnings) on their parents' incomes as they were growing up is a common (inverse) measure, as are the intergenerational correlation and rank correlation of income.[1] Early IGE estimates used "snapshots" of sons' or daughters' annual income in fixed age cohorts, typically around age 30, as a proxy for their lifetime income, while averaging parental income over several years, as their children were growing up, to reduce the downward bias stemming from classical measurement error in the right-hand variable

Note: This paper draws on earlier work with Anna Krush. It has greatly benefited from the editor's and referees' detailed comments and suggestions; from comments on earlier versions by Francois Bourguignon, Richard Burkhauser, Danny Cohen-Zada, Gary Fields, Mark Gradstein, Chris Ryan, and Jacques Silber; and from participants in seminars at Ben Gurion University and the University of Melbourne, and in EcIneq meetings in New York and Paris. Financial support from Ben Gurion University and Ruppin Academic College is gratefully acknowledged.

*Correspondence to: Moshe Justman, Department of Economics, Ben Gurion University, Be'er Sheva, Israel (justman@bgu.ac.il).

[1]There is also an extensive literature, economic and sociological, on other dimensions of intergenerational mobility, such as occupation, education, and social class; see, among others, the survey by Black and Devereux (2011). Blanden (2013) finds that mobility in earnings and education tends to be "fairly well correlated" across countries, but not social class.

(Solon, 1992).[2] However, as Jenkins (1987) first pointed out, and Haider and Solon (2006), Grawe (2006) and Nybom and Stuhler (2016, 2017) subsequently elaborated and quantified, there remains substantial bias. This is due to measurement error on the left-hand side, from using annual income at age 30 as a proxy for lifetime income, being correlated with parental income on the right hand side;[3] and to life-cycle bias arising from (non-classical) measurement error in parents' income due to variation in the age at which their income is observed. These analyses propose adjusting annual income observations for age, gender, education and other background variables, and adjustments along these lines are now routinely incorporated in IGE regressions (e.g. Hertz, 2007; Lee and Solon, 2009).

We build on these insights, but separate the two elements, first explicitly estimating lifetime income proxies for both fathers and sons, and then using them to derive multiple measures of mobility.[4] Specifically, we first use all available longitudinal income data, along with education and race indicators, to construct a set of shared age-income profiles, and derive predictions of individual annual income at age 40 that serve as our proxy for average lifetime income. Then, in the second stage, we use these proxies to derive measures of absolute, relative and positional mobility, which we average over ten-year cohort groups, recognizing that individuals compare their mobility to people of a similar—but not necessarily identical—age, while smoothing the variation of measured mobility over time.

Our two-stage method offers several advantages. It makes full use of extensive income histories available for many individuals, to reduce measurement error for individuals with only limited income histories—in particular, younger cohorts; and it is more transparent than single-stage approaches, allowing us to directly assess how well our lifetime income proxies accord with actual lifetime averages. In addition, because the first stage is not logarithmic, and we estimate common age-income profiles for large groups of individuals, our estimates are not sensitive to how we treat observations of zero or very low annual income, as we show.[5] Moreover, it allows us to explicitly measure multiple aspects of intergenerational income mobility—highlighted in Fields and Ok (1996), Fields (2010) and Jännti and Jenkins (2015)—and track their separate trajectories over time, using the same lifetime income variable.

While our approach is readily applied to a variety of income definitions, we choose to focus here on mobility in lifetime family income between fathers and sons. Parents' family income is a natural measure of the economic circumstances of a child's upbringing, and focusing on lifetime income—rather than the years the

[2]Mazumder (2005) showed that the number of observations needed to entirely remove classical measurement error in parental income is large.

[3]The slope of young adults' age-income profiles is positively correlated with their schooling, which is positively correlated with their parents' education and income.

[4]This builds directly on Justman and Krush (2013). Our approach differs from two-stage two- sample estimates (e.g. Aaronson and Mazumder, 2008), which in the absence of linked parent-child income data use linked data on parental education and/or occupation to instrument for their lifetime income. We have linked income data for parents and their children, from which we estimate their lifetime average incomes.

[5]Mitnik and Grusky (2020) elaborate on the sensitivity of previous IGE estimates to the treatment of zero and very low annual income observations. For example, Chetty *et al.*'s (2014b) IGE estimates range from 0.35 to 0.70 depending on how zero values are treated.

child was aged, say, 13 to 17—recognizes that parents are able to shift income over the life cycle. Using the same measure of income for fathers and sons conforms to the notion of the IGE of income measuring (inversely) the rate at which the origin dependence of family income dies out; and allows us also to measure absolute income mobility. We also estimate intergenerational mobility in labor earnings, for comparison. Of course, our approach could equally be applied to mothers and daughters.

We apply our method to linked longitudinal data from the United States Panel Study of Income Dynamics (PSID), up to and including its 2016 wave, on sons born between 1952 and 1981, and their fathers. We derive our first-stage lifetime income estimates over a large sample of males who meet out data requirements; and then extract a subsample of father-son pairs for whom we derive measures of absolute, relative and positional mobility, averaged over ten-year cohort groups, for three disjoint groups, 1952–61, 1962–71 and 1972–81, and for a rolling sequence of 21 overlapping groups.

Highlighting our main results, we find that absolute upward mobility, measured as the share of sons with greater lifetime family income than their fathers, declined from 67 percent for the 1952–61 cohort-group to 62 percent for the 1972–81 group. The IGE of lifetime family income increased from 0.425 to 0.532 from the oldest to the youngest cohort-groups, with similar increases in the father-son correlation of income and in the correlation of log (income), and in the slope of a quantile regressions of the conditional median, all indicating a decline in relative mobility. An integrated logarithmic regression of sons' annual income on first-stage estimates of fathers' lifetime incomes indicates that the increase in the IGE from the oldest to the youngest cohort-group is statistically significant at the 5 percent level. Positional mobility also declined, as indicated by the increase in the intergenerational rank correlation (IRC) of lifetime family income from 0.419 to 0.489, from the oldest to the youngest cohort-groups. Non-parametric regressions reveal that most of this change occurred at the top and bottom quintiles of fathers' income distribution. We also find that the IGE of individual lifetime *earnings* rose over this period while the IRC of earnings did not exhibit a trend.

We examine the sensitivity of our findings to adjusting father's family income for family size, and to bottom coding—uniform top-coding is necessary because of changes in the PSID over time—and find no effect in either the first or second stage. We also show retrospectively that the level and shape of our IRC time-series, and the shape of our IGE time-series, are robust to the addition of new waves of PSID data, suggesting that our present estimates of IRC levels and trends and IGE trends should be robust to the addition of future waves of data. This highlights the advantage of our approach for estimating mobility among younger cohorts.

The structure of the paper is as follows. Section 2 describes our first-stage method, and estimates proxies for lifetime family income. Section 3 uses these estimates to derive specific mobility measures, and considers their variation over time. Section 4 demonstrates the robustness of our findings to income specifications, sampling variation in the first stage, and the addition of new waves of data. Section 5 concludes.

2. First Stage Analysis: Estimating Lifetime Income

The first stage of our two-stage approach derives proxies for lifetime family income by estimating the following Mincer-type equation:

$$y_{it} = \alpha_{0i} D_i + \alpha_1 age_{it} + \alpha_2 age_{it}^2 + \boldsymbol{\alpha_3 educ_i} \cdot age_{it} + \boldsymbol{\alpha_4 educ_i} \cdot age_{it}^2 + \boldsymbol{\alpha_5 race_i} \cdot age_{it} + \boldsymbol{\alpha_6 race_i} \cdot age_{it}^2 + \varepsilon_{it} \quad (1)$$

where y_{it} is individual *i's* family income in year *t*; D_i is an individual fixed effect; age_{it} is *i*'s age in year *t*; $\boldsymbol{educ_i}$ and $\boldsymbol{race_i}$ are sets of dummy variables representing *i's* years of schooling and race; ε_{it} is an i.i.d. error term; and α_{0i}, $\alpha_1 \dots \alpha_6$ denote regression coefficients. (See Appendix Table A1 for education and race categories.) This estimation yields an age-income profile for each combination of schooling level and race, and individual fixed effects. We enter age_{it} as age in year *t* minus 40, so α_{0i} is predicted income at age 40, our proxy for average lifetime family income.[6] It is not actual income at age 40, but a prediction that uses information on all individuals to estimate the coefficients that shape the common age-income profiles.

Our data source is the Panel Study of Income Dynamics (PSID), from its inception in 1967 up to and including its 2016 wave, with data collected annually until 1996 and bi-annually thereafter. The PSID comprises a representative national sample drawn from the Survey Research Center (SRC), and a sample of low-income families (SEO). We follow Lee and Solon (2009), Hertz (2007) and others, and use only the SRC sample. We restrict our attention to families in which the father is the head of household. Family income includes the taxable and transfer income of all family members. We adjust all income data to 2012 prices using the personal consumption expenditures (PCE) index, and follow previous studies in top-coding annual income at \$150,000 at 1967 prices (\$814,000 at 2012 prices).[7]

Our base population comprises all 12,652 males in the SRC born between 1918 and 1981 with reported race and years of schooling. From these we extract 7,511 males with at least three annual income observations over \$1,000 (at 2012 PCE prices) between the ages of 25 and 64 and use it to estimate individuals' lifetime family income using equation (1); we refer to this as our base sample. We pool all 96,765 annual income observations in one estimation with 7,511 individual fixed effects. It accounts for 0.624 of the variance in income observations. Appendix Table A1 presents the estimation coefficients, and Appendix Figure A1 shows two age-income profiles derived from our first-stage estimation, with individual fixed effects chosen to illustrate a rank reversal in income between the ages of 30 and 40. Allowing age-income profiles to vary over time by dividing the sample into several disjoint cohort groups and estimating income within each group (as in Justman and Krush, 2013) yields nearly identical results, as we show in Section 4.1.

To demonstrate the close correlation of our income proxy with actual average lifetime family income, we compare it to average annual income between the

[6]The choice of age 40 follows Haider and Solon (2006). Our proxy is closely correlated with average lifetime family income, but larger, as income at age 40 is close to peak income.

[7]Uniform top-coding offsets the effect of changes in the PSID's top-coding ceilings, which have varied between \$99,999 and \$9,999,999. Our results are not sensitive to the level chosen.

TABLE 1
Pearson and Rank Correlations of Average Annual Family Income Between the Ages of 30 and 55 with Three Alternative Income Measures

Income Measure	Population	*N*	Correlation	Rank Correlation
Predicted family income at age 40	Full sample of males born in 1937–66, with 10 or more income observations	2,597	0.97	0.96
Average family income, age 29–31	Subsample of males with one or more income observations between the ages 29–31	2,135	0.71	0.75
Father's average family income, sons aged 13–17	Subsample of fathers with 3 or more income observations in those years	796	0.91	0.88

ages of 30 and 55 for a high-quality subset of our base sample: 2,596 males born between 1937 and 1966—and thus observed in their prime earning years—with at least ten annual income observations above $1,000 (2012 prices; 21 observations per person on average). We also compare these actual income averages to average family income between the ages 29–31, for a subset of 2,135 males with at least one annual income observation above $1,000 in this range; and to father's average family income when sons were aged 13–17, for a subset of 796 fathers with 3 or more income observations above $1,000 in those years. Table 1 presents correlations and rank correlations, and Figure 1 presents scatter diagrams, both of which illustrate the close correlation of our income proxy with actual values.

For the second stage, we identify a sub-sample of 1,536 father-son pairs, of sons born between 1952 and 1981, for each of whom we have at least 3 annual income observations above $1,000 from age 29, and their fathers for whom we have at least 5 annual income observations above $1,000 between the ages of 25 and 64; we refer to this as our "pairs sample". In the next section, we use their predicted income at age 40 to estimate multiple measures of intergenerational mobility in three disjoint ten-year cohort groups of sons, born in 1952–61, 1962–71 and 1972–81; and for the 21 rolling ten-year cohort groups in this range.[8] Table 2 presents descriptive statistics for sons and fathers in the three disjoint cohort groups, and means and standard deviations of our predicted average lifetime income proxy and its logarithm. Appendix Tables A2 and A3 present these statistics for all 21 rolling ten-year cohort-groups.

[8]Cohort-groups of 6 to 12 cohorts yield very similar results. Shorter groups yield more volatile estimates; longer cohorts leave less scope for tracking mobility over time.

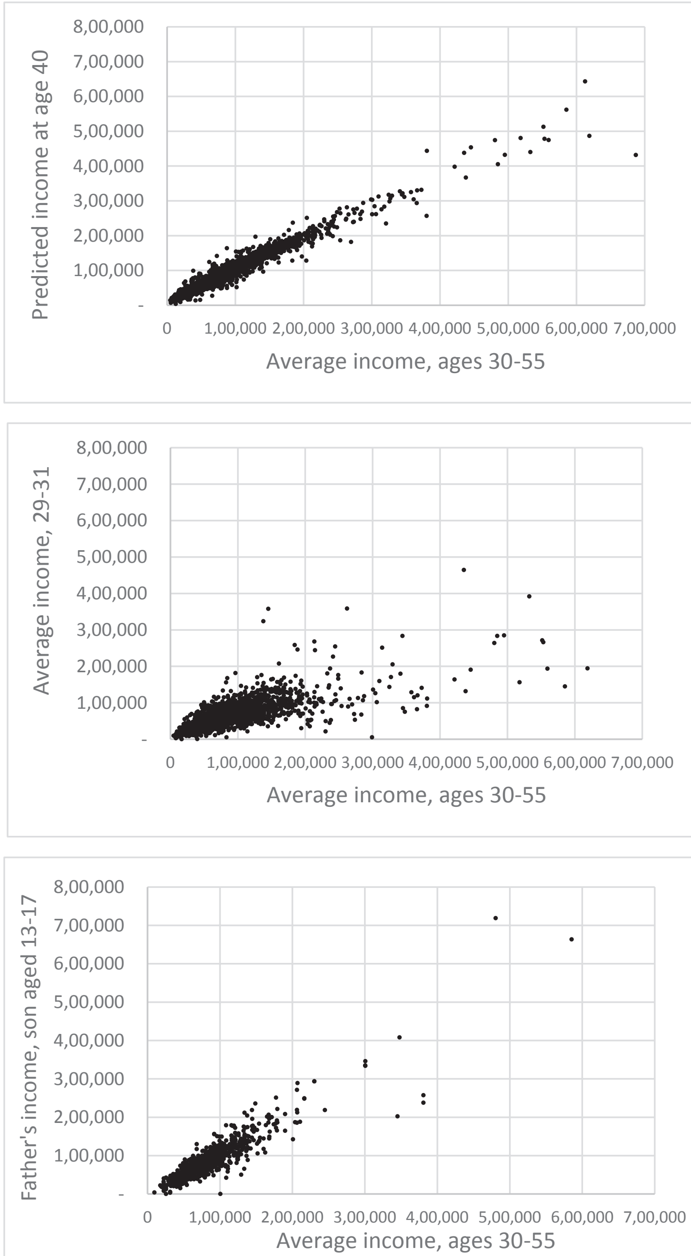

Figure 1. Average Family Income, Ages 30–55, Plotted Against: Predicted Income at Age 40 for All PSID Males Born in 1937–66, with 10 Income Observations; Average Income at Ages 29–31 for a Subsample of Males with at Least One Observation in that Range; and Average Income when Sons were Aged 13–17, for Fathers with 3 Income Observations. Income in 2012 PCE Dollars

Cohort-group medians and means of our (PCE adjusted) estimated average lifetime family incomes grew more for fathers than for sons between the 1952–61 and 1972–81 cohorts, and medians grew more than means. Thus, the mean to

TABLE 2

Descriptive Statistics, Fathers and Sons in Pairs Sample, by Sons' Ten-Year Cohort Group

Sons' Year of Birth	*N*	Mean Number of Observations		Mean Age of Observed Income		Mean Years of Schooling		Father's Mean Age at Son's Birth	% White (Sons)	% Black (Sons)
		Fathers	Sons	Fathers	Sons	Fathers	Sons			
1952–61	477	23	20	50	37	12.5	14.0	28.6	93.5%	5.2%
1962–71	468	29	12	44	35	13.2	14.0	27.7	93.0%	4.1%
1972–81	591	25	7	41	32	13.8	14.4	27.2	89.2%	6.1%

Sons' Year of Birth	Estimated Proxy for Average Annual Lifetime Family Income: Mean		Estimated Proxy for Average Annual Lifetime Family Income: Median		Standard Deviation of Log Income		Share of Variance Explained	
	Fathers	Sons	Fathers	Sons	Fathers	Sons	Fathers	Sons
1952–61	78,755	98,453	69,698	86,355	0.487	0.539	0.616	0.562
1962–71	81,672	106,592	73,010	91,449	0.458	0.558	0.581	0.622
1972–81	92,855	110,433	85,033	100,179	0.487	0.543	0.633	0.648
% change 1952–61 to 1972–81	17.9%	12.2%	22.0%	16.0%				

Notes: *N* is the number of father-son pairs in each cohort group. Our "estimated proxy for average lifetime family income" is predicted income at age 40, in 2012 PCE dollars, from equation 1. "Share of variance explained" is the share of variance in income observations explained by the first-stage regression within each cohort-group, equal to 1 minus the sum of squared residuals, divided by the sum of squared deviations from the mean.

median ratio of lifetime family income declined substantially from the oldest to the youngest cohort group, indicating a decline in inequality in lifetime income within cohort groups, in this dimension. At the same time, the standard deviation of both fathers' and sons' log income, within cohort groups, a different dimension of inequality, changed very little between the oldest and youngest cohort groups.

These cohort-based statistics on lifetime incomes have no direct counterparts in national statistics. To allow a comparison, we construct annual statistics from our pairs sample for the years 1967–2016 for average annual family income, its Gini coefficient, and its mean to median ratio. All three exhibit close correlations with national trends in PCE adjusted household income statistics (US Census Bureau, 2019a, 2019b): 0.99 for average income, 0.96 for the Gini coefficient, and 0.95 for the mean to median ratio (Appendix Figure A2 presents scatter diagrams). As our selection criteria yield a more affluent and homogeneous sample than the general population, our mean incomes are higher and our inequality measures are lower than national levels. The greater economic and racial homogeneity of our sample may result in our overestimating the level of intergenerational mobility in the population, if the tails of the income distribution, under-represented in our sample, are less inter-generationally mobile than the middle, as previous studies have found (Isaacs, 2007; Bengali and Daly, 2013; Acs *et al.*, 2016; Mitnik *et al.* 2018; Palomino *et al.*, 2018).

3. Stage Two: Mobility Measures

In the second stage, we use our average lifetime income proxy of fathers and sons to derive specific measures of intergenerational mobility, averaged over ten-year cohort groups, as they vary over time. Appendix Figure A3 presents scatter diagrams of the joint distribution of fathers' and sons' lifetime family income proxies within each of our three disjoint ten-year cohort-groups, 1952–61, 1962–71 and 1972–81. We find a rise in the intergenerational correlation, from 0.27 to 0.35 to 0.37, a first indicator of declining intergenerational income mobility.

3.1. *Absolute Upward Mobility*

Initially, in the oldest cohort-group, 67 percent of sons had greater PCE-deflated lifetime family incomes than their fathers, this share declining moderately to 62 percent for the youngest cohort-group (Figure 2). This accords with fathers' family income rising more steeply than sons' family income from the oldest to the youngest cohort-group (Table 2). The level of absolute upward mobility we find is similar to Isaacs' (2007) and Acs *et al.*'s (2016) PSID-based estimates; and both the level and downward trend we find are similar to Chetty *et al.*'s (2017, Figure 3A) PCE-deflated estimates, for the period considered here (their slightly lower baseline estimates use the CPI-U-RS deflator).

3.2. *The Distributional Incidence of Intergenerational Income Growth*

Table 3 presents fathers' and sons' average lifetime family income, within fathers' income quartiles, by sons' cohort-group. It highlights the progressive

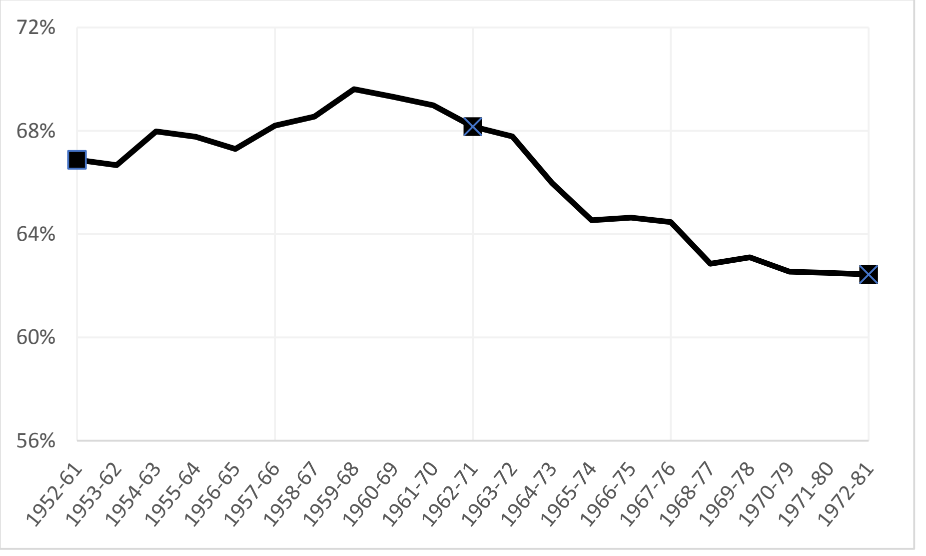

Figure 2. Absolute Upward Mobility: Share of Sons whose Lifetime Family Income Exceeded their Fathers', by Cohort Group; PCE-Deflated

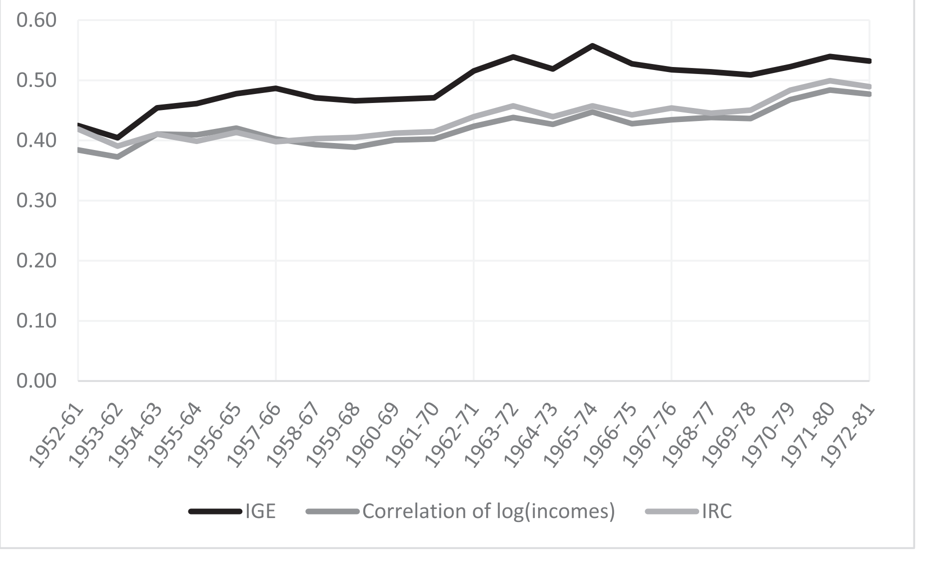

Figure 3. The Intergenerational Elasticity (IGE) of Lifetime Family Income, Correlation of Log(Income), and Rank Correlation (IRC) of Income, in Ten-Year Cohort-Groups

Note: Proxies for lifetime family income estimated in a first stage from all income observations, allowing age-income profiles to vary by education and race (equation 1).

distributional incidence (Bourguignon, 2010) of intergenerational income growth throughout the period studied. Within each cohort-group, the ratio of sons' to fathers' mean income, in column (3), falls sharply with fathers' income—a

TABLE 3

Fathers' and Sons' Average Lifetime Family Income, by Father's Income Quartile and Sons' Cohort-Group; Income in 2012 PCE Dollars

Father's Income Quartile	Fathers' Mean Income	Sons' Mean Income	Increase in Sons' Mean Income Over Fathers'	Cumulative Increase
1952–61	(1)	(2)	(3)	(4)
Lowest quartile	38,682	70,230	81.6%	81.6%
2nd quartile	61,165	95,452	56.1%	65.9%
3rd quartile	80,865	107,539	33.0%	51.2%
Highest quartile	133,846	120,405	−10.0%	25.0%
All	78,755	98,453	25.0%	
4thQ/1stQ ratio	3.5	1.7		
1962–71				
Lowest quartile	42,200	81,004	92.0%	92.0%
2nd quartile	62,982	88,393	40.3%	61.1%
3rd quartile	84,465	105,757	25.2%	45.1%
Highest quartile	137,042	151,215	10.3%	30.5%
All	81,672	106,592	30.5%	
4thQ/1stQ ratio	3.2	1.9		
1972–81				
Lowest quartile	46,088	79,810	73.2%	73.2%
2nd quartile	73,209	99,059	35.3%	49.9%
3rd quartile	95,918	119,994	25.1%	38.8%
Highest quartile	155,736	142,521	−8.5%	18.9%
All	92,855	110,433	18.9%	
4thQ/1stQ ratio	3.4	1.8		

regression to the mean that equalizes dynastic income. In all three cohort-groups, fathers' mean income in the highest quartile is more than three times fathers' mean income in the lowest quartile, while the corresponding ratio for their sons is less than two. To compare the progressiveness of intergenerational income growth over time, we calculate the cumulative increase in sons' mean income over their fathers', in column (4). It shows that the youngest cohort-group, 1972–81, is second-order dominated by the other cohort-groups, indicating that the progressiveness of intergenerational income growth has diminished. Appendix Table A4 shows a similar pattern for relative income.

Two summary indices, developed by Shorrocks (1978) and Fields (2010) to characterize intra-generational mobility, are applied here to inter-generational mobility, to quantify variation over time in the extent in which inter-generational mobility equalizes dynastic income (Table 4). Shorrocks' (1978) rigidity index compares the inequality of the sum of fathers' and sons' lifetime family income to a weighted average of income inequality in each generation, weighted by average income; Fields' (2010) index compares the inequality of the sum of incomes to the inequality of fathers' incomes. We apply both indices to the coefficient of variation in lifetime family income, as a measure of inequality. It confirms the progressiveness of intergenerational income growth: the sum of father and son incomes is less unequal than each separately, for all three cohort-groups. However, where Shorrocks' index increases by 0.03 from the first to the second cohort-group—indicating a slight decline in progressiveness—and then remains stable, Fields' index fluctuates, an initial rise of 0.09 mostly offset by a subsequent decline of 0.08. Applying these measures to Theil's index with $\alpha = 0$ shows similar patterns.

3.3. *Relative Income Mobility*

The intergenerational elasticity (IGE) of income, the most widely used (inverse) measure of relative income mobility, is estimated as the slope of a logarithmic regression of sons' average lifetime family income, y_i, on their fathers' average lifetime family income, x_i:

$$\ln y_i = \alpha + \beta \ln x_i + \varepsilon_i \tag{2}$$

We estimate equation (2) within each of our rolling ten-year cohort-groups. Table 5 reports results for our three disjoint cohort-groups, along with the intergenerational correlations of lifetime family income and of the logarithm of lifetime income, and the regression slope from a quantile regression of the conditional median. All increase over time, indicating a decline in relative mobility. The level of our estimates are well within the range of previous IGE estimates of father-son mobility in family income for a similar time frame: slightly higher than the low end of previous estimates, e.g. Lee and Solon's (2009) average estimate of 0.44 for sons born in 1952–79; but lower than Mazumder's (2018) estimate of "0.6 or higher," at the high end of previous estimates, both using PSID data. (See Appendix Table A6 for additional estimates.)

Figure 3 shows the IGE estimates and father-son correlations of log(income) for all 21 rolling ten-year cohort-groups (Appendix Table A5 presents numerical

TABLE 4
The Coefficient of Variation in Fathers', Sons' and Combined Family Income, Shorrocks' Rigidity Index and Fields' Equalization Index, by Sons' Cohort-Group

	Coefficient of Variation in:				
Sons' Birth Cohort	Father's Income	Son's Income	Sum of Incomes	Shorrock's Index	Field's Index
1952–61	0.602	0.569	0.466	0.799	0.775
1962–71	0.572	0.621	0.497	0.829	0.869
1972–81	0.596	0.539	0.468	0.829	0.785

TABLE 5
The Intergenerational Elasticity (IGE) of Lifetime Family Income, Correlations of Income and of Log(Income), and Quantile Regression Slopes, by Sons' Cohort Group

Sons' Birth Cohort	*N*	IGE Estimate	Second-Stage Standard Error	Father-Son Correlation of Log Income	Father-Son Correlation of Income	Slope of the Conditional Median
1952–61	477	0.425	0.047	0.384	0.273	0.372
1962–71	468	0.516	0.051	0.423	0.354	0.455
1972–81	591	0.532	0.040	0.477	0.372	0.548

Notes: Average lifetime family income proxies estimated in the first stage from all income observations, allowing age-income profiles to vary by education and race (equation 1). Second-stage standard errors are the IGE standard errors estimated from an OLS regression of sons' log lifetime income on their fathers' log lifetime income. As our income variables are estimates, usual significance tests do not apply.

values.) The IGE is the product of this correlation and the son-to-father ratio of standard deviations of log(income). Appendix Figure A4 shows that the rise in the IGE was initially driven by both factors rising, until the 1965–74 cohort-group; then by the rise in the ratio of standard deviations; then, for the more recent cohorts, this ratio fell, largely offsetting a renewed rise in correlations. It is tempting, if only suggestive, to relate the periodicity of the rises we observe in the IGE for the 1962–71 and 1970–79 cohorts—indicating declining mobility—to the periodicity of recessions in the early 1980s and early 1990s, about the time these cohorts entered the workforce.[9]

Previous studies that derived local estimates of the IGE at different levels of parental income found greater IGE values (less mobility) at the ends of the distribution (Palomino *et al.*, 2018, Mitnik *et al.*, 2018). To allow for such non-linearities, we estimated a local quadratic approximation of our logarithmic regression (2), pooling all thirty cohorts of sons, from 1952 to 1981, using Stata's LOWESS procedure. The results, shown in Appendix Figure A5, do not indicate such variability, possibly due to the greater homogeneity of our sample. A linear approximation explains 97 percent of the variance in our local quadratic approximation.

[9]This would lead us to expect another rise in the IGE—indicating a further decline in mobility—in younger cohort-groups scarred by the dot-com recession of the early 2000s.

TABLE 6
The Intergenerational Elasticity (IGE) of Lifetime Family Income, Estimated Simultaneously with Sons' Lifetime Family Income, by Cohort Group

Sons' Birth Cohorts	Number of Sons	Number of Income Observations	IGE	Standard Error
1952–61	477	9,672	0.394	0.055
1962–71	468	5,823	0.499	0.062
1972–81	591	4,068	0.533	0.057

Notes: $N = 19{,}563$, $R^2 = 0.158$. Standard errors clustered at the level of individual sons.

Though we find a pronounced increase in the IGE from the oldest to the youngest cohorts, as our right-hand variable is itself an estimate usual significance tests do not apply. To assess whether the increase in the IGE we observe across cohort-groups is statistically significant, we regressed sons' annual incomes on their fathers' lifetime income proxies, derived in the first stage, with separate IGE slopes and constants by sons' cohort-group, and with controls for the son's age and its interaction with his education and race. Specifically, we estimate:

$$(3)\qquad \log\left(y_{ij}\right)=\sum_{c=1}^{3} D_{ci}\left(\alpha_c+\beta_c\cdot\log\left(\hat{x}_i\right)\right)+\alpha_1 age_{ij}+\alpha_2 age_{ij}^2+\boldsymbol{\alpha_3 educ_i}\cdot age_{ij} +\boldsymbol{\alpha_4 educ_i}\cdot age_{ij}^2+\boldsymbol{\alpha_5 race_i}\cdot age_{ij}+\boldsymbol{\alpha_6 race_i}\cdot age_{ij}^2+\varphi_i+\varepsilon_{ij}$$

over all sons' income observations, where $\log\left(\hat{x}_i\right)$ is the logarithm of father's predicted income at age 40 from the first stage; educ_i, and race_i represent sets of dummy variables as in equation (1); and D_{ci} is a binary variable indicating whether son i belongs to cohort-group c, where c varies over the three disjoint cohort-groups. Thus, both the constant term, $\boldsymbol{\alpha_c}$, and the slope of father's income, $\boldsymbol{\beta_c}$, are allowed to vary across the three disjoint cohort-groups, yielding cohort-group-specific IGE estimates, presented in Table 6. They are slightly lower for our oldest cohort-group than our two-stage estimates, but similar for the youngest group, so that the increase from the oldest to the youngest is greater. A Wald test with standard errors clustered at the level of individual sons, allows us to reject the hypothesis that the IGE did not increase from the oldest to the youngest cohort groups at a significance level of 5 percent.

Previous PSID-based studies of trends in the father-son IGE of family income did not find an upward trend in the IGE (Mazumder, 2018). Thus, Lee and Solon (2009) and Hertz (2007), whose methods are closest to ours, found no significant trend, albeit theirs are negative findings—neither found a "precisely measured zero" trend. We attribute this difference to our having eight more waves of data and more stringent data requirements (their youngest cohort is observed only to age 25), and to the accuracy of our first-stage estimates of lifetime incomes. Comparing our estimates to Durlauf *et al.*'s (2017) IGE estimates for sons born between 1952 and 1975, from PSID family income observed to 2010, highlights the importance of controlling for life-cycle bias in parents' incomes, as Grawe's analysis (2006) indicated, and Mazumder (2016) and Nyborn and Stuhler (2016,

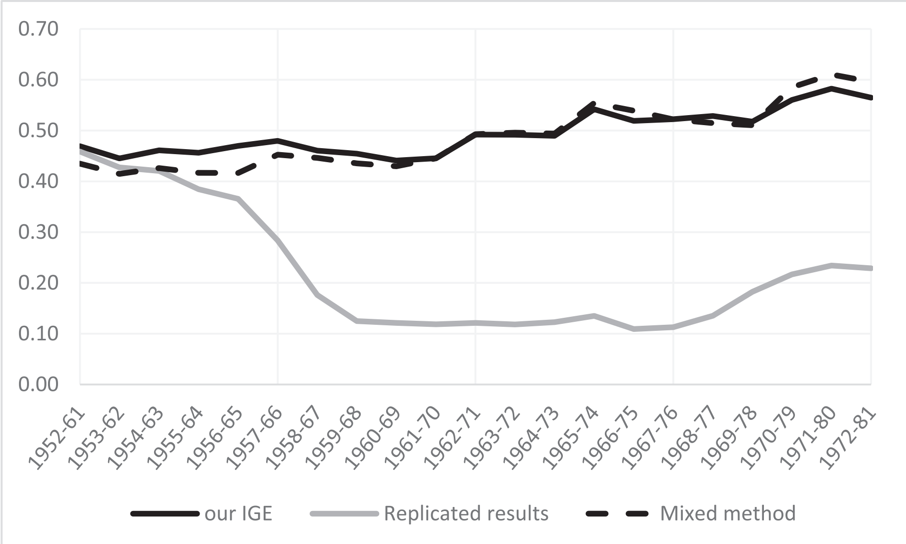

Figure 4. A Comparison of Our IGE Estimates to Durlauf *et al*. (2017)

Notes: "Replicated results" applies Durlauf *et al*.'s (2017) method to a subsample of our data that meets their criteria of at least 3 observations per father when the son is aged 13–17 and at least 3 observations per son between the ages 25 and 34, averaging sons' actual income between the ages of 25 and 34, and fathers' actual income when the son is aged 13 to 17. "Our IGE" applies our two-stage method to this subsample. "Mixed method" uses their method for sons' incomes and our method for fathers' incomes.

2017) further elaborated. Durlauf *et al*. (2017) regress sons' family income, averaged between the ages of 25 and 34, on fathers' family income, averaged when their sons were aged 13 to 17, and similarly aggregate father-son pairs in successive ten-year cohort groups. They find an initially downward trend, which we replicate in Figure 4, on a subsample of father-son pairs in our sample that meet their selection criteria (marked "replicated results"), where our method replicated on the same subsample (marked "our IGE") shows a rising trend. However, when we substitute our estimate of fathers' lifetime income for their five-year average (marked "mixed method") we find that most of the difference disappears. This suggests that even when averaging sons' income over several years is sufficient to remove life-cycle bias, averaging fathers' income while the son is growing up, leaves significant life-cycle bias, suggesting that parents are able to shift income over the life-cycle. Aaronson and Mazumder's (2008) two-stage, two-sample analysis of US census data from 1940 to 2000, found that "mobility increased from 1950 to 1980 but has declined sharply since 1980." It is difficult to align their calendar time with our cohort-based estimates, but their conclusion that mobility has declined in recent decades accords with our findings.

3.4. *Positional Mobility*

The influential work of Chetty and associates (2014a, 2014b) has focused attention on intergenerational positional mobility, measured as the intergenerational rank correlation (IRC) of income or, equivalently the rank-rank regression slope. Applying this measure to our lifetime family income estimates, we calculate IRCs within all 21 rolling ten-year cohort groups, presented graphically in Figure 3, and numerically in Appendix Table A5. We find IRC values of 0.419, 0.439 and 0.489 for our three disjoint cohort groups. All three graphs in Figure 3 are rising, indicating declines in both relative and positional intergenerational mobility over time. However, where the IRC and the intergenerational correlation of log income are very similar in both level and shape, with a correlation of 0.95 between them across the 21 cohort groups, the IGE exhibits a different dynamic pattern, with a lower correlation of 0.85 between the IGE and the IRC. Most of the rise in the IGE occurred in the earlier years, when positional mobility was less variable, while most of the increase in the IRC occurred more recently.

These findings are not directly comparable to Chetty *et al.*'s (2014a) estimates of single-cohort intergenerational rank correlations for sons aged 30, born between 1971 and 1986, using tax records on very large samples, which found no trend.[10] However, we can compare our estimates to the rank correlations in annual income that they report for three four-year cohort groups, 1971–74, 1975–78, and 1979–82: 0.299, 0.291 and 0.313 (Chetty *et al.*, 2014a, Figure 1). We re-estimated IRCs in lifetime family income for the three cohort groups 1971–74, 1975–78 and 1979–81, and found much higher values than theirs, but a similar dynamic pattern: a slight decline between the first two cohort-groups, from 0.437 to 0.430, followed by a greater increase, to 0.603, for the youngest group. We attribute our much higher elasticity values to their use of snapshot income data at age 30, which yields downward biased estimates (Mazumder, 2016; Nyborn and Stuhler, 2016; 2017; Mitnik *et al.*, 2019).

To allow for non-linearities in the IRC, we also estimate a local quadratic approximation of our rank-rank regression for each of the three cohort-groups, using Stata's LOWESS procedure (Appendix Figure A6). We find non-linearities at the ends of the distribution, with each cohort-group following a different pattern, while in the middle, between the 20th and 80th percentiles, the intergenerational rank-rank link is closer to linear, and varies less over time. This suggests that the increase in the rank correlation we find from the oldest to the youngest cohort-groups is mostly driven by declines in positional mobility at the ends of the distribution.

3.5. *Intergenerational Mobility in Earnings*

To add a further dimension to our analysis, we apply our two-stage method to estimate intergenerational mobility in labor earnings, and estimate equation (1) for labor earnings, to obtain first-stage proxies for lifetime labor earnings. First, with regard to absolute upward mobility, we find that the share of sons with lifetime incomes that exceed their fathers', increased from 55 percent for the oldest cohort-group to 59 percent for the youngest, indicating increased mobility. We then estimated logarithmic and rank regressions of sons' lifetime earnings on their fathers' lifetime earnings, within our ten-year cohort-groups, to obtain IGE and IRC estimates. These are shown in Figure 5. The two IGE series, in the top panel,

[10]For cohorts born in 1986–93, they measure intergenerational mobility with regard to college enrolment at age 19. This is outside our time frame, and measures something else.

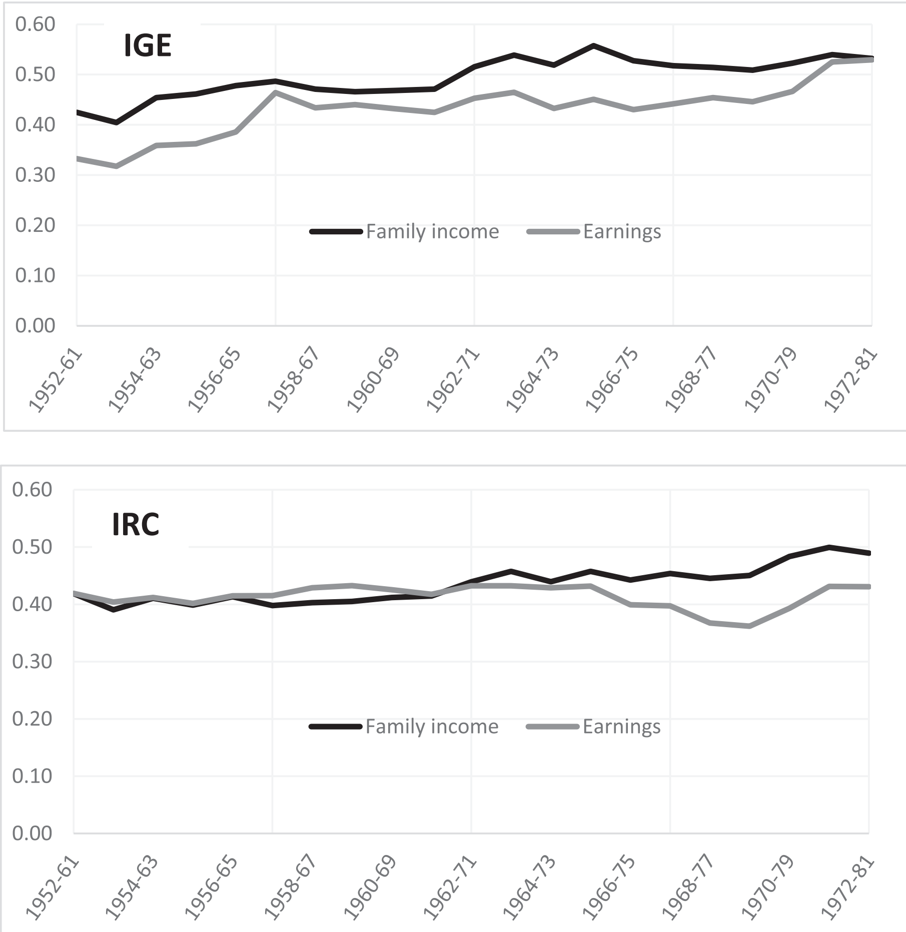

Figure 5. IGE and IRC Estimates of Mobility in Family Income and Labor Earnings, by Sons' Cohort Groups

exhibit similar increasing trends, with a correlation of 0.82, indicating that relative mobility in earnings has also declined. At the same time, the IRC series in earnings, in the bottom panel, while exhibiting a similar levels of rank mobility as family income, does not show a trend, nor are the two IRC series positively correlated. In sum, for individual lifetime earnings, absolute upward intergenerational mobility has increased over this period; relative mobility has declined; and positional mobility does not exhibit a trend.

4. Robustness to Income Specifications and Dynamic Consistency

Intergenerational mobility estimates are often highly sensitive even to seemingly small changes in the way income is specified and measured. A key advantage of our two-stage method is its robustness to such variation, and to the accumulation of new data.

4.1. *Robustness to Variation in Income Measurement*

In this section, we show that our findings are robust to allowing age-income profiles to vary over time, to variation in the age of predicted income and in bottom coding, to adjustment for family size and to first-stage sub-sampling.

Our preferred specification pools all observations in the first stage, thus imposing uniform age-income profiles, conditioned on education and race, on all fathers and sons. Now, to allow fathers' and sons' age-income profiles to vary over time, we divide our base sample into three disjoint cohort-groups of similar size—1918–48, 1949–64, and 1965–81—and estimate individuals' predicted income at age 40 from equation (1) within each group separately. The set of income proxies we obtain are almost perfectly correlated with our pooled estimates, and the corresponding IGE estimates are nearly indistinguishable (Appendix Figure A7).

Next, using predicted income at age 45 as our proxy for lifetime income—instead of predicted income at age 40—we find that the two IGE series, for our 21 overlapping ten-year cohort-groups, are almost perfectly correlated, with the level of the IGE series for age 45 slightly higher than for age 40, by a difference between 0.02 and 0.04 (Appendix Figure A8, top panel). Using predicted income at age 35 as a proxy for lifetime income produces substantially lower IGE estimates. The three IRC series (Figure A8, bottom panel) follow a similar pattern with much smaller differences in levels.

As annual incomes do not enter directly in the logarithmic second stage, our mobility estimates are much less sensitive to zero or very small annual income values than one-stage estimates, and do not require bottom coding, as in our preferred specification. To illustrate this, we follow Lee and Solon (2009), bottom-coding incomes below \$150 at 1967 prices (\$814 at 2012 PCE adjusted prices) and re-estimate our first and second stages. This yields nearly identical lifetime income proxies, and IGE and IRC series with correlations of 0.998 and 0.999 between estimates derived with and without bottom coding. Applying a higher bottom-coding value, equal to one fourth of the annualized minimum wage, produced very similar results.

Next we test whether adjusting parents' family income for family size makes a difference. A standard correction divides family income by the square root of the number of family members (e.g. Atkinson, 1996). We applied this correction to our first-stage family income estimates and found it had little effect on our estimates (Appendix Figure A9). IGE values for the adjusted series are lower by 6.5 percent, on average, with a correlation of 0.97, and the adjusted IRC series is almost identical to our preferred specification. Other corrections for family size yield very similar results. We favor the unadjusted specification because it is simpler, and preserves the symmetry between fathers and sons.[11]

[11]A previous version incorporated marital status in the first-stage estimate of the earnings profile. The correlation between the two specifications is 0.988 for the IGE and 0.993 for the IRC. Again, as the difference is very small, we prefer the simpler specification.

To test the robustness of our measures to first-stage sampling, we ran 100 repetitions of our estimation procedure. In each repetition we drew a 50 percent random subsample (with replacement) from our base sample, and used it to estimate average lifetime income from equation (1). Estimating average lifetime income for individuals in the subsample is straightforward—we use their estimated fixed effects as our lifetime income proxies. For the remaining 50 percent, not in

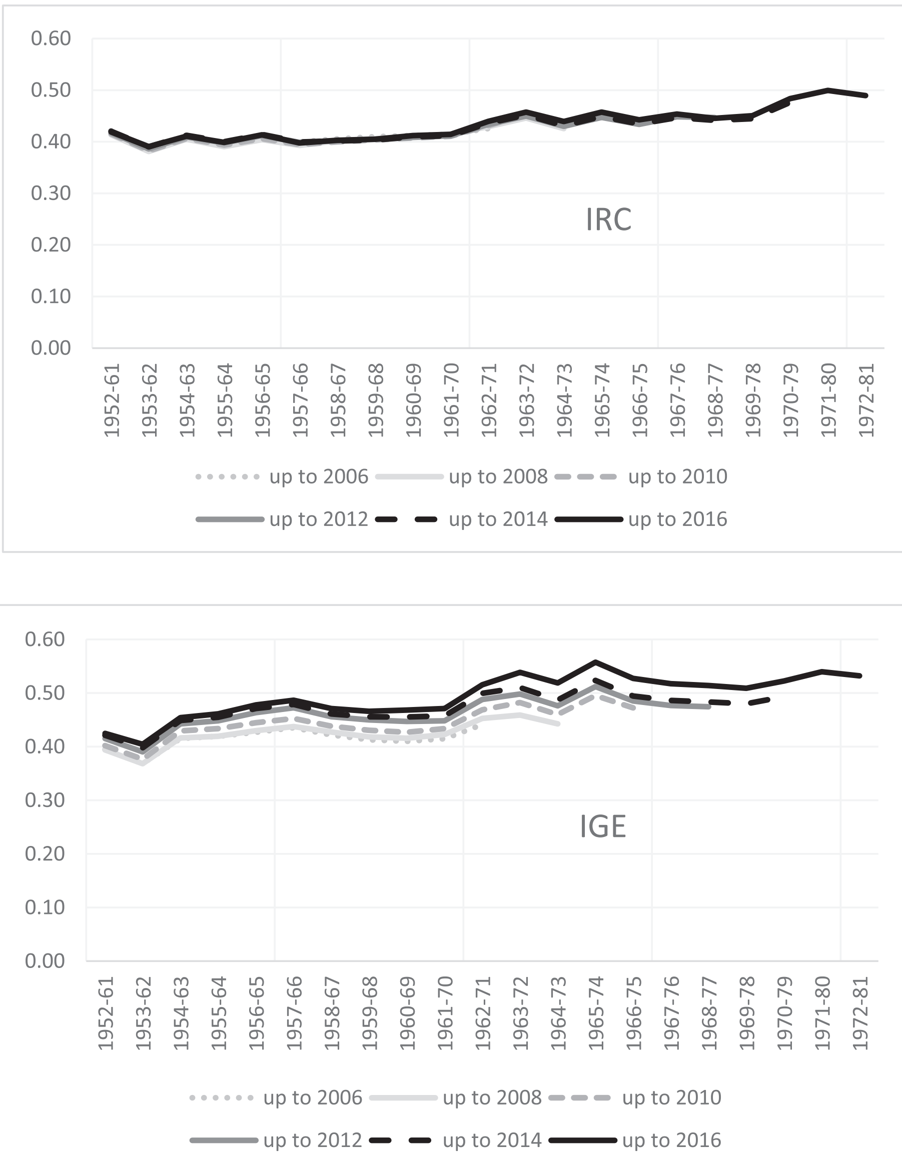

Figure 6. Intergenerational Rank Correlations (IRC; top panel) and Elasticities (IGE; bottom panel) of Lifetime Family Income, Between Fathers and Sons, Using All PSID Waves to 2016, 2014, 2012, 2010, 2008 and 2006; by Sons' Ten-Year Cohort Groups

the subsample, we find the individual fixed effect that minimizes the mean square difference between predicted and actual income over all income observations for that individual, and use it as our estimate of average lifetime income. We then use these lifetime income estimates to estimate the IGE and IRC within each of our 21 rolling 10-year cohort groups. The results are shown in the two panels of Appendix Figure 10. The top panel shows the mean IGE from 100 repetitions, with a 95 percent range, and the bottom panel shows the same for the IRC. The means closely follow our full estimation with relative differences of less than 0.006 for the IGE and 0.003 for the IRC; and the 95 percent ranges are narrow, less than 0.04 on average for both measures (slightly less in the middle of the period, slightly more at the ends).

4.2. *Dynamic Consistency: Robustness to the Addition of New Data*

A central advantage of our two-stage method is its efficient use of data in projecting income estimates into the future. To illustrate this, we analyze retrospectively the robustness of our findings to the addition of new waves of data, estimating logarithmic income and rank regressions using only data available in previous years, and comparing the results to estimates based on all currently available data. Thus, we first restrict the data to the 2006 wave and earlier, and implement our two-stage method to yield mobility estimates for ten-year cohort groups of sons born no later than 1971. We then repeat this process using data up to 2008 for sons born no later than 1973; and so on for each wave, up to the latest, which uses all data to 2016 for sons born no later than 1981.

Figure 6 presents the results of these estimations. Its upper panel presents the six IRC estimates, each based on an additional wave of data. The six graphs are virtually indistinguishable, with correlations of 0.93 and higher between pairs of IRC series, over shared cohort-groups. The level of IGE estimates, in the lower panel, increases slightly as we add more years of data, presumably due to reduced measurement error, but the graphs move very much in tandem, with correlations of 0.96 and higher between pairs of IGE series over shared cohort-groups. This close correspondence between mobility estimates derived from earlier available data and estimates derived from all available data to 2016, suggests that the trends estimated from current data for the youngest cohort groups should similarly hold up as new waves of data are collected in the future.

5. Conclusion

In this paper, we present a two-stage approach for estimating intergenerational income mobility. The first stage estimates proxies for parents' and childrens' average lifetime family income from annual income observations, education and race; and the second stage uses these proxies to derive a range of mobility measures. A key advantage of this method is its use of extensive income histories, available for many individuals, to construct first-stage age-income profiles that effectively address life-cycle bias and substantially reduce measurement error also for individuals with more limited income histories, in particular for more recent cohorts of sons. In addition, this method is less sensitive to the specifics of income measurement, and as the first stage is not logarithmic, not sensitive at all to how zero-income observations are handled. Moreover, separating the two stages allows us to directly assess the accuracy of our lifetime income proxies before applying them in the second stage to measure mobility. Finally, deriving multiple mobility measures based on the same lifetime income variable, allows us to present a consistent, multi-dimensional picture of intergenerational income mobility as it varies over time.

To demonstrate this, we apply our method to linked longitudinal data from the United States PSID, for sons born between 1952 and 1981, and their fathers. We first show that our lifetime income proxies are closely correlated with actual lifetime income averages for a subsample of males for whom longer income histories are available. We then use these proxies to derive measures of absolute, relative and positional income mobility, which we average within ten-year cohort-groups of sons, focusing on the three disjoint groups born in 1952–61, 1962–71 and 1972–81. Our findings indicate a decline in multiple dimensions of intergenerational mobility in lifetime family income over this period.

Absolute upward mobility, measured as the share of sons with greater lifetime family incomes than their fathers, declined from 67 percent for the 1952–61 cohort-group to 62 percent for the 1972–81 group. Intergenerational income growth was progressive throughout the period, dynastic income regressing to the mean, but the degree of progressiveness declined slightly. The IGE of lifetime family income increased, from 0.425 for the 1952–61 cohort group to 0.532 for the 1972–81 cohort-group, as did the intergenerational correlation of family income, the correlation of the logarithm of family income, and the slope of the conditional median, all indicating a decline in relative mobility. Rank mobility also declined, as indicated by the IRC increasing from 0.419 for the 1952–61 cohort-group to 0.489 for the 1972–81 cohort-group. Finally, we note that a retrospective analysis of the IGE and IRC using only data available in earlier years shows that estimates based on restricted data are very closely correlated with our current estimates using all available data. This suggests that our current estimates of trends in intergenerational income mobility in lifetime family income should hold up well as new waves of data are added.

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Supporting Information

Additional supporting information may be found in the online version of this article at the publisher's web site:

Figure A1: Two Age-Income Profiles, by Years of Schooling, White Males

Figure A2: Mean Annual Family Income (2012 PCE Adjusted Dollars), the Mean to Median Ratio, and the Gini Coefficient of Annual Family Income, 1967-2016, Our Pairs Sample Compared to National United States Data; Each Data Point Represents a Year

Figure A3: Fathers' and Sons' Average Lifetime Family Income, by Sons' Cohort Group, 2012 PCE Dollars, with 45° Lines

Figure A4: The Correlation Between the Logarithms of Lifetime Income of Sons and their Fathers; and the Ratio of Sons' to Fathers' Standard Deviations of Log (Lifetime Income)

Figure A5: Quadratic Local Approximation of Logarithmic (IGE) Regressions Using Stata's LOWESS Procedure; All Cohorts of Sons, 1952-81

Figure A6: Quadratic Local Approximation of Rank-Rank Regressions, Using Stata's LOWESS Procedure; by Sons' Cohort-Group

Figure A7: Estimates of the IGE of Lifetime Family Income Between Sons and Fathers, by Sons' Cohort Group, with First-Stage Age-Income Profiles Fixed Over Time, and Allowed to Vary

Figure A8: The Intergenerational Elasticity (IGE) and Rank Correlation (IRC) of Predicted Family Income at Ages 40, 45 and 35, Between Sons and Fathers, by Sons' Cohort Group

Figure A9: Mobility Estimates with and without Adjusting Income for Family Size (by Dividing Father's Family Income by the Square Root of the Number of Family Members)

Figure A10: Mean IGE and IRC with 95% Range, from 100 Repetitions. In Each, We Drew a 50% Subsample With Replacements from the Base Sample, and Used it to Estimate the First Stage

Table A1: First-Stage Estimation Result

Table A2: Mean Number and Mean Age of Income Observations, and Share of Explained Variance, by Sons' Ten-Year Cohort Group, Separately for Fathers and Sons

Table A3: Fathers' and Sons' Estimated Lifetime Family Income, by Cohort-Group

Table A4: Fathers' and Sons' Relative Average Lifetime Family Income, by Sons' Birth Cohort and Father's Income Quartile

Table A5: Estimates of the Intergenerational Elasticity (IGE) and Rank Correlation (IRC) by Rolling Ten-Year Cohort-Groups, Correlations of Log (Income), and the Son to Father Ratio of Standard Deviations in log (Income)

Table A6: IGE Estimates Between Fathers and Sons, Main Specification; Selected Papers