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Moshe Justman and Anna Krush

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Moshe Justman[†] and Anna Krush[‡]

[†] Department of Economics, Ben Gurion University, Beer Sheva, Israel; and Melbourne Institute of Applied Economic and Social Research, The University of Melbourne

[‡] Department of Economics, Ben Gurion University, Beer Sheva, Israel

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Melbourne Institute of Applied Economic and Social Research

The University of Melbourne

Victoria 3010 Australia

***Telephone* (03) 8344 2100**

***Fax* (03) 8344 2111**

***Email* melb-inst@unimelb.edu.au**

***WWW Address* http://www.melbourneinstitute.com**

Abstract

We use PSID data to 2008 to consider changes in the intergenerational elasticity (IGE) of income in fifteen successive ten-year cohort-groups of sons aged 36-45 between 1997 and 2011. Regressing sons' estimated lifetime income on fathers' income within each group, we obtain fifteen IGE estimates, which exhibit a significant rising trend, as do intergenerational correlations and rank correlations. The Gini coefficient of sons' lifetime income within these groups exhibits a correlation of 0.71 with our IGE estimates, leading us to conclude that as the United States has become economically less equal in recent years it has also become less mobile.

JEL classification: J62

Keywords: Intergenerational mobility, intergenerational elasticity of income, income inequality, attenuation bias

1. Introduction

Rising income inequality throughout the industrialized world in recent decades has triggered a heightened interest in determining whether this rise in inequality has been offset by a concomitant rise in economic opportunity, or aggravated by reduced opportunity. There is empirical evidence of a positive association between rising inequality and declining intergenerational mobility in a cross-section of OECD countries, which has been referred to as "The Great Gatsby Curve" (Corak, 2013). But the issue of greater interest is whether a co-movement of rising inequality and declining intergenerational mobility has also occurred within countries, over time. To this end a number of recent studies have sought to identify a time-trend in intergenerational income mobility, inversely measured as the intergenerational elasticity (IGE) of income.

The IGE of income quantifies the persistence of income inequality, a larger elasticity indicating slower regression to the mean and hence less mobility.[1] Estimates of the IGE at a single point in time generally regress the logarithm of sons' or daughters' income on the logarithm of their parents' income from matched parent-child data; the slope of the regression is then an estimate of the IGE. One might consider tracking the IGE over time by repeating this estimation in successive time frames, to identify a trend, and thus determine whether economic opportunity has been widening or narrowing. However, this has proved more difficult than it might seem. Recognizing that it is the elasticity of children's *lifetime* income with respect to their parents' *lifetime* income that is of interest implies that estimating the IGE of lifetime income in even a single year requires almost a century of income data: data on the

[1] There are, of course, other measures of intergenerational income mobility, some of which we apply here; and there are other dimensions of mobility—in education, occupation, social class. See Solon's (1999) seminal overview and more recent surveys by Björklund and Jantti (2008) and Black and Devereux (2011).

lifetime income of the parents of older individuals nearing retirement in that year, and data on the subsequent income of younger individuals just entering the work force, until they retire; tracking the IGE over, say, twenty years requires that many more years of data. Moreover, measurement of the IGE is highly sensitive even to small changes in data sources, variable definitions and statistical procedures, and thus requires a long, uniformly defined data panel on the incomes of parents and their children. The United States Panel Study of Income Dynamics (PSID), the longest series of longitudinal data we have, is not nearly long enough; and even if it were, the answers we would obtain from a full set of data would be largely of historical interest. The challenge is therefore to construct well-supported estimates of the IGE of lifetime income measured consistently over time from less than a full set of data, in a manner that allows inferences to be drawn on the presence of a trend.

This is the issue we address here, using PSID data to 2008 to determine existence of a time trend in recent years in the IGE of sons' lifetime family income with respect to their fathers' lifetime family income.[2] Specifically, we follow the IGE of income of sons aged 36-45 over fifteen years between 1997 and 2011, in effect estimating the IGE of lifetime income for fifteen successive rolling ten-year cohort groups, the first born between 1952-61 and the last between 1966-75. This is done in two stages: we first separately estimate, within each cohort group, fathers' and sons' predicted family income at age 40, as a proxy for their lifetime incomes; and then regress the logarithm of sons' lifetime family income on the logarithm of their fathers' family income within each group. This yields a series of fifteen successive elasticity estimates, which we examine for evidence of a trend.

[2] Our aim here is purely descriptive. We make no attempt to identify the separate effects of nature and nurture on intergenerational mobility or to distinguish between the contributions of circumstance and effort.

Previewing our results, we find an initial IGE of 0.421 (with a standard error of .047) for the first cohort-group, of sons aged 36-45 in 1997 (born in 1952-1961). IGE values then rise to a peak value of 0.516 (with a standard error of .053) for sons aged 36-45 in 2007 (born in 1962-71); and subsequently decline moderately to a final value of 0.483 (with a standard errors of 0.47) for sons aged 36-45 in 2011 (born in 1966-75). Regressing the fifteen cohort-group elasticities on a time trend, we obtain an average annual increase of .0037 with a *p*-value of 0.5%, implying a statistically significant decline in the intergenerational mobility of men's lifetime family income in this period.

To further test our findings we also estimate alternative, positional measures of intergenerational mobility for each ten-year cohort group: the intergenerational correlation (IGC) of income, which estimates mobility in relation to the standard deviation of the earnings distribution, and Spearman's *rho*, the intergenerational rank-correlation of income, a purely positional measure. The series of fifteen IGCs exhibits an average annual rise of .0027 with a *p*-value of 3%, and Spearman's *rho* increases annually with an average slope of .0048 significantly positive at a *p*-value of 0.1%. Thus both the IGC and Spearman's *rho* confirm our finding of a decline in intergenerational mobility. In addition, we apply our two-stage procedure to estimate the IGE of men's earnings and find that it increases from an initial level of .443 for the earliest cohort-group to a final level of .528 for the most recent cohort-group. Regressing the fifteen cohort-group elasticities on a time trend we obtain an estimated annual increase of .0055 with a *p*-value less than 0.1%, an even steeper decline than that found for men's family income in this period.

These findings depart from previous work similarly aimed at discerning a trend in intergenerational income mobility from PSID data: Mayer and Lopoo (2005), who analyzed variation over time in the IGE of sons' family income at age 30 for cohorts of sons born between 1949 and 1965 from income data to 1995 and found no significant trend for these years;[3] Lee and Solon (2009) who find no evidence of a trend in year-specific IGEs for sons born between 1952 and 1975, from income data to the year 2000; and Hertz (2007) whose preferred method, using the same data as Lee and Solon (2009), estimates cohort-specific IGEs and finds a weak upward trend that is not significantly different from zero.

We attribute the difference between these earlier findings and our own to the greater uniformity we apply in obtaining our successive estimates and to our longer data set. We follow a fixed number of cohorts of constant age in each year; apply our data requirements uniformly over time; separate estimation of the age structure from identification of a trend in in the IGE of income; and allow the age structure of income to vary over time. And our longer data set, which extends to 2008, provides more accurate measures for the lifetime income of the youngest cohorts studied, when much of the decline in mobility appears to have occurred.[4] When we apply our method to PSID data restricted to the year 2000 the IGE estimates for the first seven cohort-groups—though similar to those obtained from our full set of data—do not exhibit an upward trend.

Finally, we use the results of the first stage of our analysis to calculate the Gini coefficient of lifetime income within each cohort-group of sons, as a measure of inequality in lifetime income. We find that this succession of fifteen Gini coefficients exhibits a significant

[3] They find no significant trend in the period as a whole but suggest a possible nonlinear pattern in IGE that rises for sons born between 1949 and 1953 and declines subsequently. However, Lee and Solon (2009, Figure 1) and Hertz (2007, Figure 4) extend their analysis to sons born in later years, and find that this falling trend disappears as later IGE estimates rise.

[4] We consider the same set of cohorts as Lee and Solon (2006, 2009) and Hertz (2007). However, they observe the youngest cohort to age 25 where we observe the same cohort to the age of 33.

rising trend, which matches the trend of reduced mobility we found, with a highly significant correlation of 0.715 between the IGE and Gini coefficients. Increased inequality has gone hand in hand with reduced intergenerational mobility, mirroring the similar empirical association observed in a cross-section of developed countries, noted above (Corak, 2013).

The rest of the paper is organized as follows: Section 2 elaborates on the methodological issues involved and how they are addressed in the previous work of Mayer and Lopoo (2005), Lee and Solon (2009) and Hertz (2007). Section 3 describes the data we use, sets out our methodology, presents our estimates of the IGE of sons' family income and shows their declining trend. Section 4 presents concurring results for alternative measures of economic mobility. Section 5 shows that the declining trend in intergenerational mobility we found is matched by a rising trend in income inequality. Section 6 discusses our findings in light of earlier work; and Section 7 concludes.

2. Previous work in this vein

Our analysis of matched father-son PSID data to identify a time trend in men's IGE of family income in the United States builds on the previous work of Mayer and Lopoo (2004, 2005), Lee and Solon (2006, 2009), and most directly, Hertz (2007).[5] Two potential sources of bias arise in such studies. The first is classical measurement error in parental income, on the right-hand side of the regression equation, biasing estimates of the IGE towards zero (Solon, 1992;

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[5] Fertig (2004) applies a similar approach to analyzing the IGE of earnings for five rolling five-year cohort-groups, from 1952-56 to 1956-60, and finds a halving of the IGE in this shorter period, with much volatility. Two alternative approaches are Aaronson and Mazumder's (2008) estimate of the IGE from unmatched census data, using child's state-of-birth as an instrument for parental income; and Levine and Mazumder's (2002) estimate of intergenerational mobility in the United States from sibling data.

Mazumdar, 2005). The second is attenuation bias, which arises from variation in the shape of the age-income profile—and therefore in the relation between permanent and annual income—in a manner that depends on education, race and possibly other variables (Grawe, 2006; Haider and Solon, 2006; Böhlmark and Lindquist, 2006). Consequently, as several studies have found, estimates of the IGE drawn from annual income data vary systematically with the age at which income is measured (Hertz, 2007).

Solon (1992) and Mazumder (2005) demonstrated the large bias that can arise when parents' permanent income, on the right-hand side, is measured from a small number of annual observations.[6] Mayer and Lopoo (2005) control for this by averaging parental income when sons are between 19 and 25 years of age; Lee and Solon (2009) and Hertz (2007) average parental income when sons are between 15 and 17. Previous research indicates that neither is sufficient to fully remove the bias, however, as Lee and Solon (2009) note, as long as the number of observations is fixed, any remaining bias should be stable over time, and so should not distort estimates of a trend in intergenerational mobility, on which these studies focus.

The problem of attenuation bias is more difficult to resolve. Mayer and Lopoo (2005) address it by using a single cohort—30-year olds—in each year to estimate the IGE of family income in that year. As Lee and Solon (2009, p. 766) note, “the elasticity estimate for the 1960 cohort uses the cohort’s earnings only in 1990 and ignores the 1960 cohort’s earnings in

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[6] Solon (1992) shows that averaging data over three years produces IGE estimates twice as large as Becker and Tomes’ (1986) initial estimates, which were based on a single year of parental income. Mazumder (2005, Table 5) shows that increasing the number of years over which parental income is averaged from 4 to 7 increases sons’ IGE estimates by 15%; and from 7 to 10 years by a further 11%.

all other years," resulting in large sampling error and strong year-to-year fluctuations.[7] In addition, it is not clear to what extent a trend in the IGE of 30-year olds is representative of the population at large. Higher income individuals have on average a steeper gradient of income at age 30 than low-income individuals, so that income differences at age 30 are smaller than differences in lifetime income. For an estimate of a time trend in intergenerational mobility drawn exclusively from observations on 30-year olds to be valid for the population as a whole, the ratio between income differences at age 30 and differences in lifetime income, across income levels, must remain stable over time; Mayer and Lopoo (2005) do not offer evidence that this holds.[8]

Lee and Solon (2009) and Hertz (2007) base their estimates on a much broader set of data, using observations on multiple cohorts in each year, while recognizing the need to adjust sons' and daughters' income for age: as the average age of sons in the PSID increases over time, and the elasticity of lifetime income with respect to annual income increases with age, using all available income data in each year without any adjustment would introduce an upward bias in the time trend. To avoid this, both studies incorporate adjustments for sons' age and individual characteristics in their estimation procedure.[9] Lee and Solon (2009)

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[7] Annual elasticity estimates reported in Mayer and Lopoo (2004) exhibit very strong fluctuations, as do Lee and Solon's IGE estimates based on sons' income at age 30 (2009, Figure 1). To smooth their estimates, Mayer and Lopoo (2005) estimate IGEs for four-year rolling sub-samples, though these also exhibit large fluctuations.

[8] Large changes in schooling levels and in returns to schooling in this period suggest that it may not.

[9] There is also a possibility attenuation bias in parental income, as it is measured at a time determined by the son's age, and parental age has increased over time. However, this turns out to be negligible. Hertz (2007, p. 36) estimates that the increase in fathers' age, from 42 to 45, generated a bias in the estimated IGE of 0.002 over 24 years—"small enough to ignore."

assume a fixed age-structure of income for the whole period and estimate it simultaneously with the time trend in a single equation, estimating year-specific IGEs from an interactive term of parental income with year of observation. Thus they do not allow variation in the age-structure of income over time and derive their estimates of year-specific IGEs from all income observations in that year.[10] Hertz (2007) preferred method allows the age structure of income to vary over time, estimating the age structure of income for earlier cohorts separately for each cohort while pooling together younger cohorts for which fewer observations are available. Our approach, described in the following section, is closer to Hertz' preferred method in this respect. He estimates cohort-specific (rather than year-specific) IGEs.[11]

3. Data, methodology and estimation results

Our data source is the Panel Study of Income Dynamics (PSID) from its inception in 1968 until 2008, with data collected annually until 1996 and bi-annually thereafter. The advantage of the PSID for estimating the IGE of income is its long duration, which allows parental income to be measured over long periods of time, hence with greater accuracy; its principal disadvantage is that annual sample sizes are small. The PSID has two parts: a cross-sectional, national sample drawn from the Survey Research Center (SRC), and a national sample of low-income families, the Survey of Economic Opportunities (SEO). We focus on fathers' and

[10] They offer estimates of the IGE for 1977 to 2000, however their estimate for 1977 is based only on the 1952 cohort, aged 25 in that year, their estimate for 1978 is based only on the 1952 and 1953 cohorts, aged 25-26 in that year, and so on, while their estimate for the year 2000 is based on all 24 cohorts. Their estimates for earlier years are thus less representative of intergenerational mobility in those years and subject to greater sampling error than their estimates for later years.

[11] As Mayer and Lopoo (2005) consider only a single cohort in each year the distinction between cohort-specific and year-specific IGEs does not apply.

sons' family money income, which includes the taxable income of all earners in the family, from all sources, and transfer payments. Following Lee and Solon (2009) we use only SRC data to avoid over-representation of low-income families. We top-code the data at annual earnings of $150,000 in 1967 dollars, and bottom-code the data at annual earnings of $150 in 1967 dollars (annual earnings below $150 are set equal to zero). All income data is converted to 2008 dollars using the Consumer Price Index.

We extract from the PSID pairs of sons and their fathers, limiting our attention to sons born between 1952 and 1975 for whom we have at least three non-zero observations on income from the age of 29 and at least five years of non-zero observations for their father's income until the age of 64. We aggregate the data into fifteen rolling ten-year cohort groups by the sons' year of birth, separately for fathers and sons. Thus we have thirty data sets: a data set for sons aged 36-45 in 1997 (born between 1952 and 1961) and a data set for their fathers; a data set for sons aged 36-45 in 1997 (born between 1953 and 1962) and a data set for their fathers; and so on until the 36-45 cohort-group for 2011 (born between 1966 and 1975).[12] Descriptive statistics on the cohort groups are presented in Table 1. The cohort groups are similar in size, ranging from 389 to 421 with the youngest cohort-groups about 5% smaller than the earliest ones. The average age at which fathers' income is measured decreases from 51 to 44 and the average age at which sons' income is measured decreases from 38 to 33. The average number of non-zero income observations per father is uniformly high, varying between 22 and 26 among cohort-groups, while the number of income observations per son declines sharply from 15 to 5.

[12] There is, of course, extensive overlap between adjacent data sets, so that most sons and fathers have multiple estimates of permanent income.

We predict sons' family income at age 40 by estimating the following regression:

$$y_{ij} = \alpha_{1i} \cdot D_i + \alpha_2 \cdot Age_{ij} + \alpha_3 \cdot Age_{ij}^2 + \alpha_4 \cdot Race_i \cdot Age_{ij} + \alpha_5 \cdot Race_i \cdot Age_{ij}^2 + \alpha_6 \cdot Educ_i \cdot Age_{ij}$$
$$+ \alpha_7 \cdot Educ_i \cdot Age_{ij}^2 + \alpha_8 \cdot Marstat_i \cdot Age_{ij} + \alpha_9 \cdot Marstat_i \cdot Age_{ij}^2 + \varepsilon_{ij} \qquad (1)$$

where y_{ij} is son i's family income in year j; D_i is an individual dummy variable; Age_{ij} is the son's age in year j minus 40; $Race_i$ represents dummy variables for white, black and "other"; $Educ_i$ represents a set of dummy variables for less than 8 years of schooling, 8-10 years, 11-12, 13-15, 16, and 17 and over; $Marstat_i$ is a set of dummy variables for marital status; ε_{ij} is an i.i.d. error term; and α_{1i} , α_2 ... α_9 are the regression coefficients. Then α_{1i}, the coefficient of the individual dummy variable, is the predicted value of son i's income at age 40, which we take as our proxy for lifetime income.[13] Fathers' income at age 40 is predicted from a similar equation, without the interaction of race and age.

Mean and standard deviations of predicted incomes at age 40 by cohort-groups are presented in Table 2 for fathers and in Table 3 for sons. Fathers' mean incomes are stable over time while the R^2 values decline slightly from a value of 0.67 for the first cohort-group to 0.61 for the last cohort group. If this decline in precision affects our estimated elasticities it should introduce a downward bias in the time trend of IGE, which would work against the appearance of a rising trend. Sons' mean incomes (in constant dollars) increase over time as do their standard deviations, while the R^2 values are more or less constant, about 0.60.

We then regress, within cohort-groups, the logarithm of sons' predicted income at age 40 on the logarithm of fathers' predicted income at age 40 to obtain estimated IGE values for

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[13] We assume that within each cohort-group and race and education categories, sons' earnings follow the same parabolic shape over the life cycle; and the same separately for fathers.

sons aged 36-45 between 1997 and 2011. The results, presented in Table 4 and Figure 1, show a clear upward trend. Estimated IGE values range from a low of 0.421 for the oldest cohort-group of sons, aged 36-45 in 1997 (born between 1952-61), to a high of .516 for sons aged 36-45 in 2007 (born between 1962-71), with an estimated elasticity of .483 for our youngest cohort-group, aged 36-45 in 2011 (born between 1966-75). Standard errors of the estimates range between .045 and .055. Regressing cohort-group elasticity estimates on a time trend yields a rising annual slope of 0.0037 estimated with a standard error of 0.0011 and a *p*-value of 0.5%, indicating a significant decline in sons' intergenerational income mobility.

4. Other measures of economic mobility

To further test our findings we consider other approaches to estimating intergenerational income mobility.[14] We first apply two alternative measures to our family income data: Pearson's correlation coefficient between fathers' and sons' family income, and a purely positional measure, Spearman's rank correlation. We then estimate the IGE of sons' earnings with respect to their fathers' earnings, using the same two-step procedure described in the preceding section.

4.1 Correlations and rank correlations

The IGE indicates the extent to which income differences are carried over from one generation to the next and is therefore affected by change in the relative variation of fathers' and sons' income. The correlation coefficient measures movement in income in relation to the relevant standard deviation. Intergenerational correlations between fathers' and sons' predicted family income at age 40, by cohort-groups, are presented in the third column of

[14] For further discussion of the different dimensions of intergenerational mobility, see among others, Fields (2006), Van de Gaer, Schokkaert, and Martinez (2001) and Dardanoni (1993).

Table 5, and graphically as the broken line in Figure 2. They range from a low of .280 for our earliest year (our oldest cohort-group) to a high of .352 for our most recent year (our youngest cohort-group), with estimated standard errors under .05, and 5% confidence intervals of ±.095. These are smaller than the IGE values and similar to the range of values obtained by Björklund and Jantti (1997). Regressing these correlation coefficients on a time trend yields a slope coefficient of 0.0027, with a standard error of .0011 and a *p*-value of 3%.

Values of Spearman's *rho*, the rank correlation between fathers' and sons' family income, are presented in the fourth column of Table 5, and graphically as the unbroken line in Figure 2. It is an inverse measure of pure positional mobility.[15] Rank correlations range from an initial value of .415 for our earliest year to .486 for our most recent year, each with a 5% confidence intervals of ±.08. Regressing these values on a time trend yields a slope coefficient of 0.0048 with a standard error of .0008 and a *p*-value less than 0.1%. Both measures behave similarly, confirming the falling trend in intergenerational income mobility indicated by our IGE estimates.

4.2 The intergenerational elasticity of sons' earnings

Next we estimate the IGE of sons' earnings with respect to their fathers' earnings using the same two-step procedure applied in Section 3 to derive the IGE of family income. We first predict sons' and fathers' earnings at age 40 within ten-year cohort-groups using equation (1), with earnings replacing family income. Descriptive statistics are presented in Table 6, and

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[15] Spearman's *rho* is almost perfectly correlated with another positional measure, Kendall's *tau*, which offers a different formulation of positional mobility: it is the probability that a randomly chosen pair of father-son pairs exhibits a reversal in the order of lifetime family income from one generation to the next, such that father A has greater income than father B while son B has greater income than son A.

statistics on the earnings equations for fathers and sons are presented in Tables 7 and 8. The patterns are similar to those of family income, shown in Table 1, 2 and 3. Again the R^2 values are more or less constant over time for sons, but decline slightly for fathers; a decline in the precision with which fathers' income is measured introduces a downward bias in the time trend, which again works against finding an upward trend in the IGE of earnings. We then regress the logarithm of sons' predicted earnings at age 40 on their fathers' predicted earnings at age 40 within each group to obtain a series of fifteen cohort-group-specific elasticities.

These are presented in Table 9 and Figure 3, and again show a clear upward trend. Values range from a low of 0.443 for sons aged 36-45 in 1997, our earliest year, to a high of .538 for sons aged 36-45 in 2009, and an estimated elasticity of .520 for our youngest cohort-group, of sons aged 36-45 in 2011. Standard errors of the estimates range between .043 and .057. Regressing these fifteen successive elasticity estimates on a time trend yields an annual slope of 0.0055 with a standard error of 0.0011 and a *p*-value less than 0.1%, indicating a significant rising trend in the IGE of sons' earnings, steeper than the rising trend in family income. This difference in slopes between the IGE of family income and the IGE of sons' earnings may be attributable to the increase in women's labor force participation in recent decades partially offsetting the decline in men's earnings mobility.

5. Less equal *and* less mobile

These findings indicate that rising income inequality in the United States in recent years has not been offset by a concomitant rise in economic opportunity, but rather made worse by reduced opportunity. To highlight this we calculate cohort-group specific Gini coefficients from our estimates of lifetime income as our measure of income inequality. They are presented in the third column of Table 10, alongside our previously estimated values of the

IGE of income. The Gini values for lifetime income are lower than the more familiar Gini coefficients of annual income, the difference between the two reflecting *intra*-generational income mobility. The Gini coefficients exhibit a general downward trend.[16] Figure 4 plots the Gini coefficient for each cohort-group against its corresponding IGE estimate. It highlights the co-movement of declining intergenerational mobility and rising inequality over time. The correlation between IGE values and Gini coefficients is .715, statistically significant with a *p*-value of 0.3%. Rising inequality in the United States in recent years has been accompanied by a concomitant decline in intergenerational mobility: the United States has become less equal *and* less mobile, mirroring the positive association between rising inequality and declining intergenerational mobility found in a cross-section of OECD countries noted in the introduction (Corak, 2013).[17]

6. Discussion

Our finding of a significantly rising time trend in the IGE of sons' family income is at variance with the earlier studies of Mayer and Lopoo (2005) and Lee and Solon (2009), which found no such trend for sons, and Hertz (2007) which found a statistically insignificant upward trend. We attribute this to our longer data set, which extends to 2008, and to our two-stage procedure, which enables us to apply our data requirements uniformly over time;

[16] Regressing Gini coefficients on a time trend yields a slope estimate of 0.0015 with a *p*-value of 0.5%.

[17] Marks (2013) observes that the country-studies of mobility on which these cross-sectional comparisons draw use disparate methods and data sources, with the lower values observed in Canada and Scandinavian countries estimated from comprehensive administrative data, where the higher values observed in the United States and United Kingdom are estimated from panel data. This may account for some of the cross-sectional variation. Our analysis reproduces this pattern of increased inequality matched with declining intergenerational mobility within a single country, over time, using a fixed methodology and a uniform data source.

separates estimation of the age structure from identification of a trend in in the IGE of income while allowing the age structure of income to vary over time; and bases its estimate of the IGE in a given year on a fixed number of same-aged cohorts. We elaborate first on the attributes of our procedure and then on the advantage of our longer data set.

Any empirical analysis of the variation of IGE of income over time must limit its attention to parent-child pairs on which there is reliable income data over a long enough period, effectively applying its analysis to a non-random, more homogeneous sub-sample of the PSID population. This introduces both selection and attrition bias. Selection bias arises in this case because all IGE analyses based on the PSID examine samples constrained by the composition of the initial panel selected nearly half a century ago, and therefore not fully representative of the United States population in later years. Such studies, ours included, examine behaviour of the IGE within this population, implicitly assuming that the trends it exhibits are representative of the population at large. Intergenerational elasticity estimates based on the PSID are also vulnerable to bias from non-random attrition (Lee and Solon, 2009, note 11). However, if such bias remains stable over time identification of a trend in the IGE remains possible. Fitzgerald et al. (1998) did not find evidence of attrition bias in IGE estimates derived from the PSID; and Hertz (2007), applying their suggested correction to his estimations, finds it makes little difference. Our uniform application of data requirements over time promotes stability in selection and attrition biases.

In addition, by repeatedly predicting income at age 40 within ten-year cohort groups, we separate our estimation of the age structure of income from estimation of the time trend in the IGE, and allow the age structure of income to vary over time. This is closer to Hertz' (2007) preferred approach than to Lee and Solon's (2009), which assumes a fixed age-structure of income for the whole period, and estimates it simultaneously with the time trend.

Hertz (2007) allows the age structure of income to vary, but due to data limitations adopts an asymmetric procedure that estimates it separately for each of the earlier cohorts while pooling together the smaller younger cohorts. We use the same method for all cohorts.

Our study diverges from earlier work also in estimating year-specific IGEs from ten-year rolling cohort-groups of sons aged 36-45 in each year, from 1997 to 2011. Ideally, we would want to compare the IGE for the entire working population as it varies over time, by estimating for each year the lifetime IGE of, say, all men between the ages of 25 and 65 (Hertz, 2007, p. 30). This amounts to following a rolling window of forty cohorts, and estimating an IGE for each window, which is not possible with the limited data available. Moving windows of ten-year cohort groups are long enough to provide a characteristic picture of income mobility at a given time,[18] but short enough to allow us to observe variation in IGE across time. Clearly one could label our year-specific estimates in other ways—e.g., as referring to sons aged 35-44 from 1996 to 2010, but it would not change our substantive finding of a rising trend in the IGE of income concomitant with rising inequality.

Mayer and Lopoo (2005) focus on the current income of thirty-year olds as their year-specific indicator, a measure subject to large sampling error, which may not be uniformly representative of the IGE in the population at large.[19] Hertz' preferred method focuses on cohort-specific variation in IGE, similarly focusing on one cohort per year but utilizing all income data. This sharply reduces sampling error in estimating lifetime income but does not address the noise inherent in using a single cohort as representing the IGE in the population at

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[18] The narrower the window the less representative it is of intergenerational mobility in a given year and the more volatile its behavior

[19] We expanded on this in Section 2.

large. Lee and Solon (2009) directly estimate year-specific IGEs from all income observations in that year by incorporating an interactive term of parent's income with the year in which the son's income is observed. Thus their estimate for 1977 is based only on sons aged 25 in that year, born in 1952; their estimate for 1978 is based on sons aged 25-26 in that year, born in 1952-53; and so on. Estimates for earlier years are therefore less representative and more volatile than those for later years, as they recognize (Lee and Solon, 2009, p. 769). We do not estimate the IGE for these earlier years but in the short later period in which there is an overlap between our estimates and theirs—years in which their IGE estimates are based on twenty cohorts or more—we find that our estimates and theirs describe a similar rising trend (Figure 5).

Finally, we note the importance of our longer data set, extending to 2008, to our findings: it provides more accurate measures for the lifetime income of the youngest cohorts studied—when much of the decline in mobility appears to have occurred—than could be obtained from the shorter panels used by previous studies. Lee and Soon (2009) and Hertz (2007) base their findings on data that run to the year 2000, and cannot observe their youngest cohort, born in 1975, beyond the age of 25. This is too early an age from which to extrapolate a reliable proxy for lifetime income. [20] Mayer and Lopoo's (2005) data end yet earlier, in 1995. To illustrate the crucial importance of our extra years of data we apply our method to PSID data to the year 2000, and derive IGE estimates for the first seven ten-year cohort-groups from this restricted data set. In Table 11 and Figure 6 we compare these estimates to the IGE estimates derived from the full data set. The two series are very similar for the earlier

[20] We include only sons for which we have at least three income observations over the age of 29. As later PSID data is biannual, all sons in our sample have income observations at age 33 or later.

years and then diverge somewhat, with a correlation of 0.588 between them. Most relevant for our purpose, the estimates derived from the restricted data set do not exhibit an upward trend.[21]

Our finding of a rising IGE accords with Aaronson and Mazumder (2008) who use decennial census data from 1940 to 2000 to estimate trends in intergenerational mobility over a longer period. They use son’s state-of-birth as an instrumental variable, in lieu of matching fathers and sons, in a regression of sons' earnings on fathers' family income, and similarly find a decline in men's intergenerational. The cohort effects they estimate "spike up sharply" for cohorts born between 1956-60 and 1961-65, followed by a moderate decline, which leads them to conclude that "earnings are regressing to the mean more slowly now than at any time since World War II."[22] Relatedly, Blanden et al. (2002) find a decline in economic mobility in Britain from a comparison of two cohorts, born in 1958 and 1970.

7. Conclusion

In this paper we identify a trend in men’s intergenerational income and earnings mobility in the United States by using PSID data through 2008 to estimate the intergenerational elasticity (IGE) of income—an inverse measure of relative mobility—for a succession of fifteen rolling ten-year cohort-groups of sons aged 36-45, from 1997 to 2011. To allow gradual change in the age-structure of earnings, conditioned on education and race, we apply a two-stage procedure, the first stage of which predicts sons’ and fathers’ family income at age 40, as a

[21] When a linear regression is fitted to the estimates derived from the restricted data the slope obtained is negative, though not significantly different from zero.

[22] The two-sample IV estimator they use produces estimates that are upward biased; inference on a trend in intergenerational mobility assumes that this bias is constant over time.

proxy for their lifetime family income, within each of these fifteen cohort-groups. The second stage then regresses sons' lifetime income on fathers' lifetime income, in logarithm form, within each group to obtain fifteen successive group-specific estimates of the IGE of income.

We find a statistically significant rising trend in this series of fifteen successive IGEs, with a statistically significant average annual increase of .0037, indicating a decline in intergenerational income mobility. Similar significant trends are evident when intergenerational mobility is measured (inversely) by the intergenerational correlation (IGC) of income and by the intergenerational rank-correlation of income. Moreover, repeating the same procedure for fathers' and sons' earnings we again find a statistically significant rising trend in elasticities, with an average annual increase of .0055, indicating yet a steeper decline in men's intergenerational earnings mobility. This difference in slopes between the IGE of family income and of sons' earnings may be attributable to the increase in women's labor force participation in recent decades partially offsetting the decline in men's earnings mobility.

We attribute the difference between these results and earlier studies, which failed to find a significant trend, to our longer data set, which extends to 2008, and provides a better basis for estimating lifetime income for more recent cohorts; to our use of ten-year uniformly aged cohort-groups to estimate year-specific IGEs; to the greater uniformity of our method over time; and to our two-stage procedure, which separates estimation of the age structure of income from our analysis of dynamic change in the IGE of income while allowing the age structure of income gradually to vary over time.

Finally, to highlight the relationship between inequality and mobility over time, we use our estimates of lifetime income to compute the Gini coefficient of lifetime income within each cohort-group, and plot the fifteen Gini values we obtain against their matching IGE estimates. We find a significant correlation of .715 between the two series, an empirical association between declining mobility and rising inequality that mirrors a similar pattern observed in a cross-section of developed countries (Corak, 2013). These findings indicate that the United States has become less equal *and* less mobile in recent years.

The two-stage method we use here to estimate measures of intergenerational mobility and inequality from the lifetime income of rolling cohort-groups suggests several lines of further research. A similar approach could be integrated in studies that decompose IGE by income group, education, race or other individual characteristics, to try and determine how different strata of society have been affected and why the IGE of income has risen.[23] In addition, our approach should also be relevant to the growing literature on *intra*-generational income mobility and its implications for the measurement of income inequality (Burkhauser et al., 2011).

[23] The two lines of research are connected, with variation in IGE indicating its determinants. See, among others, Corak and Heisz (1999), Grawe (2004) and Cardak, et al., (2013).

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Table 1. Descriptive statistics: family income sample, by cohort-group

			Fathers		*Sons*	
Year when sons are aged 36-45	*Sons' year of birth*	*N*	*Mean number of observations*	*Mean age of observed income*	*Mean number of observations*	*Mean age of observed income*
1997	1952-61	413	22	51	15	38
1998	1953-62	417	23	51	14	37
1999	1954-63	417	23	50	13	37
2000	1955-64	421	24	50	12	36
2001	1956-65	412	24	49	12	36
2002	1957-66	401	25	49	11	36
2003	1958-67	410	25	48	10	36
2004	1959-68	406	25	48	10	36
2005	1960-69	412	26	47	9	35
2006	1961-70	397	26	46	8	35
2007	1962-71	393	26	46	7	35
2008	1963-62	390	26	45	7	34
2009	1964-73	391	26	45	6	34
2010	1965-74	389	25	44	5	34
2011	1966-75	394	25	44	5	33

Table 2. Fathers' predicted family income at age 40, by cohort-group

Year when sons are aged 36-45	*Sons' year of birth*	*N*	*Mean predicted income*	*Mean log of predicted income*	*Standard deviation*	*R^2 of the prediction regression*
1997	1952-61	413	84,111	11.203	.498	0.67
1998	1953-62	417	84,661	11.208	.501	0.66
1999	1954-63	417	85,433	11.212	.506	0.65
2000	1955-64	421	85,227	11.217	.501	0.62
2001	1956-65	412	84,711	11.205	.506	0.62
2002	1957-66	401	84,238	11.200	.499	0.64
2003	1958-67	410	83,638	11.196	.489	0.65
2004	1959-68	406	83,302	11.194	.484	0.64
2005	1960-69	412	81,788	11.182	.483	0.61
2006	1961-70	397	83,200	11.196	.495	0.60
2007	1962-71	393	83,166	11.199	.487	0.61
2008	1963-62	390	84,378	11.219	.479	0.60
2009	1964-73	391	83,863	11.214	.487	0.60
2010	1965-74	389	84,572	11.217	.500	0.62
2011	1966-75	394	85,273	11.228	.505	0.61

Table 3. Sons' predicted family income at age 40, by cohort-group

Year when sons are aged 36-45	*Sons' year of birth*	*N*	*Mean predicted income*	*Mean log of predicted income*	*Standard deviation*	*R^2 of the prediction regression*
1997	1952-61	413	97,480	11.350	.519	0.60
1998	1953-62	417	99,511	11.365	.529	0.60
1999	1954-63	417	99,537	11.371	.517	0.58
2000	1955-64	421	101,875	11.387	.520	0.59
2001	1956-65	412	102,979	11.398	.520	0.59
2002	1957-66	401	103,588	11.396	.533	0.61
2003	1958-67	410	105,216	11.407	.539	0.61
2004	1959-68	406	105,173	11.401	.550	0.62
2005	1960-69	412	105,218	11.404	.552	0.60
2006	1961-70	397	106,996	11.427	.546	0.57
2007	1962-71	393	106,777	11.414	.569	0.58
2008	1963-62	390	105,320	11.403	.563	0.59
2009	1964-73	391	109,756	11.436	.575	0.58
2010	1965-74	389	109,369	11.442	.561	0.58
2011	1966-75	394	107,722	11.438	.527	0.60

Table 4. Estimates of the intergenerational elasticity of family income between fathers and sons, by cohort group

Year when sons are aged 36-45	*Sons' year of birth*	*N*	*Estimated elasticity*	*Standard error*
1997	1952-61	413	0.421	0.047
1998	1953-62	417	0.439	0.047
1999	1954-63	417	0.461	0.045
2000	1955-64	421	0.475	0.045
2001	1956-65	412	0.442	0.046
2002	1957-66	401	0.453	0.048
2003	1958-67	410	0.456	0.050
2004	1959-68	406	0.478	0.051
2005	1960-69	412	0.481	0.051
2006	1961-70	397	0.472	0.050
2007	1962-71	393	0.516	0.053
2008	1963-62	390	0.502	0.054
2009	1964-73	391	0.478	0.055
2010	1965-74	389	0.461	0.052
2011	1966-75	394	0.483	0.047

Table 5. Estimates of the intergenerational correlation and rank correlation in family income between fathers and sons, by cohort group

Year when sons are aged 36-45	*Sons' year of birth*	*N*	*Intergenerational correlation*	*Intergenerational rank correlation*
1997	1952-61	413	0.280	0.415
1998	1953-62	417	0.290	0.425
1999	1954-63	417	0.303	0.427
2000	1955-64	421	0.323	0.437
2001	1956-65	412	0.313	0.412
2002	1957-66	401	0.287	0.411
2003	1958-67	410	0.281	0.420
2004	1959-68	406	0.284	0.430
2005	1960-69	412	0.298	0.429
2006	1961-70	397	0.314	0.448
2007	1962-71	393	0.315	0.471
2008	1963-62	390	0.295	0.468
2009	1964-73	391	0.300	0.463
2010	1965-74	389	0.340	0.471
2011	1966-75	394	0.352	0.486

Table 6. Descriptive statistics: earnings sample, by cohort-group

			Fathers		*Sons*	
Year when sons are aged 36-45	*Sons' year of birth*	*N*	*Mean number of observations*	*Mean age of observed income*	*Mean number of observations*	*Mean age of observed income*
1997	1952-61	433	22	49	14	37
1998	1953-62	432	22	49	13	37
1999	1954-63	431	23	48	13	37
2000	1955-64	431	23	48	12	36
2001	1956-65	421	24	48	11	36
2002	1957-66	405	24	47	11	36
2003	1958-67	410	25	46	10	36
2004	1959-68	404	24	46	9	35
2005	1960-69	407	25	45	8	35
2006	1961-70	390	25	45	8	35
2007	1962-71	383	25	44	7	35
2008	1963-62	377	25	44	6	34
2009	1964-73	374	25	44	6	34
2010	1965-74	369	24	43	5	34
2011	1966-75	373	24	43	5	33

Table 7. Fathers' predicted earnings at age 40, by cohort group

Year when sons are aged 36-45	*Sons' year of birth*	*N*	*Mean predicted income*	*Mean log of predicted income*	*Standard deviation*	*R^2 of the prediction regression*
1997	1952-61	433	69,203	10.982	.560	0.67
1998	1953-62	432	69,303	10.979	.570	0.68
1999	1954-63	431	69,267	10.981	.553	0.66
2000	1955-64	431	69,118	10.990	.536	0.66
2001	1956-65	421	68,798	10.984	.528	0.64
2002	1957-66	405	67,757	10.967	.526	0.63
2003	1958-67	410	66,445	10.956	.516	0.62
2004	1959-68	404	66,069	10.950	.515	0.62
2005	1960-69	407	64,483	10.937	.501	0.60
2006	1961-70	390	64,237	10.932	.503	0.60
2007	1962-71	383	63,341	10.921	.501	0.60
2008	1963-62	377	63,054	10.922	.490	0.59
2009	1964-73	374	62,395	10.914	.490	0.59
2010	1965-74	369	62,819	10.910	.502	0.61
2011	1966-75	373	62,891	10.915	.505	0.62

Table 8: Sons' predicted earnings at age 40, by cohort group

Year when sons are aged 36-45	*Sons' year of birth*	*N*	*Mean predicted income*	*Mean log of predicted income*	*Standard deviation*	*R^2 of the prediction regression*
1997	1952-61	433	68,080	10.968	.556	0.60
1998	1953-62	432	69,877	10.985	.573	0.61
1999	1954-63	431	69,742	10.988	.565	0.60
2000	1955-64	431	72,530	11.007	.587	0.61
2001	1956-65	421	73,093	11.020	.574	0.61
2002	1957-66	405	72,804	11.009	.588	0.62
2003	1958-67	410	74,733	11.029	.598	0.62
2004	1959-68	404	74,673	11.025	.602	0.61
2005	1960-69	407	74,791	11.029	.596	0.59
2006	1961-70	390	76,000	11.053	.586	0.59
2007	1962-71	383	75,349	11.036	.601	0.59
2008	1963-62	377	73,437	11.011	.598	0.62
2009	1964-73	374	75,897	11.042	.606	0.61
2010	1965-74	369	71,593	10.992	.597	0.63
2011	1966-75	373	70,447	10.971	.612	0.63

Table 9. Estimates of the intergenerational elasticity of earnings between fathers and sons, by cohort group

Year when sons are aged 36-45	*Sons' year of birth*	*N*	*Estimated elasticity*	*Standard error*
1997	1952-61	433	0.443	0.043
1998	1953-62	432	0.456	0.043
1999	1954-63	431	0.448	0.044
2000	1955-64	421	0.493	0.047
2001	1956-65	421	0.475	0.048
2002	1957-66	405	0.505	0.050
2003	1958-67	410	0.520	0.051
2004	1959-68	404	0.529	0.052
2005	1960-69	407	0.501	0.054
2006	1961-70	390	0.491	0.054
2007	1962-71	383	0.508	0.056
2008	1963-62	377	0.514	0.057
2009	1964-73	374	0.538	0.058
2010	1965-74	369	0.523	0.056
2011	1966-75	373	0.520	0.057

Table 10. Estimates of the intergenerational elasticity of income between fathers and sons, and the Gini coefficient of sons' income, by cohort group

Year when sons are aged 36-45	*Sons' year of birth*	*N*	*Estimated elasticity*	*Gini coefficient*
1997	1952-61	433	0.421	0.244
1998	1953-62	432	0.439	0.260
1999	1954-63	431	0.461	0.254
2000	1955-64	421	0.475	0.259
2001	1956-65	421	0.442	0.259
2002	1957-66	405	0.453	0.267
2003	1958-67	410	0.456	0.268
2004	1959-68	404	0.478	0.273
2005	1960-69	407	0.481	0.274
2006	1961-70	390	0.472	0.274
2007	1962-71	383	0.516	0.279
2008	1963-62	377	0.502	0.273
2009	1964-73	374	0.478	0.270
2010	1965-74	369	0.461	0.275
2011	1966-75	373	0.483	0.260

Table 11. Estimates of the intergenerational elasticity of income between fathers and sons, by cohort group, for full and restricted data sets

Year when sons are aged 36-45	*Sons' year of birth*	*Full data set, to 2008*		*Restricted data set, to 2000*	
		N	*Estimated IGE*	*N*	*Estimated IGE*
1997	1952-61	433	0.421	412	0.430
1998	1953-62	432	0.439	417	0.447
1999	1954-63	431	0.461	417	0.468
2000	1955-64	421	0.475	420	0.467
2001	1956-65	421	0.442	406	0.420
2002	1957-66	405	0.453	394	0.418
2003	1958-67	410	0.456	394	0.436

Figure 1. Intergenerational elasticity of family income between fathers and sons by year, for sons aged 36-45 in each year, with regression line

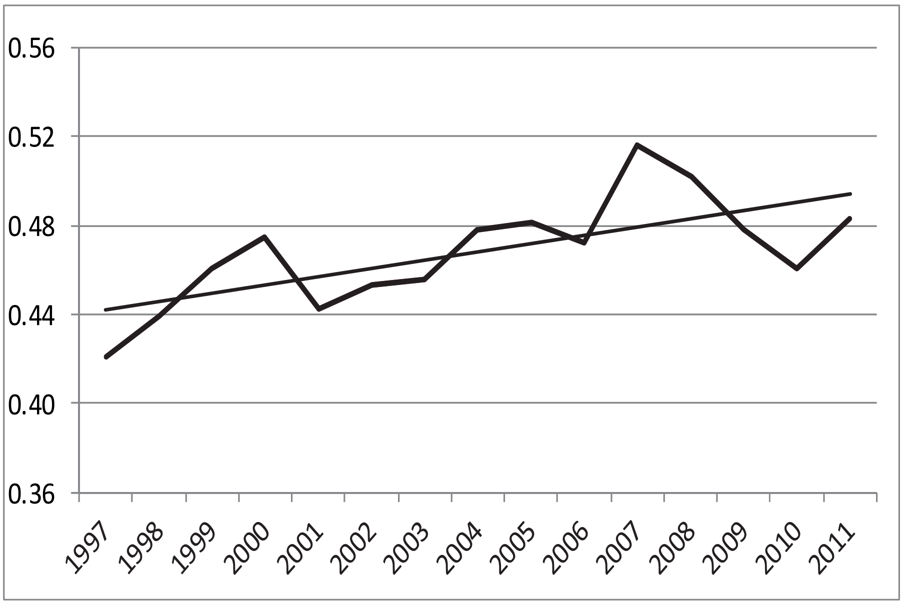


Figure 2. Pearson and Spearman correlations between fathers' and sons' family income by year, for sons aged 36-45 in each year, with regression lines

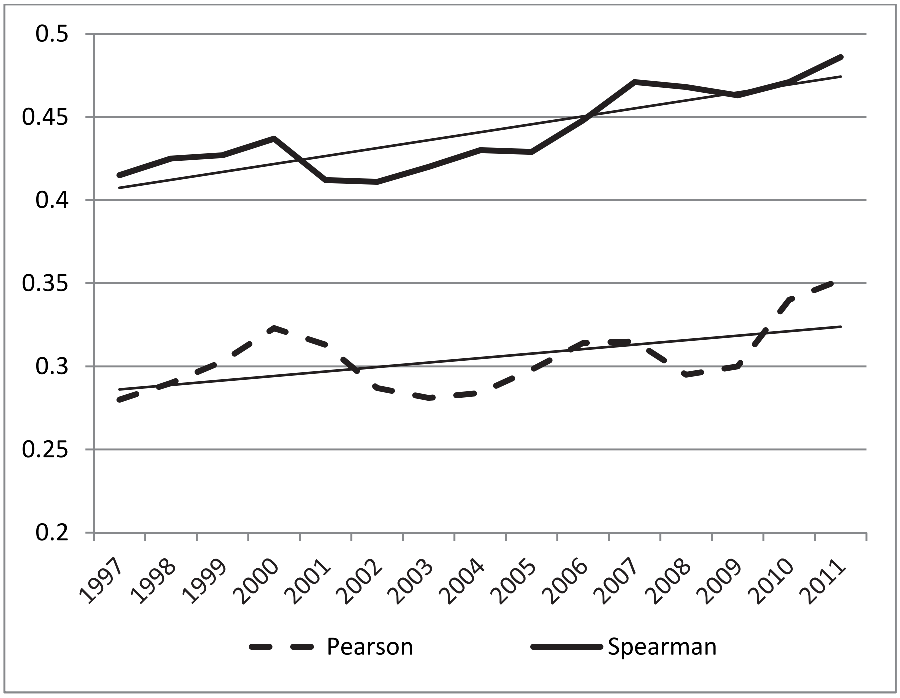


Figure 3. Intergenerational elasticity of earnings between fathers and sons, by year, for sons aged 36-45 in each year, with regression line

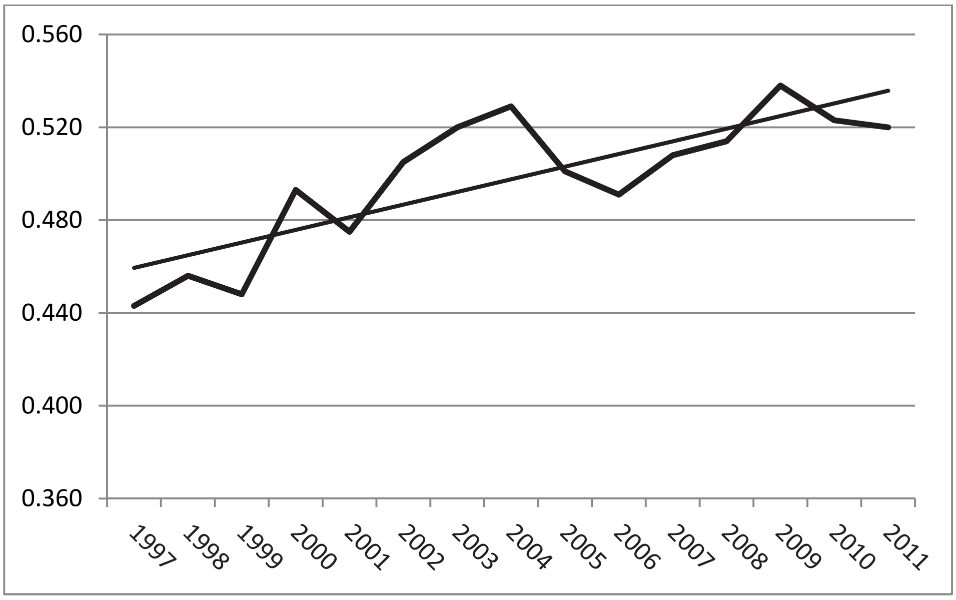
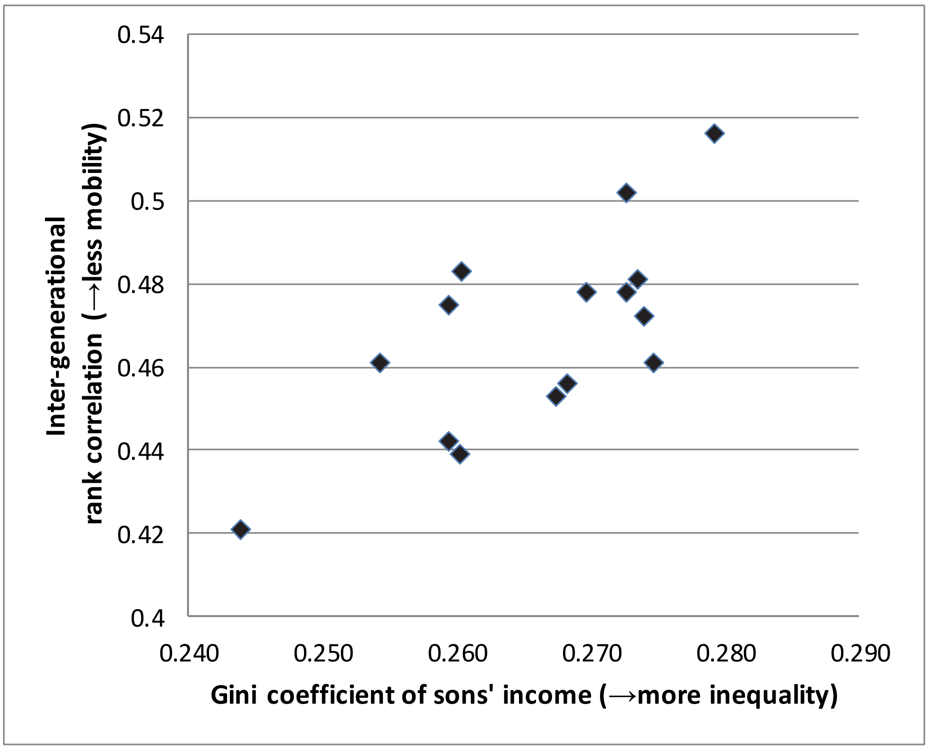
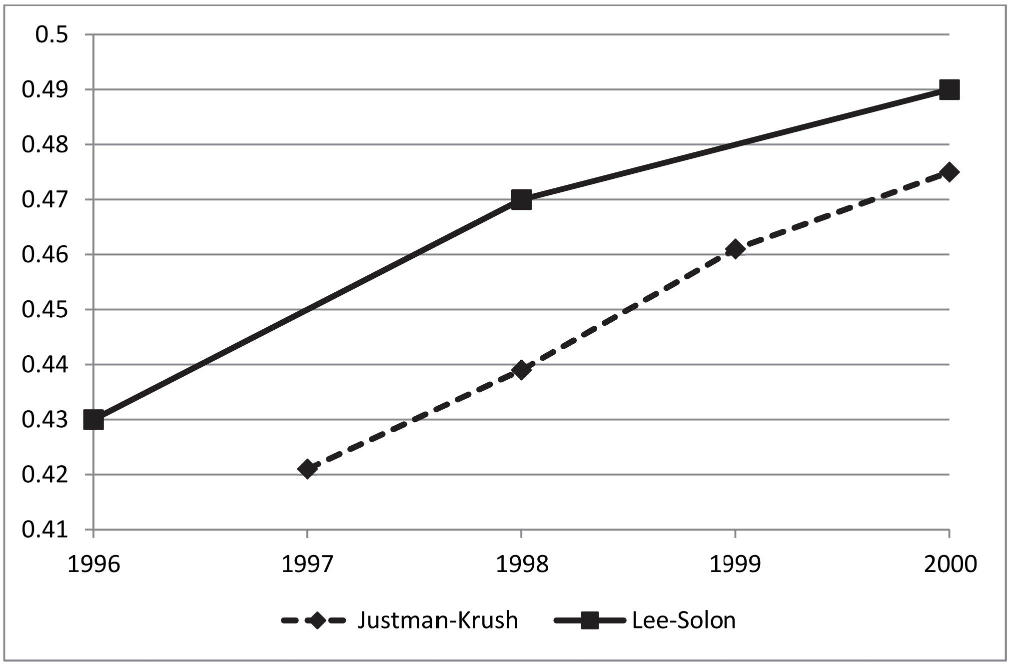


Figure 4. Inequality and intergenerational mobility in lifetime income



Each data point represents a ten-year cohort-group of sons aged 36-45 from 1997 to 2011.

Figure 5. Comparison of our estimates of sons' IGE of income for 1997-2000, to Lee and Solon's (2009) estimates for 1996, 1998, 2000



Our estimates from Table 4 above refer to sons aged 36-45 in 1997-2000;
Lee and Solon's (2009) are year-specific coefficients from their Table 1.

Figure 6. Estimated IGE of income by cohort-group, full and restricted data sets