Trends in Intergenerational Income Mobility in Sweden

Anders Björklund
Swedish Institute for Social Research (SOFI)
Stockholm University and IZA

Markus Jäntti
Swedish Institute for Social Research (SOFI)
Stockholm University

Lena Lindahl
Swedish Institute for Social Research (SOFI)
Stockholm University

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Abstract

The level of intergenerational income mobility is by now reasonably well understood across countries, but for many countries, including Sweden, little is known about its change over time. We exploit newly available data to extend existing estimates for Sweden using parental incomes starting in 1960 and offspring income up until 2007. This allows us to examine how the persistence of long-run income has changed for offspring cohorts born in 1945 to 1962 using a reasonably wide window to measure long-run incomes for both parents and offspring. By exploiting these longer time series, we are also able to study how sensitive estimated trends are to parts of the life-cycle relative to longer-run income measures. The paper thus aims to provide the first set of estimates for trends in intergenerational income persistence in Sweden, and explore their robustness to measurement problems that stem from variations in the income process across the life cycle.

1 Introduction

Economists are interested in the strength of the intergenerational association both because theoretical work in many areas predicts such associations – such as the theoretical model of Becker and Tomes (1979, 1986) – and because differences in the strength of intergenerational associations may inform judgements on social justice. If a person's economic status depends strongly on the economic status of his or her parents, we would say that life chances are unequal in the sense that economic wellbeing hinges on which family the person is born into and is therefore beyond his or her control. From a Swedish perspective, this would be regarded as a deviation from an equality-of-opportunity norm. While equality of opportunity is an important policy goal, it is unlikely that completely eradicating intergenerational associations is the socially optimal policy. Families provide the environment in which children learn many cognitive and non- cognitive skills, and both right to privacy and economic efficiency dictate that parents should largely be allowed to raise their children as they see fit. The policy implications of changes in intergenerational mobility are therefore not entirely clear. However, careful study of both the extent of mobility and how that mobility has changed across time is an important starting point for evaluating equality of opportunity (for a review, see Björklund and Jäntti, 2009).

It was earlier believed that mobility was more prevalent in societies with more inequality. We now know that, if anything, there is more intergenerational income mobility in societies where there is less income dispersion at the time the offspring were growing up. That is, mobility and inequality are negatively rather than positively correlated (see Björklund, Jäntti, and Roemer, 2011; Blanden, 2009) Sweden, like the other Nordic countries, has turned out to have more intergenerational mobility than most other countries.

While still low by international standards, Swedish income inequality is now greater than it has been since modern measurement of inequality began in 1975. For instance, the Gini coefficient of disposable income increased from a low of about 0.20 around 1981 to about 0.30 in 2009 (Jenkins et al., 2011). This increase, although starting from a low level, has received a great deal of public attention from both media and politicians over the past two decades. One natural question is whether this change has coincided with a decrease in intergenerational mobility, as the cross-country evidence suggests might be the case?

Although the recent 15-20 years have seen an upsurge of studies on intergenerational income mobility, research on changes in mobility is still in its infancy. The main reason for this is because few countries have data that allow changes

to be studied, and in those that do, longitudinal survey data tend to have so small samples that firm evidence is hard to come by. Sweden is well placed for the study of trends, as we can rely on register data to go back in time, and the datasets are also large enough to allow for reliable evidence. This research field has also been plagued by methodological problems. Recent research into the relationship between short-run and long-run income at the individual level suggests that changes in short-run income variability may lead researchers to erroneously infer that the intergenerational association of long-run income has changed even if, in truth, it is the nature of short-run variation that has changed. This means that the statistical procedures used to investigate change in mobility need to explicitly model both the short-run and long-run income processes. Again, the availability of long series of annual income from registers for both the parental and offspring generations facilitates such modeling.

This paper thus will provide the first evidence on changes in intergenerational income mobility and cross-sectional inequality in all of Sweden for men and women in three-year birth cohorts starting in 1945 and extending to 1962. This paper pays careful attention to the possibility that the relationship between short- and long-run economic status can change across time and confound the measurement of change in intergenerational mobility.

2   Literature review

The early contributions to the literature were based on US data and typically used only single-year measures of fathers' income, see for example Becker and Tomes (1979) and Behrman and Taubman (1985). The general finding in these early papers was that the associations in income between fathers and sons seemed very weak and accordingly, the US was described as a very mobile society.

During the subsequent years, a number of papers came to question this conclusion. The argument was that the use of short-run income measures by necessity makes the intergenerational association much weaker compared to measures based on long-run income. In particular, short-run income measures are thought of as being more influenced by a transitory component, which captures events such as sickness and unemployment or just good or bad luck. Taking into account the problem of measuring parental income across multiple years, based on the classical measurement error model, Solon (1992) and Zimmerman (1992) used four and five-year income averages of the father's income and estimated the father-son association to be twice as large as in the early studies. From this point on, it became

more or less standard to use averages over five income years when the aim was to capture long-run income. Results from studies on earnings dynamics, however, show that even five-year averages of earnings provide estimates of the intergenerational elasticity that are biased down by close to 30 percent, (Mazumder, 2003). Using an income average over 16 years, Mazumder (2005) estimated an intergenerational association three times as large as the ones found in the early studies. Since then, the US has ultimately been viewed as a far less mobile society than was previously thought.

Importantly, even with access to such long time series, a critical underlying assumption is that the bias is stable over time. A general critique of the generalized errors-in-variables model is given in the study by Nybom and Stuhler (2011). They argue that the model does not fully eliminate life-cycle bias due to heterogeneous income profiles and to heterogeneity being correlated with individual and family characteristics. In particular, long series of annual incomes for both fathers and sons, the find substantial remaining biases in intergenerational elasticity estimates even after applying the insights from GEIV models.

In this paper, we address this problem by taking the heterogeneous income profiles into account, i.e., by estimating both random intercepts and random income growth rates. In particular, we are able to account for how individual current income relate to long-run income at the same time as we allow for these processes to correlate between generations.

The remainder of this literature review is devoted to earlier results for Sweden and to evidence on trends in intergenerational income elasticities across countries. Starting with Gustafsson (1994), there is by now a number of studies of intergenerational income mobility on Swedish data. For example, Björklund and Jäntti (1997) and Österberg (2000) report estimates on intergenerational income mobility around .24, based on five and three income years respectively for both generations. Additional studies address particular questions, e.g Björklund and Chadwick (2003) allow the intergenerational income mobility to vary depending on the extent to which the sons lived with their fathers. The results show that having lived with the father or ot matters a lot for the results of income mobility. For example the income elasticity of sons who never lived with their biological fathers, was equal to zero. Björklund, Lindahl, and Plug (2006) use a sample of adoptees born in Sweden to investigate the origins of intergenerational income associations. In the study, the authors compare the intergenerational association between adoptive parents and their adoptive children. One of the findings is that for adoptees, both their adoptive and birth parents matter. Lindahl (2008) estimates elasticities by birth order and family size of the offspring and finally, Björklund,

Roine, and Waldenström (2010) focus on intergenerational mobility in the very top of the income distribution and find that income mobility is substantially lower compared to the rest of the distribution.

There are studies on trends in intergenerational income mobility in only a handful of countries. Several exist for the United States, and the overall results point in quite different directions. U.S. trend studies we are aware of include Corcoran (2001), Levine and Mazumder (2002), Fertig (2003), Mayer and Lopoo (2005), Hertz (2007), Lee and Solon (2009) and Hauser (2010).

Mayer and Lopoo (2005) use the Panel Study of Income Dynamics (PSID) to estimate a large but statistically insignificant downward trend in income elasticity for cohorts of sons born in 1949-1965. Levine and Mazumder (2002) find similar results using PSID, but when using the National Longitudinal Study (NLS), they find a significant increase in elasticity, from .22-.41. Fertig (2003) uses the PSID and finds a statistically significant decrease in the intergenerational elasticity of earnings between fathers and sons. Finally, Lee and Solon do not find any significant changes in mobility for cohorts born between 1952-1975. An early study by Hauser (2010) finds no trend in the intergenerational persistence of occupation, education or income from the 1960s to the 1990s. In all these studies, the authors rely on the assumption that life-cycle bias is constant over time, an assumption that is explicitly made in Lee and Solon and Mayer and Lopoo.

For the United Kingdom, Blanden et al. (2004) compare cohorts born in 1958 and 1970 to find some evidence for declining mobility. This finding was subsequently found wanting by Erikson and Goldthorpe (2010). Lefranc (2011) studies trends in intergenerational elasticities in France for cohorts born between 1931 and 1976 using two-sample IV methods, and finds persistence declines between the 1930s and 1950s cohorts, and increases as he reaches the younger cohorts. Also using TSIV methods, Lefranc, Ojima, and Yoshida (2011) find stable trends in intergenerational elasticities in Japan. Studies of change in intergenerational mobility in Norway by Hansen (2006) anbd Bratberg, Nielsen, and Vaage (2007) are inconclusive. Studies for Finland by Pekkala and Lucas (2007) and Pekkarinen, Uusitalo, and Kerr (2009) suggest a substantial declide across cohorts, driven in large part by comprehensive school reform.

For Sweden, we are only aware of a single previous study that allows for the examination of trends in intergenerational elasticity, namely the paper by Lindahl et al. (2011), which shows little or no trend. That is based on data from a single city, Malmö, however, and the focus is not so much on trends for recent cohorts as on persistence across multiple generations. Björklund, Jäntti, and Lindquist (2009) study the importance of family background using a different, more com-

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Table 1 Overview of studies containing estimates of intergenerational (IG) income (or earnings) associations on Swedish data

Study	Focus
1. Gustafsson (1994)	General IG associations
2. Björklund & Jäntti (1997)	General IG associations and a Sweden-US comparison
3. Österberg (2000)	General IG associations
4. Björklund & Chadwick (2003)	Differences among family types
5. Björklund, Lindahl & Plug (2006)	Nature and nurture
6. Hammarstedt & Palme (2006)	Mobility among immigrants
7. Jäntti et al. (2006)	Non-linearities and cross-national differences
8. Hirvonen(2008)	Gender differences and non-linearities
9. Lindahl (2008)	Differences by family size and birth order
10. Palme et al (2011)	Mobility across three generations
11. Nybom & Stuhler (2011)	Heterogeneous income profiles and life-cycle bias

prehensive measure, namely the brother correlation. Their evidence suggests that between cohorts born in the 1930s and 1960s, the importance of family background did, indeed, decline, although most of the decline occurred early on and between the late 1950s to 1960s cohorts, brother correlations increased a little. This finding is given additional support by Björklund, Jäntti, and Roemer (2011), who find that the equality of opportunity in Sweden declined marginally between cohorts born in the 1950s and and 1960s.

3 Data and raw trends intergenerational association in Sweden

Data

In order to gain insights into how income mobility has changed in in Sweden, we use data from several administrative registers, put together by Statistics Sweden, to estimate the intergenerational income elasticities for different birth cohorts.

A first and basic source is Statistics Sweden's so-called Multi-generational register. This is a register of all persons who were born 1932 and onward, and who

have ever received a unique national registration number from 1961 and onward.[1] For the Swedish population defined in this way, the register contains information about biological (and adoptive) parents and their national registration number. From this information, one can also infer which individuals are related as siblings; full siblings are those who have the same father and mother, half siblings are those who only have one parent in common. Our analysis sample is a 35 percent random sample of the Swedish male population defined in this register. We also use the Multi-generational register to identify parents.

The second source is the set of bidecennial censuses conducted from 1960 to 1980. We can identify our main sample of sons in the households of these censuses as well as other persons in the household. Thus we can determine whether our offspring generation lived with their biological parents or not in the fall of these census years.

The third source is Statistics Sweden's income register, which in turn come from the Swedish tax assessment procedure. A limitation is that such data are available for the whole population only from 1968 onwards. From that year the income register provides data on total income from all sources of income, from work, self employment, capital, real estate as well as some transfers (from 1974 onward). However, 10 percent of the population – more specifically those born the 5th, 15th and 25th in any month – income data are available 1960-66. For the early years, we are thus limited to a smaller subset of all our data when using the parent-child link. We these income data for both parents and sons. The earlier data for parents stem from their own compulsory tax assessments. In later years, when we measure sons' incomes, the source of the data is compulsory reports by employers to the tax authorities. The income concept we rely on total market income *(sammanräknad nettoinkomst)*.

With these data, we cover the association in long-run income between Swedish sons and daughters born 1945- 1962 and their parents' long-run income during their childhood. Whereas previous studies typically started observing parental income in 1970, we can go back ten more years in time. In combination with income data up to 2007, we can offer a much richer set of estimates than has been available in the previous literature.

For reasons we detail below, our preferred age range for observing incomes is 35-45 years of age for both fathers and offspring. Given that we have income data from 1960 to 2007, this limits the cohorts we can study. In particular, we report

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[1]The requirement that the persons must have been registered in Sweden from 1961 and onward implies that persons who died between 1932 and 1960 are not included.

estimates for different three-year birth cohorts of offspring, namely those born in 1945-1947, 1948-1950, 1951-1953, 1954-1956, 1957-1959 and 1960-1962. As the labour market for women has changed dramatically across the parental cohorts we examine, we limit our analysis to the study of links between offpring and father incomes although we do examine both father-son and father-daughter links. We require fathers to have been at least 20 years old at the birth of their offspring, which means that our fathers are born between 1925 and 1942.

Trends in estimated interegenerational associations

We compare intergenerational income associations for offspring, born between 1945-1962. Income mobility between generations is usually measured by the intergenerational elasticity which equals the percentage differential in children's income with respect to a marginal differential in parents' income. If the income variance is the same in both generations, the elasticity is equivalent to the intergenerational correlation in income. Because the correlation controls for changes in income distribution from one generation to the next, it comes closer to relative mobility in the sociological literature. We estimate both intergenerational elasticities and correlations using long-run income measures for both generations.

As we discussed in Section 2 and further detailed in Section 4, problems in measuring long-run incomes make simple estimates of interegenerational income elasticities and correlations subject to substantial biases. As the factors that drive those biases may well change across time, an observed change in the IGE may be driven not by a change in the slope of parent-offspring long-run incoe conditional expectation, but in those confounding factors. For instance, as discussed in Section 2, Björklund, Jäntti, and Lindquist (2009) find that the estimated brother correlation for the youngest cohorts of Swedish men increases from around 0.27 to around 0.37 when we take into account the time series structure of the underlying errors (going from white noise to AR(1) errors; Figure 2, Panel C)). While this corresponds to moving from a zero autocorrelation to the actual value of about 0.5 – a large change in the time-series structure of the random fluctuations – it does suggest that changes in the error process may drive observed differences in the importance of family background even if the true underlying parameter is unchanged.

However, in order to start the analysis, we show in Figure 1 the estimated intergenerational elasticities (Panel A) and correlations (Panel B). These numbers are estimated by taking the average of both offspring and parent income for ages 35-45, a common practice in the current literature. We work three-year intervals of

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Figure 1 Trends in estimated intergenerational income elasticity and correlation – Swedish men and women, offspring birth cohorts 1945-1962

A. Elasticity

B. Correlation
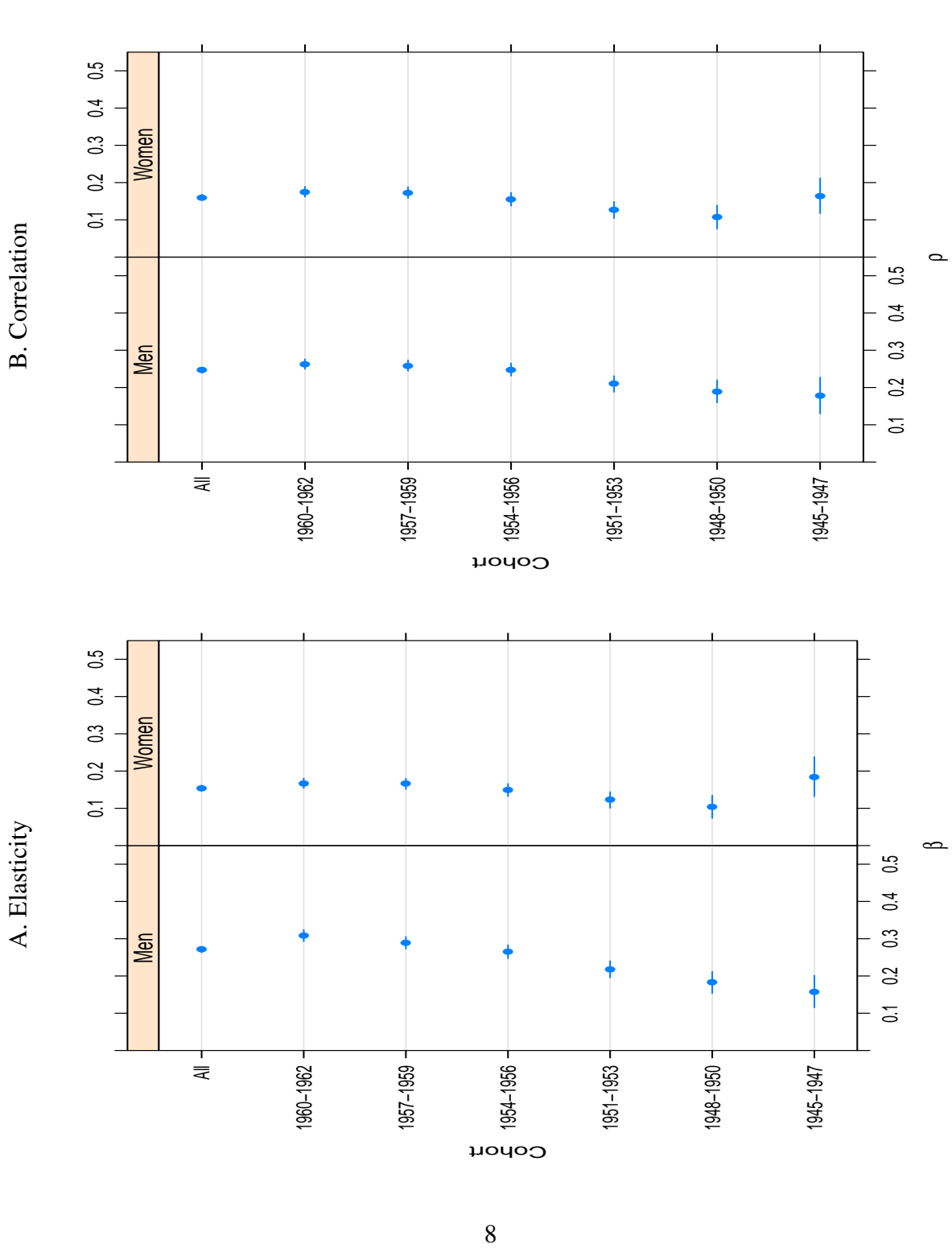
Note:

birth cohorts from 1945-1947, ..., 1960-1962. We do observe an upward trend in the estimated elasticity for father-son pairs from roughly 0.16 for the oldest cohort to about 0.31 for the youngest. For women, we do not find much of a trend – the elasticity declines substantially from the first to the secong cohort and increases again from the early 1950s to early 1960s cohorts. The estimated correlations follow the same time pattern as the elasticities.

4 Intergenerational associations with generalized errors-in-variables models and growth rate heterogeneity

The prototypical approach to estimating intergenerational elasticities is to depart from the basic Galtonian regression

$$\ln Y_{Oi} = \alpha + \beta \ln Y_{Pi} + \varepsilon_i \tag{1}$$

where Y_{ij} is a measure of the long-run or permanent income of generation $j (= Offspring, Parent)$ in family i. Denoting the natural logarithm of income $y = \ln Y$, the current practice is to use a measurement model for long-run income, based on average income, in at least the parental generation and often also in the offspring generation of

$$y_{ijt} = y_{ij} + v_{ijt}. \tag{2}$$

This is the classical measurement error model if the random fluctuations v are orthogonal to true long-run income is $y_{ij} \perp v_{ijt}$ and the v:s are identically and independently distributed.

An estimate of the IGE β using annual incomes for both parent and children has the probability limit

$$
\begin{aligned}
\mathrm{p}\lim \widehat{\beta} &= \frac{\mathrm{Cov}[y_{iOt}, y_{iPt}]}{\mathrm{Var}[y_{iPt}]} \\
&= \frac{\mathrm{Cov}[y_{iO}, y_{iP}] + \mathrm{Cov}[v_{iOt}, y_{iP}] + \mathrm{Cov}[y_{iO}, v_{iPt}] + \mathrm{Cov}[v_{iOt}, v_{iPt}]}{\mathrm{Var}[y_{iP}] + \mathrm{Var}[v_{iPt}] + 2\mathrm{Cov}[y_{iP}, v_{iPt}]}.
\end{aligned}
\tag{3}
$$

For the classical measurement error model (and assuming, additionally, that the random fluctuations v are uncorrelated across generations, the last three terms in the numerator in equation 3 are all zero. In that case, also the third term in the

denominator is zero, and only the presence of the random fluctuation is parental income in the denominator leads to downward bias.

Transitory errors in parental income have at least since Atkinson (1981) been recognized to lead to an errors-in-variables (downwards) inconsistency in the estimated intergenerational elasticity. Although few researchers have had access to the full lifetime incomes of parents, use of a multi-year average mitigate this source of inconsistency, as averaging across T years of parental income renders the source of the inconsistency $\text{Var}[v_{iPt}]/T$ instead of $\text{Var}[v_{iPt}]$. In the wake of the seminal paper to do this (Solon, 1992), many studies have been published that exploit this finding.

Recent work on so-called generalized-errors-in-variables (GEIV) model calls into question the assumption of classical measurement errors (Böhlmark and Lindquist, 2006; Haider and Solon, 2006). The GEIV model for the annual income process of an individual in family i in generation $j(= O\text{ffspring}, P\text{arent})$ (Haider and Solon, 2006)

$$y_{ijt} = \lambda_t y_{ij} + v_{ijt} \quad j = O, P. \tag{4}$$

As the process in equation 4 involves for both generations their permanent income, the parameter we are interested in estimating, the intergenerational elasticity β would, if permanent income was observed, be estimable from the population regression equation

$$y_{iO} = \beta y_{iP} + \varepsilon_i; \quad y_P \perp \varepsilon. \tag{5}$$

However, the age- or time-dependent factor loading λ_t leads to two additional source of bias in the IGE, namely the age/time point at which child incomes is measured – leading to a biased estimate of $\text{Cov}[y_{iO}, y_{iP}]$ – and when parental income is measured – leading to biased estimates of both $\text{Cov}[y_{iO}, y_{iP}]$ and $\text{Var}[y_{iP}]$.

Empirical evidence from both the United States and Sweden on the age profile of λ_t based on the GEIV model suggests that earnings early in life (even abstracting from a population age-earnings profile) are a downward-biased measure of lifetime earnings and later in life an upward-biased measure. Around age 40, at least in both the U.S. and Sweden, $\lambda_t \approx 1$ and deviations from a multi-year average are believed to be approximately classical, thus lending themselves to the analysis of intergenerational association of long-run income.

However, Nybom and Stuhler (2011) use nearly complete actual lifetime incomes for both Swedish fathers and sons. By comparing regression coefficients based on multi-year averages of sons income with that based on their full lifetime incomes, they find that the biases in the intergenerational elasticity estimates are still quite considerable. This suggests that even the GEIV model for how annual

incomes relate to permanent income is probably false. The results in Nybom and Stuhler (2011) are consistent with a model of intragenerational income processes that involve not only a random intercept (constant long-run income) but a random growth rate, so that individual "permanent" income depends on the age at which it is measured (but is deterministic for that individual).

Thus, as the results in Nybom and Stuhler (2011) are consistent with model that involves both a random intercept and a random growth rate. To capture a possible memory in the random fluctuations around such an individual income profile, we allow the fluctuationns to follow a simple ARMA structure:

$$
\begin{aligned}
y_{ijt} &= \alpha_{ij} + \beta_{ij}t + v_{ijt} \\
v_{ijt} &= \phi v_{ij,t-1} + u_{ijt} + \theta u_{ij,t-1} \\
u &\sim N(0, \sigma_u^2) \quad j = O, P
\end{aligned}
\tag{6}
$$

Here, α_{ij} is the random intercept, β_{ij} is the random growth rate and v_{ijt} is an ARMA(1,1) error term. As both the intercepts and the growth rates may be correlated across generations, we need to specify a bivariate population regression (where, for simplicity, we assume the intercepts and growth rates have their own population regressions):

$$
\begin{aligned}
\alpha_{iO} &= \gamma \alpha_{iP} + \varepsilon_{\alpha,i} \\
\beta_{iO} &= \delta \beta_{iP} + \varepsilon_{\beta,i}; \quad \alpha, \beta \perp \varepsilon_\alpha, \varepsilon_\beta
\end{aligned}
\tag{7}
$$

We assume that the parent's earnings intercept and growth rate have zero mean, positive variances $(\sigma_{\alpha_P}^2, \sigma_{\beta_P}^2)$ and may be correlated (ρ_P):

$$
\begin{bmatrix} \alpha_P \\ \beta_P \end{bmatrix} \sim F\left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sigma_{\alpha_P}^2 & \cdot \\ \rho_P \sigma_{\alpha_P} \sigma_{\beta_P} & \sigma_{\beta_P}^2 \end{bmatrix} \right)
\tag{8}
$$

Given the population regression equations 7, the child's intercept and growth rate are given by

$$
\begin{bmatrix} \alpha_O \\ \beta_O \end{bmatrix} \sim F\left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \gamma^2 \sigma_{\alpha_P}^2 + \sigma_{\varepsilon_\alpha}^2 & \cdot \\ \gamma \delta \rho_P \sigma_{\alpha_P} \sigma_{\beta_P} & \delta^2 \sigma_{\beta_P}^2 + \sigma_{\varepsilon_\beta}^2 \end{bmatrix} \right).
\tag{9}
$$

The random intercepts and growth rates are correlated across generations with regression coefficients γ_α, γ_β, δ_α and δ_β.

We are interested in the population regression $E[\ln Y_O | \ln Y_P]$ (across cohorts) which is a function of the parameters of both income generating processes. Rather

than work with the population regression parameters in equation 9, we use an approach exemplified in by Haider and Solon (2006) which is as follows. Assume, for simplicity, that an individual lives in perpetuity and that individual income in year t is given by equation 6. Then for a discount rate $r > \beta_{ij}$, the expected value of lifetime income is

$$V_{ij} \sum_{t=1}^{\infty} y_{ijt} = \sum_{t=1}^{\infty} (\alpha_{ij} + \beta_{ij}t) \approx \exp(\alpha_{ij})[(1+r)/(r-\beta_{ij})] \qquad (10)$$

The natural logarithm of the expected present value of lifetime income is thus

$$\ln V_{ij} \approx \alpha_{ij} + \beta_{ij}/r + r - \ln r. \qquad (11)$$

We measure the intergenerational elasticity in income by estimating this quantity for each offspring and parent and running the Galtonian regression in equation 1 on the resulting numbers In the absence of random growth rates, this is comparable to using an overtime average of individual income to approximate lifetime income.

Before we move on to the results, we should note that getting the income process right, or at the very least, to move away from the classical measurement error-based estimates, is important. For instance, consider what happens when, as in several recent papers, one relies on an over-time average of annual incomes. Let $\overline{y_{ij}} = 1/T \sum_1^T y_{ijt}$ be the average across ages 1 to T for generation j and let $e_{ijt} = y_{ijt} - \overline{y_{ij}}$ be the deviation of annual from that over-time-average income. If annual income follows the process in equation 6, then we have

$$
\begin{aligned}
e_{ijt} &= y_{ijt} - \overline{y_{ij}} \\
&= \alpha_{ij} + \beta_{ij}t + v_{ijt} - \frac{1}{T}\sum_1^T (\alpha_{ij} + \beta_{ij}t + v_{ijt}) \\
&= (\alpha_{ij} - \frac{1}{T}T\alpha_{ij} + \beta_{ij}t - \frac{1}{T}\beta_{ij}\frac{T(T+1)}{2} + v_{ijt} - \frac{1}{T}\sum_1^T v_{ijt} \qquad (12) \\
&= \beta_{ij}\left(t - \frac{(T+1)}{2}\right) + v_{ijt} - \frac{1}{T}\sum_1^T v_{ijt} \\
&= \beta_{ij}\tilde{t} + \tilde{v}_{ijt}.
\end{aligned}
$$

The last line of equation 12 suggests several way in which the deviation of annual from over-time-average income will display "non-classical" behavior:

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1. the deviations are strongly correlated across time, driven by three factors: the (possible) time-series structure of the deviations in equation 6, the time-average that is part of $\tilde{v}$, and the fact that the random growth rate β is present in every deviation;

2. the variance of the deviations depends on lifetime income. In particular, the variance for both the lowest and the highest income earners is larger than for those close to the average;

3. the variance of the deviations increases across time as the variance of the random growth rates is multiplied by age;

4. the deviations are correlated across generations if, as we posit in equation 7, the growth rates are intergenerationally correlated;

5. the intergenerational correlation in such deviations increases across age in both generations.

To demonstrate that there is something to these assertions, we show a plot of the empirical covariance matrix of annual deviations from long-run average income for ages 30-50 for both parents and children, where the offspring are born between 1950 and 1953. While there is some random variation in the estimates, several of the claims made above are evident in the graph. For instance, the variances are increasing in age for both offspring and parents. The autocovariances in both generations also "fan out", i.e., the autocovariances for a given lag length increases in age. Moreover, the intergenerational covariance of annual income also increase (mildly) in both parent and offspring age. Thus, we find that the evidence here also is consistent with the presence of random growth rates. It follows that we should take these into account in estimating the intergenerational associations.

5   Results

In this section, we estimate intergenerational income elasticities for Swedish men and women. The estimates are arrived as follows. First, run a regression of the ln of annual taxable income on gender, birth-cohort and outcome-year indicator variables, fully interacted. Second, we take the annual residuals for each individual (in both parent and offspring generations) for every year in the relevant age range. We estimate a model corresponding to equation 6 above using the **lme**

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Figure 2 Plot of parent-child covariance matrix
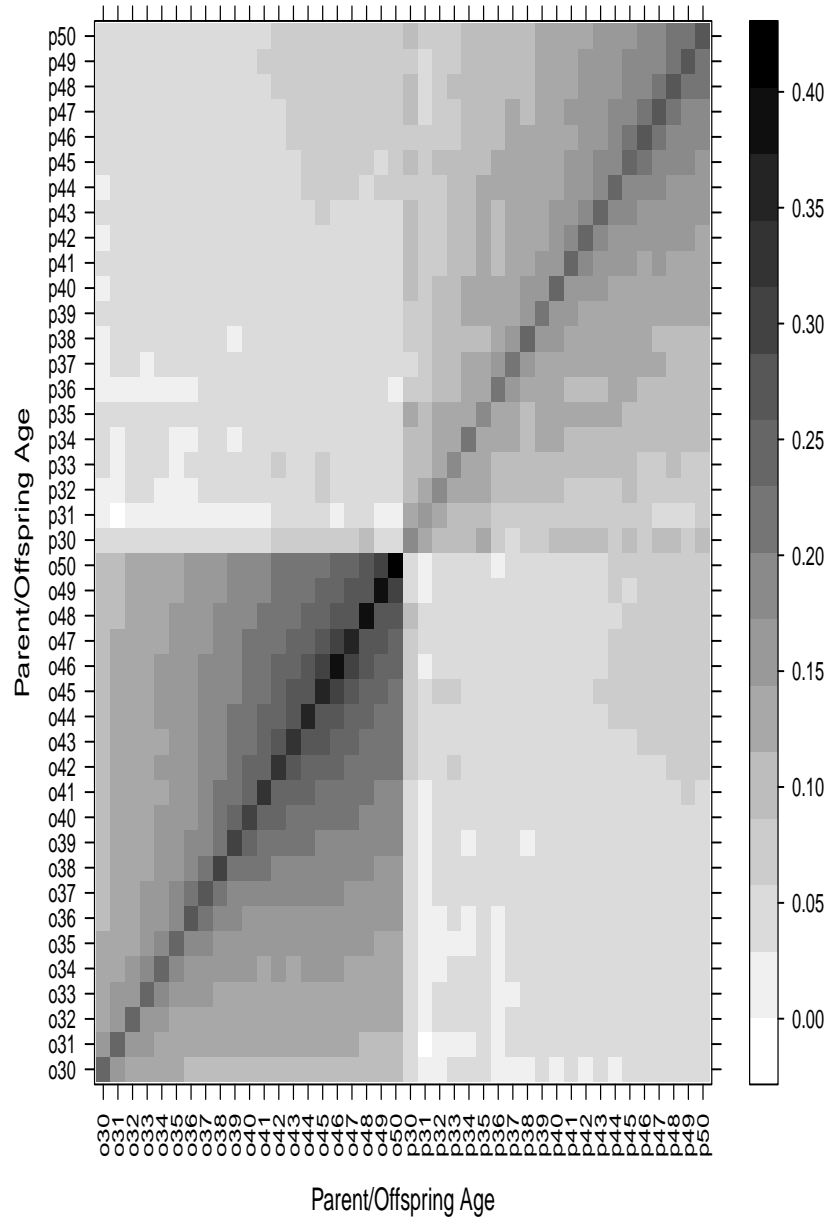
Table 2 Estimated intergenerational elasticities at different age ranges of offspring and parents

Age range	Elasticity	
	$\ln V_P$	$se(\ln V_P)$
30:55	0.347	0.040
30:40	0.155	0.040
35:45	0.338	0.054
40:50	0.202	0.041
45:55	0.242	0.045
30:45	0.332	0.056
35:50	0.279	0.036
40:55	0.202	0.041
30:50	0.231	0.040
35:55	0.349	0.036

(Pinheiro and Bates, 1999) function in the statistical package **R** (Ihaka and Gentleman, 1996) – i.e., we estimate a common ARMA(1,1) process for the whole sample, and an intercept and a growth rate for each individual. Third, we use the estimated intercept and growth rate to calculate the natural logarithm of expected lifetime income (using a discount rate of $r = .02$) for each individual. Finally, we regress the natural logarithm of expected lifetime income of offspring on that of the father, which gives us the estimate of the intergenerational income elasticity.

Since we argue the intragenerational income process involves a random growth rate, we start by examining how well the estimation procedure work when a close to full lifetime incomes are unavailable. Specifically, we start by examining elasticities for a cohort of sons and their fathers for whom we have 25 years of income observations in ages 30-55. We then compare estimated elasticities for these sons for 10-year, 15-year and 20-year intervals to see if the various estimates are close to the full period estimates.

The estimates for our test age ranges are shown in Table 2. The dependent and explanatory variables are the natural logarithm of expected lifetime income as outlined in Section 4. The point estimate for the full set of 25 years for both sons and fathers is 0.347. of the different 10-year windows, there is some variation, but the interval 35-45 is very close at 0.335. The interval 30:40, which is what many authors have used, is less than half the full period estimate. Of the 15-year periods, the age range of 30-45 comes closest to the full period estimate at 0.332.

And the 35-55 estimate, using 20 years of income data, is 0.349.

Since the purpose of this paper is to estimate changes across cohorts in intergenerational elasticities, we would like to use a short, in this case 10-year window, into the income processes of both generations. On the basis of these results, we proceed by choosing the 35-45 year age range as the basis of our estimation.

We estimate four variations of the statistical model outlined in equations 6, 7 and 11, varying whether or not we include a random growth rate or not, and if we allow the transitory fluctuations to be correlated across time (following an ARMA(1,1) process) or if they are modeled as white noise. The baseline case (strongly supported by the data) is to have ARMA errors and allow for random growth rates.

We show the baseline estimates, measuring both offspring and father income across ages 35-45, in Figure 3 and in Column A of Table 3. The remaining columns in Table 3 and the graphs in Figure 4 show the three combinations of no random growth rate and/or white-noise vs. ARMA errors. The estimated ARMA processes, the point estimates of which are quite similar across cohorts, are shown in the appendix Table 4.

In our baseline estimates for father-son pairs (column A of Panel I in Table 3), the intergenerational elasticity of expected lifetime income is 0.42 across all offspring cohorts born between 1945 and 1962. For the earliest cohort, the elasticity is 0.23, it reaches a maximum of 0.51 for the 1954-1956 cohort and is 0.45 in the 1960-1962. If we do not include a random growth rate, but do allow the transitory fluctuations to be intertemporally correlated (column B), the level of the estimates is much lower – the overall IGE is 0.27 – and the time trend is also quite different. We have a high initial estimate, followed by a decline and then a close to monotonic increase. If we further eliminate the ARMA structure from the errors (column C), the level of the point estimates is similar, the initial high estimate for the 1945-1947 cohorts disappears and we have an increasing trend. Finally, if we allow for a random growth rate, but eliminate the memory in the transitory fluctuations, the estimated elasticities are very low, around 0.10 across all cohorts. This final finding underlines the importance of distinguishing between heterogenous profiles, on the one hand, and long-memory random shocks, on the other.

The estimated baseline elasticities for women, shown in Column A of Panel II of Table 3, as well as in the right-hand panel in Figures 3, are lower, but also exhibit some tendency for increase from older to younger cohorts. The non-baseline estimates do not exhibit much of a trend at all. While these results should be viewed with some caution, they do highlight how unstable the estimated intergenerational elasticities can be to seemingly minor measurement issues.

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Figure 3 Intergenerational income elasticity in the presence of random growth rates and ARMA(1,1) errors – natural logarithm of expected lifetime income based on ages 35-45 in both generations
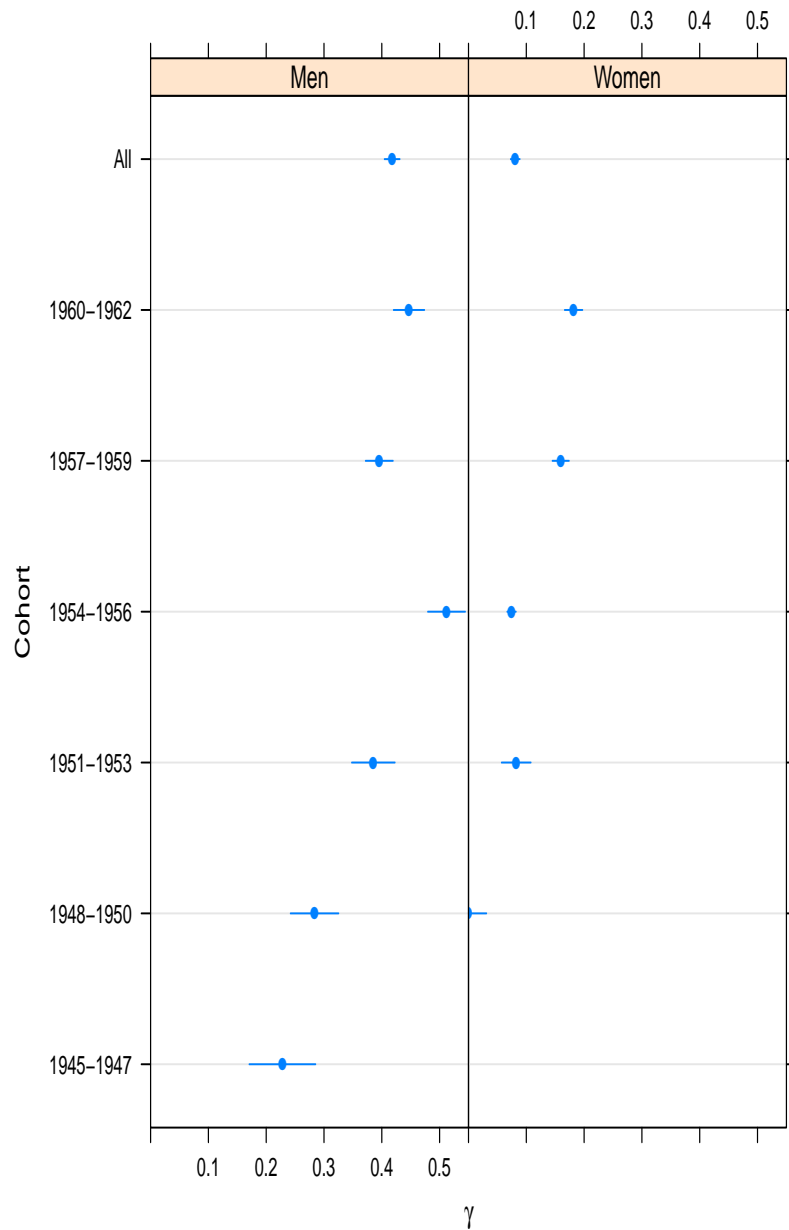
Figure 4 Intergenerational income elasticity in the presence of only a random intercept and ARMA(1,1) errors – natural logarithm of expected lifetime income based on ages 35–45 in both generations

B. Only intercept, ARMA errors   C. Only intercept, WN errors   D. Incl. random growth rate, WN errors
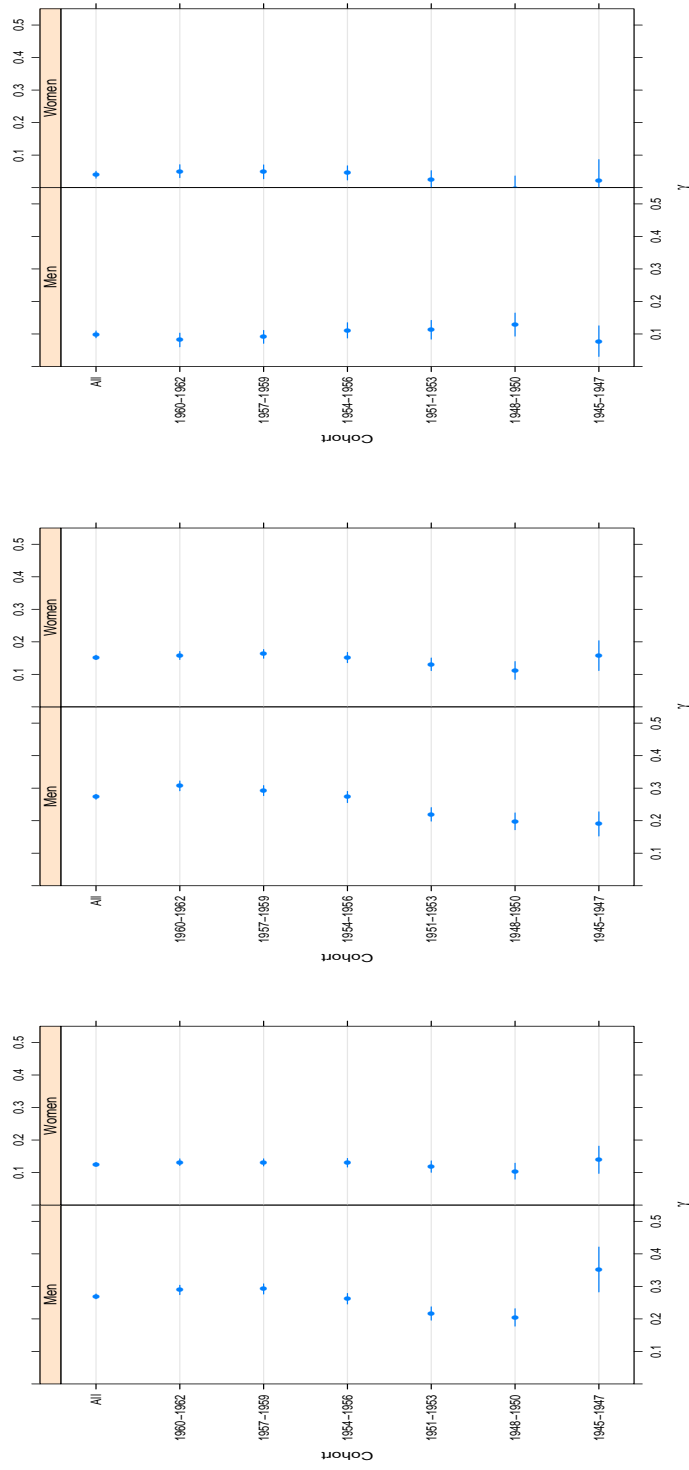
Table 3 Intergenerational elasticities for different model specifications

I. Men

A. Incl. random growth rate, ARMA errors — α, β, ARMA

Cohort	Elasticity	
	$\ln V_P$	se($\ln V_P$)
All	0.418	0.007
1945-1947	0.228	0.029
1948-1950	0.284	0.021
1951-1953	0.385	0.019
1954-1956	0.512	0.017
1957-1959	0.396	0.012
1960-1962	0.447	0.014

B. Only intercept, ARMA errors — α, ARMA

Cohort	Elasticity	
	α_P	se(α_P)
All	0.269	0.004
1945-1947	0.352	0.035
1948-1950	0.204	0.014
1951-1953	0.217	0.011
1954-1956	0.262	0.008
1957-1959	0.292	0.008
1960-1962	0.289	0.008

C. Only intercept, WN errors — α, WN

Cohort	Elasticity	
	α_P	se(α_P)
All	0.273	0.004
1945-1947	0.190	0.019
1948-1950	0.198	0.013
1951-1953	0.219	0.011
1954-1956	0.273	0.009
1957-1959	0.293	0.008
1960-1962	0.307	0.008

D. Incl. random growth rate, WN errors — α, β, WN

Cohort	Elasticity	
	$\ln V_P$	se($\ln V_P$)
All	0.099	0.006
1945-1947	0.078	0.024
1948-1950	0.129	0.018
1951-1953	0.113	0.015
1954-1956	0.111	0.012
1957-1959	0.091	0.010
1960-1962	0.082	0.011

II. Women

A. Incl. random growth rate, ARMA errors — α, β, ARMA

Cohort	Elasticity	
	$\ln V_P$	se($\ln V_P$)
All	0.081	0.004
1945-1947	-0.065	0.023
1948-1950	-0.001	0.016
1951-1953	0.083	0.013
1954-1956	0.075	0.004
1957-1959	0.160	0.007
1960-1962	0.182	0.008

B. Only intercept, ARMA errors — α, ARMA

Cohort	Elasticity	
	α_P	se(α_P)
All	0.125	0.003
1945-1947	0.139	0.022
1948-1950	0.104	0.013
1951-1953	0.118	0.009
1954-1956	0.130	0.007
1957-1959	0.131	0.006
1960-1962	0.132	0.005

C. Only intercept, WN errors — α, WN

Cohort	Elasticity	
	α_P	se(α_P)
All	0.152	0.004
1945-1947	0.158	0.024
1948-1950	0.112	0.014
1951-1953	0.131	0.010
1954-1956	0.152	0.008
1957-1959	0.163	0.007
1960-1962	0.158	0.007

D. Incl. random growth rate, WN errors — α, β, WN

Cohort	Elasticity	
	$\ln V_P$	se($\ln V_P$)
All	0.040	0.006
1945-1947	0.021	0.033
1948-1950	-0.003	0.020
1951-1953	0.025	0.014
1954-1956	0.045	0.011
1957-1959	0.049	0.011
1960-1962	0.051	0.010

6   Concluding remarks

We have examined changes across cohorts of Swedish men and women from 1945 to 1962 in intergenerational income mobility. We find little trend in the intergenerational mobility of women, but we do find an increase in income persistence for men. We have paid close attention to problems associated with estimating intergenerational income associations in the presence of non-classical measurement error models. In particular, we allow the annual income of both fathers and offspring to contain both an intercept and a random growth rate, and we also allow the transitory fluctuations to be intertemporally correlated.

For a few cohorts of offspring who are born in 1950-52, we can observe 25 years of income for both fathers and offspring. In order to choose the age range to examine, we compare estimates based on different 10-, 15- and 20-year age ranges. In order to be able to make most of our data, we choose to work with 10-year windows. Based on this comparison, our methods appear to work best in the age range 35-45, in the sense of coming closes to the benchmark of 25 years from ages 30-55.

The level of intergenerational persistence, as measured by the elasticity between the expected long-run income of fathers and their offspring is higher than has previously been estimated in Sweden. While the level in these estimates is higher than of the "raw" income elasticities, the trend in the more refined measure is similar to that of the raw estimates. The evidence we have uncovered is thus that at least for cohorts born between 1945 and 1962, the intergenerational persistence of economic status has increased.

Table 4 Estimated parameters of ARMA process for transitory errors

A. Men

	Offspring			Parent		
	$\sigma^2_{v,O}$	ϕ_O	θ_O	$\sigma^2_{v,P}$	ϕ_P	θ_P
All	0.098	0.719	−0.176	0.138	0.715	−0.183
1945-1947	0.144	0.876	−0.327	0.103	0.682	−0.152
1948-1950	0.106	0.725	−0.198	0.107	0.668	−0.163
1951-1953	0.096	0.703	−0.196	0.134	0.730	−0.199
1954-1956	0.104	0.715	−0.140	0.156	0.762	−0.226
1957-1959	0.099	0.731	−0.184	0.156	0.717	−0.184
1960-1962	0.095	0.714	−0.176	0.175	0.703	−0.159

B. Women

	Offspring			Parent		
	$\sigma^2_{v,O}$	ϕ_O	θ_O	$\sigma^2_{v,P}$	ϕ_P	θ_P
All	0.104	0.735	−0.172	0.151	0.693	−0.074
1945-1947	0.113	0.828	−0.314	0.166	0.715	−0.054
1948-1950	0.107	0.715	−0.154	0.140	0.678	−0.064
1951-1953	0.097	0.686	−0.140	0.126	0.647	−0.069
1954-1956	0.095	0.689	−0.147	0.167	0.769	−0.139
1957-1959	0.104	0.748	−0.179	0.163	0.726	−0.089
1960-1962	0.104	0.732	−0.171	0.177	0.678	−0.063

Appendix tables

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