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Abstract

We survey the literature on income mobility, aiming to provide an integrated discussion of mobility within- and between-generations. We review mobility concepts, descriptive devices, measurement methods, data sources, and recent empirical evidence.

Keywords: Intragenerational mobility, intergenerational mobility, income mobility, earnings mobility.

JEL Classification: D31, I30.

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1. Introduction

Most of the information that we have about the income distribution is cross-sectional in nature: there are statistics about for example income levels, poverty rates, and the extent of inequality for a given year or for a series of years. The data sources used to provide estimates for the different year refer to different samples of individuals. In this chapter, we discuss a different but complementary perspective on income distribution to the cross-sectional one. We take an explicitly longitudinal perspective, one that is based on tracking over time the fortunes of the same set of individuals. We are interested, broadly speaking, in how individuals' incomes change over time in a society. 'Income mobility' is a shorthand label for this topic. In this chapter, we address questions such as: what exactly do we mean by mobility and why should we be interested in it? How should mobility be measured? What is the evidence about income mobility for rich industrialised nations?

The period of time over which income mobility is assessed is a fundamental issue and different choices have led to two relatively distinct literatures. On the one hand, there is the subject of how an individual's income changes between one year and another during their lifetime; on the other hand, there is the subject of income change between generations of parents and children. We use this distinction between intragenerational and intergenerational income mobility as an organisational device in this chapter, reflecting the division in existing literature, but we shall also attempt to draw out the features of the measurement of income mobility that are common to both topics while also highlighting dimensions of them for which different approaches to analysis are appropriate.

Conceptual issues are addressed first because clarification of them is an essential preliminary to any discussion of measurement principles, data sources, and assessment of empirical evidence. In Section 2, we review the reasons why and how income mobility is said to be of interest. There are several distinct reasons and this is because, as we also discuss, there are multiple concepts of mobility, each of which arguably has normative validity. This situation contrasts with assessments of an income distributions at a point in time, in which case there is greater consensus about what is meant by income inequality, and how it might be accounted for in social welfare evaluations.

We review the measurement of income mobility in Section 3, focusing on the generic case in which there are data on income at two points in time, whether this be two years (as in the intragenerational mobility literature) or two generations (as in the intergenerational mobility literature). This is the most commonly-examined situation. Thus we are interested in not only summarising a single bivariate joint distribution of income but also comparing such distributions across time or countries in order to say whether mobility is greater or smaller. We explain various descriptive methods for situations in which income data are either continuous or grouped into categories. First we discuss graphical devices and methods that may be used to undertake mobility comparisons without resort to choice of a particular mobility index (so-called dominance checks). Second, we consider scalar indices of mobility ranging from regression coefficients and correlations through to other more specialist developments.

By considering measurement from a generic point of view, we aim to show how there might be greater cross-fertilisation between the intra- and intergenerational mobility literatures in approaches to measurement. At the same time, we highlight how the different measurement approaches relate to different concepts of mobility identified in Section 2.

Evidence about income mobility is the subject of the next two sections: Section 4 considers intragenerational mobility; Section 5 considers intergenerational mobility. In each case, our strategy is to build a bridge linking concepts and measurement principles to empirical evidence by first discussing data sources, as well as issues of empirical implementation including data comparability and quality more generally.

The final section provides brief concluding remarks and makes some proposals concerning where the returns to future research efforts are the greatest.

Earlier research on income mobility has typically focused on either within- or between-generation topics. For surveys of intragenerational measurement issues, we build on Jenkins (2011a) who, in turn, draws heavily on other surveys such as by e.g. Atkinson et al. (1992), Burkhauser and Couch (2009), Fields and Ok (1999a), Jenkins and Van Kerm (2009), and Maasoumi (1998). For intergenerational mobility, important earlier reviews are provided by Solon (1999), Björklund and Jäntti (2009), Black and Devereux (2011), and Piketty (2000). Many of the reviews just cited appear in volumes with 'Handbook' in their title. Indeed extensive surveys of cross-sectional approaches to income distribution were provided throughout the *Handbook of Income Distribution, Volume 1* (Atkinson and Bourguignon, 2000). It is timely and appropriate to give income mobility similar attention.

While the chapter draws heavily on the work of others, it also has some distinctive features besides simply being more up-to-date. One aspect is our goal to try and integrate the discussion of intra- and intergenerational mobility in so far as this is possible, while also highlighting what aspects of each topic are intrinsically different and deserving of separate attention. Other aspects include our coverage from conceptual issues through to data, issues of empirical implementation and evidence.

The emphasis of this chapter is on the *measurement* of income mobility, broadly defined. Of course, it is also of interest to not only describe how individuals' incomes change from one time period to another but also to explain the patterns observed. We have deliberately chosen not to systematically review *models* of mobility in order to make our task manageable.

There is some discussion of intragenerational models of earnings dynamics, nonetheless, in Section 3 because estimates from 'variance components' models have been used to derive measures of mobility in the form of income risk. Other types of modelling approach are reviewed by Jenkins (2000), who also discusses more general issues concerning the modelling of intragenerational income dynamics. These are further elaborated by Jenkins (2011a, chapter 12).

One important distinction is between reduced-form and structural empirical models, each of which has different strengths and weaknesses. The former are:

> empirically grounded rather than derived from a well-developed theoretical model that implies specifications, the parameters of which are estimated from the data. ... The advantage of a structural approach is that there is a close relationship between parameter estimates and behavioural model parameters and so interpretation is improved and one may be able to say more about underlying causes. The problem with a structural approach is that clear cut implications for model specification and proofs of relationships can often only be derived by massive simplification – simplification that compromises claims that the

> model describes empirical reality. The tension between reduced form and structural approaches has existed for a long time and is likely to remain ... The reason for the tension is obvious – approaches combining structure, practicality, and feasibility are very difficult to develop ... The problem is that a model is needed not only for the dynamics of labour earnings for an individual but also the earnings and possibly other income sources of other individuals in a multi-person household, and the dynamics of household structure itself also needs to be modelled. (Jenkins, 2011a, 368–369.)

Exactly the same tension has arisen in empirical modelling of intergenerational income dynamics, where there is also a need to consider not only multiple income sources but also demographic factors. The structural ('optimizing') approach is epitomized by Becker and Tomes (1986) and the reduced-form ('mechanical') approach by a series of papers by Conlisk (1974, 1977, 1984).[1] The relative merits of the two approaches are lucidly discussed by Goldberger (1989), with a 'reply to a skeptic' provided by Becker (1989).

[1]See also Solon (2004) for a simple model highlighting the key ingredients of an optimizing model and Mulligan (1997) for a monograph-length treatment of the theoretical literature.

2. Mobility concepts

Writers on income mobility have long emphasised that mobility has multiple dimensions. For example, a leading survey from a decade ago commented that:

> the mobility literature does not provide a unified discourse of analysis. This might be because the very notion of income mobility is not well-defined; different studies concentrate on different aspects of this multi-faceted concept. At any rate, it seems safe to say that a considerable degree of confusion confronts a newcomer to the field (Fields and Ok, 1999a, 557).

The systematic reviews by Fields and Ok and others, have done much to reduce the potential confusion. But they cannot banish mobility's multiple facets, and so newcomers continue to require guided tours of the concepts and literature. This section explains what the multiple dimensions of mobility are. We address the question of whether more mobility is socially desirable in each case, arguing that the answer depends on which mobility concept is the focus. A review of the implications of mobility's various facets for social welfare is used to illustrate trade-offs between different types of mobility. We also point out how different concepts have received different emphasis in studies of mobility within- or between-generations.

2.1. Mobility's multiple dimensions

Consider first the case in which there are observations on income for N individuals for two periods. In the first period, the income distribution is $\mathbf{x}$, in the second period, the distribution is $\mathbf{y}$; there is a bivariate joint density $f(\mathbf{x}, \mathbf{y})$. Overall mobility for the population can be thought of as the transformation linking marginal distribution $\mathbf{x}$ with marginal distribution $\mathbf{y}$.

In this section, we distinguish four concepts (Jenkins, 2011a): positional change (which comes in two flavours), individual income growth, reduction of longer-term inequality, and income risk.[2] The different concepts 'standardise' the marginal distributions $\mathbf{x}$ and $\mathbf{y}$ in different ways in order to focus attention on the nature of the link $\mathbf{x} \rightarrow \mathbf{y}$.

Positional change refers to mobility that arises separately from any changes in the shapes of the marginal distributions in each period, for example a rise in average income or in income inequality or, more generally, a change in the concentration of individuals at different points along the income range in $\mathbf{y}$ compared to in $\mathbf{x}$. Standardisation for such changes is most easily accomplished by summarising each person's position not in terms of their income per se but in terms of their rank in the population normalised by the population size. (The marginal distribution of these 'fractional' (or 'normalised') ranks is a standard uniform distribution for both $\mathbf{x}$ and $\mathbf{y}$.) Thus positional change mobility refers to the pattern of exchange of individuals between positions, while abstracting from any change in the concentration of people in a particular slot in each year. The latter change is 'structural mobility', whereas the former is 'exchange mobility': see for example Markandya (1984). Changes in income affect positional mobility only in so far as these changes alter each person's position relative to the position of others. Equiproportionate income growth or equal absolute additions to income for everyone raise incomes but there is immobility in the positional sense.

[2] This classification is similar to that employed by Fields and Ok (1999a) and Fields (2006). See also Van de gaer et al. (2001).

There are some distinctive characteristics of the concept of mobility as positional change. Mobility for any specific individual necessarily depends on other people's positions as well, which is not true for every mobility concept as we shall see. The definition of each person's origin and destination position depends on the positions of everyone else in the society: it is these taken altogether that define a hierarchy of positions. Second, and related, if one person changes position then so too must at least one other person. It is not possible for everyone to be upwardly mobile or, indeed, downwardly mobile. Third, the situation corresponding to 'no mobility' is straightforwardly defined: maximum immobility occurs when every person has the same position in **x** and in **y**. If income mobility is summarized using a transition matrix (see below) in which cell entries a_{jk} show the probability that an individual in income class *j* in period 1 is found in income class *k* in period 2, then maximum immobility is the case in which $a_{jk} = 1$ for all income classes (all individuals are on the leading diagonal). However, fourth, there are two different ways of thinking about what reference points to use when there is mobility, one focusing of lack of dependence and the second focusing on movement.

One situation is when one's destination is completely unrelated to one's income origin ('origin independence'). For example, the chances of being found in the richest tenth in period 2 are exactly the same for people who were in the poorest tenth in period 1 as for the people who were in the richest tenth in period 1. In transition matrix terms, this is the case in which $a_{jk} = a_{mk}$ for all origin classes *j* or *m* (each row of the transition matrix has identical entries). Another view is that the reference case when there is mobility is if destination positions are a complete reversal of origin positions ('rank reversal'), emphasising positional movement per se. For example, the poorest person in period 1 is the richest person in period 2, and the richest person in period 1 is the poorest person in period 2, and so on. All entries in the transition matrix lie on the diagonal going from bottom left (richest origin class and poorest destination class) to top right (poorest origin class and richest destination class).[3]

Mobility as *individual income growth* refers to an aggregate measure of the changes in income experienced by each individual within the society between two points in time, where the individual-level changes might be gains or losses. Income growth is defined for each individual separately and income mobility for society overall is derived by aggregating the mobility experienced by each and every individual.[4] This mobility concept contrasts sharply with the positional change one in several ways. No distinction is made between structural and exchange mobility: it is gross (total) mobility that is described. It is possible for everyone to be upwardly mobile or, indeed, to be downwardly mobile. Positive income growth for everyone may count as mobility even if relative positions are preserved. Thus, standardisation of the marginal distributions is not an essential feature of the concept.

In the individual economic growth case, it is natural to define mobility for each individual in terms of 'distance' between origin and destination income, and to think of the maximum immobility case for the population as being when the measure of distance equals zero for every individual ($x_i = y_i$ for all i). Mobility is greater if the distance between origin and destination is greater for any individual,

[3]The two reference points are sometimes referred to as cases of 'perfect' or 'maximum' mobility, but we resist these. The language in the former case makes potentially unwarranted assumptions about the optimality of particular mobility configurations (to be discussed below), and it is difficult to argue that origin independence represents 'maximum' mobility in the literal sense.

[4]Observe that this is an assumption, albeit commonly made. It is what Fields and Ok (1996) call the 'individualistic contribution' axiom.

other things being equal. This is similar to the idea of greater movement meaning more mobility according to the 'reversals' version of positional mobility. Again there is no natural maximum mobility reference point as distance has no obvious upper bound.[5] Defining the metric for 'distance' in terms of the income change for each individual is of course vitally important for the concept, and the main distinctions have been measures of 'directional' and 'non-directional' growth. In the first case, income increases over time are treated differently from income decreases; in the second, an income increase and an income decrease of equal magnitude are attributed the same distance and the measure summarizes income 'flux' (more on this shortly). For more precise definitions, see Fields and Ok (1999a).

The third mobility concept defines income mobility with reference to its *impact on inequality in longer-term incomes*. The longer-term income for each individual is defined as the longitudinal average of incomes in each period (variations on this are considered below). In the two period case, longer-term income equals $\frac{1}{2}(x_i + y_i)$ for each i. Averaging across time smooths the longitudinal variability in each person's income and, in addition, the inequality across individuals in these longitudinally-averaged incomes will be less than the dispersion across individuals in their incomes for any single period. Mobility can therefore be characterized in terms of the extent to which inequality in longer-term income is less than the inequality in marginal distributions of period-specific income. (See Shorrocks, 1978a) and below for further details. The zero mobility reference point is when the income of each person in every period is equal to their longer-term income:

[5]Observe that individual income growth cannot be represented using a transition matrix, since the mobility concept in this case is intrinsically individual- rather than group-based. However, income growth can be represented using a mobility matrix in which category boundaries are defined in real income terms.

there is complete rigidity. At the other extreme, maximum mobility occurs when there is inequality in per-period incomes but no inequality at all in longer-term incomes. The issue of whether everyone can be upwardly (or downwardly) mobile does not arise with this mobility concept because it defines mobility using inequality comparisons, and inequality is measured at the aggregate (population) level. There are similarities between this concept of mobility and the rank reversal flavour of the positional change concept since both are concerned with movement, but they use different reference points to assess this (longer-term incomes versus base-period positions respectively). We return to this issue later.

The fourth concept of mobility, as *income risk*, is related to the third. The previous paragraph expressed each person's period-specific income as the sum of a 'permanent' component (the longer-term average) and a 'transitory' component (the period-specific deviations from the average). Suppose now that the longer-term average is given a behavioural interpretation: it is the expected future income per period given information in the first period about future incomes. From this ex ante perspective, the transitory components represent unexpected idiosyncratic shocks to income, and the greater their dispersion across individuals each period, the greater is income risk for this population. The measure of mobility cited in the previous paragraph, i.e. the inequality reduction associated with longitudinal averaging of incomes, is now re-interpreted as a measure of income risk and has different normative implications (see below). Income movement over time represents unpredictability. This is essentially what Fields and Ok (1999a) refer to as income 'flux' (non-directional income movement). Despite their apparent similarities in construction, the concepts of mobility as inequality-reduction and as income risk diverge in practice when the process describing income generation is not a simple sum of a fixed individual-level permanent component and an idiosyncratic transitory component. Econometric models have been developed with more complicated descriptions of how the permanent and transitory components evolve over time and these imply, in turn, different calculations of expected income and transitory deviations from it. However the distinction between predictable relatively fixed elements and unpredictable transitory elements of income is maintained and hence so too is a link between mobility as transitory variation and income risk.

2.2. *Is income mobility socially desirable?*

In what ways are these various mobility concepts of public interest over and above providing useful descriptive content? Does having more mobility represent a social improvement or is it undesirable? The answers depend on the mobility concept employed, and that the support for the different concepts has depended on whether one is assessing within- or between-generation mobility.

Greater mobility in the sense of less association between origins and destinations has long been linked with having a more open society: if where you end up does not depend on where you started from, there is greater equality of opportunity. For example, a classic statement by R. H. Tawney, originally from 1931, is that equality of opportunity

> obtains in so far as, and only in so far as, each member of a community, whatever his birth, or occupation, or social position, possesses in fact, and not merely in form, equal chances of using to the full his natural endowments of physique, of character, and of intelligence (Tawney, 1964, 103–5).

More recently, a UK government advisor's report on Social Mobility stated that 'Social mobility matters because ... equality of opportunity is an aspiration across the political spectrum. Lack of social mobility implies inequality of opportunity' (Aldridge, 2001). For more about equality of opportunity, see Chapter 5 in this volume by Roemer and Trannoy.

From this perspective, greater mobility is socially desirable since equality of opportunity is a principle that is widely supported, regardless of attitudes to inequality of outcomes. This is relevant because independence of origins and destinations is consistent with inequality of outcomes being relatively equal or unequal. The argument just rehearsed is, however, typically made in the context of intergenerational mobility rather than intragenerational mobility, and origins refer to parental circumstances, such as 'birth, or occupation, or social position' referred to by Tawney. The appeal to fairness in this context is based on the meritocratic idea that someone's life chances should depend on their own abilities and efforts rather than on who their parents were. At the same time, it is important to appreciate that the degree of intergenerational association is an imperfect indicator of the degree of inequality of opportunity.

The degree of origin independence is a direct measure of inequality of opportunity only if two rather special conditions apply (Roemer, 2004). First, the advantages associated with parental background (over which it is assumed that an individual had no choice) are entirely summarised by parental income. Second, the concept of equality of opportunity that is employed views as unacceptable any income differences in the children's generation that are attributable to differences in innate talents (which might be partly genetically inherited). This is what Swift (2006) describes as a 'radical' interpretation of the equality of opportunity principle, and likely to command much less widespread assent than what he refers to as the 'minimal' and 'conventional' definitions (respectively, access and recruitment processes to life chances are free of prejudice and discrimination; and outcomes achieved depend on 'ability' and 'effort' but not on family background).

The social desirability of mobility as independence of origins has less force in the intragenerational context. The reason is that incomes are measured at a point within the life course. By that stage, period-1 incomes are likely to reflect differences in peoples' abilities and efforts (in addition to family background and other factors), and period-2 incomes to reflect the persisting effects of these factors. To the extent that abilities and efforts do play this role (or are seen to) and also viewed as fair on the grounds of merit or desert, the reduction of dependence between origins and destination has less appeal as a principle of social justice.

More common in the within-generation context are statements that income mobility is desirable because it is a force for reduction in the inequality of longer-term incomes. The most famous statement in this connection was by Milton Friedman six decades ago in his *Capitalism and Freedom* (though observe that he also refers to equality of opportunity in this context):

> A major problem in interpreting evidence on the distribution of income is the need to distinguish two basically different kinds of inequality; temporary, short-run differences in income, and differences in long-run income status. Consider two societies that have the same annual distribution of income. In one there is great mobility and change so that the position of particular families in the income hierarchy varies widely from year to year. In the other there is great rigidity so that each family stays in the same position year after year.

> The one kind of inequality is a sign of dynamic change, social mobility, equality of opportunity; the other, of a status society (Friedman, 1962, 171).

Similar views are apparent across the political spectrum in the USA. The Chairman of President Obama's Council of Economic Advisors recently stated that

> Higher income inequality would be less of a concern if low-income earners became high-income earners at some point in their career, or if children of low-income parents had a good chance of climbing up the income scales when they grow up. In other words, if we had a high degree of income mobility we would be less concerned about the degree of inequality in any given year (Krueger, 2012).

Although both authors are referring to the distributions of incomes within generations, one could extend the same inequality-reduction idea to the intergenerational context, by summarising mobility in terms of the extent to which dynastic inequality (referring to incomes averaged over generations of the same family) is less than the inequality in any given generation. But this is rarely done, perhaps because the normative appeal of the dynastic average income is much less than that of a multi-period average within generations, and data for more than two generations are rarely available.

According to the arguments about longer-term inequality reduction, income mobility is socially desirable for instrumental reasons rather than for its own sake. That is, society is assumed to care about income inequality (less is better, other things being equal), but inequality is assessed using longer-term incomes and year-to-year mobility means that the inequality of this distribution is less than the inequality of incomes in any particular year. The normative content of the mobility principle therefore hinges on views concerning the nature and validity of the benchmark that is provided by the distribution of longer-term incomes. As Shorrocks points out,[6] there is

> the presumption that individuals are indifferent between two income streams offering the same real present value. This might be true if capital markets were perfect (or if there was perfect substitutability of income between periods), but it seems likely that individuals are concerned with both the average rate of income receipts and the pattern of receipts over time. We may go further and suggest that individuals tend to prefer a constant income stream, or one which is growing steadily, to one which continually fluctuates (Shorrocks, 1978a, 392).

Thus, the argument is not only about the feasibility of smoothing incomes to achieve the longer-term average, but also the undesirability of the uncertainty associated with a fluctuating income stream.

This brings us to the fourth concept of income mobility, as income risk. To illustrate this, Shorrocks defines for each individual a 'constant income flow rate generating receipts which gives the same level of welfare as the income stream he currently faces' (Shorrocks, 1978a, 392), and he argues that

> [r]eplacing actual recorded incomes with this alternative income concept in the computation of inequality values introduces a new dimension into the discussion of mobility. No longer is mobility necessarily

[6] Shorrocks also draws attention to the assumption that the same measure is used to summarise both the dispersion of longer-period incomes and the dispersion of per-period incomes.

> desirable. Changes in relative incomes still tend over time to equalise the distribution of total income receipts, and to this extent welfare is improved. But greater variability of incomes about the same average level is disliked by individuals who prefer a stable flow. So to the extent that mobility leads to more pronounced fluctuations and more uncertainty, it is not regarded as socially desirable. A more detailed examination of these two facets of mobility will provide a better understanding of the impact of income variability and the implications for social welfare (Shorrocks, 1978a, 392–393).

Thus, even though income mobility has an inequality-reducing impact, mobility is not necessarily socially desirable if mobility represents transitory shocks. In this case, mobility is a synonym for not only income fluctuation but also unpredictability and economic insecurity. Fluctuating incomes are undesirable because most people prefer greater stability in income flows to less, other things being equal, if only because it facilitates easier and better planning for the future. But, more than this, by definition, transitory income variation is an idiosyncratic shock which cannot be predicted at the individual level: greater transitory variation corresponds to greater income risk, and greater risk is undesirable for risk-averse individuals. The definition of the 'alternative income concept' from which transitory shocks deviate is of course crucial, and we return to this.

What about the social desirability of individual income growth (the second mobility concept)? The answer is not clear cut because it depends on the nature of the income growth and who receives it. An increase in income for any given individual is a social improvement and an income fall is socially undesirable. The main issue, then, is how to aggregate gains and losses in the social calculus. Evaluation of the impact of individual income growth on the welfare of society as a whole requires a weighing up of the gains and losses for different people, and opinions are likely to differ about how to do this. An egalitarian may weight income gains for the initially poor greater than income gains for the initially rich because this will contribute to reducing income differences between them over time. (On the progressivity of income growth, see e.g. Benabou and Ok (2000) and Jenkins and Van Kerm (2006).)

Arguments to the contrary appealing to principles of desert or incentives might also be made. It might be argued, for instance, that differential income growth rates are of less concern if income gains among the rich reflect appropriate returns to entrepreneurial activity or to widely-acclaimed talents. The rise in bankers' bonuses in the manner observed in many Anglophone countries in recent years may not count as an example of the former. But as an example of the latter, we note the views of the UK's former Prime Minister Tony Blair expressed in an interview asking him whether it was acceptable for the gap between rich and poor to get bigger. His response referred instead to individual income growth:

> the justice for me is concentrated on lifting incomes of those that don't have a decent income. It's not a burning ambition for me to make sure that David Beckham earns less money... [T]he issue isn't in fact whether the very richest person ends up becoming richer. ... the most important thing is to level up, not level down (Interview on BBC Newsnight, 5 June 2001).[7]

Another concept of desert may also be relevant when assessing mobility. This

[7] Transcript at `http://news.bbc.co.uk/1/hi/events/newsnight/1372220.stm`.

is the argument concerning 'distressed gentlefolk' – people who were previously well-off, but experience a significant fall in resources through no fault of their own. Thus income gains and income losses for an individual may not be assessed symmetrically but, again, relate to why income changed. (See also the discussion of 'loss aversion' below.)

We end this subsection with two observations. First, our discussion of the social desirability or otherwise of income mobility has referred to income movement from throughout the range of base-period income origins to all potential final-period income destinations. There has been no particular focus on persistence at the bottom or at the top. In part, this is because such a focus arguably does not raise additional conceptual issues, except where to draw the cut-offs demarcating the poor and non-poor, or rich and non-rich. Indeed if the bivariate joint distribution is summarised using a transition matrix, then suitable definition of the income groups reveals the movement at the top and the bottom. However, we do discuss selected aspects of the measurement of high- and low-income persistence in the next two sections.

Second, our discussion of the social desirability of mobility has focussed on its normative aspects. We ignore the positive political economy arguments about public support for mobility. On this, see e.g. the analysis by Benabou and Ok (2001) of the 'prospect of upward mobility' (POUM) hypothesis, which is that individuals who currently have low income may not support high levels of redistribution because of their aspiration that they or their children will become rich in future.

2.3. Income mobility and social welfare

The discussion so far demonstrates that the impact on social welfare of greater income mobility is not clear cut, and depends on the mobility concept that is emphasised. A natural question for an economist to ask is whether there are explicit welfare foundations for the various mobility concepts that have been discussed so far. For inequality measurement, the use of an explicit model of social welfare is known to yield dividends: see, notably, Atkinson's (1970) demonstration of how the 'cost' of income inequality can be summarised in social welfare terms and how inequality comparisons based on Lorenz curves are intimately linked to orderings by social welfare functions that are additive, increasing, and concave function of individuals' incomes. The corresponding literature on the social welfare foundations of mobility measurement is small, with contributions including Atkinson (1981a), reprinted as Atkinson (1983), Atkinson and Bourguignon (1982), Markandya (1984), and Gottschalk and Spolaore (2002). In this section, we focus on the nature of the social welfare functions employed in the mobility context; how these functions relate to mobility dominance results is discussed later.

The social welfare function (SWF) used in the multi-period context is a straightforward generalization of the one-period case discussed by Atkinson (1970). Overall social welfare, W, is the expected value (average) of the utility-of-income functions of individuals. In the two-period case, the utility-of-income function is $U(x,y)$, and weighted by the joint probability density $f(x,y)$. That is,

$$W = \int_0^{a_y} \int_0^{a_x} U(x,y) f(x,y) dx dy \tag{1}$$

where $U(x,y)$ is differentiable and a_x and a_y are the maximum incomes in periods

1 and 2. It is assumed that increases in income in either period are desirable, other things being equal (so positive income growth raises utility): $U_1 \geq 0$ and $U_2 \geq 0$.

Research in this tradition concentrates on the case in which the marginal distributions **x** and **y** are identical. In other words, the economic context is the same as the one used earlier to characterize positional mobility. All relevant mobility is encapsulated by the changes in individuals' ranks or by the transition matrix when individual incomes are classified into discrete classes. Atkinson and Bourguignon (1982) show that if the SWF is additively separable across time periods (so that $U_{12} = 0$, then income mobility is irrelevant for social welfare: only the marginal distributions matter.[8] If, instead, $U(x,y)$ is a concave transformation of the sum of the per-period utilities, then $U_{12} < 0$.

How does one interpret this sign? Atkinson and Bourguignon (1982) discuss the class of least concave functions associated with a particular preference ordering and the special case in which preferences are homothetic. In this situation, the utility function $U.$ is neatly characterized by two parameters: $\varepsilon > 0$ summarizing *aversion to inequality of multi-period utility*, and $\rho > 0$ summarizing the inverse of the elasticity of substitution between income in each period, i.e. the degree of *aversion to inter-temporal fluctuations in income* (Gottschalk and Spolaore, 2002, 295). The case $U_{12} < 0$ corresponds to the situation in which $\varepsilon > \rho$, i.e. in the social welfare assessment, multi-period inequality aversion offsets aversion to inter-temporal fluctuations (which are of course reducing multi-period inequality). Observe that when $\rho = 0$, an increase in income mobility must increase social welfare. With perfect substitution of income between periods, one is only interested in the reduction of multi-period inequality.

[8] See also Markandya (1984) and Kanbur and Stiglitz (1986).

Gottschalk and Spolaore (2002) point out that origin dependence has no role in the Atkinson-Bourguignon model.[9] In transition matrix terms, if there is any preference at all for income reversals ($\varepsilon > \rho$), not only does an increase in mobility represent a social welfare gain, but the complete reversal scenario is preferred to the origin independence one. This feature has relevance to the application of the social welfare framework to mobility measurement using stochastic dominance checks (discussed in the next section). The irrelevance of origin dependence suggests that the approach is less applicable to intergenerational mobility comparisons, since origin independence is the principle most commonly espoused in that context (see earlier).

However, an important contribution of Gottschalk and Spolaore (2002) was to show that greater origin independence can be social welfare improving if the SWF is generalized to take account of aversion to future income risk. In the two-period context, they drop Atkinson and Bourguignon's assumption that period-2 income is known with certainty in period 1. Individuals take conditional expectations of period-2 incomes based on observed period-1 incomes and the joint density of outcomes. With homothetic preferences, the utility function is now characterized by a third parameter, γ, summarizing the degree of aversion to second-period risk. As Gottschalk and Spolaore demonstrate,

> Origin independence reduces both multi-period inequality and intertemporal fluctuations, but increases future risk. Individuals will positively value origin independence as long as aversion to multi-period inequality and aversion to fluctuations dominate aversion to future risk (ε and

[9] See also similar remarks by Fields and Ok (1999a, 578–579).

ρ are not smaller than γ, and at least one of them is larger) (Gottschalk and Spolaore, 2002, 204).

In summary, evaluation of income mobility in terms of social welfare has payoffs. There is a single unifying framework. Within this, whether an increase in income mobility is social welfare improving depends on the priority given to different mobility concepts. For instance, reversals are less likely to be valued the greater the aversion to intertemporal fluctuations and to future income risk, but more likely to be valued the greater the aversion to multi-period inequality. Nonetheless one limitation of the SWF framework discussed so far is that it does not incorporate evaluations of mobility in the form of individual income growth – apart from aspects of this that overlap with the other concepts. One leading exception is the research by Bourguignon (2011) who shows that the Atkinson and Bourguignon results can be applied to comparisons of alternative 'growth processes' in the case in which the pair of marginal distributions relating to the first period are identical. However, this is a severe constraint on the applicability of the results.

An alternative strategy is to define SWFs explicitly in terms of income mobility – income changes rather than income levels. For example, one may assume that individual-level mobilities are represented by some measure of 'distance' between first and second period incomes for each individual i, $d(x_i, y_i)$, where the distance function is common to all individuals, and a social weight. Overall social welfare is the weighted sum over individuals of the d_i. King (1983) and Chakravarty (1984) assume that d_i is a function of period-1 and period-2 income ranks (the positional mobility case), and that re-ranking is desirable ($\partial W/\partial d_i > 0$) and the social weight is increasing in period-2 income. By contrast, for Van Kerm (2006,

2009) and Jenkins and Van Kerm (2011), d_i is a directional measure of individual income growth, and the social weight depends on base-year income ranks. For a more general discussion, see Bourguignon (2011), who discusses how the Atkinson-Bourguignon utility-of-income function, $U(x,y)$, can be re-written as $V(x,y-x)$ with the same properties on the differentials of the second (income change) argument. This framework would lead one to question, for example, the approach of Fields et al. (2002), whose SWF is the simple average of the d_i (equality of social weights), and so $\partial V/\partial x = 0$: mobility evaluations do not depend on initial income at all.

The main advantage of defining SWFs in terms of mobility directly is that there is great flexibility in the specification of the distance function d_i. The disadvantage of the approach is that it runs the risk of being ad hoc rather than a general unifying framework like the Atkinson and Bourguignon (1982) one. In particular, how should the social weights be specified? Unfortunately, the Bourguignon (2011) framework provides no simple answers.

The social welfare approaches described so far assume that W is a form of expected utility evaluation, though modified to context: Atkinson and Bourguignon (1982) incorporated preferences that were not time-additive and in addition Gottschalk and Spolaore (2002) abandoned complete predictability of income. A different approach altogether is to suppose that evaluations are based not on expected utility but prospect theory. Jäntti et al. (2013) explore this idea, utilising a utility function that incorporates reference-income dependence and loss aversion. The latter feature means that, over and above any preference for smooth rather than fluctuating incomes over time, fluctuations lower individuals' welfare directly since losses outweigh gains of equal size. There is therefore an asymmetric treatment of income decreases and decreases, as for the 'distressed gentlefolk' argument cited earlier but rather differently motivated. This approach is a promising area of research, and chimes with more popular expressions of the problem of growing income risk. Hacker and Jacobs (2008), for instance, specifically cite loss aversion as one of the factors related to the growth of income risk in the USA.

3. Mobility measurement

This section is about measuring mobility. First we discuss descriptive devices, by which we mean graphical and tabular methods for summarizing patterns of mobility. We consider them in more detail than other surveys because we think it is important to 'let the data speak' (though there are limits to which this is possible, as we show). Second, we describe how descriptive devices also have normative implications, being linked to dominance checks for mobility comparisons. Third, we consider scalar indices of mobility. Throughout the section we relate the descriptive devices and measures to the different concepts of mobility identified earlier. Most of the examples that we use are drawn from the intragenerational literature, reflecting their greater use in that context. But one of the lessons to be drawn is that the same methods could also be applied to the intergenerational context.

3.1. Describing mobility

In the two-period case, the bivariate joint distribution of income contains all the information there is about mobility, so a natural way to begin is by summarizing the joint distribution in tabular or graphical form.[10] How one proceeds depends on the nature of the data to hand, and the mobility concept of interest. We have been assuming that income distributions are continuous but in practice it is often convenient to represent the data in grouped form, or the data may intrinsically discrete as in the case of 'social classes'. In addition the information content of the descriptive device is related to the way (if any) in which the analyst

[10] We consider a summary device for mobility as equalization of longer-term income in the case when there are more two periods in the next subsection.

standardises the marginal distributions of any one bivariate distribution and, when making comparisons of bivariate distributions, makes further adjustments, e.g. to control for differences in average income between the bivariate distributions for two countries. If one is solely interested in pure exchange mobility (changes in relative position), then both issues are dealt with by working with the fractional rank implied by an individual's income rather than the income itself. In this case, all the marginal distributions are standard uniform variates and the same across time periods and countries.[11] But if the focus is on other mobility concepts, other standardisations may be used.

A mobility matrix, M, is constructed by first dividing the income range of each marginal distribution into a number of categories (which need not be the same in each period, but typically is) and cross-tabulating the relative frequencies of observations with each matrix cell: typical element m_{ij} is the relative frequency of observations with period-1 income in range (group) i and period-2 income in range j. The graphical representation of the discrete joint probability density function is the bivariate histogram. Alternatively, the mobility process may be represented by the transition matrix and the marginal distributions. Borrowing notation from Atkinson (1981a), suppose that there are n income ranges, with the relative number of observations in group k in period-1 is m_1^k for $k = 1, ..., n$, and correspondingly in period 2. The marginal (discrete) distribution in period-1 is summarized by the vector $m_1 = (m_1^1, m_1^2, ..., m_1^n)$ and correspondingly for period-

[11] Fractional (or 'normalised') ranks range between zero and one, with a mean of 0.5. Particular care needs to taken in their estimation when there are tied income values to ensure that these conditions are met. See e.g. Lerman and Yitzhaki (1989).

2. Hence,

$$m_1^k = m_2^k A \tag{2}$$

When the focus is on pure exchange mobility, the ranges typically refer to quantile groups. For example, in the case of decile groups, each group contains one tenth of the population. The transition matrix is then bistochastic. Mobility is entirely characterized by the transition matrix A.

An illustrative example is shown in Table 1. Mobility refers to changes in the relative positions in the USA between 1979 and 1988, and 1989 and 1998, with each individual's income defined as the equivalised real annual family disposable income of the family to which the individual belongs. The USA in the 1980s and the 1990s is a long way from the total immobility scenario (in which every cell percentage would equal to zero, except those on the leading diagonal which would equal 100%). Clearly, there is also neither origin independence (every cell entry equal to 10%) nor total reversal of positions. The general pattern is one of much short-distance mobility with long-distance mobility being rare. For example, of those individuals in the poorest tenth in 1989, around 42 per cent are also in the poorest tenth in 1998 with fewer than one per cent making it to the richest tenth. Of the richest tenth in 1989, around 46 per cent stay in that group, and less than 2 per cent are in the poorest tenth in 1998. More generally, the largest transition proportions are on or close to the matrix diagonal (Hungerford (2011) reports that 73 per cent of individuals remained in the same tenth or moved at most two deciles), and upward and downward mobility appears to be broadly symmetric. Since the US situation described in Table 1 is not particularly close to the standard mobility reference points, it is not straightforward to say whether there is a large or small amount of mobility. It is also of interest to say assess whether mobility increased between the 1980s and 1990s. Methods for mobility comparisons are discussed in the measurement section that follows. Further empirical evidence about within-generation mobility is presented in Section 4.

If the interest is in mobility other than of the positional kind, changes in the marginal distributions are also of interest. A particular example might be when the income class boundaries are defined as fractions of median income, or as fractions of the poverty line and there is interest in poverty rate trends as well as movements into and out of low income.[12] More generally, defining income group boundaries that are fixed in real income terms over time provides indications about individual income growth for individuals of different origins; if each period's incomes are standardized by period-average income, the information refers to income growth relative to the average.[13] (We say 'indications' regarding this mobility concept because its essence refers to income changes at the individual rather than group level.) Similarly, the dispersion across origin groups of individuals from a common income origin may be indicative of income risk, but the connection is not altogether obvious. Neither mobility matrices of this kind or conventional transition matrices are directly informative about mobility as longer-term inequality reduction.

Graphical summaries can complement and sometimes be more effective than tabular presentations: visual impact matters. Even transition matrices and comparisons of them can be visualised. We refer, for instance, to the use of transition probability colour plots introduced by Van Kerm (2011). Suppose individuals are

[12]For examples, see e.g. Hungerford (1993, 2011) for the USA and Jarvis and Jenkins (1998) for the UK.

[13]For examples, see Hungerford (1993), Hungerford (2011), and Jarvis and Jenkins (1998).

Table 1 Decile transition matrices: USA, (a) 1979–1988 and (b) 1989–1998 (percentages)

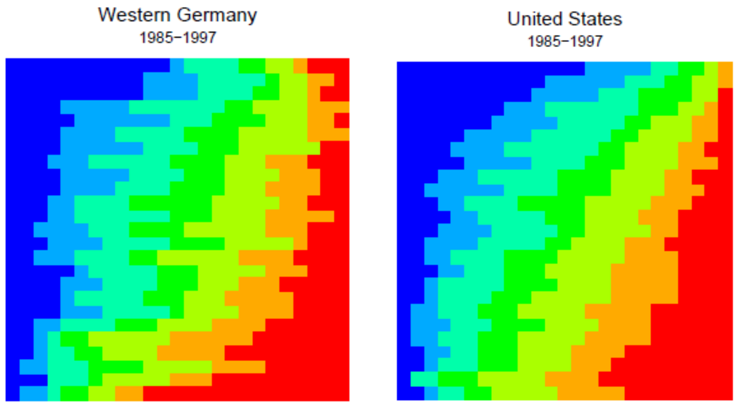
	Destination group									
Origin group	1	2	3	4	5	6	7	8	9	10
1979					**1988**					
1	44.3	18.3	12.4	9.2	7.1	3.0	1.8	2.0	0.7	1.3
2	18.1	25.3	21.0	11.7	7.5	5.4	4.7	3.2	1.9	1.1
3	10.6	18.2	15.3	16.8	11.6	9.0	8.8	4.9	3.1	1.7
4	7.2	8.9	14.0	14.0	14.7	15.7	12.0	5.6	6.0	2.1
5	6.1	9.2	10.9	12.8	13.3	16.9	12.3	7.5	7.7	3.4
6	4.1	5.2	8.8	10.3	11.8	10.0	14.2	16.9	12.6	6.2
7	3.5	6.5	6.9	8.6	10.4	13.4	13.3	16.8	13.4	7.2
8	3.1	4.6	3.2	7.7	12.3	9.5	12.6	15.7	17.7	13.6
9	1.2	2.2	4.8	6.3	6.9	10.2	12.2	14.7	18.0	23.5
10	2.1	1.5	2.8	2.5	4.2	7.0	8.5	12.8	18.6	40.0
1989					**1998**					
1	41.9	21.6	13.7	7.0	4.6	3.7	2.7	2.2	1.9	0.7
2	20.4	22.5	15.4	11.6	11.0	8.1	4.0	4.0	1.7	1.2
3	12.5	20.8	17.1	16.4	10.9	10.3	5.2	3.2	1.7	1.9
4	6.9	11.6	15.5	16.9	14.5	11.4	10.1	7.7	2.3	3.1
5	4.8	6.2	12.2	13.8	16.0	14.2	12.4	7.1	7.5	5.8
6	3.2	3.7	9.1	11.6	16.0	14.4	15.7	11.7	7.7	6.9
7	3.2	4.5	7.6	9.3	8.7	12.2	16.3	15.6	16.8	5.8
8	3.0	4.7	5.2	5.4	7.9	12.1	17.2	17.0	19.3	8.3
9	2.5	3.1	4.0	4.9	7.5	7.1	10.7	18.2	21.8	20.3
10	1.7	1.0	0.4	3.2	3.0	6.3	6.0	13.1	19.3	46.1

Note: Income refers to equivalized real annual family disposable income, distributed among all individuals (adults and children). The decile groups are ordered from poorest (1) to richest (10).
Source: Hungerford (2011, Tables 2 and 3), based on PSID data.

classified into vingtile groups in each of period-1 and period-2. For the visualisation, individuals are classified according to their income group in period-2, and lined up in rows with the poorest twentieth in one row at the top, the next twentieth in the row beneath, and so on down to the final row containing the richest twentieth. Each person is also tagged with their period-1 group membership using a colour coding system. Suppose the poorest twentieth in period-1 is represented by blue and the richest twentieth by red, and the intermediate groups are represented by the colours of the rainbow in between. If there were no changes in relative position over time, every one would remain in their period-1 income group: there would be a one-to-one correspondence between rows and colours. (Rows would consist of full blocks of the same colour.) If there no association between income origin and income destination, every colour would form an equal-sized block in each and every row. If there were complete rank reversal, the original colour scheme would be reversed, with the richest period-1 group (red) in the top row and the poorest period-1 group (blue) in the bottom row.

Examples of such representations, due to Van Kerm (2011), are shown in Figure 1 below for individuals' household income mobility between 1987 and 1995 in Western Germany (left) and the USA (right). It is immediately apparent that, over this twelve year period, there is substantial income mobility in both countries, and throughout the income distribution, including a small fraction of the richest twentieth falling to the poorest twentieth, and vice versa. But there is clearly no origin independence in either country, let alone complete rank reversal. Interestingly, however, it also clear that the main differences in patterns of mobility are at the bottom of the income distribution (more changes in relative position in Western Germany than in the USA). We return to this finding in the next section.

Figure 1 Transition colour plot examples



Source: Van Kerm (2011).

The particular advantage of the transition colour plots is their visual immediacy. However colour is not always available. The colour transition plots summarising income mobility in the book by Jenkins (2011a, Figure 5.1) were reproduced in black and white, and this reduced their effectiveness.

What about alternative devices? Perhaps the most straightforward way to summarize a bivariate joint distribution is using a scatterplot of period-2 incomes against period-1 incomes. Figure 2 provides a within-generation example using British income data for 1991 and 1992.

The advantages of the scatter plot are that it is very easy to produce and provides an immediate impression about the degree of immobility of incomes (the clustering around the 45° line), as well as the nature of the marginal distributions. For a focus on changes in relative position alone, the corresponding scatter plot would be of individuals' normalised ranks in each of the two periods. The main disadvantage is that potentially important detail is lost since the bivariate density

Figure 2 Scatterplot example

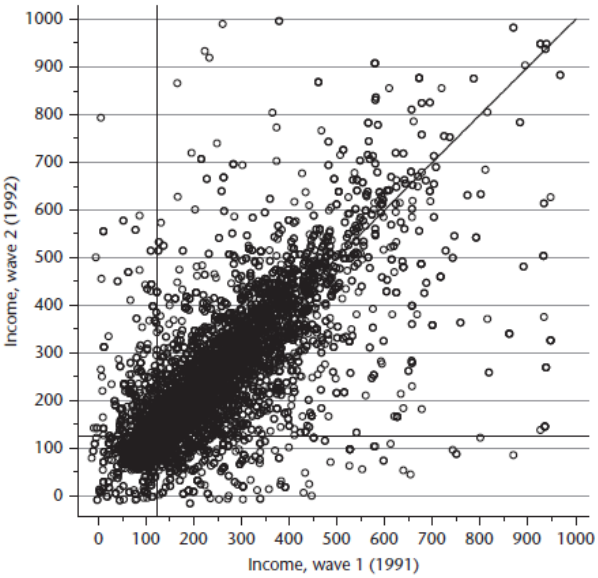


Figure 1.2. Scatter plot of 1991 and 1992 incomes

Notes: Sample of individuals (adults and children) present at BHPS waves 1 (1991) and 2 (1992) with incomes less than £1,000 per week. Each circle represents the incomes for the two years for each individual. The definition of income is given in the text (the adjustment for differences in household size and composition uses the Modified OECD equivalence scale). Incomes are expressed in pounds per week (January 2008 prices). The dark horizontal and vertical lines correspond to an income equal to 60% of contemporary median income (£123 per week for wave 1; £126 per week for wave 2).

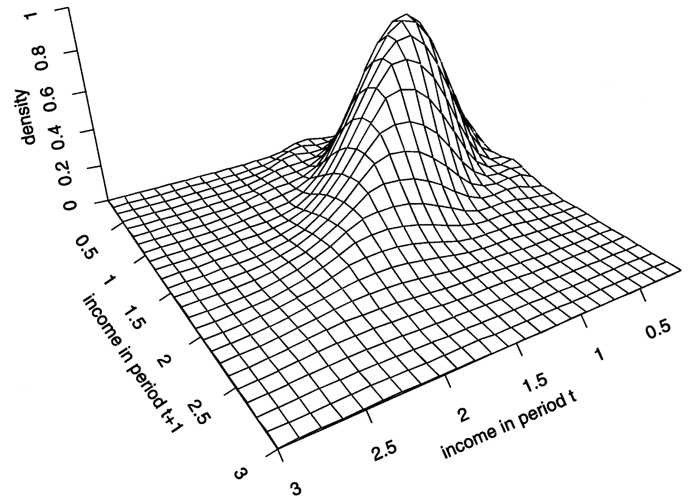
Source: Jenkins (2011a, Figure 1.2).

is not estimated: there is no difference to the eye between 10 observations with a particular combination of period-1 and period-2 incomes and 100 observations with the same pair of incomes.

One way to proceed is derive and plot the joint density. The simplest estimates to produce are those of the bivariate discrete density (essentially plotting the bivariate histogram – see above). However, there are well-known disadvantages of such discretization: as in the univariate distribution case, the estimates are sensitive to choice of income class boundaries and, of course, information within the ranges is lost with the grouping. Kernel density estimation methods avoid the problem because of the way in which they smooth data within a moving window rather than within fixed categories. Figure 3 shows a 'typical' joint bivariate density for West German family incomes for two consecutive years over the period 1983–89.[14] Note that incomes in each year are normalized by the contemporaneous median but, otherwise, the marginal distributions are not constrained to be same (so this a representation of exchange mobility alone). Compared to the scatterplot, the concentration of individuals on and around the 45° representing perfect immobility is readily apparent. However the fine detail remains difficult to ascertain, partly because the three-dimensional representation has to use a specific projection. What a reader perceives may change if the estimates are viewed from a different angle. Related, differences in marginal distributions are difficult to examine; so too is individual income growth. A further issue, shared with the scatterplot and bivariate histogram, is that it is difficult to compare a pair of bivariate distributions, e.g. for two different countries, even if the plots to be compared

[14]The source does not state which specific pair of years over the period 1983–89 was used for the calculations.

Figure 3 Bivariate density plot example



Note: the charts shows a 'typical' kernel density estimate for incomes in two consecutive periods.
Source: Schluter (1998, Figure 1).

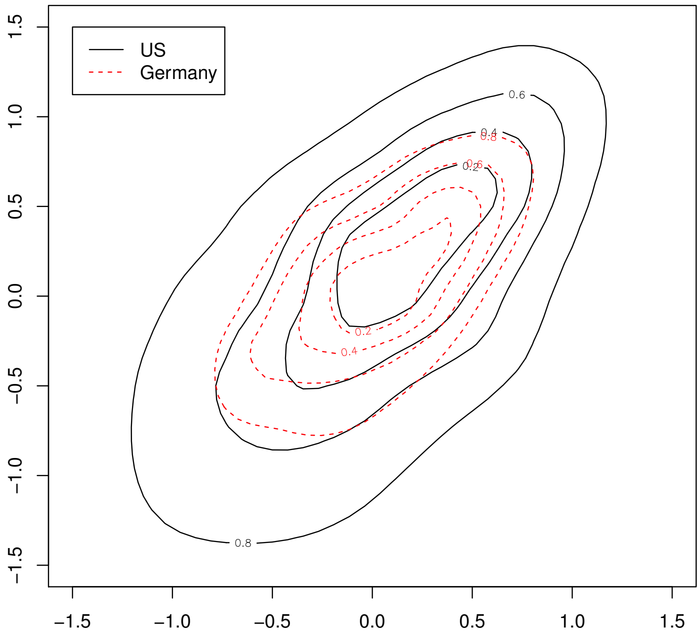
are placed adjacent to other. Overlaying one plot on another is far too messy but, without some form of overlay, detailed comparisons are constrained.

Both issues are resolved to some extent by summarizing the density estimates using contour plots in which contour lines connect income pairs with the same density. An example is provided using US and West German income data for 1984 and 1993 in Figure 4. Income refers to the log of equivalized family income expressed as a deviation from the national contemporaneous mean. Contour lines are drawn at values that separate the quintile groups for each country (the 20th, 40th, 60th, and 80th percentiles). The solid lines are for the USA, the dotted lines are for West Germany. As Gottschalk and Spolaore (2002) comment, the plot reveals multiple features of the joint distribution. Each contour line for Germany lies inside its US counterpart indicating greater cross-sectional inequality in the

USA. Clustering around the 45° immobility line is apparent for both countries but is greater for the USA. Also, the contour lines are generally flatter for Germany, meaning that expected period-2 income (conditional on period 1 income) varies less with period 1 in West Germany than it does in the USA. Gottschalk and Spolaore (2002) comment that this suggests a lower cross-period correlation in the USA, and they also point to a greater variation around the conditional means in the USA. Contour plots are also used in the US-West German comparisons by Schluter and Van de gaer (2011, Figure 2).

Just as contour plots for continuous income distributions correspond to mobility matrices, there are also devices for continuous incomes corresponding to the transition matrix. One requires estimates of the conditional density $f(y|x)$ which is straightforwardly estimated in principle using the fact that $f(y|x) = f(y,x)/f(x)$. Estimates of the numerator and denominator are derived across a grid of values of x and y using kernel density estimation. See Quah (1996) who refers to this concept as a 'stochastic kernel'. Applications to income mobility include Schluter (1998) and Schluter and Van de gaer (2011). Compared to unconditional joint density plots, the conditional density plots allow a more direct comparison of expected income growth across the base year income range. Examples are provided in Figure 5 based on data for the USA (top chart) and Western Germany (bottom chart) for 1987 and 1988. Income is equivalized net household income expressed relative to the 1987 median. Schluter and Van de gaer (2011, 11) point to not only the greater spread of contours in the USA indicating differences in marginal distributions, but also that the 'particular … feature of the conditional densities is the greater upward mobility of low-income Germans' compared to low-income Americans. Note the more distinct upturn of the con-

Figure 4 Contour plot example



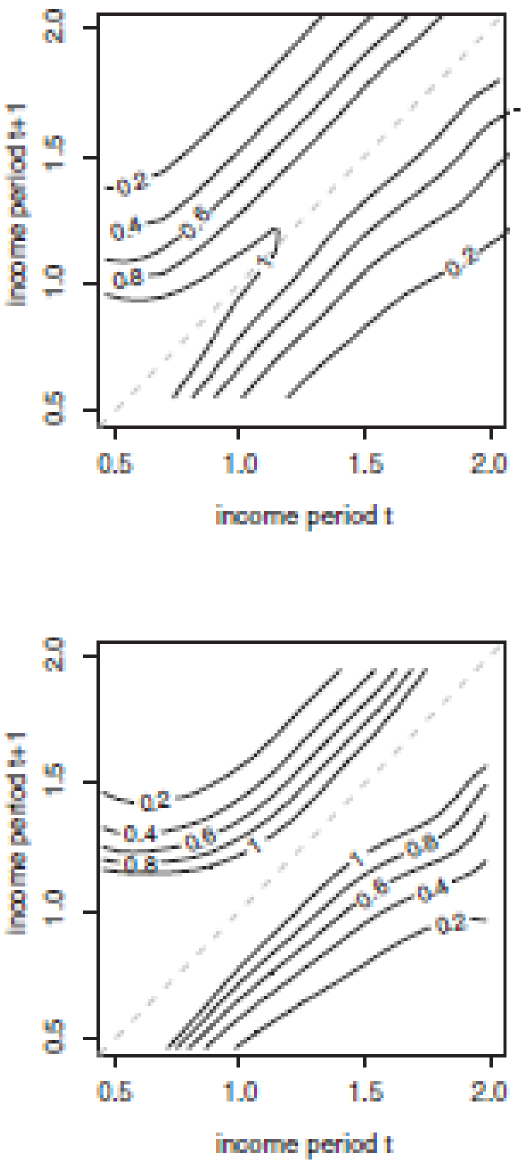
Note: the chart shows the kernel-smoothed joint density of income in 1984 and 1993 for the USA and West Germany, where income is post-tax post-transfer family income equivalised by the PSID equivalence scale, and income for each year is expressed as a deviation from the year-specific mean.
Source: Gottschalk and Spolaore (2002, Figure 1), redrawn by the authors.

tours in the top left of the Western German chart compared to the shape of the corresponding US contours.

Observe that conditional densities are not the same as conditional probabilities, which is what constitute the transition matrix. Estimation of the conditional (cumulative) probability density $F(y|x)$ requires integration over the marginal distribution of y. As Trede (1998) explains, estimates of $F(y|x)$ can be inverted to give the probabilities for second-period income conditional on particular values of first-period income ('p-quantiles'). Trede's device for 'making mobility visible' is a plot of these p-quantiles against first-period income values. Figure 6 shows one of these non-parametric transition probability plots using data for West German equivalized family incomes in 1984 and 1985. Incomes are normalised by the 1984 median, so 'growth mobility is not excluded from the analysis' (Trede, 1998, 80). In the extreme case of origin independence, each transition probability contour would be horizontal. If, instead, there were complete immobility so that second period incomes were completely determined by first period incomes, the contours would lie on top of each other. (In particular, if there were no change in median income, the contours would lie on the 45° line.) The greater the gaps between the contour lines, the greater is inequality in the second period. The slope of the contours is generally less than 45°, indicating some regression to the median. Figure 6 shows that, among individuals with median income in 1984, around 10 per cent have an income less than 0.7 and about 10 per cent have an income of at least 1.7 of the 1984 median in 1985. Methods closely related to Trede's are used by Buchinsky and Hunt (1999) to derive non-parametric estimates of transition probability estimates, which the authors report in tabular rather than chart form.

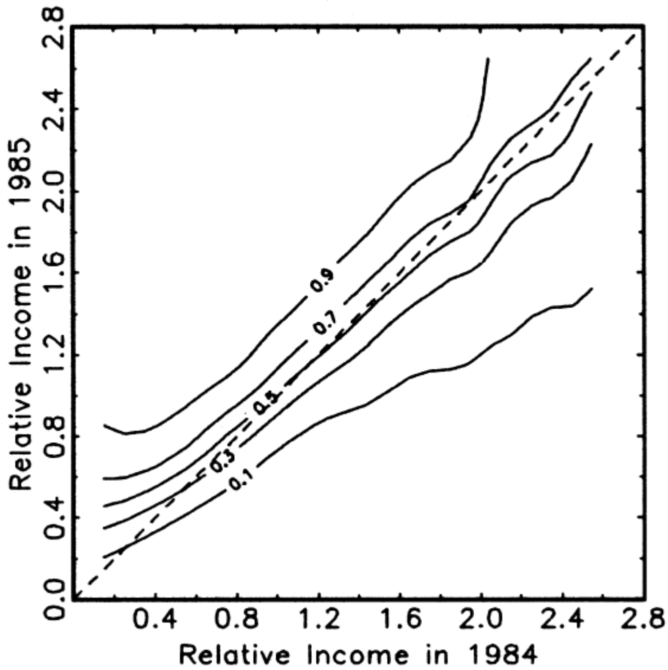
Patterns of mobility in the form of individual income growth are not shown di-

Figure 5 Conditional density plot example



Note: Year t refers to 1987; year $t+1$ refers to 1988. The top chart refers to the USA; the bottom chart to Western Germany.
Source: Schluter and Van de gaer (2011, Figure 2).

Figure 6 Non-parametric transition probability plot example.



Note: Relative income in each year equal to income divided by the 1984 median income.
Source: Trede (1998, Figure 1).

rectly in the devices discussed so far. The simplest way to focus on this aspect to define income growth at the individual level between the two periods using some measure of directional income growth (Fields and Ok, 1999b), thereby converting the bivariate joint distribution to a univariate distribution of income changes. Then all the devices commonly used for summarizing univariate income distributions are available with one important proviso. Income changes may be negative or zero and not restricted to positive values (and the mean change may also be zero or negative). However, the ratio of second-period income to first-period income is positive (assuming incomes are positive), and it is often convenient to use this metric. Schluter and Van de gaer (2011, Figure 2) present kernel density estimates of the distribution of income ratios. Comparisons based on plots of CDFs of income change distributions are also presented by Chen (2009, Figure 4) and Demuynck and Van de gaer (2012, Figure 1).

Observe that a CDF plot of this type is based on an ordering of individuals' income *changes* from smallest (most negative) to the largest. One is often interested in the extent to which individual income growth is 'pro-poor', that is whether income growth is greater for those at the bottom of the first-period income distribution relative to those at the top. In particular, pro-poor growth between two periods is a factor reducing the the inequality of second period incomes relative to first period incomes.[15] See also the discussion of social welfare functions in Section 2. Fields et al. (2003) plot the average change in log per capita income between two time points against income in the base year, for four countries. Com-

[15]But pro-poor growth does not guarantee inequality reduction, because it also leads to reranking which may have an offsetting effect. See Jenkins and Van Kerm (2006) for a fuller explanation and empirical examples.

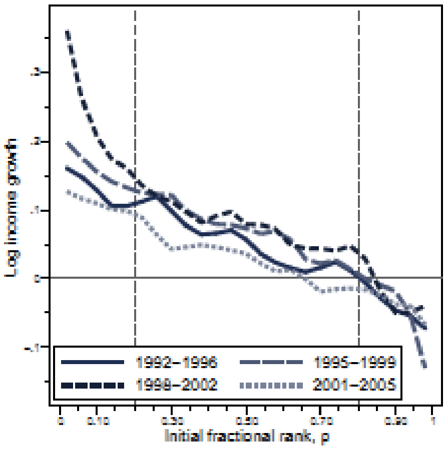
parisons across countries are constrained by the fact that income range on the horizontal axis (base-year income) varies tremendously. Comparability is enhanced if, instead, one plots individuals' average income change against their normalised (fractional) rank in the base-year distribution (with individuals ordered from poorest to richest). The horizontal axes in this case are bounded by 0 and 1. Such plots were developed by Van Kerm (2006, 2009) and independently by Grimm (2007). Extensive empirical examples are provided by Jenkins and Van Kerm (2011) for four five-year periods in Britain during the 1990s and 2000s, from which Figure 7 is taken. (Individual income growth refers to the change in the log of individuals' household income between two years.) It is clear that income growth is distinctly pro-poor in each of the subperiods, especially 1998–2002.[16]

In sum, we have reviewed a portfolio of tabular and graphical devices for summarising income mobility between two periods. By standardizing marginal distributions in different ways, different aspects of the mobility process can be focused on and, for individual income growth, there are separate devices.

Within-generation income mobility analysis has tended to use graphical summaries and comparisons rather more than between-generation mobility analysis, which has mainly relied on transition matrix tabulations for detailed summaries of the mobility process. In part, this emphasis is because the mobility concept most associated with intergenerational mobility is pure positional change totally separate from any changes in the marginal distributions. Nonetheless, there do appear to be opportunities forgone to use other methods to describe the distribution.

[16]As the authors explain, the negative slope to each curve is driven by 'regression to the mean', and so the substantive interest is mostly in the changes in slopes of the curves rather than the slopes themselves (as well as their heights).

Figure 7 Individual income growth and mobility profiles



Source: Jenkins and Van Kerm (2011).

Our final observation here is that there appear to be no straightforward descriptive summaries that directly highlight the concepts of mobility as longer-term inequality reduction or as income risk. We consider the former case below. In the latter case, one wants something analogous to the mobility profile but, instead, of summarising expected (average) income growth conditional on base year income or income position, one would summarise conditional income dispersion.

3.2. Mobility dominance

Dominance checks are a widely-used part of the analyst's toolbox for comparing univariate distributions of income. To what extent can and should this be the case for mobility comparisons? We identify three main approaches.

The most well-known dominance results are those of Atkinson and Bourguignon (1982). The results are derived with reference to the social welfare framework discussed earlier. Social welfare is the expected value of individuals' utility-of-income functions defined over period-1 and period-2 income, where individual utility is a concave transformation of the per-period utilities of income, and also increasing in each income.

Welfare comparisons of differences in mobility for bivariate distributions f and f^* are based the difference

$$\Delta W = \int_0^{a_y} \int_0^{a_x} U(x,y) \Delta f(x,y) dx dy \tag{3}$$

where $\Delta f(x,y) = f - f^*$ is the difference in bivariate densities and the same $U(.)$ is used for the social evaluation of each distribution. Cf. equation (3).

Analysis has focused on the case in which the marginal distributions **x** and **y** are identical, and social welfare functions satisfy the conditions $U_1 \geq 0$, $U_2 \geq 0$, and $U_{12} < 0$ (guaranteed if $U(x,y)$ is a concave transformation of the sum of the per-period utilities). Atkinson and Bourguignon (1982) show that a necessary and sufficient condition for a welfare improvement $\Delta W \geq 0$ is that $\Delta F(x,y) \leq 0$ for all x and y. That is, differences in the cumulative bivariate distribution are lower at each point (a first-order stochastic dominance condition).

What sorts of differences between joint distributions are associated with such conditions being satisfied? Atkinson and Bourguignon (1982) discuss the case of a 'correlation-reducing transformation' which leaves the marginal distributions unchanged but reduces the correlation between x and y:

$$\left\{ \begin{array}{ccc} & x & x+h \\ y & \text{density reduced by } \eta & \text{density increased by } \eta \\ y+k & \text{density increased by } \eta & \text{density reduced by } \eta \end{array} \right\} \text{where } \eta, h, k > 0.$$

When the bivariate distribution is represented using a transition matrix, this transformation is equivalent to shifting probability mass away from the matrix diagonal.[17] The cumulative density can be straightforwardly derived by cumulation across cells of the transition matrix starting from the lowest origin and destination group. For comparisons of two transition matrices, first-order welfare dominance exists when the difference in cumulative densities in corresponding cells is everywhere of the same sign. Atkinson (1981a,b) demonstrates the approach in action using intergenerational income data for Britain. Further examples are provided later in this chapter.

The dominance result is a notable addition to the tool box for comparisons of bivariate distributions but, perhaps surprisingly, has not been widely used. There are several reasons for this. The first is that, although relevant to evaluations of pure positional change mobility, the Atkinson-Bourguignon social welfare function is primarily sensitive to mobility as reversals rather than mobility as origin dependence (see the earlier discussion).[18]

Second, the first-order dominance checks have not provided clear cut rankings in practice (cf. Atkinson (1981a,b)). A natural reaction in this case is to seek unanimous mobility rankings according to more restricted classes of social welfare functions using second- and higher-order dominance checks. Atkinson and Bourguignon (1982) provide the theoretical results. The problem, however, is that the additional restrictions on the SWF are hard to interpret. They involve the

[17] But see also Jenkins (1994) and Fields and Ok (1999a). Both articles question the intuitive attraction of linking correlation-reducing transformations with more mobility if the transformations are made off the diagonal.

[18] Gottschalk and Spolaore (2002) modify the social welfare function but do not derive dominance results.

signs of third- and fourth-order partial derivatives of $U(x,y)$. Although Atkinson and Bourguignon point out that in the case of homothetic preferences, 'the signs of higher derivatives depend on the relation between the degree of "inequality-aversion" ... and the degree of substitution' between periods (Atkinson and Bourguignon, 1982, 18), i.e. the relation between parameters ε and ρ discussed earlier, they do not elaborate. It is difficult to understand what the sign conditions mean in everyday language.

Third, analysts may be interested in alternative concepts of mobility besides positional change. Individual income growth is the most prominent example of this situation. As discussed earlier, researchers have used social evaluation functions that are increasing functions of a measure of 'distance' between first and second period incomes for each individual i, $d(x_i, y_i)$, and defined social welfare as the socially-weighted sum over individuals of the d_i. For instance, Fields et al. (2002) undertake checks based on comparisons of pairs of cumulative distribution functions of d_i, where d_i is defined in six different ways in their empirical application. However, as remarked earlier, their social welfare function has unappealing properties. The challenges involved in the derivation of stochastic dominance results for Fields and Ok (1999b)-type measures of non-directional income movement are discussed by Mitra and Ok (1998). Van Kerm (2006, 2009) explicitly derives dominance results for two classes of social welfare function defined over the d_i. The first is when the social weights are simply assumed to be positive. Van Kerm shows that unanimous rankings by this evaluation function are equivalent to non-intersections of mobility profiles (the graphical device discussed earlier), a first-order dominance result. If one also assumes that the social weights are non-increasing functions of base-year income ranks (poorer individuals receive higher weights), unanimous social welfare rankings are equivalent to non-intersections of cumulative mobility profiles. Bourguignon (2011) shows that dominance conditions can be derived for social welfare functions more closely related to Atkinson and Bourguignon (1982) ones but the conditions are difficult to interpret intuitively and, in any case, are restricted to the case in which marginal distributions in the initial year are identical.

Dardanoni (1993) derives stochastic dominance results for rankings of mobility processes that are summarised by transition matrices, focusing on pairs of monotone matrices with the same steady-state income distribution.[19] The social welfare function is defined on a vector containing each individual's lifetime expected utility (the discounted sum of per-period utility values, where each income class has a common utility value associated with it; there is no within-class inequality in utility). Overall social welfare is not the average of the individual lifetime expected utilities, since linearity combined with anonymity would imply that mobility is irrelevant for social welfare assessments (as discussed earlier). Instead, Dardanoni's social welfare function is 'a weighted sum of the expected welfares of the individuals, with greater weights to the individuals who start with a lower position in the society' (Dardanoni, 1993, 371). Thus there is a direct parallel with the social weight system employed in the welfare function used by Van Kerm (2006, 2009).

Dardanoni shows that unanimous social welfare rankings by this evaluation

[19]Monotone transition matrices are those in which each row stochastically dominates the row above it. Essentially, being in a higher income class in the initial period means improved prospects in the second period. Most empirically-observed transition matrices are monotone or approximately so (Dardanoni, 1993). If a regular transition matrix characterises a first-order Markov chain, there is a constant long run steady-state marginal distribution corresponding to that matrix.

function can be checked by comparisons of the cumulative sums of the ‘lifetime exchange’ matrices corresponding to the two transition matrices. (A lifetime exchange matrix summarises the joint probability that an individual starting in some income class i is in lifetime income class j.) These matrices depend on the discount factor underlying them: although in general mobility processes which improve the position of initially poorer individuals are more highly valued, the timing of utility receipt also matters. Dardanoni (1993) provides additional results for checking the robustness of dominance results to the choice of discount factor. The fact that actual societies may not be in steady state and transition matrices may imply different steady-state distributions limits the applicability of the dominance results. Dardanoni (1993) acknowledges this, but also points out that this could be remedied by focusing on bistochastic quantile transition matrices (as Atkinson (1981a,b) did, in which case attention is restricted to changes in relative position). The orderings derived differ from those of Atkinson (1981a,b), however, because the social welfare function is different. For instance, Dardanoni (1993) points out that maximal mobility according to his ordering corresponds to the situation of origin independence, not rank reversal. Finally, we observe that Dardanoni’s dominance results appear to have been rarely used. As with the results of Atkinson and Bourguignon (1982), we suspect that is because applied researchers have found them relatively complicated to interpret and implement.

In sum, we have shown that there are dominance results for mobility comparisons, but the ‘toolbox’ is much less settled than it is for comparisons of univariate income distributions. In part, the reason comes back (again) to the fact that there is a multiplicity of mobility concepts, and (related) a lack of consensus about how to specify the social welfare function function in the bivariate case.

3.3. Mobility indices

In this sub-section, we review indices that might be used to summarise intra- and intergenerational income mobility. After a brief discussion of generic properties of indices, we discuss some commonly-used measures of bivariate association – what Atkinson et al. (1992) refer to as 'intuitive' measures – and then move on to more specialist indices, i.e. ones more directly corresponding to the various mobility concepts identified earlier. Whether an index focuses on positional change, individual income growth, longer-term inequality reduction, or income risk, accounts for many of its properties. There are general features on which we contrast indices.[20]

First, there are different normalisations. Although all indices equal zero in the case in which there is complete immobility, there is no shared maximum mobility value and, indeed, some measures have no maximum value imposed (principally the indices of income growth and income risk). Second, there is a distinction between 'pure' measures of positional change and other indices. The former indices, of exchange mobility, are sensitive only to the (re)ordering of individuals and hence with values unaffected by any monotonic transformation of each income between time periods (or, equivalently, also unaffected by changes in the marginal distributions of income). By contrast, structural measures register mobility even if ranks are constant but the income values associated with those positions change over time.

Third, and related, indices differ in how they reflect income changes that are common to all persons, whether by the same proportion or by the same absolute amount. Measures are 'strongly relative' ('intertemporally scale invariant') if

[20]This discussion draws on Jenkins and Van Kerm (2009).

equiproportionate income growth does not affect the mobility assessment. Measures are 'weakly relative' (or 'scale invariant') if the units in which income are measured are irrelevant but, by contrast with strongly relative measures, equiproportionate income growth may count as mobility.[21] There are also translation invariance counterparts of these properties. Again, the principal distinction is between measures of pure positional change (exchange mobility) – which satisfy both intertemporal translation and scale invariance – and the other indices. For example, most indices of longer-term inequality reduction are scale invariant but not intertemporal scale invariant. Most indices of individual income growth are neither intertemporal scale or translation invariant.

Fourth, there is the issue of directionality, which refers to the roles played by the base year and current year in mobility assessments. An index is directional if it matters whether a particular income change refers to a change from a base year to a current year or vice versa. This is relevant if one wishes to take the temporal ordering of changes into account, and this is particularly important for measures of individual income growth, as one would want to treat differently an income change from 100 to 150 and an income change from 150 to 100. One would want the former to represent an improvement in circumstances, and the latter a deterioration.

Fifth, indices may satisfy various decomposability properties. Mobility indices may be (additively) decomposable by population subgroup, as inequality indices are, according to which total mobility can be written as the weighted sum of mobility within subgroups defined by an exhaustive non-overlapping partition of the population in question according to some characteristic (e.g. sex, age, or

[21] For more on this distinction, see Fields and Ok (1999a).

education) plus (possibly) a term representing between-group mobility. Most indices of longer-term inequality reduction can be decomposed thus, and so too can individuals of income growth though there is typically no between-group mobility term in that case.[22] Measures based on changes in ranks are not decomposable, because in general there is no one-to-one correspondence between an individual's rank in his/her subgroup and in the population as a whole.

A second type of decomposability is into structural and exchange components. Unlike decompositions by population subgroup, these decompositions are not additive, and rely for their derivation on the use of counterfactual income distributions representing the situations when there is an absence of exchange mobility or of structural mobility. On this, see e.g. Markandya (1984), Ruiz-Castillo (2004), and especially Van Kerm (2004).

A third decomposition idea, most commonly exploited in measures of individual income growth or income flux refers to inter-temporal consistency – whether mobility calculated for income changes between times t and $t+s$ is the sum of the mobility between times t and $t+r$ and between $t+r$ and $t+s$ (with $r<s$) or, alternatively, the product. This is the concept of additive (alternatively, multiplicative) path separability or path independence. A fourth mobility-related decomposition relates changes between two years in *inequality* measured by the change in a generalized Gini coefficient to the sum of two mobility indices – one of progressive income individual income growth and the other of reranking. See Jenkins and Van Kerm (2006) for details.

[22] Such decompositions have mostly been used to provide anatomies of mobility during a single time period, rather than for accounting for the correlates of changes in mobility between two time periods in terms of the relative importance of changes in subgroup sizes and mobilities and between-group mobility changes.

We refer to these features at several points in what follows. We now turn to consider the most commonly-used 'statistical' or 'intuitive' measures of (im)mobility are the Pearson (product moment) correlation, r, between the log of incomes at two time points or its close sibling Beta (β), the slope coefficient from a least-squares linear regression of log(period-2 income) on log(period-1 income):

$$r = \beta \frac{\sigma_1}{\sigma_2} \tag{4}$$

where σ_1 and σ_2 are the standard deviations of log incomes in periods 1 and 2. Put differently, r is β scaled by the changes in inequality in the marginal distributions as assessed by the variance-of-logs inequality index, and it measures the degree of regression to the (geometric) mean in income between periods 1 and 2. $H = 1 - r$ is the Hart (1976) index of mobility, the properties of which are discussed in detail by Shorrocks (1993), and often used in the intergenerational mobility context. H ranges between -1 and 1, and $H = 0$ in the case of complete immobility.

Beta, as we shall discuss later, has been used in almost every empirical study of intergenerational income mobility ($1 - \beta$ is an index of mobility). This is perhaps surprising because it is the positional mobility concept that has been of the greatest interest in this context, and yet Beta and r (or H) reflect structural as well exchange mobility. A perfect linear relationship between period-2 and period-1 incomes ($r = 1$, $H = 0$) is consistent with unchanged ranks but also income growth. It is sometimes argued (see Section 5) that r is more suitable than Beta as a measure of income (im)mobility when undertaking cross-national comparisons on the grounds that r controls for differences in marginal distributions. But such controlling is only done to a rather limited extent, since changes in inequality are only one distributional feature (and uses one particular inequality measure to do so). Differences in marginal distributions would be fully controlled, however, were analysts to employ the Spearman rank correlation rather than r (because both marginal distributions would be standard uniform distributions), and this would also have the advantage in the intergenerational context of focusing on positional change. Note also D'Agostino and Dardanoni (2009a) who provide an axiomatic characterisation of the Spearman rank correlation as an measure of exchange mobility, thereby taking it beyond being a mere 'statistical' index.

A second question regarding Beta and r is why they should be calculated using log incomes rather than incomes. To be sure, Beta is a unit-free measure (an elasticity) but this begs the question of whether we are interested in immobility as the lack of a linear or log-linear relationship.[23]

All in all, there are probably two reasons for the continuing widespread use of Beta and r in the intergenerational mobility literature. The first, as we discuss in Section 5, is that various methods to assess the impact of measurement error, and discussions of the relationship between Beta, r, and sibling correlations, rely on properties of regression and moments. The second reason is simply inertia: researchers continue to use Beta because they want to compare their estimates with those of others before them. The main problem with Beta as a measure that intergenerational mobility researchers have noted is its scalar nature rather than more fundamental concerns about the mobility concepts that are reflected in it. Their developments of 'vector' measures take us some back towards the more detailed graphical summaries of bivariate distributions discussed earlier.

For example, instead of fitting a single log-log regression, researchers have estimated quantile regressions of period-2 incomes on period-1 incomes (see e.g.

[23]It might also be that log(income) is viewed as a measure of the utility of income, but we have not seen that argument stated explicitly in the income mobility literature.

Eide and Showalter, 1999). (The periods refer to offspring and parental generations, rather than years within a generation.) However, it is not immediately clear what the estimates tell us about (im)mobility. At a technical level, the answer is clear. A quantile regression of, say, the tenth percentile of son's income on father's income allows the researcher to express the tenth percentile of son's income as a function of father's income. The quantile regression coefficient on father's log income then measures the elasticity of the particular quantile of son's income with respect to father's income. Differences in estimates across the quantiles tell us how sensitive different parts of the son's distribution conditional on father's income are to small changes in father's income. However, why these marginal changes, measured by slopes of the conditional quantiles, are of interest is not obvious.

One way to interpret the information provided by the vector of quantile regression coefficients is in terms of the full conditional distribution of son's income. A picture that is familiar to most students of regression analysis is the fitted regression line from a regression with one explanatory variable, with the distribution of the error term around that line drawn in a a few different levels of the explanatory variable. In the classical case, all those distributions are the same, or at least have the same variance, i.e. the error term is homoscedastic. If the distribution of y_2 (for sons), conditional on y_1 (for fathers), is homoscedastic, all estimated quantiles would have the same slope coefficient (save for random error). If the regression slopes are greater for higher percentiles of the son's distribution, suggests that the conditional variance of son's may be increasing with father's income.

Comparisons of the quantile regression estimates with the Beta from the log-linear regression can also reveal some further aspects of the distribution of period-

2 incomes conditional on period-1 incomes. The log-linear regression line gives the conditional expectation and the regression slope for the 50th percentile gives the expected median. If we find that for most of the range of father's incomes that the conditional mean for sons is lower than that of the relevant father's income, this suggests that, conditional on father's income, son's income is skewed to the left, rather than skewed to the right (as is usually true for income distributions). Observe also that one can use the predicted percentiles for different values of father's income to generate summary distributional statistics for the conditional distribution. For instance, one can derive the (discrete) cumulative distribution functions for period-2 (son's) income, conditional on a set of period-1 (father's) income percentiles. One could then check e.g. whether the distributions first-order stochastically dominate each other. (This relates to Monotonicity assumption of Benabou and Ok (2000).) Any conditional summary statistics can be generated this way, including inequality statistics such as percentile ratios. Individual income growth summaries such as in Figure 7 refer to conditional expectations (means) of distributions like these (except that they are typically drawn at different base-year ranks rather than different base-year income levels). In sum, there are close connections between some of the 'vector' measures of mobility and the graphical devices discussed earlier. In what follows, we return to focusing on scalar measures.

The second most common type of intuitive measure is an 'Immobility Ratio' (IR). IRs summarise how much clustering there is on (or, sometimes, also around) the leading diagonal of a transition matrix – and hence summarize positional change. For example, for a decile transition transition matrix, an IR might be defined as the percentage of all persons who remain in the same decile group between the two periods. (A variant would be to calculate the percentage remaining in the same decile group or one either side of it.) Clearly, the IR equals 100 per cent in the complete immobility scenario. If, instead, there is complete independence of origin, the IR for a decile transition matrix is 20 per cent (52 per cent in the variant). An index of mobility can easily be calculated as $1 - IR$.

Shorrocks (1978b) proposed a mobility index closely related to the IR, a Normalized Trace measure, equal to $[n - \text{trace}(A)]/(n-1)$, where A is the transition matrix with n income classes and $\text{trace}(A)$ is the sum of the transition proportions on the leading diagonal of A. This provides a neatly normalized index: with complete immobility, $\text{trace}(A) = n$ and the Normalized Trace equals 0; with complete origin independence $\text{trace}(A) = 1$ and so the Normalized Trace equals 1.

By construction, an IR and the Normalized Trace are insensitive to any differences between transition matrices aside from those in the respective diagonals. Bartholomew's (1973) Average Jump index is a positional mobility measure that addresses this aspect. It is equal to the number of income class boundaries crossed by an individual (whether upwards or downwards), averaged over all individuals (and equal to 0 in the complete immobility case). One feature of the Average Jump index is that it generalizes to the situation when the researcher has individual-level data on incomes rather than simply grouped data (a transition matrix). The index is then the population average of the absolute changes in fractional ranks (i.e. the ranks normalized to range from 0 to 1 rather than from 0 to the population size).

The transition matrix also offers, as a byproduct, measures of low and high income persistence defined in terms of rank order immobility. Mobility matrices, in which class boundaries are defined in real income terms, also do so. For example, if the lowest class boundary in each period is the poverty line, then the mobility matrix shows the proportion of individuals who are poor in a base period who are still poor in some later period (or who escape poverty). And with repeated longitudinal data for multiple periods, it is straightforward to define 'survival probabilities' – the chances that a person remaining poor for τ years, where $\tau = 1, 2, \ldots$. One can also define measures of high income persistence analogously (and we report some estimates in the next section).

This way of thinking about low-income persistence provides a link with the more well-known literature on poverty persistence, especially the approach pioneered by Bane and Ellwood (1986) in which consecutive periods spent poor are aggregated into spells (summarizing the total time spent poor). Rather than looking at spells of poverty (or affluence) one may simply count the number of times each person is poor (or rich) over some fixed time horizon and summarize that distribution. Low income persistence statistics of this nature are published by e.g. the UK Department for Work and Pensions (Department for Work and Pensions, 2009) and the Statistical Office of the European Communities (Eurostat).

There is also a nascent literature developing indices of poverty persistence that focuses attention on the way in which people's experience of poverty over time is aggregated, and hence how to compare, say, a history of three consecutive years in poverty followed by three years of non-poverty, with a history in which the person was poor every second year in the six year period. Research in this tradition includes Bossert et al. (2012), Dutta et al. (2013), Foster (2009), Gradín et al. (2012), Mendola et al. (2011), Mendola and Busetta (2012), and Porter and Quinn (2012). This literature works with a time horizon of fixed length and summarizes individuals' experiences within that window, ignoring whether whether poverty spells were already in progress at the beginning of the window, or remained in progress at the end of the window. If one wants to derive the shape of the poverty spell distribution in the population (rather than simply the sample), these issues of left-and right-censoring of poverty spell data (which are ubiquitous) need to be accounted for. They are given great attention in the spell-based literature on poverty persistence following Bane and Ellwood (1986) which, on the other hand, ignores longitudinal aggregation issues.

We now turn to a selection of more specialized indices of positional change that have been less commonly-used than the ones mentioned so far. The first is the Gini Mobility index of Yitzhaki and Wodon (2005). It is based on the idea of mobility as (lack of) correlation but, instead of using the Pearson or Spearman correlations, it uses the Gini correlation which, like other Gini-based measures, focuses on ranks rather than income levels per se. The Gini correlation between the income distributions in periods 1 and 2 is

$$\Gamma_{12} = \frac{\text{cov}(y_1/\mu_1, F_2)}{\text{cov}(y_1/\mu_1, F_1)} \tag{5}$$

where y_1/μ_1 is period-1 relative income, i.e. income divided by the period-specific mean income, F_1 and F_2 are the fractional ranks in the two periods, and cov(.) means covariance. Since $1-\Gamma_{12}$ is a directional measure of mobility ($\Gamma_{12} \neq \Gamma_{21}$ in general), the overall Gini Mobility index is defined as a weighted average of the two possible directional measures, where the weights depend on the inequalities in each marginal distribution, measured using the Gini coefficient (G). That is,

$$\text{Gini Mobility index} = \frac{G_1(1-\Gamma_{12}) + G_2(1-\Gamma_{21})}{G_1+G_2}. \tag{6}$$

Yitzhaki and Wodon (2005) show that if there is no positional change, the Gini mobility index equals 0, it equals 1 if there is complete origin independence, and equals 2 if there is complete rank reversal.[24]

[24]The index of reranking used in Jenkins and Van Kerm's (2006) decomposition of inequality

The Gini mobility index uses a particular weighting function when aggregating changes in individuals' ranks but one that is not immediately clear. By contrast, the King (1983) index takes an explicit welfarist approach in which differences in social weights across ranked are defined and tuned parametrically. The basic building block is the 'scaled order statistic' for each individual i, s_i, equal to the absolute magnitude of the difference between i's period-2 income and the period-2 income that i would have had, were s/he had maintained the same rank in period 1, all expressed relative to mean period-2 income. There is complete immobility if $s_i = 0$ for all individuals. Using an approach analogous to that of Atkinson (1970), King defines his mobility index as the proportion of total period-2 income that society would be prepared to forego in order to have the mobility observed rather than complete immobility (positional change is socially valued). Assuming a homothetic form for the social welfare function leads to a mobility index depending on two parameters – the degree of aversion to period-2 income inequality and the degree of aversion to income immobility (larger values of which give greater social weight to mobility, other things being equal). For generalizations of and commentary on King's approach, see Chakravarty (1984) and Jenkins (1994).

On the one hand, the systematic welfarist approach used by the King index (and others like it) has much to recommend it. On the other hand, it relies on a rather special characterization of what counts as mobility at the individual level (the scaled order statistic), in which the implications for social welfare of changes in ranks are summarized by income values and a particular no-mobility thought experiment. Observe also that the social welfare function does not depend directly on incomes in period-1, except in so far as they characterize s_i. Compare this with

change into reranking and income growth components is a directional Gini correlation.

the Atkinson and Bourguignon (1982) social welfare function defined over incomes in periods 1 and 2 that was discussed earlier. Gottschalk and Spolaore (2002) embrace (and extend) the latter in the first part of their article, but when they later define specific mobility indices, they use an approach that is similar to King's (1983) in that the social gains from mobility are all expressed relative to a complete immobility reference point, and this is defined in the same way as in the King index. To develop their indices, Gottschalk and Spolaore (2002) also assume homotheticity in their SWF, and the resulting class has three parameters, representing aversion to multi-period inequality, rank reversal, and origin dependence. Although each parameter has a clear interpretation when taken individually, thinking about the implications of different combinations of values is more complicated. We are aware of no use of the Gottschalk and Spolaore (2002) indices other than by the authors themselves.

We move now to consider measures of individual income growth. As mentioned in the Introduction, these incorporate two basic ideas: (i) income increases for an individual count positively in the social calculus and income decreases count negatively, and (ii) total income growth is a function of income growth values for each individual (and the measure of each person's income growth depends only on their incomes in the two periods, and not the incomes of other people). The first idea refers to the directionality of the income growth measure. The second is a form of decomposability property across individuals, and also leads to aggregate measures that are decomposable by population subgroup. Although the empirical applications of these measures have all been to intragenerational income mobility, the indices could also be applied to intergenerational income mobility when there is interest in structural mobility over and above exchange mobility.

Fields and Ok (1999b) provide the most well-known aggregate measure of directional income growth in this tradition.[25] They show that directional measures of individual income growth that satisfy the properties of scale invariance, subgroup decomposability, and multiplicative path separability must take the form

$$D1 = c[\frac{1}{N}\sum_{i=1}^{N}(\log(y_i) - \log(x_i))] \quad (7)$$

where c is a normalizing constant which may be set equal to one, and N is the population size. That is, overall income growth is the average of individuals' proportional income growth. This is the case in which (directional) distance between incomes, $d(x_i, y_i) = \log(y_i) - \log(x_i)$. Observe that the social weighting scheme treats all individuals the same, regardless of their base-year income and regardless of how much income growth each experiences. Both these aspects, and some other generalizations, have been incorporated in later work.

Demuynck and Van de gaer (2012) use axioms similar to Fields and Ok (1999b), but explore the implications of assuming additive as well as multiplicative path separability, and also of imposing an axiom of 'priority for lower growth' that builds in aversion towards inequality in the individual growth rates. The axiom states that 'aggregate growth increases more when additional income growth is allocated to individuals with lower income growth than when it is allocated to individuals with higher income growth' (Demuynck and Van de gaer, 2012, 750). The authors prove that the measure satisfying their axioms is of the form:

$$S = \frac{1}{N^{\delta}}\sum_{i=1}^{N}(i^{\delta} - (i-1)^{\delta})\widetilde{d_i}, \text{ with } \delta \geq 1. \quad (8)$$

[25] We review their non-directional indices of income 'flux' later.

Given a measure of individual-level income growth for each person, d_i, $\widetilde{\mathbf{d}}$ is the vector of such income 'distances' ordered from largest to smallest. If multiplicative path separability is among the axioms, then $d_i = (y_i/x_i)^{\pi}$ or if, instead, additive path separability is assumed, then $d_i = \pi(\log(y_i) - \log(x_i))$, with $\pi > 0$ in both cases.

When $\delta = 1$, the general indices reduce, in the first case, to the directional measure of Schluter and Van de gaer (2011) and, in the second case, to the Fields and Ok (1999b) measure described above (with $\pi = c$, and also normalized to 1). In Schluter and Van de gaer's (2011) index, π is a sensitivity parameter, with higher values increasing the 'distance' measured between incomes in period-1 and period-2 but keeping ranks the same. Demuynck and Van de gaer (2012, 754) remark that when $\delta = 1$, correlation-reducing transformations to incomes in either period of the kind discussed earlier increase mobility according to S but $D1$ is insensitive to such changes. In the more general case, with $\delta > 1$, more weight is given to individuals with smaller values of d_i. When $\delta = 2$, the weights are like the weights used to characterize the Gini coefficient of inequality and when $\delta \rightarrow \infty$, only the smallest d_i counts. In these more general cases, S is no longer additively decomposable by population subgroup, and it is possible for correlation-decreasing transformations to reduce mobility. The larger question, however, concerns the social desirability of 'priority for lower growth': why should we be concerned about the inequality of individual growth rates (the d_i) independently of incomes in the initial or final period? Because of this issue, and (related) the greater complexities involved with using a two-parameter index, we conjecture that empirical researchers will be more likely to use S with $\delta = 1$ than the more general case.

The directional measures of income growth of Jenkins and Van Kerm (2011) are built using a different approach, and relate to a social welfare function defined as the weighted average of the the d_i (see Section 2), in which the social weights are a decreasing function of period-1 income ranks, defined using a single-parameter generalized Gini scheme.[26] Put differently, Jenkins and Van Kerm (2011) build in a social preference for pro-poor income growth, and the choice of different parameter values provides indices ranging from limiting cases in which aggregate growth is the simple average of the d_i values (as with $D1$) or in which only the growth rate for the poorest period in the initial year counts.[27] Palmisano and Van de gaer (2013) provide an axiomatic characterization of the Jenkins and Van Kerm (2011) class of measures. The usefulness of these indices rests largely on the extent to which the concept of pro-poor income is viewed as a desirable normative principle: see the discussion in Section 2 about the link between progressive income growth and inequality reduction.

The pioneering paper on mobility as reduction in the inequality of longer-term income is by Shorrocks (1978a). The essential insight is that, were one to longitudinally average each person's income over a number of years (T, say), the inequality in these averaged incomes would be less than average annual inequality because each individual's income fluctuations would be smoothed out and no longer contribute to aggregate cross-sectional dispersion in incomes for the T-year accounting period. Shorrocks (1978a) defines a measure of income rigidity, $R(T)$, equal to the ratio of inequality among T-averaged incomes ('longer-term'

[26]This scheme is like Demuynck and Van de gaer's (2012), except that the weights are applied to period-1 ranks, and not to d_i values.

[27]Jenkins and Van Kerm's (2011) classes of measures focus on the cases in which income growth rates are defined in proportional or absolute terms, i.e. $d_i = \log(y_i) - \log(x_i)$ or $d_i = y_i - x_i$.

inequality) to the weighted average of single-year inequality values:

$$R(T) = \frac{I[Y(T)]}{\sum_{k=1}^{k=T} w_k I[Y^k]}. \tag{9}$$

$I[Y(T)]$ is the inequality in T-averaged incomes, and $I[Y^k]$ is inequality in period-k incomes calculated using the same inequality index (e.g. $I[Y^1]$ is inequality in period-1 incomes). The weights w_k are the proportion of aggregate T-averaged income received in period k, i.e. $w_k = \mu_k/\mu$, and the weights sum to unity. Shorrocks shows that if one restricts attention to conventional relative inequality indices, then R is bounded above by 1. When there is complete rigidity in relative incomes, inequality in each period corresponds to inequality for the longer accounting period.[28] The more frequent or larger that income changes are, the less rigid the income system, and thus one may define a measure of mobility: $M(T) = 1 - R(T)$.

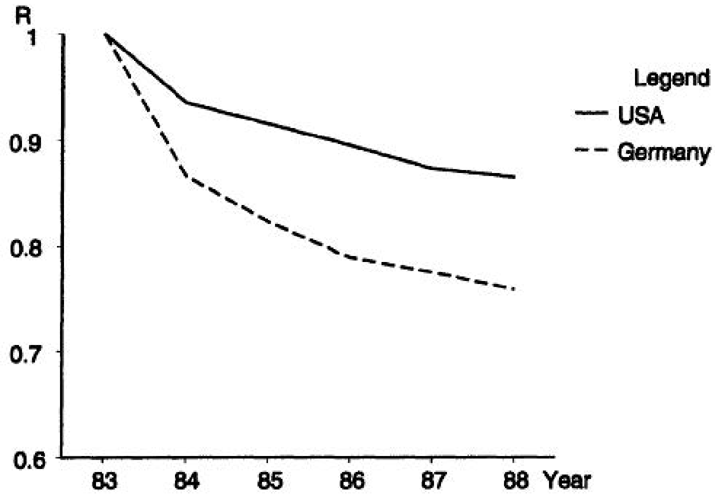
As Shorrocks (1978a, 178) puts it, 'mobility is regarded as the degree to which equalisation occurs as the observation period is extended'. In terms of the properties discussed earlier, $M(T)$ is a non-directional index and scale invariant (because it is defined in terms of relative incomes), but not inter-temporal scale invariant (given the way in which the per-period weights are defined). Although R and M are usually used to describe within-generation mobility, in principle they could also be used to describe mobility between generations. R and M are distinctive in that they are well-defined when there are data for many periods, but they can also be calculated if there are only two (the typical situation with intergenerational data).

[28]By conventional relative inequality indices we mean all those that are convex functions of relative incomes (incomes expressed relative to the mean income), i.e. all those that satisfy the Principle of Transfers. This excludes indices such as the variance of log incomes. R is bounded below by zero, assuming all incomes are positive.

A nice feature of the Shorrocks approach is that it can be used in two ways. The first is to calculate a single index value conditional on a particular value of T (and inequality index). This fixed-window calculation can be employed, e.g., to examine trends in income mobility over time in a country using moving fixed-width windows. Second, one can examine how $R(T)$ changes as T is increased from its minimum value of 1 to some larger maximum (i.e. there is one window, the width of which is varied). The resulting rigidity and mobility profiles provide a straightforward graphical device for comparisons of the extent of mobility within a country, and also comparisons across population subgroups and countries. Rigidity profiles for the USA and Western Germany from a pioneering cross-national study of income mobility discussed further in the next section are shown in Figure 8. Observe that the profile for Western Germany lies everywhere below that for the USA: whatever the accounting period used, mobility is greater in Western Germany than in the USA.

Clearly, the values derived for the Shorrocks indices are conditional on the inequality index employed for the calculations. It is also well-known that inequality indices differ in the sensitivity to income differences in different parts of the income distribution (Atkinson, 1970). So, it is important to know how estimates of rigidity and mobility relate to choice of inequality index, and how differences in inequality index sensitivity translate into mobility index sensitivity. It has been found as an empirical regularity, from Shorrocks (1981) onwards, that using different indices can make a big difference to the estimates of R derived and also that the Gini coefficient tends to show greater R values than other inequality indices. The explanation is that '[since] the main effect of cumulating income is to average out incomes that are temporarily high or low, the strongest egalitarian trend

Figure 8 Income rigidity (longer-term inequality expressed as a fraction of total inequality) falls as the time period is lengthened



Note: Income is post-tax post-transfer income. The Shorrocks rigidity index *R* is computed using the Theil index of inequality. ‘Germany’ refers to the federal states of Western Germany.
Source: Burkhauser and Poupore (1997, Figure 2).

will be found in the tails. The distribution of relative incomes in the middle range is not substantially affected by cumulating incomes over time' (Shorrocks, 1981, 182). Combine this information with the fact that the Gini coefficient is relatively insensitive to income transfers in the tails of the income distribution and we have the result.

The sensitivity of R to the choice of inequality index is examined more systematically by Schluter and Trede (2003). For the two-period case, they show that the global rigidity measure, R, can be expressed, to a good approximation, as the weighted average of 'local' rigidity comparisons at each point along the income range of a value for the longer-term averaged income, and the average of the per-period distributions. Differences in global measures arise, therefore, from a combination of differences in the way the different inequality indices summarize local comparisons at each point along the income range, and the different weighting systems that they incorporate. Schluter and Trede (2003) show that the sensitivity of mobility measure to choice of inequality index is partly dependent on data, but they also show some clear empirical regularities. For example, the weighting functions for commonly-used generalized Entropy indices and the Gini coefficient are broadly similar around the middle of the distribution (relative income = 1), and tend to place greater weight on mobility at the tails of the distribution. In addition, the overall U-shape for the weighting function is distinctly shallower for the Gini than for the other indices (as Shorrocks argued). Given the ready availability of longitudinal data on incomes nowadays (see the next two Sections), it is straightforward for researchers to examine sensitivity empirically.

Refinements to the Shorrocks approach have gone in two main directions. The first addresses the assumption that individuals are able to smooth incomes across time: see the discussion in Section 2. This aspect is relaxed by Maasoumi and Zandvakili (1990) and Zandvakili (1992), building on Maasoumi and Zandvakili (1986). See also the survey by Maasoumi (1998). The basic idea is allow for different degrees of substitutability between incomes in different periods. Thus, rather than defining longer-term income for each individual as the simple arithmetic average, it is defined as a generalized mean for which the choice of a parameter tunes the degree of substitutability. Observe that a common parameter is used for each individual, and yet one would expect the ability to smooth income over time to vary with e.g. income level. Incorporating such heterogeneity into an index would be a rather complicated exercise and has not been done, as far as we are aware. As it is, researchers wishing to implement the Maasooumi-Zandavakili variant on R need to choose a substitutability parameter as well as inequality index. Of course, estimates can easily be derived for a number of combinations but the volume of results produced is probably one reason the approach is not commonly used. Also, the empirical illustrations provided by Maasoumi and Zandvakili (1990) and Zandvakili (1992) tend to suggest that the more general index tended to provided qualitatively similar results to that of the Shorrocks approach.[29]

The second refinement to the Shorrocks approach is to reconsider the reference point against which longer-term inequality values are compared. The main argument of Fields (2010, 410) is that '[w]hat we as empirical researchers would want to know in a given context is the extent is the extent to which the mobility

[29] Maasoumi and Zandvakili (1990) and Zandvakili (1992) were also the first to provide decompositions of total mobility calculated using Shorrocks-Maasoumi-Zandvakili indices into components representing within-group mobility and between-group mobility. They did not provide formulae for the decomposition, however. For these, see Buchinsky and Hunt (1999, 354).

that takes place works to *equalize* longer-term incomes relative to base, *disequalizes* longer-term incomes relative to base, or has no effect' (emphasis in original). This leads to Fields' proposal that the denominator in the expression for R be changed from the weighted average of the per-period inequalities to the inequality in first period income. Chakravarty et al. (1985) emphasize rather different aspects in the derivation of their mobility index: they are concerned with 'ethical' indices of relative income mobility which are derived from social welfare functions and measure changes in welfare. Mobility is the percentage change in social welfare (measured by the equally-distributed equivalent income, defined in the Atkinson (1970) sense) of the actual distribution of longitudinally-averaged incomes compared to what social welfare would have been in the completely immobile benchmark distribution – taken to be the observed period-1 distribution.

If the same welfare function is used to evaluate both distributions, and that SWF is homothetic, then the mobility measure 'has a natural interpretation; it is the percentage change in equality of the aggregate distribution compared with the first-period benchmark' (Chakravarty et al., 1985, 6). Although the authors go on to state that appear to no convincing ethical argument for applying the same welfare function to both distributions, all empirical applications that we are aware of have applied the same welfare function. The class of mobility indices for the two-period case is then defined as (Chakravarty et al., 1985, 8):

$$C = \frac{1 - I[Y(T)]}{1 - I[Y^1]} - 1. \tag{10}$$

where I is a relative inequality index equal to one minus an index of relative equality (as is the case with the Atkinson (1970) class of inequality indices). It turns out that the Fields (2010) mobility index, $1 - [I[Y(T)/I[Y^1]$, equals κC where $\kappa = (1 - I[Y(T)])/I[Y^1]$, and so the measures are closely related (assuming the same inequality index is applied in each case). But it is possible for them to differ about whether mobility has increased or not: the value of κ matters. In short, ethical index C always evaluates mobility as welfare-increasing (but of different degrees), whereas the more descriptive Fields (2010) index allows mobility to be positive or negative. A more fundamental issue, common to both indices, is whether one agrees with the proposal to accord special normative status to period-1 incomes relative to incomes in other periods – which is an issue that has arisen with other mobility measures as well.

The concept of comparing short- and longer-term incomes has been used to examine poverty persistence in particular as well as income mobility in general. The basic building block is again 'longer-term income', a measure of longitudinally-averaged income for each individual, and people are defined as 'chronically' poor if their longer-term income is less than the poverty line. Chronic poverty in aggregate is the poverty in the population calculated using a poverty index that is additively decomposable over people and time (e.g. a member of the Foster et al. (1984) class). Transitory poverty is Total Poverty (poverty calculated over individuals and separate time periods) minus Chronic Poverty. The main papers to date in this tradition are Rodgers and Rodgers (1993, 2009), Chadhuri and Ravallion (1994), and Jalan and Ravallion (1998). See also the development by Duclos et al. (2010) which takes a more explicitly welfarist approach. As with the Shorrocks mobility measures, there is an important issue concerning how longer-term incomes are calculated and (related) the assumptions made about abilities to income smooth. See e.g. the discussion by Rodgers and Rodgers (1993, 34–35).

The final group of more specialist mobility indices we discuss are those that summarize notions of income risk. These can be classified in two main ways.

On the one hand, there are measures of the transitory variance of (log) income, calculated using either model-based or non-parametric approaches and generally requiring income data for multiple periods. On the other hand, there are measures of income flux, income movement and volatility, generally defined over incomes in two periods only. We consider the approaches in turn and discuss the relationships between them and the measures of longer-term inequality reduction.

To fix ideas,[30] suppose that the dynamics of income for each individual can be described using the canonical random effects model

$$\log y_{it} = u_i + v_{it} \tag{11}$$

where y_{it} now refers to the income for person i in year t. It consists of a fixed 'permanent' random individual-specific component, u_i, with mean zero and constant variance σ_u^2 (common to all individuals), and a year-specific idiosyncratic random component with mean zero and variance σ_v^2 (common to all individuals) that is uncorrelated with u_i. Thus total inequality as measured by variance of log incomes is equal to the sum of the variance of 'permanent' individual differences plus the variance of 'transitory' shocks:

$$\sigma_t^2 = \sigma_u^2 + \sigma_v^2. \tag{12}$$

Assuming that permanent differences are relatively fixed over time, changes over time in income inequality (σ_t^2) arise mostly through changes in the variance of the transitory component. The interpretation of this latter component as idiosyncratic unpredictable income change leads to the association of changes in its variance with changes in income risk.

[30] The exposition in the next few paragraphs draws heavily on Jenkins (2011a, chapter 6).

This canonical model is patently unrealistic in several respects and three types of extension have been incorporated.[31] The first additional factor allows the relative importance for overall inequality of the permanent and transitory components to change with calendar time. For example, if there is an increase in the demand for skilled labour, and permanent component of income represents relatively fixed personal characteristics related to skills (for example human capitals of various kinds), then greater inequality resulting from widening differences over time in returns to skilled versus unskilled labour can be represented as the growing importance of the permanent component. In contrast, a secular trend towards greater labour market flexibility can be represented as a growth in the importance of transitory variations. The second additional feature is persistence in transitory shocks. The factors leading to a temporary fall (or rise) in income in one year are likely to have effects that last longer than a year: a transitory shock persists but with diminishing impact and eventually dies out. An example might be an accidental injury leading to a reduction in work hours that diminishes over time. This is usually characterized using an autoregressive moving average process for v_{it}.

The third modification to the canonical model is to allow the fixed individual component to change over time. Two main approaches have been followed, originally distinct but now commonly combined. One is to allow u_i to vary over time via a 'random walk': this year's value is equal to last year's value plus or minus a random element. The second approach allows for individual-specific rates of

[31] For surveys of model specification and estimation methods, see inter alia, Baker and Solon (2003), Guvenen (2009), Meghir and Pistaferri (2011), Haider (2001), Shin and Solon (2011), Moffitt and Gottschalk (2012), and the references cited in these sources and in Section 4 below. Note that these variance components models have usually been applied to data on men's earnings and only rarely to household income. See Jenkins (2011a, chapter 6) for further discussion.

growth in income (he 'random growth' model). The expression for the permanent component is modified so that it also varies linearly with time but with heterogeneity in this slope. Both a random walk and random growth lead to a fanning out of the income distribution over time, other things being equal. Rankings are preserved: those at the bottom stay at the bottom but fall further behind those at the top, who stay at the top. It is increases in the transitory variance that increase mobility in the sense of reranking.

The estimation of transitory variances (mobility) and permanent variances using these models is common, but has also been criticized on the grounds that estimates are sensitive to the particular model specification employed, and there are potential identification issues with the relatively short household panels used to estimate the models: see e.g. Shin and Solon (2011), Guvenen (2009), and Doris et al. (2013). This has led to simpler non-parametric methods also being regularly used.

The most common non-parametric method for deriving estimates of variance components is the window-averaging method first employed by Gottschalk and Moffitt (1994), also known as their 'BPEA' method (the acronym refers to the journal in which their work was published). The BPEA method works by first calculating the longitudinal average of each person's log income over a time window of fixed width, say T years. This provides an estimate of the person's 'permanent' income for that period, and is directly analogous to the longer-term income concept used to derive R except that it refers to averaging of log incomes. (If equation 11 describes the income generation process, the longitudinal average is an estimate of u_i.) The transitory incomes for each individual within the window are derived as a difference between this permanent income and observed log income, from which can be calculated the individual-specific transitory variance. The overall sample transitory variance is the average of these variances. The sample permanent variance for each window is calculated from the differences between each person's permanent income and the sample grand mean of these, with an adjustment to account for the fact that the mean contains a proportion of the transitory component that has not been fully averaged to zero over the T-year window. See Gottschalk and Moffitt (2009, 7) for full details of the formulae, and Kopczuk et al. (2010, 98) for a small variation on the same theme. The BPEA method is known to provided biased estimates of the transitory variance and its trend if the permanent component's contribution changes over time (see e.g. Shin and Solon, 2011). Using shorter-width windows for the calculations (smaller T) reduces the potential impact of this problem but at the cost of reducing the statistical reliability of the estimate of each person's permanent income.

It is inevitable that measures derived using methods like the BPEA one will reflect the variability from permanent shocks and not only from transitory shocks. Shin and Solon (2011, 9) argue that this is a virtue of such measures: 'The recent interest in volatility trends stems in large part from a concern about whether earnings risk has increased. Because permanent shocks, such as those experienced by many displaced workers, are even more consequential than transitory ones, it makes good sense to include them in the measurement of earnings volatility. Their own calculations use instead a measure of 'volatility' that will be discussed shortly.

Both of the two main methods for estimating transitory variances have potential weaknesses, and there are virtues in using both as well as other measures (such as of volatility) as a sensitivity check. (This is increasingly done, as the next section shows.) Regardless of estimation method, we would point out a distinction between measures of mobility which are based on the transitory variance itself and measures that are based on the transitory (or permanent) variance expressed as a proportion of the total variance. Most discussion uses the former as the definition of mobility in the form of income risk.

Some authors also present estimates of the permanent variance expressed as a proportion of the permanent variance, and note that, if estimated using the BPEA method, there is a close relationship with the estimates of the Shorrocks measure of income rigidity R. See e.g. Burkhauser and Couch (2009) and also Chen and Couch (2013, 202) who state that they prove that ‘under one testable condition a measure of economic mobility formed by the ratio of permanent to total variance employing the methods of Gottschalk and Moffitt (1994) is equivalent to the Shorrocks R constructed with a Theil General[ized] Entropy Index’. It is clear that there must be some relationship, but we believe that it is not as close as stated by these authors, for the simple reason that the BPEA method calculation uses log incomes, and calculations of R invariably use incomes expressed in levels rather than logs. Evidence showing that a BPEA-estimated ratio of permanent to total variance and Theil-based estimate of R can move in opposite directions appears in Bayaz-Ozturk et al. (2013, Figure 2). For related discussion, see also Shorrocks (1981, Section 6) who considers the shape of the profile for $M(T)$ in the case in which incomes – not log incomes – follow the basic canonical random effects model (cf. equation 11) and inequality is calculated using half the squared coefficient of variation. He shows that were the model to hold, $M(T)$ would converge to its limiting value fairly rapidly. Slow convergence is evidence that the canonical model is inappropriate.

Income volatility in a given year t, V_t, is commonly measured by the standard deviation (sd)) of the distribution of individual changes in log income between one year and an earlier year:[32]

$$V_t = \mathrm{sd}[\log(y_{it+\tau}) - \log(y_{it})]. \tag{13}$$

Changes are typically measured over a one- or two-year horizon: $\tau = 1$ or 2. The Fields and Ok (1999b) index of individual income growth ($D1$ discussed earlier) is the mean of the distribution of log-income changes. Volatility is therefore a measure of dispersion of the same distribution using one specific index of inequality. There are further connections: if the Gottschalk-Moffitt BPEA method is used to calculate the transitory variance in the two-period case, the resulting estimate is equal to one-quarter of the variance of the change in earnings (i.e. V_T^2 with $T = 2$). See Moffitt and Gottschalk (2012, 218) who also point out that the relationship no longer holds if the data window is longer than two periods.

This brings us to the measures of income flux, most commonly associated with the names of Fields and Ok (1996, 1999b) who proposed a number of measures of non-directional income movement for the two-period case. Although such indices are rarely related to the measure of income risk discussed so far, their inventors had this application in mind: 'A measure of income movement ... identifies how unstable the incomes of individuals have been throughout the time period. Since income instability may cause economic insecurity, ..., measure of income movement are useful complements to the traditional measures of relative income mobility' (Fields and Ok, 1999b, 455). In their 1996 paper, Fields and Ok consider what

[32] See e.g. Shin and Solon (2011). Other variants use a different definition of proportional income change, most often the arc percentage change: see e.g. Dynan et al. (2012). This has the advantage of allowing for zero income values in the estimation of volatility.

they label absolute measures of income movement. First, they propose a number of axioms to describe measures for a fixed population of N individuals: linear homogeneity (equi-proportionate increases in all incomes, in base and final year, lead to the same proportionate increase in the measure); translation invariance; a normalization axiom; decomposability (total mobility for the N individuals is a symmetric function of the income changes for each individual); growth sensitivity (if two bivariate distributions are identical except that in one distribution an individual experiences more income movement than in the other distribution, total mobility differs in the two distributions); and, finally, the axiom of individualistic contribution (the contribution of each individual's mobility to total mobility does not depend on how other people's incomes change).

Fields and Ok (1996) prove that the measure satisfying these seven axioms is the sum over the N individuals of the absolute differences between period-1 and period-2 incomes, i.e. $|y_i - x_i|$, for each individual $i = 1, \ldots, N$. The final step is to consider versions of these measures that would enable comparisons across populations of different size. Specifically, their per capita measure of absolute measure of absolute income movement is:

$$D2 = \frac{1}{N}\sum_{i=1}^{N} |y_i - x_i|. \tag{14}$$

Their 'percentage' measure is the same as $D2$ except that denominator is total income in period-1 rather than population size, and has been less commonly-used perhaps because it is unclear that the base year should be used as the reference point (see our earlier discussion).

In their 1999 article, Fields and Ok take a similar similar set of axioms but also consider scale-invariant measures of movement as well as translation-invariant ones. This leads to the per-capita relative movement index: given by

$$D3 = \frac{1}{N}\sum_{i=1}^{N} |\log(y_i) - \log(x_i)|. \tag{15}$$

Both $D2$ and $D3$ are additively decomposable by population subgroup: total income movement can be expressed as the weighted sum of the movement within each subgroup, where the weights are the subgroup population shares. Fields and Ok (1996, 1999b) show that $D2$ and $D3$ also satisfy a different sort of decomposition: in each case, aggregate income movement can be expressed as the sum of a component representing income 'growth' for individuals and a residual component that can be interpreted as income 'transfers' between individuals. (Slightly different versions of the decomposition apply depending on whether the average of the first component is positive or negative.) It turns out in the case of $D3$ that the growth component of this decomposition is the directional measure of proportionate income growth ($D1$) discussed earlier.

To return to the remarks earlier about the links between measures of income flux and other measures of income risk, observe that the variance of log-income changes between two periods can be written as $\mathrm{E}(d_i)^2 - \mathrm{E}^2(d_i)$ where E is the expectation operator, and $d_i = \log(y_{it+\tau}) - \log(y_{it})$. That is, volatility-squared is equal to the average of the squared log-income changes, minus the square of the average log-income change. The first term is a measure of income flux in which the distance concept used to record income changes is Euclidean distance. Thus, there is a close relationship between orderings by this measure and a volatility measure volatility when average log-income changes are 'small'. This Euclidean distance measure is characterized axiomatically by D'Agostino and Dardanoni (2009b), who also compare their approach with that of Fields and Ok (1996, 1999b). In fact, a measure incorporating the Euclidean distance concept is also characterized axiomatically by another pioneering paper on the measurement of income movement, by Cowell (1985). (See also Cowell and Flachaire (2011).) The axiom set is rather different in Cowell's (1985) paper, however, and also leads to parametric classes of subgroup decomposable measures of 'distributional change'. These indices have rarely been used in empirical applications, however, perhaps because their properties (in particular the implications of choosing different parameters) are rather opaque in comparison with the overt transparency of measures like $D2$ and $D3$.

This completes our review of the many measures of income risk. A question that could be asked about all of them is whether they actually measure income 'risk' in a more fundamental sense, namely the ex ante uncertainty aspect drawn attention to by e.g. Gottschalk and Spolaore (2002). As Creedy et al. (2013, 236) remind us, this requires a model of expectations formation based on observed income dynamics. There are also additional complications for welfare evaluations such as the extent to which observed income changes reflect voluntary decisions of individual and families, and the extent to which these are insurable (and how these aspects differ across people). These complicated underpinnings are absent from the measures we have discussed. At the other extreme are more structural models such as proposed by Blundell et al. (2008) and Cunha et al. (2005). Our overall assessment is that the measures we have discussed are useful descriptive measures despite these flaws. Their relative simplicity facilitates transparency and interpretation as well as empirical implementation. But they should be interpreted cautiously.

Our final remarks concern the applicability of the mobility measures to intragenerational and intergenerational data. As we have noted, different mobility concepts may be more relevant in one context than another. For example, positional mobility concerns appear of particular relevance to discussions of intergenerational mobility, and income growth of particular relevance to discussions of intragenerational mobility. But structural intergenerational mobility is also of interest, and so too is the identification of intragenerational reranking along with income growth. By providing a unified treatment of mobility measures, we hope that some cross-context fertilization may be facilitated. In principle (and data permitting), all the measures we have discussed could be used in either context. In the two sections that follow, reviewing empirical evidence about intra- and intergenerational mobility, we will reveal which measures have been used to date.

4. Intragenerational mobility: evidence

This section assesses evidence about within-generational income mobility. It first considers definitional issues, the nature of the longitudinal data available and issues of empirical implementation, and then turns to the evidence itself. Our review of the topics is selective. We draw on and refer readers to Jenkins (2011a, Chapters 2 and 3) for a much more extensive discussion of data sources for within-generation mobility and related empirical issues, as well as extensive references to other literature. Our survey of evidence concentrates on findings emerging over the last two decades, and gives greatest attention to the USA, with examination of trends over time and cross-national comparisons between the USA and (Western) Germany, but studies for other countries are also considered. Our focus reflects the emphasis in research to date and this, in turn, is related to the availability of suitable data (as we explain). Also, in order to make the review manageable, the focus is on mobility of household income rather than of individual labour earnings (though selected earnings studies are referred to). We show how conclusions about trends over time and cross-national differences vary with the mobility concept chosen.

Issues of statistical inference are ignore here. On these, see e.g. Biewen (2002) and Chapter 7 of this volume by Cowell and Flachaire.

4.1. Data and issues of empirical implementation

Any study of income mobility faces three ‘W’ issues: mobility of What, among Whom, and When? Studies of trends over time or across countries add another issue, that of comparability. The choices that researchers can make under these headings are much constrained by the sources of longitudinal data that are available. But the data situation has improved substantially over the last two decades. (Contrast the situation described below with the discussion by Atkinson et al. (1992, Chapter3) which focuses on earnings.) Although many of the 'W' issues arise in any study of income distribution, looking at mobility adds some extras twists to those arising in cross-sectional analysis.

Mobility of 'What' refers to which income sources are included in the definition of 'income'. Definitions typically range from measures with only a single source (typically earnings from employment) to a broader measure such as household income which includes multiple sources. Many variations are possible: e.g. labour earnings may refer to employment earnings only, or earnings from all jobs that an individual has, and may also include self-employment earnings (thought often not). There are multiple definitions of income as well. The most common distinction in empirical work is between measures of pre-tax pre-transfer income, pre-tax post-transfer income and of post-tax post-transfer (also often labelled original or market or pre-government income; gross income; and net, disposable, or post-government income, respectively). Pre-government income typically includes labour earnings, income from savings and investments, and transfers received from non-government sources. Taxes usually refer to taxes on income (typically at national level, sometimes also including local taxes) and contributions levied for public pensions. 'Transfers' usually refer to cash benefits received from the state.[33]

Mobility among 'Whom' refers to the definition of the income-receiving unit. Clearly this is closely related to the issue of What. For example, it is individuals that receive labour earnings. Benefits are assessed and income taxes levied

[33] For a comprehensive discussion of the various definitions, and recommendations for measurement, see Expert Group on Household Income Statistics (The Canberra Group) (2001).

on families and households. Individuals not in paid work such as stay-at-home mothers, or children, often do not receive income in their own right, but benefit from income sharing with families and households. Putting things another way, note that analysis of earnings mobility is typically restricted to workers with earnings, excluding those without earnings many of whom are women, children or of retirement age. In contrast it is typically assumed that each individual receives the (equivalized) total income of family (or household) to which he or she belongs. Since total household income is rarely zero, all individuals regardless of age or labour market attachment, can in principle be included an analysis of income mobility. There is no universally correct definition of the income unit, and which should be used depends on the goals of the mobility analyst. E.g. in a study of labour market flexibility, a focus on individual earnings is appropriate (though there remain questions about whether women can and should be included in such analysis – much empirical analysis is of men only). On the other hand, if the interest in mobility is stimulated by a desire to describe and summarize important features of society as a whole, then there is a strong case for using more inclusive samples. As we show below, some empirical studies focus on individuals of working age (variously defined), others on all individuals, and this can complicate cross-study comparisons.

'When' mobility issues refer to two aspects related to time. The first is the the length of the period to which income refers to. For instance, is it the hour, week, month, or year? Economists often argue in favour of longer reference periods (e.g. a year) on the assumption that grounds that temporary variations and measurement error are smoothed out, thereby providing a more accurate measure of living standards. There is relatively little empirical evidence available about the veracity of this hypothesis because analysts rarely have income data for the same people over both a shorter- and longer-period. Böheim and Jenkins (2006) survey the literature and, from their analysis, argue that income mobility calculated using current (monthly) and annual income definitions are similar, and they provide a number of data-related reasons. Cantó et al.'s (2006) analysis is more comprehensive, based on comparisons from quarterly and annual income data for Spain, show that use of the longer assessment period leads to higher estimates of poverty prevalence, lower inequality, and less mobility.

A second 'When' issue relates specifically to mobility analysis in particular rather than income distribution analysis in general. For much mobility analysis, the data refer to a bivariate income distribution in which the marginal distributions refer to two years t and $t+\tau$, and empirical analysis of longer-term inequality reduction over requires a definition of how many years constitutes the longer-term. In both cases, how far apart the base- and final-years will affect the conclusions because the longer the interval, the greater the possibilities of mobility (as we illustrate below).[34] Choices about what interval to use have implications for the analysis that one can undertake too because data sets cover a time period of particular length (rarely more than 20 or 30 years), so researchers can only look at mobility trends if they use relatively short time windows for their measures. The constraint becomes acute with longitudinal data sets like EU-SILC (discussed below) in which the maximum time period is four years.

How researchers can address the three 'W' issues is much constrained by the data that they have available to them, and this raises issues of comparability over time and country. Longitudinal data sources suitable for within-generation income

[34]This issue is of course closely related to the issue of income reference period discussed above.

mobility analysis are of two main types.

First, there are household panel surveys in which nationally-representative samples of the private household population are interviewed about their incomes and many other domains of their lives in an initial year and then re-interviewed thereafter at regular intervals (usually a year). Second, there are administrative registers (e.g. tax files) in which income records for individuals are linked longitudinally. Household panel surveys typically utilize income definitions (i.e. resolve the 'What' and to 'Whom' issues) that are consistent with definitions accepted as being of good quality in large cross-sectional surveys. By contrast, administrative record data are typically designed for administration of the tax and benefit system, and the definitions used of income and the income-receiving unit, and the population that is represented, are determined by the needs of administration than by research. But register data also have advantages relative to surveys: their samples are very much larger, issues of respondent drop-out or measurement error do not arise in the same way (see the discussion below), and coverage of the very richest income groups is much better (they are typically not reached by surveys).

The clinching argument for empirical researchers in favour of household panel surveys over administrative registers is that the former became widely available for many countries, especially from the mid-1980s onwards, with cross-nationally harmonised versions of the data following few years later. Administrative registers with longitudinal income data have remained rare until recently in most countries, with the exception of Scandinavian countries which have a rather longer history of use.

The longest-running household panel is the US Panel Study of Income Dynamics (PSID) which began in 1968 and still continues, though it changed from annual interviewing to biennial interviewing after 1997. Panels started in the early 1980s in the Netherlands and Sweden, but the most well-known European panel is the German Socio-Economic Panel (SOEP) which started in the 1984 and is still running. Other country panels include the British Household Panel Survey (BHPS) which started in 1991 and finished in 2008. (The BHPS was recently replaced, after a break, by a new and very much larger panel (Understanding Society) which incorporates most of the original BHPS sample.) The Household, Income and Labour Dynamics in Australia (HILDA) survey began in 2001 and is on-going. There is also Survey of Labour and Income Dynamics (SLID) for Canada is a rotating panel operational between 1998 and 2011.

As shall be seen below, it is the household panels cited in the last paragraph that have provided most of the empirical evidence about income mobility over the last two to three decades, both in their native format (often to examine trends over time within a country) or in a harmonized form (to undertake cross-national comparisons). The production of cross-nationally comparable household panel data with harmonized labour earnings and household income variables has been one of the major successes in social research infrastructure creation over the last few decades.

The Cross-National Equivalent File (CNEF) began in 1991 with harmonization of data from the US PSID and German SOEP, incorporated the BHPS and SLID in 1999, and HILDA in 2007. (Data for more countries have been added subsequently.) It should be stressed that the project does more than simply harmonize variables; it adds value. One important example of this is the derivation of comparable post-tax post-transfer household income variables. The original PSID family income variable refers only to pre-tax post-transfer income and the government transfers do not include income deriving from non-refundable tax credits (the EITC) or near-cash benefit income in form of Food Stamps (now called SNAP). The CNEF uses the NBER TAXSIM model to simulate taxes. Similarly, involvement in the CNEF project was a stimulus for the SOEP to develop and maintain a similar model in-house. (Other CNEF members also use such models.) For a more detailed discussion of the CNEF, see Frick et al. (2007).[35]

Another important initiative providing cross-nationally comparable panel data on incomes was the former European Community Household Panel (ECHP), though this has been used less often for mobility analysis than the CNEF and its constituent panels. The ECHP relied on 'input' harmonization by contrast to the CNEF's 'output' harmonization. That is, household panel surveys with the same design and questionnaires including the same variables were fielded in a number of countries, so that harmonization was built-in from the start. Data from a maximum of 8 annual interview rounds are available, covering the period 1994–2001. Twelve EU member states participated in the ECHP initially, with two more joined shortly thereafter. The ECHP never realized its full potential because, for many years, researcher access to the data was constrained and financially costly. This is by contrast with the CNEF which, from the start, has had a much more open data access policy and been more research(er)-driven.[36]

The ECHP was replaced – after a gap – by the European Statistics on Income

[35]For documentation and user access information, see `http://cnef.ehe.osu.edu/` or `http://www.human.cornell.edu/pam/research/centers-programs/german-panel/cnef.cfm`.

[36]The Survey of Health, Ageing, and Retirement in Europe (SHARE) is another multi-country longitudinal study proving input-harmonized income data, and also research-driven. Its focus, however, is on older individuals, and so it cannot be used to study income mobility in the wider population.

and Living Conditions (EU-SILC), from 2005. EU-SILC is explicitly designed to deliver data on a set of social indicators that include income distribution statistics. This is output harmonization again, though the target variables are pre-defined by the needs of EU policy-making rather than by researchers. Some member states use administrative registers to produce the data, others use panel surveys, an aspect which has led to questions about data comparability (see below). The longitudinal data in the publicly-released EU-SILC data sets track individuals for a maximum of four years (by design), and so the scope for longer-run mobility analysis is ruled out. The great advantage of the EU-SILC longitudinal data is that, when mature, they will cover all EU member states. Understandably the EU-SILC has not been much used for income mobility to date, and this is reflected in our review of evidence below.

This review of data sources suggests that there has been a substantial increase over the last three decades in the volume of high quality longitudinal data available to researchers. But there remain a number of important issues of empirical implementation that need to be kept in mind when assessing the value of a particular mobility study. So, before turning to discuss empirical evidence, we briefly review these issues.

There are generic issues associated with longitudinal surveys, notably the potential problem of survey attrition. Over time, some respondents to a panel survey drop out from the data, either no longer with to participate or unable to be tracked down for interview. Attrition has two potentially adverse effects. The first is reduction of sample size, with consequences for the precision of estimates. The second potential effect, more commonly-discussed, is on the representativeness of the sample. Particular groups such as young people tend to be more likely to drop out, in which case estimates may be biased.[37] Note that differential attrition may be related to both observed and unobserved characteristics of individuals and families. For the former case, data producers routinely produce and release sets of weights that can be used to maintain the representativeness of estimates, and virtually all the studies cited in our evidence review use these weights. By definition, it is harder to assess the effects on estimates of differential dropout related to unobserved characteristics; it requires modelling of the attrition process. For an extensive discussion of attrition in US household panel surveys, see Fitzgerald et al. (1998) and other papers in the Summer 1998 issue of the *Journal of Human Resources*.

The likely impact of attrition is associated with the type of mobility analysis undertaken. Attrition between successive waves of a household panel is typically relatively low (around 5 per cent) with the exception that drop out rates are noticeably greater between the initial and second waves. Estimates of mobility over short periods (one or two years, say) are likely to less affected by attrition, than estimates based on long runs of data.

Respondents may remain in a longitudinal survey, but not provide complete responses to particular questions, either because they don't understand the question, or don't know or don't wish to provide the answer. This is the issue of 'item' non-response leading to missing data for some respondents and, as with attrition, may be associated with both observed and unobserved respondent characteristics.

[37]Representativeness typically refers to the ability of the sample to represent the private household population in the first wave of the panel. If a country experiences significant migration or immigration, a panel inevitably becomes unrepresentative of the population in later years. Sample refreshment has been employed to counter this problem but, if mobility estimates are required for time points spanning the old and new population structures, refreshment cannot improvement representativeness.

Item non-response is particularly prevalent for questions about income sources by comparison with items such as e.g. a respondent's age. In the public-use panel data sets used by mobility researchers, missing income values are typically replaced by an imputed value (together with a flag that enables identification of such observations) generated using procedures allocating similar values to respondents with similar sets of (observed) characteristics. Imputation is very useful for analysts but can potentially have effects on analysis because, by comparison with non-imputed data, extra 'noise' is added by the inevitable imperfection of the process.[38] These can have particular effects on mobility analysis, because some of the changes in a person's income over time may simply reflect the imputation process in the different years. But if one simply drops the imputed observations, there may be a critical loss of sample size and use of a potentially non-representative subsample. In most of the income mobility studies discussed later, analysts have routinely used imputed data on household income. By contrast, in studies of earnings volatility, it is more common practice to drop imputed observations. Researchers tend to find that this reduces observed volatility but the effects are relatively small. Again, the likely effects will depend on whether the particular mobility measure employed requires, say, two years relatively close together, or many years over a longer interval.

The problems raised by imputation are closely related to the more general issue of measurement error in earnings and income data. Even if survey participants respond to a question, their answer may be incorrect either because the respondent does not want to give the true answer or simply doesn't know what

[38] The imputation of households' tax payments when deriving measures of post-tax post-transfer income are another important example of useful imputation that may also add noise.

it is. Key questions are whether observed responses are systematically under- or over-reports of the (unobserved) true value or simply random, and how errors are correlated across successive years of data for the same respondent. Clearly, the answers to these questions may differ by income source. The largest body of research on measurement error has been about labour earnings, and used validation studies in which linked administrative record data are used to provide a picture of each worker's 'true' earnings. (See e.g. the survey by Bound et al. (2001).) Few studies have looked at the effects of measurement error on measures of earnings mobility.

The perhaps surprising finding of Gottschalk and Huynh (2010) is that estimates of men's earnings mobility, defined in terms of the Pearson correlation between log earnings in one year and the next, are much the same in the survey data and their administrative data set. The result arises because measurement errors are not important; rather, it is because they are 'non-classical' in nature, i.e. mean-reverting and correlated across years, and these various features happen to offset each other. See also Fields et al. (2003) who use a non-classical measurement error model similar to that of Gottschalk and Huynh (2010) to put bounds on estimates of income change. For the case considered, they argue that the effects of measurement error are 'relatively minor' (Fields et al., 2003, 90). Dragoset and Fields (2006) calculate a large portfolio of mobility measures from both survey and linked administrative record data on US men's earnings. They conclude that most of their qualitative results are the same in both data sources, and that the estimates from the administrative source were neither systematically above or below the corresponding survey estimates. Overall, this small body of research might be taken to imply that measurement error has relatively unimportant effects on measures of mobility in practice. We would caution against this interpretation, convenient as it is for empirical researchers; the situation is more that we know rather little at present. All the studies cited refer to earnings for US men, and results may differ for household income and in other countries. (The only similar study for household income that we are aware of is by Rendtel et al. (2004) who also report finding mean reversion and serial correlation.) There is also a more fundamental question of whether administrative record data can be assumed to provide error-free representations of the truth (Abowd and Stinson, 2013).[39]

A rather different sort of measurement error arises in the case of outlier observations, for example very high or very low observations. These may be genuine but may also represent errors of e.g. transcription leading to additional zeros being added. The problem is that, even if the number of observations with this kind of data is very small, they may have a big influence on the estimates that are derived. This lack of robustness is undesirable. See Cowell and Schluter (1999) for a discussion of this problem in the context of income mobility analysis. Empirical analysts' response to this issue is usually to simply drop a fraction (e.g. 1%) of the very richest and of the very poorest income values in each year. This procedure, known as 'trimming', or similar algorithms directed at removing potential outliers, has been applied in virtually every study cited in our discussion of empirical evidence.

[39] Aside from analysis based on validation studies, there have been a small number of model-based assessments of the impact of measurement error on estimates of poverty transition rates: see e.g. Breen and Moisio (2004) and references therein. Longitudinal data on observed transitions are combined with an assumption that the 'true' transition probabilities are stable over time, so that difference between them is attributed to measurement error. In technical terms, the statistical approach involves fitting latent class models with a Markov structure. For further discussion, see Jenkins (2011a, 53–55).

A final empirical issue is whether income changes over time represent genuine mobility or, instead, systematic changes associated with lifecycle patterns such earnings following an inverse-U shape with age. Many income mobility studies do not adjust for this factor; they look at observed incomes. Some other studies, mostly of earnings mobility, have regressed observed earnings against variables such as age, and then the mobility analysis is of the earnings residuals: see below.

4.2. Intragenerational income mobility in the USA: levels and trends

We take as our initial reference point the estimates of income mobility for the USA provided by Hungerford (2011), as he uses good quality comparable data from PSID (as released via the CNEF) and provides a range of mobility summaries. (Transition matrices from the study were presented in Table 1 earlier.) Hungerford compares mobility over two 10-year intervals, 1979–1988 ('1980s') and 1989–1998 ('1990s'). The measure of income is annual disposable (post-tax post-transfer) family income adjusted for differences across families in household size and composition using the equivalence scale proposed by Citro and Michael (1995). His samples include all individuals within households. The 1980s sample includes the PSID's SEO low-income sample; the 1990s sample does not (about half the SEO sample was dropped in 1997). All estimates are derived using the PSID's weights. We noted earlier that, in both periods, there appeared to substantial short-distance mobility over a ten-year period, but long-distance moves were relatively rare. Moreover, the chances of upward mobility from the bottom and downward mobility from the top appeared symmetric. We now compare mobility in the two decades in greater detail, in particular considering whether mobility increased or decreased according to various mobility concepts and measures.

To assess changes in positional mobility, a natural first approach is to apply the

Table 2 Differences in cumulative density: USA, 1979–1988 versus 1989–1998

	Destination group									
Origin group	1	2	3	4	5	6	7	8	9	10
1	0.2	–0.1	–0.2	0.0	0.3	0.2	0.1	0.1	–0.1	0.0
2	0.0	0.0	0.4	0.6	0.5	0.2	0.2	0.1	0.0	0.0
3	–0.2	–0.5	–0.2	0.0	0.0	–0.5	–0.1	–0.1	0.0	0.0
4	–0.2	–0.7	–0.6	–0.6	–0.7	–0.7	–0.2	–0.3	0.1	0.0
5	0.0	–0.3	–0.3	–0.5	–0.7	–0.5	0.0	–0.1	0.4	0.0
6	0.1	–0.1	–0.1	–0.4	–1.1	–1.3	–0.9	–0.5	0.4	0.0
7	0.1	0.2	0.0	–0.3	–0.8	–0.9	–0.8	–0.3	0.3	0.0
8	0.1	0.2	–0.2	–0.2	–0.3	–0.7	–1.1	–0.7	–0.3	0.0
9	0.0	–0.1	–0.3	–0.2	–0.4	–0.4	–0.7	–0.6	–0.6	0.0
10	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0

Note: The estimates are in percent, rounded to one decimal place, and show in each cell the cumulative discrete density for the 1980s minus the corresponding cumulative discrete density for the 1990s.
Source: Authors' calculations from (Hungerford, 2011, Tables 2 and 3), based on PSID data.

dominance check of Atkinson and Bourguignon (1982) based on the differences in the discrete cumulative densities implied by the decile transition matrices in Table 1. See Table 2 for the estimated differences. Observe that first-order dominance does not hold: there is a mixture of positive and negative differences.[40] There is an interesting pattern, however. Most of the positive differences (greater cumulative density in the 1980s) are found in cells corresponding to movements out of or into the poorest fifth of the distribution. Put another way, there is greater movement in the 1980s than the 1980s into and out of the richest 80 per cent, broadly speaking.

[40]The density estimates and conclusions drawn from them need to be interpreted cautiously, not least because they are susceptible to measurement error and sampling variability. If the estimates in Table 2 are rounded to 2 d.p. to reflect this (rather than 3 d.p. as reported), then many matrix entries become zero, and there is now dominance: positional mobility is greater in the 1980s than the 1990s.

Saying conclusively that mobility increased or decreased in the USA between the 1980s and 1990s, and by how much, requires additional assumptions about the weighting of mobility in different parts of the distribution. Also, the answers depend on the mobility concept. These points are illustrated by the mobility index estimates reported by Hungerford (2011) and summarized in Table 3. The first three rows of the table provide estimates of positional mobility (reranking), and all the indices show a small decline between the 1980s and 1990s. In contrast, the Shorrocks and Fields equalization indices record a increase, and so too do the two measures of income flux shown in the bottom two rows. For the last four indices, the estimated increase is small, with the exception of the Fields equalization measure, for which the large change reflects the increase in (cross-sectional) income inequality over the period. The general lesson is that conclusions about whether mobility increased or decreased between the 1980s and 1990s depends on the mobility index employed.

Mobility as individual income growth is also summarized by Figure 9, which shows the median real income growth for each base-year decile group, by period. (This is a grouped data version of Figure 7 discussed in the previous section.) Clearly income growth is pro-poor in the USA (consistent with regression to the mean), but the patterns differ between the 1980s and 1990s. Income growth was greater in the 1990s than the 1980s for the richest eight base-year decile groups, but no different for the two poorest base-year decile groups.

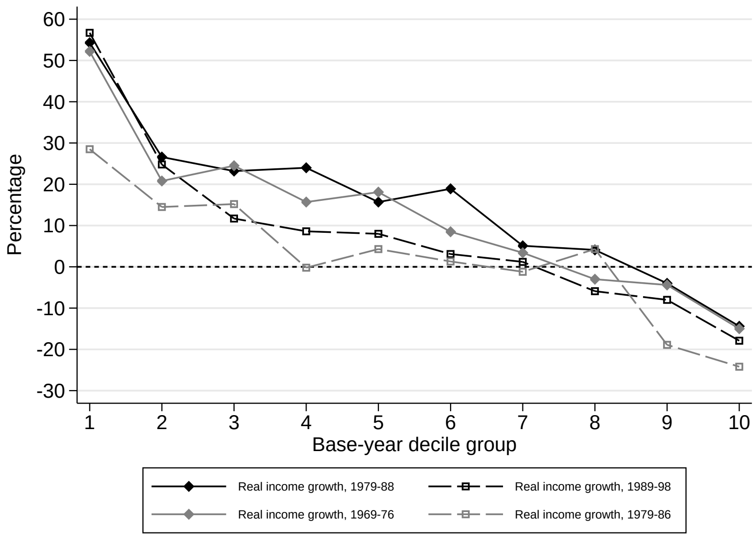
The extent to which US mobility comparisons can be extended to periods before the 1980s and after the 1990s is restricted by data availability (e.g. the PSID only started in 1968), because different studies use different income variables and estimation samples and often do not report the same mobility statistics.

Table 3 Selected mobility indices (%): USA, 1979–1988 versus 1989–1998

Index	1979–88	1989–98
Decile mobility	79.1	77.0
Normalized trace	87.9	85.6
Gini mobility	36.2	34.4
Equalization (Shorrocks, Gini-based)	10.9	11.1
Equalization (Fields, Gini-based)	2.1	8.2
Average of absolute income changes ($D1$)	11,368	13,878
Average of absolute income share changes	0.421	0.459

Note: The estimates are in percent, rounded to one decimal place, apart from those in the last two rows (in constant-price dollars). Decile mobility is the proportion of persons changing at least one decile group. The normalized trace is the Shorrocks (1978b) index calculated from the decile transition matrix. The Gini mobility index is the index of Yitzhaki and Wodon (2005). The Equalization indices are those of Shorrocks (1978a) and Fields (2010). On the average of absolute income and income share changes, see Fields and Ok (1996) and Fields (2010). See text for more details.
Source: Authors' calculations from Hungerford (2011, Tables 4 and 8, and p. 97), based on PSID data.

Figure 9 Median real income growth, by base-year decile group: USA, by period



Note: The estimates show median income growth for each base-year decile group over the relevant period.
Source: Hungerford (1993, Table 9) and Hungerford (2011, Tables 5 and 6).

For example, Hungerford (1993) provides much information about US income mobility in the 1970s and 1980s, but the estimates are not fully comparable with those in Hungerford (2011) because the earlier study uses a different income definition (pre-tax post-transfer income rather than equivalized post-tax post-transfer income) and the interval between base- and final-years differs (8 years rather than 10; e.g. 1979–86 rather than 1979–88). The relevance of definitional differences is illustrated by the estimates for the '1980s' from the two studies of the proportions of individuals remaining in the poorest tenth and remaining in the richest tenth: 44.3 per cent and 40.0 according to Hungerford (2011), but 49.0 per cent and 42.1 per cent according to Hungerford (1993, Tables 1 and 2). Look also at the different estimates of real income growth rates for the 1980s for the two periods in Figure 9. Using the Hungerford (1993) definitions, the overall growth rate for the 1980s is smaller (which is unsurprising since aggregate income growth was positive throughout the mid-1980s (Hungerford, 2011, Table 1)) but observe that the estimates of pro-poorness in income growth also differ (the income growth curves from the two studies do not have the same slope).

One can compare mobility in the 1970s with the 1980s, however. If we examine differences in cumulative densities using Hungerford's (1993) estimates, again there is no clear cut mobility ordering (authors' calculations) and there is a broadly similar pattern of differences to that described earlier. Hungerford (1993) does not report summary indices to compare with those in Table 3 but two statistics based on the transition matrices (Cramér's V) and the contingency coefficient 'are the same ... suggesting that the degree of association between a person's decile rank in one year and another was the same in the 1970s and 1980s' (Hungerford, 1993, 407). Fields and Ok (1999a) used exactly the same data as Hungerford (1993) and report that their measure of income flux, the average of the absolute changes in log income increased from 0.498 in the 1970s to 0.528 in the 1980s.[41] So, again, changing the mobility concept leads to a different conclusion about trends.

Hungerford's (1993) study is also useful because it analyzes whether the estimated mobility patterns are robust to adjustment for transitory income variation. Specifically, Hungerford calculates each individual's five-year longitudinally-averaged income (centred on the year in question) and uses these 'permanent' incomes instead of the single-year incomes to define base-year and final-year income positions. Interestingly, the patterns of mobility revealed are remarkably similar, though with perhaps less movement at the top and bottom of the distribution.[42] For example, according the annual income calculations for 1979–86, 12.9 per cent of the poorest fifth remain in that group and 11.0 per cent of the richest fifth remain in that group. According to the permanent income calculations, the corresponding estimates are 11.5 per cent and 9.6 per cent (authors' calculations from Hungerford (1993, Tables 2 and 4)).

To examine trends in US income mobility further, we turn to Bradbury (2011). She provides estimates using consistent definitions for the period 1969–2006, and for a large portfolio of mobility indices. Her estimates are not fully comparable with Hungerford's, however. Although she and Hungerford (2011) both use post-tax post-transfer real family income measures from the CNEF version of the PSID, they use different samples. Bradbury focuses on adults who are a family

[41] Fields and Ok's (1999a) decompositions reveal that the increase in income movement is entirely accounted for by persons with education to high school level or above, and by young adults rather than prime-age adults.

[42] One potential non-comparability is that the estimation samples differ slightly: the permanent income estimates are based on balanced samples with valid data for all five years within the relevant period.

head or spouse rather than all individuals within families, both head and spouse (if present) are required to be of working age (16–62 years), and the time interval spans 11 years rather than 10. She uses the square-root-of-household-size equivalence scale rather than the Citro and Michael (1995) one.

Trends in three general indices of positional mobility are displayed in Figure 10: the fraction of individuals changing decile group ('decile mobility'), one minus Spearman's rank correlation, and Yitzhaki and Wodon's (2005) Gini mobility index. All three indices are broadly constant over the 1970s, and decline over the 1980s (11-year intervals starting at the end of the 1970s), with the rate of decline perhaps slowing from the late 1980s onwards. The fall in mobility over the 1980s is consistent with Hungerford's estimates of trends based on only two intervals during this period, but is rather larger in magnitude. The Gini mobility index fell by about a sixth between the intervals starting in 1979 and 1989 (but only about 5 per cent according to Hungerford (2011)). One minus the rank correlation fell by about one-fifth over the same period, and so the decline in positional mobility is relatively large. It is unclear what lies behind the secular decline in mobility, but we note that it was at the end of the 1970s that US family income inequality also began to increase (Burkhauser et al., 2011), suggesting that inequality and positional mobility share some common drivers. There is no very obvious association between series' turning points and the business cycle (there were recessions at the beginning of the 1970s and 1980s).

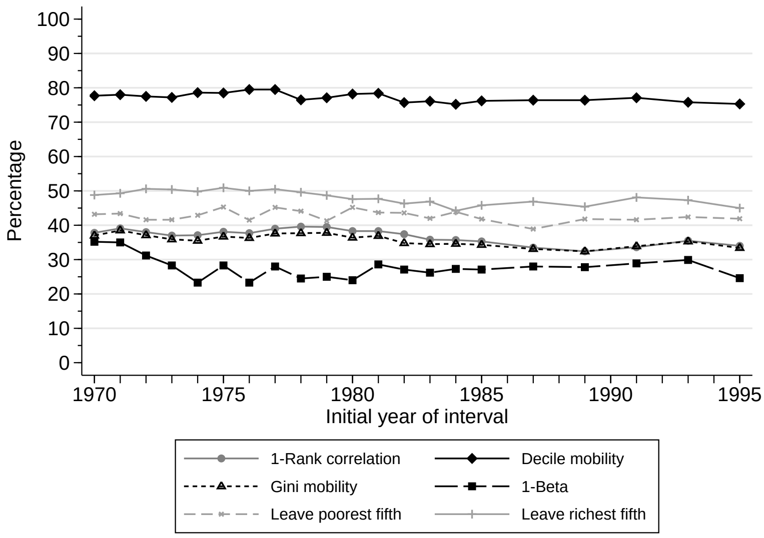
The conclusions about trends cited so far refer income changes over an interval of 10 or 11 years, and it is of interest to know how results change if rather different interval lengths are used. The research of Gittleman and Joyce (1999) suggests some sensitivity. Using PSID data for 1967–91 and, like Bradbury (2011), focusing on working-age adults and employing a broadly similar income definition[43], they calculate Immobility Ratios, defined as the percentage of individuals remaining in the same fifth, for intervals of one year, five years, and ten years. Gittleman and Joyce (1999, Table 1, Figure 2) show that the level of positional mobility increases (the immobility ratio falls) as the interval width is widened. But conclusions about mobility trends are also affected. For the ten-year interval case, there is a small downward trend during the 1980s consistent with Bradbury's (2011) estimates. However, five-year immobility ratios exhibit no similar trend, and one-year immobility ratios generally decline from the end of the 1960s until the end of the 1970s and increase in the following decade (though the changes are not large in absolute magnitude).

To provide a comparison with another commonly-used mobility index, we also show trends in one minus Beta. It follows a different trend, which is perhaps unsurprising given that it is not a purely positional measure (see Section 3). Compared to the trends shown by the three positional indices, the decline during the 1970s is earlier and sharper, and there is no decline during the 1980s.

The final two measures shown in Figure 10 are two 'corner probabilities' from a quintile transition matrix (cf. Section 3), specifically the proportion of individuals in the poorest fifth in the base-year who are in a different fifth in the final year and, analogously, the proportion leaving the richest fifth over the relevant interval. These statistics pick up on particular aspects of positional mobility. Interestingly, it appears that the trend in the percentage leaving the richest fifth tracks the trend in overall positional mobility better than does the trend in the proportion leaving the poorest fifth. The estimates also bear on our earlier comments that the US decile

[43] But see below for more about differences.

Figure 10 Indices of positional income mobility: USA, 1970–1995



Note: The estimates refer to 11-year intervals, with incomes in base- and final-year averaged over two years. For example, the estimates labelled as 1970 refer to incomes longitudinally-averaged over 1969 and 1970 (base year) and 1979 and 1980 (final year). See text for index definitions.
Source: Bradbury (2011, Tables 2 and 3).

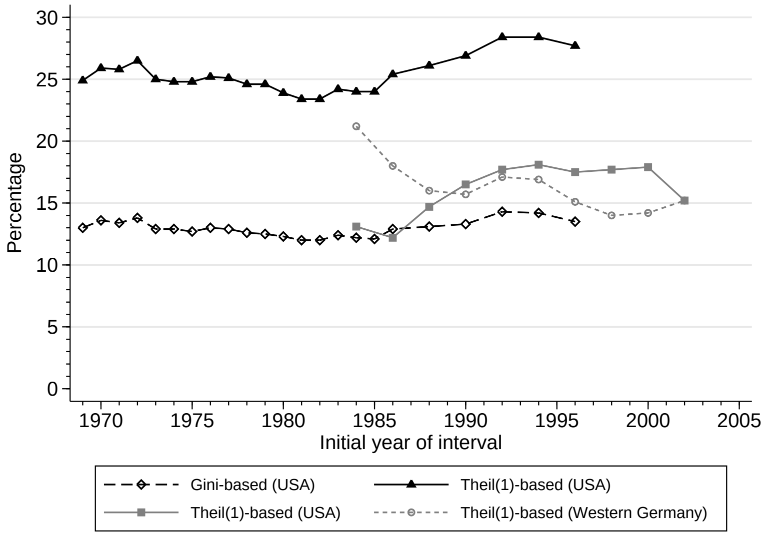
transition matrices for the 1980s and 1990s suggest that there is a broad symmetry to upward and downward mobility. We now see that asymmetry is more apparent if mobility is summarized using quintile rather than decile groups. In particular, it appears from Bradbury's (2011) estimates that the chances of downward movement from the top (richest fifth) are typically several percentage points greater than the chances of upward mobility from the bottom.

This asymmetry finding also may be contingent on the particular samples and other definitions used. For example, Bradbury and Katz (2002, Annex A) report quintile transition matrices for 1969–79, 1979–89, and 1988–98 using similar PSID samples to Bradbury (2011) except that 'working age' now refers to a wider age range (head and spouse (if present) less than 66 years), and family income is pre-tax post-government family income, equivalized using the PSID scale. The two probabilities are approximately equal in each matrix (50 per cent in the first two periods, 47 per cent in the last one). In contrast, Gittleman and Joyce (1999, Table 5) report quintile transition matrices for 1967–79 and 1979–91 using a similar income definition (but equivalized using the US poverty line) and 'working age' refers to head and spouse between 25 and 65 years. According to this study, the chances of leaving the poorest fifth are distinctly smaller than the chances of leaving the richest fifth (around 50 per cent compared to around 60 per cent).

Trends in mobility defined as equalization of longer-term incomes are summarized by Figure 11 using Shorrocks's (1978b) measure $M = 1 - R$. The long series (shown in black) are derived from Bradbury (2011); we discuss the series in gray shortly. Although mobility levels differ substantially depending on which inequality index is used – there is much greater mobility according to the Theil index compared to the Gini – the patterns of change over time are the same according to the two series. There was a decline in mobility between the early 1970s and the mid-1980s, followed by a rise over the following decade, with levelling off around the mid-1990s. Although the changes are small in absolute terms, they are relatively large in proportionate terms. For example, between the mid-1980s and mid-1990s, the Theil-based measure increased by some 15 per cent, and the Gini-based measure by almost 13 per cent. The results are consistent with Hungerford's (2011) finding of only a small increase in a Gini-based measure between the 1980s and 1990s, but Figure 11 shows that this is partly a consequence of the timing of measurement; Hungerford's two intervals lie on either side of the bottom of a U-shaped series. Observe also that the turning points in the these two series differ from those for the positional measures shown in Figure 10, suggesting that the different aspects of mobility have different underlying causes. In addition, mobility according to the Shorrocks measure is much the same (Gini-based index) or greater (Theil-based index) in the mid-1990s than in the early 1970s, whereas mobility is lower according to the positional mobility indices shown in Figure 10.

The research of Bayaz-Ozturk et al. (2013) allows to consider what happened to mobility as equalization after the mid-1990s. Although they also use a Theil-based measure, similar income measures, and the same data source, their series are not directly comparable with Bradbury's (2011): Bayaz-Ozturk et al. (2013) include all individuals in families in their analysis samples (not only working-age adults) and use a five-year rather than eleven-year interval. As a consequence, mobility levels are estimated to be substantially lower in all years (compare the gray line for the USA with the black one). Reassuringly, however, the series show broadly similar trends (and turning points) over the period for which they overlap. Bayaz-Ozturk et al.'s (2013) estimates indicate that mobility changed little in the

Figure 11 Mobility as longer-term income inequality reduction: USA, 1970–1995



Note: The estimates refer to the Shorrocks equalization measure, $M = 1 - R$, calculated using the Gini and Theil inequality indices. The Bradbury (2011) calculations are based on eleven-year intervals with longer-term average incomes calculated using every second year's income in order to handle the PSID's change to alternate-year interviewing in the late-1990s. The Bayaz-Ozturk et al. (2013) calculations use five-year intervals, with interval base-years two years apart.
Sources: Bradbury (2011, Table 4) for the series shown in black and Bayaz-Ozturk et al. (2013, Table A1) for the series shown in gray. Both use PSID (CNEF) data.

second half of the 1990s, with a suggestion that it fell again in the 2002–06 period.

All estimates of trends in household income mobility presented so far in this section are based on PSID data, and it is of interest to know whether the evidence from other data sources tells a similar story. The main reference point on this issue is Auten and Gee's (2009) work based on income data from tax administration records covering the two decades between 1987–2005. The data and definitions used are not fully comparable with those in the PSID studies, but there are advantages from having much larger sample sizes and much better coverage of top incomes. The analysis focuses on tax filers and their spouses (if present), excluding taxpayers aged under 25 years. An individual's income is the income of the his/her tax filing unit, divided by the square root of household size. Income is a measure of pre-tax income, and includes all taxable income sources reported on tax returns supplemented with data about Social Security benefit income provided to the Internal Revenue Service.

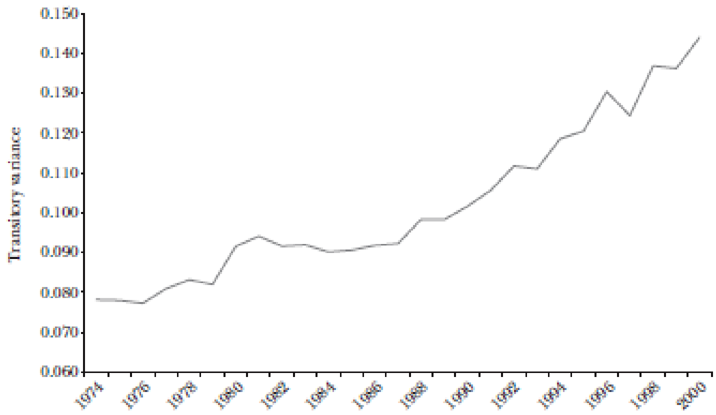
The first part of Auten and Gee's (2009) article describes mobility between 1996 and 2005 in terms of positional mobility (transition proportions) and income growth (by base-year income group). The results are broadly consistent with the studies cited earlier in terms of pointing to substantial movement between quintile groups but with short-distance moves the most prevalent, and real income growth is greater the poorer the base-year income group. The distinctive feature of the study is the information about mobility at the very top of the distribution with mobility statistics also provided for the very top income groups. The authors report that there is a large amount of turnover at the top and that 'the incomes of many taxpayers at the highest levels are very volatile' (Auten and Gee, 2009, 311). For example, among the richest 0.01% in 1996 only 23 per cent remained in the group in 2005. Although over 80 per cent were still in the top 1%, 6 per cent dropped out of the richest fifth (Auten and Gee, 2009, 311).

The second part of Auten and Gee's (2009) article assesses changes in mobility between 1987–96 and 1996–2005 using the same measures, and the authors state with regard to positional mobility that 'the basic finding . . . is that [it] is approximately the same in the last 10 years as it was in the previous decade' (Auten and Gee, 2009, 311). Also, although overall real income growth was around 23 per cent in the first decade compared to 8 per cent in the second, its pro-poor pattern was similar across most of the distribution. (Median real income increased by about 15 percentage points for the top four quintile groups and by about 10 percentage points for the poorest base-year fifth). Things were different at the very top, however. Real income growth was –32 per cent for top 1% in 1987, and –31 per cent for top 1% in 1987 (Auten and Gee, 2009, Table 7).

Further information about persistence in the top 1% is provided by Auten et al. (2013) for tax filers aged 25–60. Their Table 3 shows survival rates in the top 1%, i.e. taking taxpayers in this group in some base year t, what proportion of them in the top 1% in each and every subsequent year $t+\tau$ where $\tau = 1,2,3,4,5$. Base years run from 1991–2009. The five-year survival rates range between 21 per cent and 36 per cent and the one-year survival rates between 52 per cent and 70 per cent. The authors point out that lower persistence rates tend to occur in recessionary periods (1991, 1999 through 2001 and 2007), and they suggest that income sources of particular relevance for the richest groups such as capital gains and net business income are relatively sensitive to the business cycle.

The body of evidence on trends in measures of mobility as family income risk is much smaller than for the other concepts, and also is difficult to synthesize because a wide range of descriptive and model-based measures has been used. One set of PSID-based estimates derived by Gottschalk and Moffitt (2009), is shown in Figure 12. The estimates refer to all individuals in families, and income is the PSID pre-tax post-transfer measure equivalized using the US poverty line for the family type in question. The chart shows that the transitory variance of log annual family income increased substantially, by around 70 per cent, between the mid-1970s and 2000, though this included a period during the 1980s when there

Figure 12 Transitory variance of log annual family income: USA, 1974–2000



Note: Transitory variances computed using the Gottschalk and Moffitt (1994) window-averaging method, with rolling 9-year windows.
Source: Gottschalk and Moffitt (2009, Figure 5), based on PSID data.

was little change. Other PSID-based studies report a similar rise taking the period as a whole (and concur on the increase during the 1990s), though use different measures, time periods, and analysis samples. See inter alia Hacker and Jacobs (2008), and especially Dynan et al. (2012) who also include a useful review of earlier studies for the USA.

There is on-going debate about the robustness of the PSID-based estimates, notably for the 1990s onwards. This is illustrated by the findings of Dahl et al. (2011). They assess household income volatility using data in which responses to the Survey of Program Participation are linked to earnings data from Social Security Administration records ('SIPP-SSA' data).[44] Household income is calculated

[44]Between 10% and 20% of respondents were not matched with SSA records and up to 40% in the 2001 SIPP panel (Dahl et al., 2011, 755). This is a potential source of bias and one that the authors were unable to address.

as the sum across household members of earnings from the SSA records plus the survey reports of non-labour income (but income is apparently not equivalized), and the analysis samples refer to individuals in households with heads aged 25–55 years. Using multiple SIPP panels, the authors derive one-year volatility estimates at 8 time points between 1985 and 2005. The headline finding is that there is no upward trend in volatility and in particular there is little change over the 1990s. Dahl et al. (2011, 769) conclude that they cannot reconcile their results with the divergent set of results from the PSID and other survey data sources, but draw attention to the potential roles of differences in the data per se (rather than the summary measures applied to them). Reconciliation of results is an important task for future research.

The recent study of DeBacker et al. (2013) is a helpful contribution in this respect. It is based on a 1/5000 sample of the US taxpayer population with panel data covering 1987–2009, analyzing individuals aged 25–60 years. There are no potential issues arising from matching or imputation for missing values as in the SIPP-SSA data. The definition of household income is similar to the Auten and Gee (2009) one (see above). The authors calculate one- and two-year volatility measures and the transitory variances using descriptive and model-based estimates. According to all three measures, there was a small rise throughout the period considered (Figures VI, VII, A.1(e)). DeBacker et al. (2013) attribute the rise in the transitory variance primarily to changes in spousal labour earnings and investment income.[45]

[45] Although the transitory variance increased, DeBacker et al. (2013) emphasize that it was the increase in the permanent variance that contributed most to the increase in inequality over the period.

We finish this discussion of US mobility trends with reference to evidence about the mobility of individual labour earnings. The recent literature on trends is dominated by analysis of what we have described as measures of income 'risk', as summarized by the transitory variance and volatility of earnings, and is almost entirely about men's earnings. (The estimates for household income risk cited earlier are usually byproducts of this analysis.) Most analysis is of earnings residuals rather than raw earnings. That is, researchers first run regressions to control for differences in education, age, and work experience, and work with the residuals from the fitted models.

Most studies show that men's earnings instability increased during the 1970s, but then levelled off somewhat through to the early- to mid-1980s or fell slightly. Findings about what happened in the 1990s and 2000s depend on the data set and measure used. This is particularly so when measures of volatility are used. Estimates derived from the PSID suggest a rise in volatility (Celik et al. (2012), Shin and Solon (2011), Moffitt and Gottschalk (2012)) whereas those derived using linked-CPS data, administrative record data or survey data linked to administrative record data, suggest that volatility either remained flat (Ziliak et al. (2011), Celik et al. (2012), Dahl et al. (2011), DeBacker et al. (2013)) or at least appear not to have risen (Juhn and McCue, 2010). There appears to be more agreement across studies and data sets about what happened in the 1990s and afterwards if the focus is on the transitory variance of men's earnings rather than volatility, namely that the earlier rise levelled off in the 1990s and thereafter: see e.g. Gottschalk and Moffitt (2009), Moffitt and Gottschalk (2012), and DeBacker et al. (2013). This is consistent with a finding that it is the variance of the permanent component of men's earnings that has grown most over this period, and note that measures of short-term volatility reflect permanent as well as transitory shocks. For further discussion of the different findings across measures and data sets, see Moffitt and Gottschalk (2012, Section V).

For analysis of trends in earnings mobility using other measures, we refer to Buchinsky and Hunt (1999) and Kopczuk et al. (2010). (There are few other relatively recent studies.) Buchinsky and Hunt (1999) is a detailed study of mobility of in wages and annual labour earnings over the period 1981–91 using the cohort of young people in the National Longitudinal Study of Youth (aged 14–24 in 1979), excluding military personnel and individuals who are self-employed or in education. Mobility is summarised using the Shorrocks equalization measure (M, using multiple inequality indices), and transition probabilities estimated using the non-parametric density method cited in Section 3. The main result about trends is that mobility declined between 1981 and 1991, regardless of which inequality index M is calculated with and using window lengths of one, four, or six years (Buchinsky and Hunt, 1999, Table 2). Positional mobility also declined: the chances of remaining in the same quintile group, and the average jump and normalized trace indices also fell. The decline in mobility as equalisation is the opposite trend from what we discussed earlier for household income. One potential reason relates to the fact that this is a youth cohort, and Buchinsky and Hunt (1999) discuss the difficulties of separately identifying time and age effects.

Kopczuk et al. (2010) is a landmark study of earnings mobility because of its rich data. They use longitudinal Social Security Administration data on earnings stretching from 2004 right back to 1937. The focus is on men and women aged 25–60 years with annual earnings from employment in the commerce and industry sectors greater than a minimum threshold (one-fourth of the full-time full-year minimum wage in 2004 indexed forward and back).[46] Kopczuk et al. (2010) exploit their long series to examine trends in mobility with multiple short- and long-term measures: they variously use longitudinally-averaged earnings over five- and eleven-year windows, and look at measures defined for intervals of various length between base- and final-year. With their large samples and coverage of the tax data, they can also analyse mobility at the top of the earnings distribution.

Short-term mobility is summarised using three measures, the rank correlation for earnings one year apart, and a Gini-based Shorrocks rigidity measure ($R = 1 - M$) and transitory variance of log earnings (calculated using a method similar to the BPEA one) each derived using income averaging over moving five-year windows. According to the first two measures (Kopczuk et al., 2010, Figures IV, V), earnings mobility for all workers increased sharply over the years of World War II and then fell, reaching pre-war levels by around 1960. Thereafter, there was remarkably little change. The transitory variance for log earnings was also roughly constant from around 1960 until the mid-2000s. This result is at odds with the PSID estimates for 1970s discussed earlier (see e.g. the increase shown by Gottschalk and Moffitt, 2009, Figure 1) but consistent with the IRS-data based study from 1987 onwards by DeBacker et al. (2013) (which also, like Kopczuk et al.'s (2010), emphasises the increase in the permanent rather than transitory variance).

From 1978 onwards when earnings data were no longer top-coded, Kopczuk et al. (2010, Figure 6) examine the probabilities of remaining in the top 1% over

[46] The authors undertake extensive checks of the sensitivity of their findings to different assumptions about sample selection, top-coding, coverage of the various administrative sources, etc., and report that their conclusions are robust.

one, three, and five year intervals. There is remarkable stability in these series: e.g. the one-year probability ranges between 72 per cent and 79 per cent, and the five-year probabilities between 60 per cent and 65 per cent. These staying probabilities are greater than those shown by Auten et al. (2013, Table 3) for pre-tax income (for 1991–2009). It is the pre-tax income components other than labour income that are apparently sensitive to the business cycle (and note also that Kopczuk et al.'s (2010) series pre-date the onset of the Great Recession in 2007/8).[47]

To summarise long-term (im)mobility, Kopczuk et al. (2010) use the rank correlation between long-term earnings in years t and $t+\tau$, where $\tau = 10, 15, 20$. For each year, earnings positions are measured by the 11-year average earnings centred around the year in question. The results suggest, first, mobility is greater the larger that T is, which is unsurprising, and yet even after 20 years, the correlation is relatively large (around 0.5 for all workers). Second, for all workers, the rank correlation decreased (mobility increased) between the early 1950s and the early 1970s and was then broadly constant. The trends differ for men from those for all workers: the mobility increase is much less pronounced and appears to rise again slightly from the early 1970s (Kopczuk et al., 2010, Figure VIII).

4.3. Is there more income mobility in the USA than in (Western) Germany?

Perhaps the most well-known 'stylized fact' about income mobility is that mobility is greater in Germany than in the USA. One of the reasons for it being well-known is that it is surprising: many people expect more mobility in the

[47] Auten et al.'s (2013) staying probabilities for years t to $t+\tau, \tau > 1$, are also greater than the corresponding Kopczuk et al. (2010) ones, because the latter's refer to presence in the top 1% in each year rather than simply in the base-year and final-year. For a brief discussion of persistence in the top 1% of the Canadian earnings distribution, see Saez and Veall (2005).

USA because, compared to Germany, the USA has the more flexible labour market and less comprehensive social safety net to cushion income shocks. What is often forgotten is that the original finding refers to one particular mobility concept (equalization of longer-term incomes) and to one particular time period (the 1980s, prior to German re-unification in 1990).

In this sub-section, we review the evidence about income mobility in USA compared to Western Germany. Unless stated otherwise, the data source for the USA is the PSID. We use the term Western Germany ('WG') to refer to the states included in the Federal Republic of Germany before re-unification. The German data source, the SOEP, surveyed the former Eastern German states as well from 1990 onwards, but few mobility studies to date have included these data (see below). We focus on studies that examine household income mobility (which, as it happens, form the vast majority of US-WG comparative analyses). In Table 4, we refer to 11 studies, and summarize them in terms of the time period covered, the mobility measure(s) employed, and the main findings relevant to our question.

The pioneering study by Burkhauser and Poupore (1997) is the source of the stylized fact that we referred to earlier. It was the first major cross-comparative study of household income mobility using the new generation of comparable household panel survey data becoming available in the 1990s.[48] The period covered is 1983–88, a time of upswing in the economic cycle in both countries. Income immobility was summarized in terms of equalization of longer-term incomes using the Shorrocks *R* measure computed with three inequality indices (the

[48]Duncan et al. (1993) and Fritzell (1990) are examples of earlier cross-national studies of poverty dynamics and income income mobility using data that were not as comparable. For an earlier cross-national study of earnings mobility across 8 countries, see OECD (1996).

Table 4 Studies comparing household income mobility in the USA and Western Germany (WG)

Study	Time period covered	(Im)mobility measure(s)	Remarks
Burkhauser and Poupore (1997)	1983–88	Shorrocks *R*	First finding that mobility greater in WG than in USA
Burkhauser et al. (1998)	Year pairs $t, t+\tau$, $\tau = 1, \ldots, 5$, 1983–88	Quintile transition matrices	Slightly more income mobility in WG
Maasoumi and Trede (2001)	1984-89	Maasoumi-Shorrocks *R*	Greater mobility in WG; statistically significant
Gottschalk and Spolaore (2002)	1983, 1993	SWF-based indices	WG–USA difference depends on index parameters
Schluter and Trede (2003)	Year pairs $t, t+1$ between 1984–92	Shorrocks *R*	WG's greater mobility arises from greater mobility in low-income ranges
Van Kerm (2004)	1985, 1997	Portfolio of indices	More income movement in USA; otherwise varies by index
Jenkins and Van Kerm (2006)	Year pairs $t, t+5$: USA 1981–93, WG 1985–99	Indices of re-ranking, progressivity	Reranking and pro-poorness of income growth greater in WG
Schluter and Van de gaer (2011)	Year pairs $t, t+1$ between 1984–92	Index sensitive to upward structural mobility	US 'typically' has more mobility
Allanson (2012)	Year pairs $t, t+5$: USA 1981–96, WG 1985–04	Indices of re-ranking and structural mobility	Reranking and pro-poorness of income growth greater in WG
Demuynck and Van de gaer (2012)	1984–85, 1996–97	Indices of 'inequality-adjusted' income growth	USA–WG ranking depends on weight given low-income-growth individuals
Bayaz-Ozturk et al. (2013)	5-year windows, alternating years, 1984–2006	Shorrocks *R*, ratio of permanent to total variance, log incomes	More mobility in USA from around 1990 onwards

Note: Studies are listed in order of publication year. Each study measures income as equivalized post-tax post-transfer household income (using various equivalence scales), analysis samples are all individuals in households (except Burkhauser et al. (1998), all individuals aged 25–55). Western Germany: the states included in the Federal Republic of Germany before re-unification. Data sources: PSID (USA) and SOEP (WG).

Gini coefficient and the two Theil indices). The base year is 1983 and *R* is calculated as the time period is lengthened from one to a maximum of five years (corresponding to 1988). The headline results were summarized earlier in Figure 8 and refer to estimates based on the Theil index. (The other two indices yield similar profiles and orderings: see Burkhauser and Poupore (1997, Figure 3).)

There is greater longer-term income equalization (less rigidity, lower *R*) in WG than in the USA in each year: the curve for the USA lies everywhere above that for WG. In numerical terms, inequality of six-year-averaged income is 86 per cent of average annual inequality in the USA, compared to 76 per cent in WG, i.e. some 13 per cent larger. The authors show that this mobility ordering is preserved if one uses different income concepts and analysis samples, including labour earnings (for all workers, workers aged 25–50, and the subsets of full-time workers in each case), and equivalized pre-tax pre-transfer ('pre-government') household income.[49] For example, among full-time workers aged 25–50, the six-year *R* for annual labour earnings is 88 per cent for the USA and 79 per cent for WG. For the subset of men, the corresponding estimates are 86 per cent and 78 per cent; for women, 87 per cent and 66 per cent (Burkhauser and Poupore, 1997, Table 4).

The mobility of labour earnings over the same period is analyzed in greater detail by Burkhauser et al. (1997) using different summary methods: statistics based on quintile transition matrices, the rank correlation, and regression-based variance components modelling. Interestingly, given the subsequent focus by researchers on the US-WG differences in household income mobility, Burkhauser

[49] The equivalence scale for all measures of household income in this study is derived from those in the US official poverty line thresholds.

et al. (1997) emphasized the similarities in earnings mobility:

> While we have found evidence of differences in the dynamic earnings movements of workers in the United States and Germany, it is perhaps the similarities of the 'end results' of the two labor markets, despite substantial differences in their institutions, that highlight our multi-period look at these two industrial giants (Burkhauser et al., 1997, 793).

Burkhauser et al. (1998) supplement the two earlier studies from the Burkhauser team. As in the first study they use multiple measures of income (and associated samples), but analyze individuals aged 25–55 years; like the second study, (im)mobility is summarized in positional terms using quintile transition matrices, not *R*. Again, the conclusions point more to cross-national similarities rather than differences: '[i]ndividual mobility patterns in the two countries are remarkably similar' (Burkhauser et al., 1998, 143–4). For example, the proportion of individuals in the same quintile group of post-tax post-transfer household income in 1983 and 1988 is 44.7 per cent in the USA compared to 41.4 per cent in WG; for labour earnings mobility, the corresponding proportions are 52.6 per cent and 53.8 per cent (Burkhauser et al., 1998, Tables 6.2, 6.5).

It is the cross-national difference in *R* that receive the most attention in the later studies, with most authors concerned with the robustness of the conclusion to use of different mobility indices. And all the subsequent studies that we are aware of have focused on household income, not labour earnings. Schluter and Trede's (2003) article is rather different in that they aim to examine the Burkhauser-Poupore result in greater detail. As discussed earlier, their methodological contribution was to explain how *R* reflected the aggregation of distributional changes, differently weighted, at each point along the income range from poorest to richest, and to explore how the aggregation function differed by inequality index. Using a moving two-year window over the period 1984–92 for the calculation of *R*, Schluter and Trede (2003) confirm that mobility is greater in WG than the USA. But their main substantive contribution was the finding that this difference in aggregate reflected a combination of greater mobility in low-income ranges combined with greater local weight given to these changes by the mobility index. The cross-national differences in mobility at the bottom are reminiscent of those revealed in Section 3 by graphical devices such the transition color plot (Figure 1) albeit for a different period (1985 compared with 1997).

Maasoumi and Trede's (2001) article built on earlier work by Maasoumi and Zandvakili (1986) which modified the Shorrocks *R* measure to use different measures of longer-term income (essentially a generalized mean rather than a simple arithmetic average). Maasoumi and Trede (2001) examine USA-WG mobility differences using these Maasoumi-Zandvakili-Shorrocks indices and essentially the same household income data as Burkhauser and Poupore (1997), and also derive the sampling distribution of the indices thereby allowing consideration of whether mobility differences were statistically significant. The substantive findings are threefold: mobility is greater for WG than the USA regardless of the indices (i.e. regardless of the measure of longer-term income, or the inequality index); that cross-national differences were statistically significant; and mobility is greatest among 16–25 year-olds but for all six age groups considered, mobility is statistically significantly greater in WG than the USA.

Gottschalk and Spolaore (2002) is the first (and only) paper that we are aware of that undertakes US-WG comparisons using an explicit SWF-based approach (the application considers mobility between 1984 and 1993). As indicated in Section 2, their approach allows for different weights to be placed on mobility as reversal and as time independence (as well as incorporating inter-temporal inequality aversion of varying degrees). If the reversals and time independence aspects are ignored, so that the SWF reflects inequality-aversion considerations only, Gottschalk and Spolaore (2002) report that the USA 'gains more' from mobility than does WG. But 'this reflects similar gains from reversal in the two countries but greater gains in the U.S. from origin independence. The introduction of aversion to intertemporal fluctuations and aversion to future risk makes the impact of mobility in the two countries more similar' (Gottschalk and Spolaore, 2002, 191). Put simply, conclusions about mobility differences depend on the mobility concept(s) taken and how they are weighted.[50]

Van Kerm (2004) was the first to use Fields and Ok (1999b) indices of income movement to compare the US and WG amongst a portfolio of measures of household income mobility. (He also studies Belgium.) Changing the mobility concept leads to a reversal in the country ranking: the average absolute change in log incomes between 1985 and 1997 is 0.523 in the USA but only 0.392 in WG (and 0.335 in Belgium). Van Kerm remarks that '[d]ifferent concepts of mobility may indeed lead to completely different rankings of economies .... In all cases, mobility is higher in Western Germany than in Belgium, but the USA can stand at any of three positions depending on the index considered' (Van Kerm, 2004, 233). Van Kerm's decompositions highlight that the importance of distinguishing

[50]Gottschalk and Spolaore (2002, Table A.4) provide quintile transition matrices 'for comparison to other studies'. Unfortunately, the scope for doing e.g. dominance checks is limited by the fact that the matrices are not bi-stochastic. The column sums differ greatly from 100% in several cases.

between mobility measures sensitive to positional change and those also reflecting individual income growth and changes in the marginal distributions. The 'exchange' factor of distributional change is greater for WG than the USA, whereas the 'growth' and 'dispersion' factors are greater for the USA (Van Kerm, 2004, Table 4).

Parallel research by Formby et al. (2004) comparing mobility in individual annual labour earnings in WG and the USA between 1985 and 1990 underlines the relevance of the mobility and income concepts chosen. Using measures based on quintile transition matrices, the authors show that there is more positional mobility in the USA than WG according to four out of five indices and there is no dominance in the Atkinson and Bourguignon (1982) sense. However, when origin and destination earnings groups are defined as fractions of mean or median earnings (so the mobility matrices reflect real income growth as well), all five summary indices show greater mobility in the USA. The fact that US-WG positional mobility differences are less pronounced (or reversed) for individual earnings compared to household income underlines the conclusions of Burkhauser et al. (1997) cited earlier.[51]

A range of different mobility indices and time periods is used in the remainder of the studies cited in Table 4. Jenkins and Van Kerm (2006) show that indices of both re-ranking and of progressive individual income growth are greater in WG than in the USA. Using related methods and data, Allanson (2012) confirms the greater re-ranking in WG but also highlights other dimensions of mobility differences. Schluter and Van de gaer (2011) and Demuynck and Van de gaer

[51]The main focus of Formby et al.'s (2004) article is methodological – to derive statistical inference procedures for transition matrices and summary mobility indices based on them.

(2012) propose classes of mobility indices that are sensitive to individual income growth, with different indices reflecting differences in the weights given to income changes of different sizes. Unsurprisingly (in the light of our earlier discussion), both papers report that mobility from this perspective is generally greater in the USA than WG but, also, the ranking can be reversed for some weighting functions.

The final article cited in Table 4 brings us full circle because Bayaz-Ozturk et al.'s (2013) research is in effect a re-analysis of the original Burkhauser and Poupore (1997) study, but using more up-to-date data (1984–2006).[52] The main mobility index is the Shorrocks *R* calculated using the Theil inequality index, but now also supplemented with estimates of the transitory variance of log income expressed as a proportion of the total variance (calculated using the Gottschalk and Moffitt (1994) 'BPEA' method). If the two indices are calculated taking 1984 as the base year and extending the period over which longer-term incomes are calculated to the full 23 years (i.e. also restricting analysis to a sample with fixed structure), income mobility is greater in WG than the USA in each year. The profile of *R* (and for the other measure) for the USA lies above that for WG thoughout, though the gap between them gets smaller over time (Bayaz-Ozturk et al., 2013, Figure 1). In this sense, the results are consistent with the Burkhauser and Poupore (1997) finding (see also Figure 8). However, when the indices are calculated using a moving five-year window (and hance also different samples) in order to examine mobility trends, an interesting result emerges which is illustrated in Figure 11. We remarked earlier in an apparent increase in mobility in the USA in the late-1980s

[52] Alternating years are used to account for the change to bi-annual interviewing in the PSID. The authors report that using contiguous years over the periods where it was feasible leads to similar results.

(though Bayaz-Ozturk et al. (2013) report that the changes in their estimates are not statistically significant). Figure 11 shows that mobility in WG fell between the late 1980s and 1990s (the changes are statistically significant). The result is that, compared to the late 1980s when the WG-USA mobility differences were statistically significant, they were no longer so in the period thereafter.

An interesting substantive question is why WG mobility fell, and to what extent it reflects changes in the (West) German labour market and economy associated with re-unification or with other structural factors (observe that the downward trend apparently started before 1990). Bayaz-Ozturk et al. (2013) report that, when they applied their methods and samples to examine the mobility of labour earnings for men aged 25–59, they found similar patterns of change over time and cite Aretz (2013) as also finding a downward trend in earnings mobility when using administrative record data covering 1975–2008. Interestingly, Aretz's (2013) work shows that the downward trend in WG was broadly U-shaped between the mid-1970s and late-1980s, but did decline again sharply from around 1990. The decline in mobility the former Eastern Germany (measured only after 1990) fell even more rapidly, down to around WG levels by the mid-2000s.[53] See also Riphahn and Schnitzlein (2011), who point to the role of increasing job stability in Eastern Germany.

In sum, although income mobility in the USA and Germany has received much attention, there remains plenty to learn. The sensitivity of conclusions about cross-national differences suggests the need for a more comprehensive analysis using a portfolio of measures within the same study, and using up-to-date data. The

[53]This is found for both men and women, and using the average jump index of positional mobility as well as the Shorrocks *R*.

income concept also matters: researchers have highlighted WG-US differences in household income mobility, and the similarities in earnings mobility have received less attention. Looking at earnings mobility is also informative for tracing the sources of changes in household income mobility.

4.4. Intragenerational income mobility: selected other evidence

The remainder of our discussion of evidence about intragenerational income mobility reviews cross-national comparative studies for wider set of countries and selected country studies analyzing trends over time. The focus remains on household income mobility. We consider work done in the last two decades rather than earlier studies.

A natural place to begin is with the analysis of Aaberge et al. (2002) and Chen (2009) because both includes mobility comparisons between the USA and with other countries. In the former case, the comparisons are with three Scandinavian countries (Denmark, Norway, and Sweden) in the 1980s. In the latter case, they include Canada, Germany, and GB, over the 1990s. Chen (2009) also provides some information about mobility trends.

Aaberge et al.'s (2002) research is based on diverse sources of longitudinal data. For Denmark and Norway, the income data and samples come directly from registers; for Sweden, incomes refer to register data linked to respondents to the Level of Living Survey (the analysis sample is survey- rather than register-based), and for the USA, the source is the PSID (sample and income data come a survey). This diversity leads to some compromises in the search for comparability. For instance, the post-tax post-transfer income concept in the main analysis refers not to a household total but an aggregate across two adults (in the case of a legally married couple) or one adult (all other cases), and equivalized by the number of adults (two or one, respectively). The constraint is what is possible with the Swedish data: no account can be taken of cohabitation, and the number of children is unknown. As it happens, when the authors re-ran their analysis using more conventional definitions (but excluding Sweden), mobility levels change for all countries but 'the mobility ordering of countries is unaffected by this sensitivity check' (Aaberge et al., 2002, 457).

The Aaberge et al. (2002) study provides analysis for 1986–91 and 1990–91, with the end chosen because a major Swedish tax reform in 1991 made later income data non-comparable (the registers covered a different combination of income sources). Mobility is measured using a Gini-based Shorrocks M index and summaries of the directional income movement in the Fields and Ok (1999b) sense. The perhaps surprising finding is that, across the four countries, and despite the substantially greater cross-sectional income inequality in the USA than the three Scandinavian countries, 'the pattern of mobility turns out to be remarkably similar in the sense that the proportionate reduction in inequality from extending the accounting period for income is much the same' (Aaberge et al., 2002, 443). This finding arises whether the analysis is of individual labour earnings or disposable income. The 'remarkable similarity' is also reported by Fritzell (1990) in an earlier study of income mobility in Sweden and the USA. Clearer cross-national differences are apparent, however, when Aaberge et al. (2002) look at the distribution of changes in relative incomes changes between one year and the next over their sample period (relative income is the ratio of income to the year-specific mean; relative income change is a directional summary of individual income movement). As it happens the distribution of relative income changes is more dispersed in the USA than in the Scandinavian countries for both individual earnings and disposable income. Once again, the conclusions about mobility that are drawn, depend on the measure employed.

Chen's (2009) article is based on data from the CNEF, covering from the early 1990s to around 10 years later. Income refers to post-tax post transfer household income, equivalized using the square root scale; the analysis is of all individuals in households with positive incomes. Observe that Germany refers here to the unified country, not WG as earlier. Comparisons with the USA in the late 1990s are complicated by the move to alternate-year interviewing by the PSID.

Chen (2009) summarizes short-term positional mobility in terms of two- and five-year immobility ratios for decline transition matrices, calculated over moving time windows. Choice of measure matters. For example, over the 1990s, around 40 per cent of British individuals remained in the same tenth between one year and the next, compared to nearer 50 per cent in Canada, with Germany's rate in between. With a five-year interval, the cross-national differences become much smaller, with the proportion remaining in the same decile group falling to between 25 per cent and 30 per cent for all the countries. Chen's summary refers to 'a high degree of similarity in relative income mobility across nations' (Chen, 2009, 81) rather than to differences.

Chen (2009, Table 1) presents estimates of the Fields and Ok (1999b) index of income flux, the average absolute log-income change calculated over five-year intervals using between 1991 and 2002. The USA and GB have broadly similar income flux over the period, Germany's is the lowest, and Canada's is in between. Only for the USA is a trend over time apparent (slightly upwards). Assessment of these patterns is complicated because the estimates reflect a combination of differences in overall national income growth rates and changes in how pro-poor the income growth is. Chen (2009, Table 1) shows that economic growth accounts for an increasing share of total income flux in each country (all four countries were in an economic upswing over the period) but does not discuss pro-poorness.

Chen's final set of estimates refer to mobility as equalization of longer-term incomes, summarized using the Shorrocks measure $M = 1 - R$, with 1993 taken as the base year and time periods of up to 6 years (Canada), 10 years (GB and Germany), and 8 years for the USA (1995 and 1997 are excluded). The finding is that mobility is greatest in GB and least in Canada for all time periods, with the profiles for Germany and the USA in between and very similar to each other. Chen (2009, Figure 5) shows this for the case in which M is calculated using the mean logarithmic deviation index, but his Table A2 shows that the result is the same if calculations are done instead with the Theil or Gini index. (If half coefficient of variation squared is used, the US profile is closer to Britain's.) These results echo Bayaz-Ozturk et al.'s (2013) finding of similar longer-term income equalization in the USA and WG after 1990 (see earlier). In his discussion of Burkhauser and Poupore's (1997) results, Chen comments that his results suggest that 'income mobility has increased considerably in the United States between the 1980s and 1990s, while it has declined in Germany' (Chen, 2009, 88).

Leigh (2009) extends comparisons to include Australia, using estimates of R for periods of two and three years. using CNEF data for Britain, Germany and the USA, plus data from the Australian household panel HILDA (HILDA data were not included in the CNEF at the time). He finds that '[a]round 1990, the US was more immobile than either Britain or Germany ... During the 1990s, Germany became somewhat less mobile, and the US somewhat more mobile' Leigh (2009, 16) and that Australia was more mobile than all three other countries in the early

2000s.

A different set of countries is included in the cross-national analysis of Ayala and Sastre (2008), based on ECHP data covering 1993–97: Great Britain, France, Germany, Italy, and Spain. Income is post-tax post-transfer household income equivalized by the modified-OECD scale, and mobility is examined for all individuals using a balanced five-wave panel for each country. According to the Fields and Ok (1999b) index of income flux (Ayala and Sastre, 2008, Table 2), the average absolute log-income change, and looking at income changes between 1993 and 1997, Spain, Great Britain, and Italy have relatively high income flux (index values of 0.390, 0.373, and 0.360, respectively) whereas Germany and especially France are low income flux countries (0.309 and 0.250). Income flux is shown to be greater among individuals in single-parent households, and relatively stable among older persons (as might be expected). A second set of estimates relates to mobility as equalization of longer-term incomes assessed using the ethical indices proposed by Chakravarty et al. (1985), and calculated using multiple inequality indices and for an interval of two years only (individuals' base-year income is the average of their 1993 and 1994 incomes; their final year income is the average of their 1995 and 1996 incomes). The country mobility ranking changes. Regardless of the inequality index used, Italy has the greatest mobility, but Spain slips down the ranking and Germany rises up to second place. As the authors comment, the 'results show that cross-country comparisons of income mobility can be dependent on the approach used' (Ayala and Sastre, 2008, 470). They also refer to potential issues related to differences in national samples (including e.g. a relatively high attrition rate in the Spanish data), and the particular time period covered.

Gangl (2005) was more ambitious in that his mobility comparisons involve eleven EU countries (data from the ECHP) and the USA (PSID). The periods covered are 1994–99 (ECHP) and 1992–97 (PSID). Income is equivalized post-tax post-transfer household income samples are restricted to individuals aged 25–54 years. Gangl calculates two principal measures, namely Shorrocks *R* for a six-year period, and the transitory variance of log income expressed as a proportion of total inequality (derived using a regression decomposition). Discussing *R*, Gangl emphasises similarities across countries rather than differences: e.g. using a Theil-based index, 'about 75% to 80% of observed income inequality has been permanent over the 6-year observation period in most countries' (Gangl, 2005, 149–51). Nonetheless Germany, Ireland, and the USA are relatively immobile countries and the Netherlands and Denmark the most mobile ones. Interestingly, 'low-inequality countries . . . also tend to be the countries exhibiting the lowest degree of persistence in income inequality over time' (Gangl, 2005, 151). Germany is an exception to this description: it is a relatively low inequality country but also with relatively high immobility. Note that this description of Germany also fits with the findings of Aaberge et al. (2002) discussed earlier (for an earlier period). In sum, and on balance, it is unclear whether there is a positive relationship between cross-sectional inequality levels and rigidity of longer-term incomes.

Gangl's (2005) results for household income are consistent with those of Gregg and Vittori (2009) who examine the mobility in labour earnings of individuals aged 20–64 in Denmark, GB, Germany, Italy, and Spain, also using ECHP data. Using *R* calculated for different inequality indices, they find that longer-term earnings inequality reduction is greatest in Denmark followed by Italy, and Germany is the least mobile, with GB and Spain in between. Applying the methods of

Schluter and Trede (2003), Gregg and Vittori (2009) also find that most of the cross-national mobility differences are accounted for by differences in mobility patterns in the lowest earnings ranges.

With his variance components measure, Gangl (2005) finds that

> most (i.e., 65%–70%) of the observed total income inequality for any single country is permanent income inequality with countries like Denmark, the Netherlands, Spain, or Italy at the low end and Ireland, Portugal, the United States, and Germany at the upper end of the scale. Still, cross-national variations in relative income persistence are small, and the country ranking in terms of permanent income inequality in consequence almost exactly mirrors the country ranking for total income inequality (Gangl, 2005, 152).

However, if one focuses on the variance components in absolute levels rather than as expressed as a share of the total variance, the picture changes somewhat. For example, the countries with the lowest transitory variances are Denmark, Germany and Ireland, and the largest are for Italy, the USA, and Spain.[54]

The most comprehensive analysis of income mobility to date using the new EU-SILC longitudinal data, is by Van Kerm and Pi Alperin (2013), who also point to a number of important issues concerning the cross-national comparability of the constituent data sources and the short period covering by the data. On these issues, see also Jenkins and Van Kerm (2013).

[54]The estimates are not directly comparable with those reported for the USA earlier because Gangl's (2005) decomposition uses a model specification that is non-standard. For example, he does not allow for persistence over time in transitory shocks. Also, Gangl includes the variance of heterogeneous income trends in the transitory component rather than permanent component.

We now turn to country studies of income mobility with a focus on trends, of which there are few. Jenkins's (2011a) book contains a comprehensive study for GB, using BHPS data covering from 1991 through to 2006, and examining trends in various concepts of mobility.[55] The headline finding is that, for all but one concept of mobility, there is virtually unchanged mobility throughout the period. This is found for a portfolio of measures, including one-year positional mobility, Shorrocks R measures calculated over moving six-year windows, and the transitory variance of log household income (Gottschalk and Moffitt (1994) 'BPEA' method, using moving seven-year windows).[56]

Jenkins (2011a) reports the same lack of trend if one looks at the earnings of prime-age men and women: see also Jenkins (2011b) and Cappellari and Jenkins (2013). (These studies also cautiously suggest that transitory variances of household income and men's earnings in Britain are larger than their counterparts for the USA.) The lack of change in earnings mobility during the 1990s found in BHPS data is also found by Dickens and McKnight (2008) using administrative record data on earnings covering the period between financial years 1978/9 and 2005/6.[57] They summarize mobility using the Shorrocks equalization measure $M = 1 - R$, calculated using multiple inequality indices, and over moving windows of 2, 4, 6, 8 and 10 years, and each series tells the same story. Interestingly, Dickens and

[55] The book also reviews earlier British studies of income mobility, most of which are by Jenkins and collaborators, and based on shorter spans of BHPS data. For example, Jarvis and Jenkins (1998) used four waves; Jenkins (2000) used 6 waves.

[56] Using Jenkins's (2011a) data, we have compared the decile transition matrices for 1991–98 and 1999–2006 and found that there is no stochastic dominance. Nor is there if we compare the British matrix for 1991–8 with the decile transition matrix for the USA for 1989–98 shown in Table 1.

[57] They make pioneering research use of the Lifetime Labour Market Database (LLMDB), a 1% sample of individuals identified by National Insurance numbers, and originally designed to estimate worker's National Insurance contributions and State retirement pension entitlements.

McKnight's (2008) research also finds that mobility was on a downward trend between 1978/9 and the beginning of the 1990s (though this trend is less pronounced for women than for men).

Jenkins (2011a) observes that the lack of change in British income mobility between the early-1990s and mid-2000s is surprising given significant changes over the period considered in tax-benefit policies, and the upswing in the macroeconomy from trough to peak. Jenkins (2011a, Chapter 6) adduces some evidence to suggest that the lack of trend in aggregate may reflect a balance between changes in mobility associated with different income sources comprising total household income, but he concedes the exploratory nature of the analysis.

The exceptional measure for which some (relatively small) changes are observed is in the pattern of individual growth. Jenkins and Van Kerm (2011) show that income growth between 1998–2002 was more pro-poor than in earlier periods (1992–1996 and 1995–1999) but not so compared with 2001–2005. (An extract from their results was shown earlier in Figure 7.) The authors suggest that the propoor nature of individual income growth in the 1998–2002 period arose because the economy was buoyant, with unemployment rates continuing to fall relatively rapidly from their early-1990s peak, and the incoming Labour government had an explicit anti-poverty agenda, unlike the preceding Conservative governments. It is speculated that the subsequent fall in the progressivity of income growth to do with the slow-down in the economy from around 2000.

Trends in transitory (and permanent) earnings variances of earnings in Western Germany are studied by Bartels and Bönke (2013). Bartels and Bönke work with samples of man aged 20–59 years over the period 1984–2009, calculating variance components using the Gottschalk and Moffitt (1994) 'BPEA' method, using moving five-year windows). The striking finding (Figure 2) is that, although the transitory variance of log earnings rose over the period as a whole, the transitory variance for equivalized post-tax post-transfer household income (for the same sample) does not change at all over the period, pointing to important roles played by the German welfare state and by families in offsetting shocks to men's earnings. When the same methods are applied to Britain (BHPS data for 1991–2006), Bartels and Bönke (2013) find, like Jenkins (2011a), that the transitory variance for equivalized post-tax post-transfer household income does not change over time, unlike him, they also report (Figure 6) a rise in the transitory difference of men's earnings (and higher levels). These differences are traced to differences in samples: Jenkins (2011a) considered men aged 25–59 (as in most similar US studies) whereas Bartels and Bönke (2013) also include younger men (down to age 20) and they argue that transitory earnings shocks are more important for this group. Overall, the authors conclude from their analysis that 'redistribution and risk insurance provided by the welfare state is more pronounced in Germany than in the United Kingdom' (Bartels and Bönke, 2013, 250). Whether this also applies to other groups beyond prime-aged men requires examination.

Mobility in top incomes in Germany over the period 2001–6 is studied by Jenderny (2013) using tax administrative data, a 5% balanced sample of all tax filers in those years. Income is the tax unit's gross pre-tax income (i.e. including tax-exempted income, but not realized capital gains). One-year probabilities of remaining in the top 1% are about 78 per cent, and thus larger than the estimates of around 70 per cent reported by Auten et al. (2013) for non-recessionary periods in the USA (see above). Five-year survival rates are also larger in Germany than in the USA.[58] Jenderny (2013, 32) concludes that the increase in top income concentration in Germany since the 1990s described by Bach et al. (2009) is unlikely to be offset by high or rising top income mobility.

4.5. Summary and conclusions

Empirical studies of income mobility show that, in all countries, there is a substantial degree of longitudinal flux in incomes, whether looking at incomes one year apart, or five or ten years apart, resulting in changes in relative position and a reduction in the inequality of longer-term incomes. It is also clear, however, that most income changes are relatively small so that, even after many years, relative positions are quite highly correlated and substantial inequalities in longer-term incomes remain.

To the big questions of whether income mobility in country *A* has increased or decreased over time, or is greater or less than in country *B* (or *C* or *D* or ...), we have found few clear cut conclusions – apart from a general finding that the answers to the questions depend on the mobility concept that is used, and other issues such as the time period considered and the measure of income are relevant.

This is illustrated by the comparisons of the USA with West Germany. Early research suggested that income mobility in the 1980s was (surprisingly) greater in WG than in the USA (Burkhauser and Poupore, 1997) when mobility is measured in terms of equalization of longer-term income. But more recent research (Bayaz-Ozturk et al., 2013) for the 1990s using the same measure suggests that mobility in the two countries is now similar. And it is often forgotten that the Burkhauser team had long argued that earnings mobility in WG and USA were remarkably similar.

[58] The stayer rates for the top 0.1% are also slightly higher than the Canadian estimates reported by Saez and Veall (2005, Figure 2).

Moreover, when one switches the mobility concept to one of income movement (or individual income growth), mobility in the USA shows up as greater than in most other countries – the ranking consistent with many people's expectations given the nature of the US economy, labour market, and welfare state.

It remains an open question, as well, whether there is a systematic cross-national relationship between levels of income mobility and cross-sectional income inequality. The evidence is mixed, and the issue deserves to be revisited. (Note the wide-spread interest too in whether there is a corresponding relationship for intergenerational income mobility – see the discussion of the 'Great Gatsby' curve in 5). Because the evidence we have reviewed suggests similarities across countries in the extent of mobility (positional and longer-term income equalization) rather than marked differences, we are inclined to conclude that there is no obvious relationship between mobility and inequality since cross-national differences in equality are pronounced.

Looking at trends over time in income mobility within countries, the picture is one of diversity and depends on the mobility concept, and the length of time period over which trends are assessed. Mobility changes are observed in the USA over the 30 years since the early 1970s and in Germany between the late 1980s and the 1990s, though whether these count as large or small changes partly depends on the eye of the beholder. For Britain, there is a clearer case that income mobility in Britain changed hardly at all in the 1990s and 2000s (again with the exception of mobility as individual income growth). Relatively large changes in mobility are more apparent to most eyes once trends are assessed over a relatively long period. The US study of earnings mobility by Kopczuk et al. (2010), with data going back to 1937, is the best example we have of this.

In sum, our review of evidence about income mobility suggests that there is much to learn. The advent of cross-nationally comparative household panel surveys over the last three decades facilitated a relative boom in intragenerational mobility analysis. There are signs that the next generation of studies will make greater use of administrative register data or surveys linked to administrative data, at least for analysis of trends over time. As we have discussed, data from sources such as tax administrative records provide the advantages of huge samples with good coverage of top incomes and can provide long historical series as well. On the other hand, these benefits come at the potential costs of having income definitions that are not as useful for mobility analysis as those now in comparative survey collections such as the CNEF (and may change over time as tax laws change), and data access and undertaking the analysis are also non-trivial issues. For cross-national comparisons, administrative record data also have potential, but the problems of comparability are an order of magnitude greater, and data may simply be unavailable for countries of key interest.

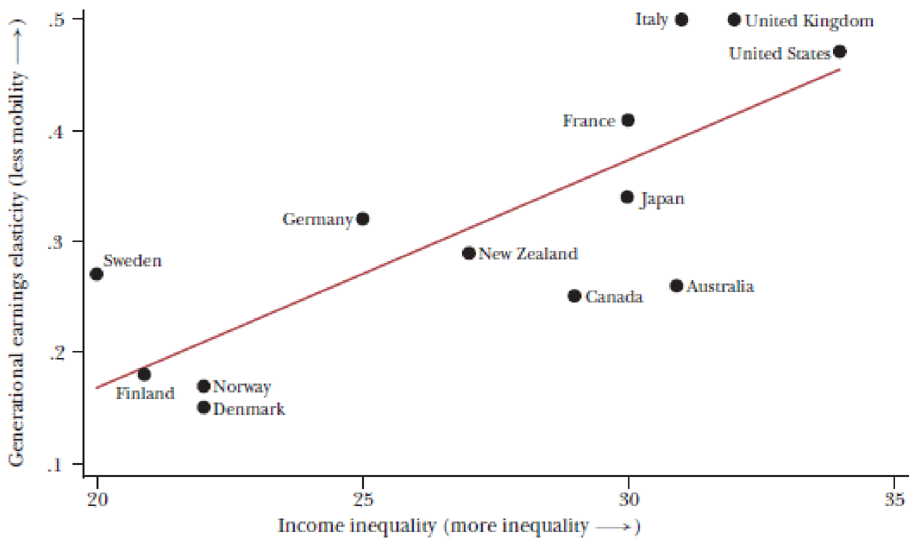
5. Intergenerational mobility: evidence

Section 3 presented a set of measures by means of which we can describe not only intra- but also intergenerational associations in a society. This section reviews evidence on such associations.

There are several earlier reviews of intergenerational income mobility. Solon (1999) reviews intergenerational labour market with a focus on long-run earnings, while Solon (2002) focuses on a subset of that literature, namely cross-national differences in mobility. Björklund and Jäntti (2009) build on and extend the empirical evidence assembled by Solon (1999). Black and Devereux (2011), who examine intergenerational links in income and education, emphasize evidence on causal links in intergenerational mobility. Blanden (2013) contrasts the cross-national evidence on intergenerational income, earnings, and education mobility with mobility in social class. Corak (2006, 2013a), in turn, emphasizes policy implications. Corak (2013a) also draws on recent research about both socioeconomic gradients in child development and the emergence of economic persistence in labour markets.

Several recent reviews present international evidence on intergenerational income persistence in a scatterplot, plotting the estimated persistence in different countries on the vertical axis and estimated income inequality, often in the parental generation, on the horizontal, adding a linear bivariate regression line (Corak, 2013a; Blanden, 2013; Björklund and Jäntti, 2009). Labelled the 'Great Gatsby' curve by the then Chairman of the US Council of Economic Advisors (Krueger, 2012), such plots are interpreted to suggest countries with higher persistence are also countries with greater inequality. Figure 13 reproduces the most recent such graph, from Corak (2013a, Figure 1). Although the precise estimates used by dif-

Figure 13 The Great Gatsby curve: the relationship between intergenerational earnings persistence and cross-sectional income inequality



Note: Income inequality is measured by the Gini coefficient of disposable household income in 1985 taken from the OECD. Persistence is measured as the Beta of parental and son earnings. Sons are born in early 1960s and outcomes for them are measured in late 1990s. See Corak (2013a,b) for further detail.
Source: Corak (2013a, Figure 1).

ferent authors vary, the results are broadly similar. The Nordic countries have low persistence and low inequality, the USA, the UK along with France and Italy, have high persistence and reasonably high inequality.

There are theoretical models that can account for the positive association between inequality and persistence. For instance, in Solon's (2004) version of the Becker and Tomes (1979, 1986) model, the factors that drive intergenerational persistence, such as the heritability of human capital endowments, the returns to education, and the progressivity of public education expenditure, affect cross-sectional inequality with the same signs. In Hassler et al.'s (2007) model, which examines links between inequality and mobility under different kinds of labour market institutions, some institutional arrangements have mobility and inequality being inversely related (and hence persistence and inequality positively correlated). Checchi et al.'s (1999) model of beliefs about own ability, educational choice and mobility can also generate positive as well as negative associations between inequality and mobility depending on the model parameters. As we shall see, however, it is far from clear that intergenerational persistence and inequality are, in fact, as clearly positively correlated as Figure 13 suggests.

This part of the chapter proceeds as follows. In Section 5.1, we discuss data requirements and special problems that come up in estimating intergenerational and family associations. In Section 5.2, we review studies of intergenerational persistence and mobility in the USA. The focus on this country is motivated, as in intragenerational mobility case, by the sheer amount of evidence about mobility in the USA relative to that in other countries. First, we examine evidence on the level of the intergenerational elasticity (IGE) of earnings or income – first for father-son pairs, and then widen the scope to look at broader pairings of parents and offspring – and then examine evidence about trends in the IGE over time. (The IGE is the Beta measure discussed in Section 3; we use both terms interchangeably in this section.) We then examine evidence that is based on measures that go beyond the simple log-linear Galtonian regression Beta (IGE), product-moment correlation coefficient r: for example, quantile regressions, transition matrices, non-parametric conditional mean functions. In Section 5.3, we examine evidence on intergenerational mobility from other countries, following the same structure as for the USA. In Section 5.4, we examine evidence on another way to measure the importance of family background, the sibling correlation, and in Section 5.5, we discuss other approaches to intergenerational mobility, old and new. Section 5.6 concludes.

5.1. Data and issues of empirical implementation

As discussed in Section 4.1, any study of income mobility faces three 'W' issues: mobility of What, among Whom, and When? For intergenerational mobility, each question must be answered twice, once in each of the parental and offspring generations. As with intragenerational mobility, researchers choices are constrained with the available data.

At one level, just as with intra-generational income mobility, mobility of 'What' refers to the income concept that is used. The overwhelming majority of studies we review below use the labour market earnings of the parent and the offspring with several variations, discussed in Section 4.1. Other choices might add non-labour income sources from the market such as capital income to measure factor or market income. If the goal is to examine the intergenerational association of living standards, it would make sense to study disposable income, i.e. to add public transfers and deduct income taxes paid. It would seem reasonable to have identical answers to the 'What' question in both generations. It is frequently the case that available data do not support such a choice: it is not unusual for 'income' to be family income in the parental generation and to be earnings in the offspring generation.[59]

The aim of early research on this topic was to measure the intergenerational association of 'permanent' income, which was believed to be captured quite well

[59]This is the case in US studies that rely on the NLSY and UK studies that rely on the BCS and NCDS. Note that the Galtonian regression of child height on parent height also used mid-parent height on the right hand side, see Galton (1886) and Goldberger (1989).

by labour market earnings. It has long been recognized (see Atkinson, 1981b) that short-run income measures are different from longer-run measures because of transitory fluctuations, and that the association sought were those of the more stable or permanent measures of living standards.

As with intragenerational mobility, 'Whom' refers to the definition of the income-receiving unit in both the parent and child generation. Most studies that are modelled on Solon (1992) examine mobility of father-son pairs, ignoring the incomes of other household members. Many departures from this are due to data-related reasons. For instance, studies such as that of Zimmerman (1992) which relies on data from the US National Longitudinal Survey of Youth (NLSY), uses family income in the parental generation as that is the only income concept available in that data source.[60]

The Whom question becomes more complex when the intergenerational associations of women's incomes are studied, and compared with those of men. Over the last four to five decades, women's labour market attachment has increased substantially in most developed nations, with female labour force participation rates increasingly resembling those of men. However, around the age commonly believed to be appropriate for measuring men's long-run income (around age 40), women often have breaks from employment due to child birth and caring. Studies that examine women's intergenerational mobility are more likely to examine family or household income as a better gauge of their living standard than individual incomes. Comparing mobility across men and women would then naturally also need to examine family or household income for men (Chadwick and Solon,

[60]The Galtonian regression Beta that is the most often used measure of (im)mobility originally related offspring height to 'midparent' height. See Goldberger (1989) and Galton (1886).

2002; Raaum et al., 2007).

There is an added dimension to the Whom question, namely the nature of the parent-offspring relationship. In the early intergenerational studies of Atkinson (1981b), Solon (1992) and Zimmerman (1992), the parent-child association was more or less driven by the survey design – 'children' were the children of the sample parents who were followed up in adulthood. However, children can have multiple parents – step parents, adoptive and foster parents in addition to birth parents. Common choices are to restrict the population to those parent-offspring pairs where the offspring was observed as living with the parent at some age, say 10 or 16, or to birth parents. One aspect of this Whom dimension is the role of separated families. Should one focus on associations of offspring income with the head in lone-parent families or on father-child associations? Some studies, especially based on register data, have examined the sensitivity of the population of parent-child relationships and found differences across definitions and family types to be relatively small.[61]

As with the two other 'W' questions, the When question for intergenerational mobility analysis is mostly a superset of that for intragenerational mobility. Most of the same questions addressed in Section 4.1 need to be resolved for both the parent and offspring generations. The underlying data record income for a specific period: often annual income data but in some cases 'current' income data are available. But, in contrast to the case of intragenerational mobility, shorter-run fluctuations are noise that make more difficult the uncovering of the more interesting underlying longer-run incomes. This leads directly to the issue of over

[61] See Björklund and Chadwick (2003), and, using mobility in education, Holmlund et al. (2011) and Björklund et al. (2007b).

what periods, and over what ages, incomes should be studied (and aggregated) to give reasonable measurements of longer-run economic status? And if, due to data limitations, ideal measurements cannot be made, how are mobility measurements affected? The two main issues that have been addressed are transitory variation in observed income measures, and 'lifecycle' bias (Jenkins, 1987; Grawe, 2006). We discuss these in turn.

Since at least Atkinson (1981b), it has been recognized that transitory errors in parental income lead to an errors-in-variables (downwards) inconsistency in the estimated intergenerational elasticity. Since the seminal paper to empirically address this issue, (Solon, 1992), many studies have exploited this finding.[62] Solon's estimate of intergenerational persistence for the USA, based on averages across five-years of parental income, resulted in point estimates of Beta that are between 10 per cent to 70 per cent larger than the estimates derived using a single-year of parental income.

Recent work on so-called generalized-errors-in-variables (GEIV) model calls into question the assumption that transitory income variations have the same properties as classical measurement errors (Haider and Solon, 2006; Böhlmark and Lindquist, 2006). The GEIV model for the annual income process of an individual in family i in generation $j(= O\text{ffspring}, P\text{arent})$ at age t relates permanent income y and transitory errors v to annual or current income by (Haider and Solon, 2006)

$$y_{ijt} = \lambda_{jt} y_{ij} + v_{ijt} \quad j = O, P. \tag{16}$$

[62] It is much less common to correct for measurement error in other ways to asses mobility, such as transition matrices. For the econometrics involved and evidence based on simulations, see O'Neill et al. (2007).

The key advance here is the introduction of the age-dependent parameter λ which 'loads' underlying permanent income onto annual income and is hypothesized to be lower than one early in the life cycle, equal one at some point and be higher than one thereafter. Note that we allow for the λ parameters to differ across generations.[63]

The measurement error model in equation 16 is the same as the classical measurement error model if (i) $\lambda_{jt} \equiv 1$, and (ii) the random fluctuations v are orthogonal to true long-run income ($y \perp v$), and the vs are identically and independently distributed within a generation. An estimate of the IGE β using annual incomes for both parents and children has the probability limit

$$\begin{aligned} \operatorname{plim}\widehat{\beta} &= \frac{\operatorname{Cov}[y_{iOt}, y_{iPt}]}{\operatorname{Var}[y_{iPt}]} \\ &= \frac{\operatorname{Cov}[y_{iO}, y_{iP}] + \operatorname{Cov}[v_{iOt}, y_{iP}] + \operatorname{Cov}[y_{iO}, v_{iPt}] + \operatorname{Cov}[v_{iOt}, v_{iPt}]}{\operatorname{Var}[y_{iP}] + \operatorname{Var}[v_{iPt}] + 2\operatorname{Cov}[y_{iP}, v_{iPt}]}. \end{aligned} \tag{17}$$

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 17 are all zero. In that case, also the third term in the denominator is zero, and only the presence of the random fluctuation in parental income in the denominator leads to downward inconsistency, a point made in this context first by Atkinson (1981b). Denote the variance of permanent earnings in generation $j = 0, P$ by $\operatorname{Var}[y_{ji}] = \sigma^2_{y_j}$ and similarly the variance of transitory earnings by $\operatorname{Var}[v_{jit}] = \sigma^2_{v_j}$. The most common empirical solution to deal with the inconsistency is to diminish it by taking multi-year averages of parental income,

[63] The current exposition treats the transitory errors as being white noise. In case they are autocorrelated, the attenuation factor involves also the parameters of that process. If v follows and AR(1) process, such autocorrelation worsens the errors-in-variables inconsistency. See e.g. Mazumder (2005b, Table 1).

in which case the measurement error variance in the denominator is $\sigma^2_{y_P}/T < \sigma^2_{v_P}$, an approach first used by Solon (1989, 1992) and Zimmerman (1992) and now standard in the literature.[64]

However, the age- or time-dependent factor loading λ_{jt} leads to two additional sources 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 $\text{Var}[y_{iP}]$.

If only parental income is measured with error, we would have

$$\text{p}\lim \widehat{\beta} = \frac{\text{Cov}[y_{Oit}, y_{Pi}]}{\sigma^2_{y_P}} = \theta_{Ps}\beta, \tag{18}$$

where s is the age at which parental income is measured and

$$\theta_{Ps} = \frac{\text{Cov}[y_{Pis}, y_{Pi}]}{\text{Var}[y_{Pis}]} = \frac{\lambda_{Ps}\sigma^2_{y_P}}{\lambda^2_{Ps}\sigma^2_{y_P} + \sigma^2_{v_P}} \tag{19}$$

is the linear projection coefficient of y_{Pi} on y_{Pis} (Haider and Solon, 2006). If parental income is measured at an age at which $\lambda_{Ps} \approx 1$, θ is the standard errors-in-variables attenuation factor. If only *offspring* income were characterised by the GEIV process, the probability limit of the IGE estimated using annual income would be

$$\text{p}\lim \widehat{\beta} = \frac{\text{Cov}[y_{Oit}, y_{Pi}]}{\sigma^2_{y_P}} = \lambda_{Ot}\beta. \tag{20}$$

Whenever $\lambda_{Ot} \neq 1$, the IGE is inconsistently estimated (Haider and Solon, 2006). Thus, when in the lifecycle offspring incomes are measured matters for estimates of intergenerational persistence. This is the issue of 'lifecycle bias'.

[64] A consistent estimator can also be constructed if there is a valid instrument for permanent income, an approach that has also been used (see e.g. Dearden et al., 1997).

If both offspring and parental incomes are characterised by the GEIV model, which is plausible, the estimated IGE is (Haider and Solon, 2004; Gouskova et al., 2010)

$$\operatorname{plim}\widehat{\beta} = \lambda_{Ot}\theta_{Ps}\beta. \tag{21}$$

Note that λ can be either below and above unity, and θ is constrained to lie in the unit interval, so β can be underestimated (when $\lambda < 1$), be correctly estimated (when $\lambda_{Ot}\theta_{Ps} \approx 1$) and be overestimated (when $\lambda_{Ot}\theta_{Ps} > 1$). Finally, the Pearson correlation coefficient r has in this case the probability limit

$$\operatorname{plim}\hat{r} = \theta_{Ot}\frac{\sqrt{\lambda_{Ot}^2\sigma_{y_O}^2 + \sigma_{v_O}^2}}{\sigma_{y_O}}\theta_{Ps}\frac{\sqrt{\lambda_{Ps}^2\sigma_{y_P}^2 + \sigma_{v_P}^2}}{\sigma_{y_P}}r. \tag{22}$$

The probability limit of the correlation coefficient depends on the θ in both generations, on the ratio of the observed standard deviation to that of long-run income in both generations, as well of course on the true r.

Empirical evidence from both the USA 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-inconsistent measure of lifetime earnings and later in life an upward-inconsistent measure (Haider and Solon, 2006; Böhlmark and Lindquist, 2006). Around age 40, at least for men in both the USA and Sweden, $\lambda_t \approx 1$ in which case deviations from a multi-year average are approximately classical, thus lending themselves to the analysis of intergenerational association of long-run income *under the assumption* that the λs in both generations are approximately equal, i.e., $\lambda_{Pt} \approx \lambda_{Ot}$ (or at least that they equal unity at about the same age).

Grawe (2006), building on insights in Jenkins (1987), examined the extent of both attenuation and lifecycle bias in Betas estimated in several different countries and datasets. He finds, using data for Canada, Germany and the USA, that lifecycle biases in fathers' age are an important source of bias and proposes several rules of thumb to diminish it: either to use points in time at which measurement errors are roughly classical, as suggested above, or at least to use observations on income for parents and children at similar points in their lifecycle.

There are several caveats, however. First, as implied above, the λs may well change from one generation to the next. Second, the λs that apply to, say, earnings, may differ from those that apply to, say, disposable household income. Third, the λs that apply to men may be quite different than those that apply to women depending, for instance, on patterns of labor force withdrawal and re-entry due to child bearing. Finally, the λs can be quite different in different countries. Without access to estimates of these, cross-country differences in the IGEs can be driven not by differences in the underlying βs but by different values of λ_{Ot} and θ_{Ps}, even if the ages at which incomes are measured are kept constant.[65]

It may be interesting to know how large the bias in Beta or r is in a given population. However, we are often interested in comparing these parameters in two different populations, for example across time in a country, or between two countries. Denoting the two populations by A and B and focusing on Beta, and assuming for simplicity we are measuring both parents and offspring at the same ages, we have

$$\widehat{\beta}^A - \widehat{\beta}^B \simeq \lambda_{Ot}^A \theta_{Ps}^A \beta^A - \lambda_{Ot}^B \theta_{Ps}^B \beta^B. \tag{23}$$

[65]Moreover, Nybom and Stuhler (2011) use nearly complete actual lifetime incomes for both 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 may mean that more complex models that link short-run to 'permanent' income need to be explored.

Unless we have estimates of λ and θ in both countries, we must assume that $\lambda^A_{Ot}\theta^A_{Ps} \approx \lambda^B_{Ot}\theta^B_{Ps}$ to be able to deduce from the estimated difference that the underlying βs are different, also. A similar argument applies, of course, to r. We can in that case infer the sign of their difference without bias (but can not know its size unless we know λ and θ), or we can estimate their ratio.

The almost exclusive focus on permanent income (which, in some sense, involves both the What and When questions) can be questioned in light of the more complicated measurement models that link short-run to long-run income. The focus on permanent income is based on the notion that differences between short- and long-run income are transitory and largely classical, i.e. positive and negative shocks are roughly as likely (low or not autocorrelation) and the magnitude of the shocks does not vary by either permanent income or other characteristics (shocks are homoscedastic and orthogonal to permanent income). If capital markets are well functioning and individuals have a fair idea of what their permanent income is (so they know if they have been hit by a negative or positive shock) they smooth their consumption by relying on saving and borrowing. These demanding conditions would justify the focus on permanent income. (See the discussion in Section 2.) In this view, it is permanent income, not short-run fluctuations, that best captures the distribution of well-being.

It follows that, if the assumptions are violated, even short-run fluctuations are interesting from a well-being perspective. Jäntti and Lindahl (2012) demonstrate that income volatility in Sweden is strongly but non-monotonically associated with the level of long-run income. Moreover, analysis of intergenerational associations in income suggests that not only long-run incomes but also income volatility is associated across generations (Shore, 2011; Jäntti and Lindahl, 2012).

Thus, a focus on long-run incomes alone probably understates the extent to which economic well-being is associated across generations.

Before we discuss commonly-used data sources for intergenerational analysis, we point to an additional complication in interpreting the evidence. The most commonly used measure of intergenerational mobility is the persistence as measured by Beta (the IGE). Arguably, we would like to abstract from the marginal distribution of offspring and use the correlation, r, related to Beta by the ratio of parental to offspring standard deviation (see equation 4 in Section 3). In 'steady state', the two are equal but, when inequality increases (decreases) across generations, r is lower (higher) than Beta. Extra care should then be taken in comparing Betas across countries, as different Betas may be consistent with the same, or at least more similar r, depending on how marginal distributions have changed across generations in the two countries. Björklund and Jäntti (2009) cite evidence that suggests the ratio σ_P/σ_O in the US is less than one – inequality increased – and in Sweden greater than one – inequality decreased – suggesting the difference in r is likely to be less than that in the Betas. (Note, however, that is only changes in inequality in the marginal distributions that are controlled for, and in a particular way.)

Suitable data for intergenerational analysis need to meet two basic criteria. The data need to be able to identify and link parent-child pairs.[66] They need to include measurements of the incomes in both generations at comparable points in the lifecycle and preferably multiple such measurements to allow for the effect of measurement errors.

[66] Intergenerational persistence can be estimated using two-sample methods (Björklund and Jäntti, 1997), which we discuss below.

There are three main types of data that are used. Many studies rely on longitudinal household surveys that have been running long enough to allow for the offspring to be observed living with their parent(s) as a child or youth and then followed up as adults, having often formed households of their own. These data sources, discussed in greater detail in Section 4.1, include the US PSID and the German SOEP. The UK BHPS has only recently been used for intergenerational analysis. Cohort studies are another type of data that are commonly used for intergenerational analysis. Such datasets, including the US National Longitudinal Study of Youth (NLSY) and the UK National Child Development Study (NCDS; cohort of 1958) and British Cohort Study (BCS, cohort of 1970) have been specifically designed to collect data on children, and follow them across time as they grow older. The income and other information used is mostly gathered through interviews with the parents and children in such studies, and the parent-child link is ascertained from both information about both birth and living arrangements.

A variation of the survey-based approach was taken in the UK, where Atkinson (1981b); Atkinson et al. (1983), using data originally collected by Seebohm Rowntree and collaborators (Rowntree and Lavers, 1951) for the study of cross-sectional poverty in York, built an intergenerational dataset by interviewing the adult children of the original survey households, creating a longitudinal dataset from what was originally a cross-sectional dataset.

Register-based datasets are another important source of data. Such data, which underlie intergenerational mobility estimates in Canada and the Nordic countries, and increasingly also in the USA, rely on administrative records, often drawn from data originally collected for purposes of taxation or social security, to measure income, and identify parent-child links based either on administrative records that link parents to children or on census data.[67] The key to the use of such data is the use of personal identifiers and the presence of a reliable parent-child link.

A third approach to data is to use synthetic parent-child links. One way to do this is to use two-sample methods, i.e. to estimate the Beta using empirical moments based on different datasets. This requires a sample of 'parents' to provide information on the unconditional distribution of income in the parental generation and of the distribution conditional on a few key predictors of income, as well as a sample of 'offspring' to provide information on both their income distribution and on the predictors for their parents. Two-sample methods, first used in the intergenerational context by Björklund and Jäntti (1997) for a comparison of the USA and Sweden, have later been used in several countries, including Britain, Italy, France, Brazil, and Australia.[68]

Each of these three types of data are subject to measurement challenges (see Section 4). Measurement errors in income and other data used in the analysis are an issue, and not only in survey- but also in register-based sources, although the nature of the errors are likely different (e.g., recall error in survey data and underreporting due to tax evasion in register data).[69] Attrition, especially selective attrition, is a concern in longitudinal surveys, a problem that may be compounded in intergenerational follow-up. The reliability of the identification and linking of parents and children is a concern when that is done using administrative data.

[67]For instance, Canadian father-son pairs in studies such as Corak and Heisz (1999) rely on tax records for of only the earnings information, but also the father-son link. In the Nordic countries, parent-child links are from either Census data or from birth records (Bratsberg et al., 2007).

[68]Two-sample methods were independently developed by Angrist and Krueger (1992) and Arellano and Meghir (1992). Methods to estimate the variance of such estimators were derived by (see Inoue and Solon, 2010).

[69]For a discussion, see e.g. Ehling and Rendtel (2004).

Before we delve into the evidence, we should note that the overwhelming majority of studies, in the USA as well as in other countries estimate elasticities, i.e. estimates of the Beta measure discussed in Section 3. When correlations are available (either directly, reported by the authors, or derived using equation 4 based on Betas and the standard deviations), they are product-moment (Pearson) rather than rank (Spearman) correlations. Fully controlling for the marginal distributions would require the latter. Moreover, we are unaware of any study that explicitly recognizes the implications of the GEIV model for estimated elasticities that attempts to control for those effects.

One reason most analysts may estimate Betas rather than Pearson or Spearman correlation coefficients, or transition matrices, is convenience: controlling for systematic lifecycle effects in both generations is simple in a multiple regression framework. Moreover, the impact of transitory errors, when classical, is well understood and simple to mitigate. Estimation of the Pearson correlation is subject to the same errors-in-variables inconsistency as the Beta, but transitory errors in both offspring and parent income cause r to be underestimated, so reducing the inconsistency requires time-averaging income in both generations. Transitory errors lead to inconsistent estimates of rank correlations also. O'Neill et al. (2007), who present simulation evidence based on bivariate normal distributed parent-offspring income that are subject to a range of different kinds of measurement error, suggest robustly that intergenerational persistence is underestimated and mobility overestimated in the presence of measurement error. Finally, in many cases, Beta and r are estimated using instrumental variables, often using sample moments from different samples. These techniques are well understood for in moment-based estimation, but less so for rank correlations and non-parametric techniques.

5.2. Intergenerational persistence in the USA

Although there are many studies of intergenerational mobility in the USA as well as in other countries, the literature is characterized by a surprising number of omissions. For instance, we have been unable to locate transition matrices for different cohorts of parent-child pairs, so we are unable to examine the change across time in mobility using the dominance approach.[70] Most US researchers report only Betas, not *r* or rank correlations, so standardization for changes in the marginal distribution of earnings or income is incomplete at best. And yet this is a period in which there have been pronounced increases in inequality in the USA (and many other countries).

By the late 1980s, two longitudinal datasets in the USA, the Panel Study of Income Dynamics (PSID) and the National Longitudinal Study of Youth (NLSY) had been running for sufficiently long to allow the study of the incomes of parents and children at economically active ages. Around that time, three papers were published in reasonably close succession that made use of these data, by Solon (1992), Zimmerman (1992) and Altonji and Dunn (1991). The papers by Solon (1992) and Zimmerman (1992), which appeared prominently in the same issue of the *American Economic Review*, made two major contributions.[71] First, they pointed out some of the statistical problems involved in estimating the relationship between 'long-run' incomes of members of the same family. Most earlier studies used single-year measures of permanent earnings and were based on non-representative, homogeneous samples. Their analyses suggested that the estimates

[70]Fertig (2003) shows evidence based on transition matrices for multiple cohorts of offspring but does not report the full transition matrices.

[71]Altonji and Dunn (1991) received far less attention, in part because it was published in a less well-known journal.

of intergenerational correlations in previous studies most likely were considerably downwards biased. By using multiple years of father's earnings, this downward bias could be reduced. Solon also presents an estimator that most likely overestimates the correlation and thus produced a range within which the true correlation must lie. Second, their results suggested Betas between father's and son's long-run incomes as high as 0.4 or 0.5, numbers that are much larger than those in the previous studies surveyed by Becker and Tomes (1986). Solon and Zimmerman obtained similar results using two different data sets, which lends additional credibility to their findings.

Solon (1992) also estimated the correlation by use of an instrumental variables (IV) method, arguing that this produces an *upwards* inconsistent estimate of Beta. The argument, in brief, is to treat parental education as an omitted variable, but also to use parental education as an instrument, so it is an invalid instrument. If the true direct effect of parental education is positive and it is positively correlated with parental income, such an IV estimate produces an overestimate of the intergenerational income elasticity. Thus, the OLS estimator using time-averaged parental income underestimates and the IV estimator overestimates the elasticity, bounding the parameter from below and above. Moreover, in pointing to the possibility of using an IV estimator, Solon (1992) also opened the door to estimating elasticities, reliably as it turned out, in cases where actual father-son pairs are unavailable.[72]

[72]The first example of such work is that by Björklund and Jäntti (1997), which used an IV estimator based on two independent samples to estimate the intergenerational elasticity in Sweden, and constructed a similar estimate for the USA. Betas estimated using actual Swedish father-son pairs are almost exactly the same as the two-sample IV estimate. There have been a large number of studies subsequently that use two-sample IV estimators including for Italy (Checchi et al., 1999), Brazil (Dunn, 2007), Australia (Leigh, 2007) and France (Lefranc et al., 2009).

Mazumder (2005b), using earnings information from the US Social Security Administration (SSA), examines intergenerational earnings Betas for US sons and daughters with respect to father's earnings. His focus is on variations in the number of years across which father's earnings are averaged, along with several other measurement issues such as whether or not to require fathers to have positive earnings in all years, if zero earnings due to non-coverage of social security by registers are imputed or not, as well as whether or not zero earning offspring are included in the analysis. The results demonstrate, among other things, that attenuation from transitory variation in earnings remains substantial even after averaging fathers' earnings over up to 16 years, especially if transitory errors are characterized by autocorrelated errors. In a striking demonstration of the the impact of averaging across multiple years of parental income, Mazumder (2005b, Table 4) reports an elasticity of 0.253 (se 0.043) when only two years (1984–85) of father's incomes are used, increasing to 0.553 (se 0.099) and 0.613 (se 0.096) when averages across 1976–85 and 1970–85 are used, instead. The US estimates he reports thus encompass the majority of the estimates reported in Figure 13, excluding only Peru at the top end and Canada, Finland, Norway and Denmark at the low end. Note, however, that in extending the number of periods over which father's income is averaged conflates to types of effects, namely transitory errors (whose variance is reduced) and lifecycle effects (which become averaged). In the absence of estimates of λ_{Ps} and θ_{Ps}, it is hard to tell which of these is empirically responsible for the change in the elasticity.

Dahl and DeLeire (2008), also using data from the SSA but with data on non-covered years as well, and using even longer time spans for fathers earnings than-Mazumder (2005b), estimate Betas for father-son and father-daughters pairs. The father-son estimates vary between 0.259 and 0.632, spanning, again, much of the observed range in the cross-country evidence on display in Figure 13. The father-daughter Betas range from –0.041 (which is not significantly different from zero) to 0.269. Naga (2002) uses father-son pairs observed at the same point in the life cycle and estimates elasticities using three methods – OLS on time-averaged data, IV and a MIMIC latent variable estimator – and finds elasticities that range from 0.297 to 0.7.

Chadwick and Solon (2002), Minicozzi (2002) and Fertig (2003) also examine the Betas for women. Chadwick and Solon (2002) highlight the importance of using family income when comparing Betas for men and women (in which case they are quite similar; when using individual earnings, Betas for women tend to be much lower). The sensitivity of Betas to sample selection rules was examined by Couch and Lillard (1998) and Minicozzi (2003). In both of those papers, sample selection issues are found to be very important for the Betas. Hertz (2005) examines racial differences in the elasticity. Estimates of Beta for the USA can also often be found in research that is either comparative or primarily about other mobility in other countries, Examples include the studies of Germany by Couch and Dunn (1997), Australia by Leigh (2007), Sweden by Björklund and Jäntti (1997), and Singapore by Ng et al. (2009).

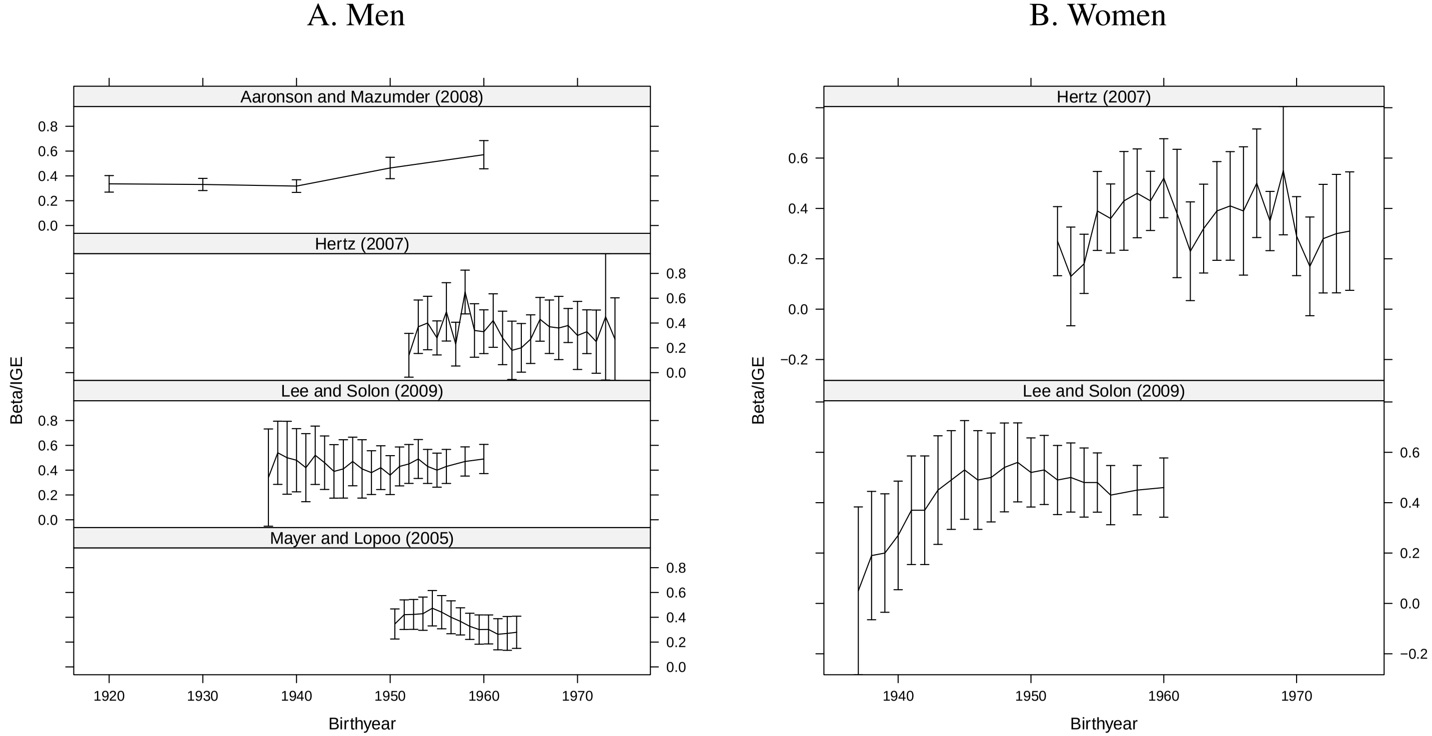
Two US papers that drew attention early on to the possibility that estimates of intergenerational persistence may be subject to not only attenuation inconsistency from transitory errors in fathers' earnings or income, but also to lifecycle effects in the offspring generation, were Buron (1994) and Reville (1995). Instead of adjusting for the average life cycle effects, Buron (1994) allows earnings profiles to vary across demographic groups, which leads to a higher estimated persistence than when using the same adjustment. Reville (1995) in turn investigates how varying the age and outcome year of sons changes the estimated persistence. For instance, by following the same cohort of offspring as they age from 26–30 to 34–38 and using a four-year average of their earnings (keeping father's earnings constant), the Pearson correlation r increases from 0.296 to 0.423 (Reville, 1995, Table 5). Hertz (2007), Lee and Solon (2009), Gouskova et al. (2010), Chau (2012) all try to take into account biases from both transitory errors and lifecycle effects.

Gouskova et al. (2010), applying the insights of both Haider and Solon (2006) and Grawe (2006), estimate earnings elasticities for father-son pairs using data from the PSID where the fathers and sons are of the same age. Using age ranges 25–34, 35–44, and 45–54, regressing a three-year average of sons' earnings on a five-year average of fathers' earnings, they find elasticities of 0.29, 0.41 and 0.42, respectively. These estimates, especially the low value for the 25–34 age range, are consistent with the patterns for λ in Haider and Solon (2006). Another recent study considering the implications of the results in Haider and Solon (2006), Chau (2012), models the income processes of both fathers and sons using heterogeneous growth profiles and autocorrelated errors. Intergenerational elasticities are then estimated based on data simulated using the parameter estimates. The US estimates, based on PSID data, show an estimate of Beta of 0.392, but elasticities are as high as 0.662 when the earnings processes of sons and fathers are allowed to be different.

Muller (2010) tackles another complication with estimating the measurement of permanent income, namely if the elasticity varies because of shocks income to parental income that take place when the offspring was living in the parental home. Parental income earned in childhood years are associated with much higher elasticities than either before the child was born or after he had left home, a result that is broadly robust with respect to standardizing the stage of the lifecycle at which incomes are measured in the two generations. The results are consistent with the view that transitory shocks in childhood do affect offspring income. While the purpose of the literature on intergenerational mobility reviewed here is not to uncover causal effects of income, this finding lends weight to the view, discussed above in Section 5.1, that income risk may also be intergenerationally correlated.

Trends over time in intergenerational mobility in the USA, as measured by changes in Beta, have been estimated by Hertz (2007), Mayer and Lopoo (2005), Lee and Solon (2009) and, using two-sample methods, by Aaronson and Mazumder (2008). We show a selection of estimates in Figure 14, indexed by the birth year of the offspring, ranging from men born in the 1920s to men and women born in the early 1970s. The elasticities are evaluated at somewhat different ages, but the picture that emerges is one which suggests little systematic trend among men, with the possible exception that persistence may have increased among men from the 1940s to 1960s, mainly on display in the Aaronson and Mazumder (2008) estimates and weakly supported by both Hertz (2007) and Lee and Solon (2009). The estimates for women in Hertz (2007) and Lee and Solon (2009) suggest increasing persistence for the early cohorts but little change from around 1960 onwards. The differences across studies suggest care must be taken in interpreting trends based on but a few data points and sets of definitions. The large confidence intervals around each point estimate also highlights the importance of statistical inference. Indeed, all of the confidence intervals in the series from Hertz (2007), Lee and

Figure 14 Trends in US intergenerational income persistence



Note: The estimates in Lee and Solon (2009) are the elasticities for different outcome years at age 40, presented here by subtracting 40 from the outcome year, amd are derived using a three-year average of parental income. Mayer and Lopoo (2005) estimate elasticities for four-year birth cohorts which are centered here, and observe offspring at age 30, and use a seven-year average of parental income (at ages 19–25). Hertz (2007) presents elasticities at age 25 and uses a three-year average of income. His estimates further control for panel attrition. Aaronson and Mazumder (2008) uses two-sample methods applied to (IPUMS) census data, with elasticities applying to 35–44 year olds, here centered at age 40.
Source: (Aaronson and Mazumder, 2008, Table 1, column 6) Hertz (2007, Table 4), Mayer and Lopoo (2005, Table A1) and Lee and Solon (2009, Table 1).

Solon (2009) and Mayer and Lopoo (2005) overlap. While this does not mean there cannot be significant differences between point estimates, it does warrant some caution.

The IGE (Beta) is related to a 'global' log-linear regression, forcing the slope of the conditional expectation of offspring log income to be a linear function of parent log income. There are many ways to relax the assumption that the slope is the same everywhere. Differences in the slope at different levels of parental income can be motivated by theoretical concerns. A commonly cited concern is the potential presence of borrowing constraints with respect to parental investments in child human capital (Becker and Tomes, 1986; Grawe, 2004b; Bratsberg et al., 2007). Bratsberg et al. (2007) fit a polynomial in parental income to the data for the US drawn from the NLSY to allow for a flexible shape between offspring and parental income. They find that a second-order polynomial in parental income provides a reasonable fit for US data. The IGE based on a log-log regression is 0.542, while those based on the polynomial imply elasticities of 0.489, 0.575 and 0.646 at the 10th, 50th and 90th percentiles of parental income, respectively. Couch and Lillard (2004) demonstrate that these results are highly sensitive to the procedure applied. Using both second- and third-order polynomials in both the log and the level of parental income, they estimate elasticities in the first, third, and fifth quintile groups of fathers' income to be 0.124, 0.234 and 0.292 using a quadratic, and 0.219, 0.230 and 0.171 for the cubic polynomial, compared to 0.158 in the log-log. Thus, using the second-order polynomial, elasticities increase monotonically across father's income but, using the third-order polynomial, they increase to decline at higher levels. Another option is to estimate the conditional mean (and, by implication, its slope) non-parametrically, for instance using kernel regression.

The elasticity is a measure of average *persistence* of income rather than of *mobility*. In other words, the regression coefficient on father's log (permanent) earnings tells us how closely related, on average, an offspring's economic status is to that of his or her parent. It is quite possible for two distributions to have highly similar average persistence, but for one to have substantially more mobility around that average persistence. The elasticity can thus be the same, but arguably the distribution with a greater residual variation – variability around the average persistence – is the one with greater mobility. (See the discussion of the Gottschalk and Moffitt 'BPEA' measure in Section 3.) Moreover, two distributions with the same regression slope may have quite different, and varying, conditional variances around that slope. For instance, a distribution with a 'bulge' in the variance at low levels of fathers' earnings, that is, a pear-shaped bivariate distribution, will exhibit relatively more mobility at the low end of the distribution than will a distribution with a constant conditional variance.

One approach is to examine both the regression coefficients and residual variances. Other approaches, such as non-parametric bivariate density estimates, similar to Figure 4 in the intragenerational case in Section 3, would in principle be available (see e.g. Bowles and Gintis, 2002). Very few studies take that route, however. Quantile regression (Koenker, 2005) can also be used to examine the conditional distribution of offspring income, conditioned on parental income. While the slopes of the conditional quantiles of offspring income can be of interest in and of themselves, we tend to find what they say about the full conditional distribution of greater interest than the slopes of individual quantiles (cf. the discussion of this in Section 3.3). In the prototypical homoscedastic regression, where the variance (or indeed, any higher moments) of the residual does not depend on the explanatory variable, the quantile regression slopes should all be straight lines with slopes equal to the conditional mean and median. Deviations from these patterns are informative of variations in the shape of the conditional distribution.

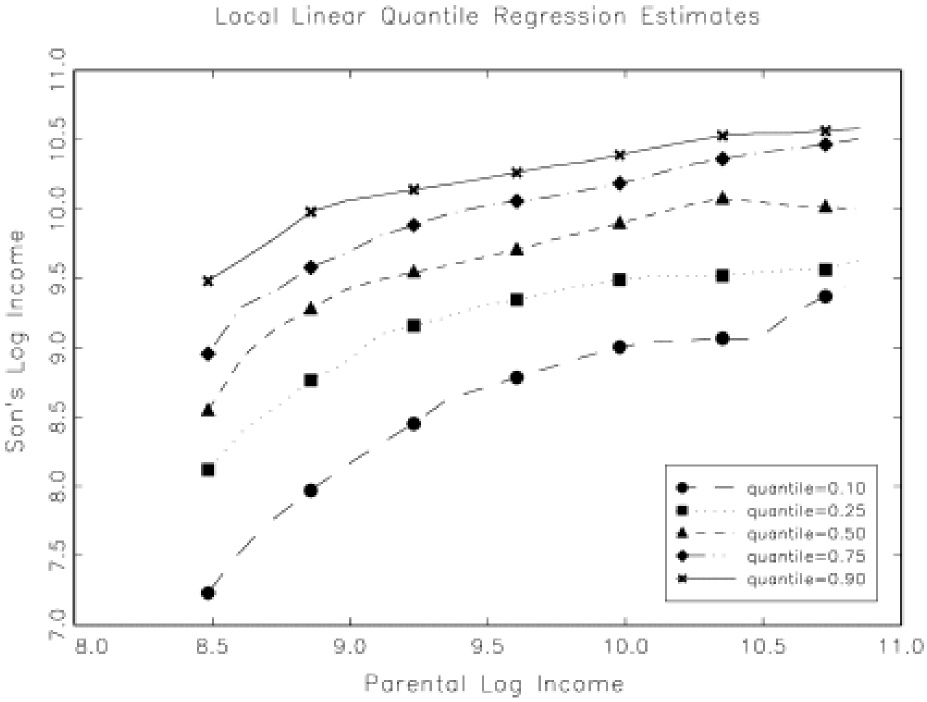
Eide and Showalter (1999) estimate quantile regressions for several percentiles using PSID data on father-son pairs where the sons are 25–34 years old, using a three-year average of parental income and seven-year average of sons' earnings. They find a Beta of 0.34, and slopes of the conditional quantiles with respect to parental income of 0.77 at the 5th percentile, 0.47 at the 10th percentile, 0.37 at the 50th percentile (median), 0.17 at the 90th percentile, and 0.19 at the 95th percentile. That is, they (mostly) find the slope to be decreasing in the percentile but also that the Beta is lower than the slopes of the quantiles up to the 75th percentile.[73]

Conditional quantiles can be combined with non-parametric techniques to allow for the shape to change flexibly. We illustrate this in Figure 15 from Lee et al. (2009), who use PSID data for US sons and fathers to non-parametrically estimate the conditional quantiles of sons income conditioned on fathers. We can see that the slopes of lower quantiles tend to be steeper at low parental income than for the higher quantiles, and that the slopes tend to level of as parental income increases.

Asymmetries in intergenerational mobility can be straightforwardly described using transition matrices, a simple but under-used device for illustrating intergenerational mobility. In allowing for fairly general patterns of mobility, mobility or transition matrices offer the additional advantage of allowing for asymmetric pat-

[73] See also Grawe (2004a) for additional US estimates.

Figure 15 Intergenerational income persistence: non-parametric quantile regression for US father-son pairs



Note: Estimates based on PSID father-son pairs as prepared by Minicozzi (2003). Sons' income is the average of labour income at ages 28 and 29 and parental income is predicted parental income as defined by Minicozzi (2003).
Source: Lee et al. (2009, Figure 1).

terns, for example more mobility at the top than at the bottom. To illustrate, we show in Panel A of Table 5, a decile transition matrix for US fathers and sons.

The cell entries show, for each decile group of origin, i.e., fathers' decile group, the percentage of sons in each destination decile group. Specific aspects of the transition matrix tend to be highlighted. For instance, the main diagonal shows the percentage of sons who remain in their father's decile group. One descriptive statistic is the sum of the main diagonal probabilities (the matrix trace), in this case 165. With ten income classes, there is origin independence if each entry in the table is 10 per cent, which implies an average 'excess' immobility relative to origin independence of 6.5 percentage points. Conversely, 83.5 per cent of US sons are in a different decile group than their fathers. The Normalized Trace index (Shorrocks, 1978b) for this matrix is $(10-165/100)/(10-1) = 0.93$.

The corner probabilities are often of special interest also. In this case, 22 per cent of the sons of the poorest tenth of fathers are in the poorest tenth themselves, whereas 26 per cent of the sons of the richest tenth of fathers are in the richest tenth. Conversely, upward mobility from the lowest tenth is $100-22 = 78$ per cent and downward mobility from the highest tenth is $100-26 = 74$ per cent. By contrast, 7 per cent of sons of poorest fathers and 3 per cent of the richest end up in the top and bottom decile groups, respectively. Somewhat to our surprise, we are unable to illustrate an application of dominance analysis to examine the change across cohorts in US intergenerational mobility. We are unaware of a comparable transition matrix for a later or earlier cohort.

A final observation can be made regarding the 'shape' of the transition matrix. Transition matrices for bivariate normal data, such as the simulated data in O'Neill et al. (2007) or the illustrations of the consequences of different r in Björklund and

Jäntti (1997), are symmetric. For instance, the two corners on the main diagonal are equal as are the corners on the antidiagonal, and the upper triangle is the mirror of the lower triangle. The US father-son transition matrix clearly exhibits very little symmetry of this sort. The lack of symmetry implies that both mobility and persistence may be different across the distribution, and of course that the data are unlikely to be well described by a bivariate log normal distribution.

Recently, Bhattacharya and Mazumder (2011) proposed a set of measures based on the bivariate percentile distribution, focusing specifically on upward and downward mobility relative to a parameter τ that specifies the number of percentiles one needs to move up to be considered upward mobile, illustrating their approach by comparing mobility differences between racial groups in the USA using data on men from the NLSY. Whites are found to be distinctly more likely to move upward than blacks.

5.3. Cross-national comparative evidence on intergenerational associations

We now turn examining evidence on intergenerational income mobility in other (mostly rich) countries. To illustrate the importance of how mobility is measured for cross-country rankings, we start this subsection by reporting results from two recent papers, each of which compares three countries. Corak et al. (2013) compare earnings mobility between fathers and sons in Canada, Sweden, and the USA. Their focus is on comparing upward and downward mobility, but we rely here on their three estimates of persistence: Beta (IGE), the Pearson correlation r, and the Spearman rank correlation, are reported in Table 6 along with the ranking of the three countries in each case. The estimated Betas are in line with those found in previous research and show intergenerational income persistence to be the greatest in the USA, followed by Canada and Sweden. The ranking by the

Table 5 Intergenerational decile transition matrices for earnings, father-son pairs, Canada and the USA

A. USA

Father	Son 1	2	3	4	5	6	7	8	9	10
1	22	18	10	10	11	11	5	5	2	7
2	9	15	16	15	9	9	9	5	9	5
3	9	10	12	17	15	9	9	7	7	5
4	17	9	10	12	3	15	9	11	7	7
5	12	7	12	6	14	9	12	10	12	8
6	7	11	6	10	11	13	13	11	7	11
7	8	7	12	9	11	9	16	13	9	5
8	8	8	8	11	10	7	11	15	13	8
9	4	8	8	5	9	11	7	9	20	19
10	3	8	6	7	7	5	10	16	11	26

B. Canada

Father	Son 1	2	3	4	5	6	7	8	9	10
1	16	14	12	11	10	9	8	7	7	7
2	13	13	12	12	11	10	9	8	7	6
3	11	11	12	12	12	11	10	8	8	7
4	10	10	11	11	11	11	11	10	8	7
5	9	10	10	10	11	10	11	11	10	8
6	9	9	10	10	10	11	11	11	10	9
7	8	9	9	9	10	10	11	11	11	11
8	8	8	8	9	9	10	11	12	12	12
9	8	8	8	8	8	10	10	12	13	15
10	8	8	8	8	8	9	10	11	13	18

Note: The cell entries show, for each decile group origin (referring to fathers), the percentage of sons in each destination decile group. US estimates are based on SIPP matched to social security earnings. Fathers' earnings are averaged across 1979–85 and sons' across 1995-98. Canadian data are based on tax records. Fathers' earnings are averaged across 1978–82 and sons' earnings across 1993–95.
Source: Mazumder (2005a, Table 2.2) and Corak and Heisz (1999, Table 6).

Table 6 Intergenerational earnings mobility in Canada, Sweden and the USA: Beta, *r*, and the rank correlation

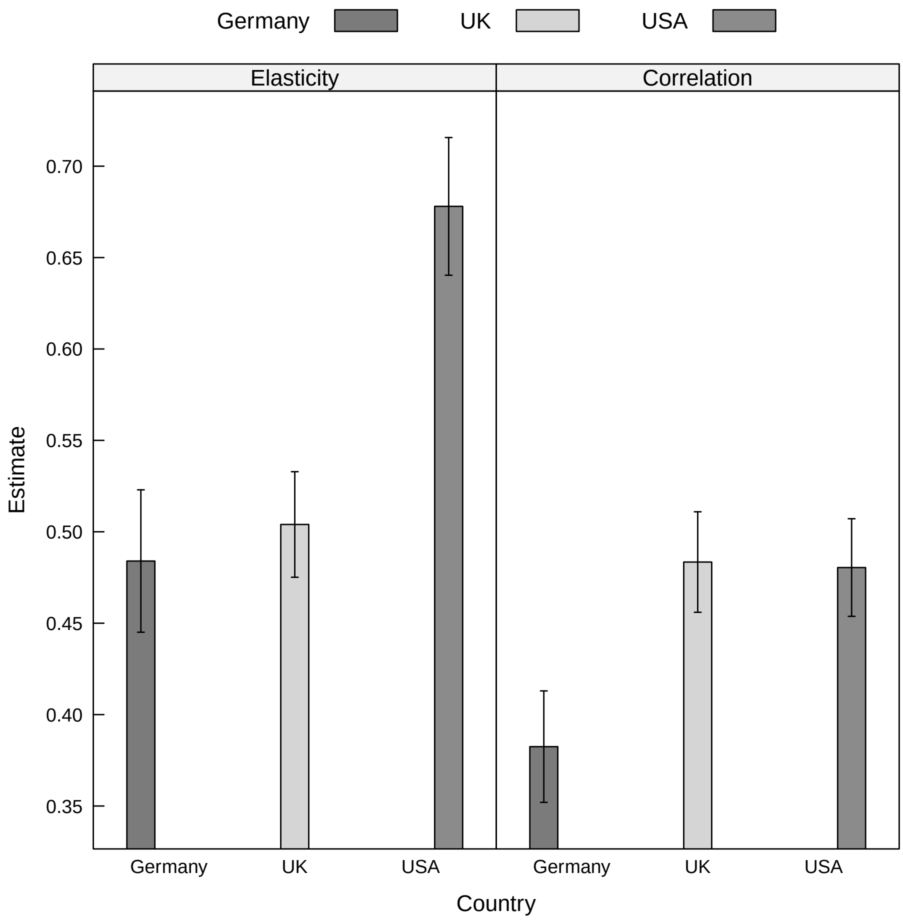
Country	Beta		*r*		Rank correlation	
	Estimate	Rank	Estimate	Rank	Estimate	Rank
Canada	0.26	(2)	0.23	(2)	0.24	(1)
Sweden	0.25	(1)	0.21	(1)	0.30	(2)
USA	0.40	(3)	0.26	(3)	0.30	(2)

Note: Canadian estimates rely on tax records. Father's earnings are a five-year average and son's a three-year average 1997–1999 when they were 31–36 years old. Swedish estimates, also based on tax records for earnings, rely for fathers on 20 years of earnings data measured at ages 30–60 and for sons on an 11-year average across ages 30–40. The US estimates stem from the Survey of Income and Program Participation panels using earnings from Social Security records. Fathers earnings are a nine-year average between 1979–1986 when they were 30–60 years old. Sons' earnings are a five-year average between 2003–7 in years they were at least 28 years old.
Source: Corak et al. (2013, 10–11).

product-moment correlation *r* is the same, but now the US point estimate is much closer to those of Canada and Sweden. By contrast, according to the rank correlations, Canada has the lowest persistence and Sweden and the USA are tied. This, arguably the preferred scalar index of persistence (as it most clearly abstracts from differences in marginal distributions), suggests a very different ordering of countries with respect to intergenerational mobility than that on display in the 'Great Gatsby' curve of Figure 13.

Eberharter (2013) estimates persistence in terms of Betas for disposable income among men and women in Germany, the UK, and the USA, using data from the US PSID, the German SOEP, and the UK BHPS. The elasticity estimates are reported in the left panel of Figure 16 together with the 95 per cent confidence intervals. This is a rare study because it presents estimates for several countries us-

Figure 16 Intergenerational persistence of disposable income: elasticities versus correlations



Note: Error bars show 95% confidence intervals. Estimates are for post-tax, post-transfer income for all individuals (for sons and daughters combined). Offspring incomes are observed for those older than 24 who are out of full-time education and are averaged across 2005–9 (Germany), 2003–7 (USA) and 2004–8 (UK). Parental income are observed as offspring were 14–20 years old and are averaged across 1988–92 (Germany), 1987–91 (USA), and 1991–95 (UK). Eberharter (2013) reports standard deviations in parental and offspring generations for full samples rather than the estimation samples, so the estimated implied correlations, obtained using $\rho = \sigma_P/\sigma_O\beta$ are approximate only.
Source: Authors' elaborations based on Eberharter (2013, Tables 1, 2).

ing measures of disposable income. It is also unusual to pool sons and daughters, although that choice is arguably well-motivated when the purpose is to examine the persistence in living standards.

Although Eberharter (2013) does not report rank correlations, these results bring out quite forcefully the importance of being wary of changes in marginal distributions across the cohorts, especially when comparing estimates from different countries.[74] As can be seen by comparing the left panel of Figure 16, which plots the elasticities, with the right panel, which reports the implied (Pearson) correlations *r*, the results are dramatically different in the two cases. The USA has a substantially higher elasticity than either Germany or the UK (0.68 as opposed to 0.48 and 0.50), but when we derive the correlations, the UK has a correlation that is higher than that in the USA, and Germany's is substantially lower than either of those.[75] It is not possible, of course, to infer what the rank correlations are from the Betas and *r*.

Thus, even confining ourselves to scalar measures of mobility, switching between Beta and the two correlations leads to rank reversals. The fact that Sweden and the USA, two countries that inhabit very different regions in the 'Great Gatsby' curve diagram, have equal mobility as measured by the rank correlation, is particularly notable.

Most studies of intergenerational income persistence and mobility were inspired by the US studies of Solon (1992), Zimmerman (1992) and Altonji and

[74] See e.g. Björklund and Jäntti (2009).

[75] The orderings are statistically robust: the confidence interval for the US elasticity does not overlap those for Germany or the UK and, while the intervals for the correlations between Germany on the one hand and the USA and the UK overlap a little, pairwise t-tests reject the null that the correlations are the same.

Dunn (1991). An exception is the study of intergenerational mobility in the UK Atkinson (1981b); Atkinson et al. (1983) (cited several times by Solon (1989, 1992), possibly serving as inspiration for the US studies). Intergenerational income persistence in the UK, especially the question of whether it has changed, has been subject to substantial controversy recently. We therefore start our discussion of single-country studies with UK evidence.

The early estimates of father-son Betas in Atkinson (1981b); Atkinson et al. (1983) using a geographically limited and truncated sample were around 0.44. Atkinson et al. (1983, 111) discuss the impact of measurement error in parental income, finding that for plausible value of the signal-to-noise ratio, the true Beta might well be at least 0.5. Dearden et al. (1997), using data from the cohort study of 1958-born children (the NCDS), estimated Beta to be around between 0.29 (OLS) and 0.58 (2SLS). Later studies by Blanden and Machin (2008). Blanden et al. (2010) and Blanden et al. (2013) have generated a reasonably wide range of UK estimates.

One particularly contested UK finding is that mobility has decreased, based on the finding that the IGE estimated for the cohort born in 1958 (NCDS) is greater than the IGE for the cohort born in 1970 (BCS). Depending on estimation method, the elasticity increased from 0.31 to 0.33 (OLS) or 0.33 to 0.50 (2SLS), both measured for sons at age 34 (Blanden and Machin, 2007). Most recently, using a single year measure of parental income and no controls for parental age, Blanden et al. (2013) report an increase in the IGE between NCDS and BCS cohorts from 0.211 to 0.278 for parent-son pairs, corresponding to a difference of 0.067 (se 0.034). These estimates have been widely discussed in UK public policy debates about social mobility, as discussed recently by Goldthorpe (2013).

The UK debate provides several lessons. First, two estimates provide little evidence about the existence of a trend. US estimates for different birth cohorts vary quite substantially: see Figure 14, where there is no apparent trend. Moreover, different data sources and estimation methods may generate different results. For example, Nicoletti and Ermisch (2007) derive Betas for Britain using two-sample methods applied to BHPS data. They estimate relatively stable elasticities and correlations for cohorts born between 1950 and 1960. For the cohorts born between 1961 and 1972, elasticities rise somewhat over time but correlations are stable. These results are only partially consistent with the estimates derived from the BCS and NCDS cohorts. Second, data quality has serious implications for public policy. Part of the UK controversy centers around whether the two cohort studies in question, the NCDS and BCS, have sufficiently comparable data. Third, measures of intergenerational *income* mobility may change over time in a different way from measures relating to other concepts of intergenerational economic and social mobility. These differences may in turn be informative of the nature of societal change.

The possibility that intergenerational income persistence in the UK has increased, but class mobility has not, led Erikson and Goldthorpe (2010) to examine mobility in the earnings/income and class spaces. They conclude that problems with the measurement of income in the parental generation render the finding of an increase in income persistence suspect, and they emphasise the stability of social class mobility over time as indicating that there has been little change in intergenerational mobility in the UK.

More recently, Blanden et al. (2013) use an approach proposed by Björklund and Jäntti (2000) to decompose the *r* (strictly speaking, the partial correlation) into the correlation of 'class-predicted' incomes, the correlation of deviations of actual from class-predicted incomes, and their cross-correlations. Their results are consistent with there being stable class mobility, as suggested by there being no contribution (but in fact, a small negative one) from the 'class-predicted' income correlation to the change, whereas all three correlations involving the residuals contributed to an increased partial correlation. The results can be interpreted as saying income and class mobility decreasingly capture the same phenomena, as the relationship between income and class appears to be different in the later than in the earlier cohort. The discussion of these results by Blanden et al. (2013), Erikson and Goldthorpe (2010), and Goldthorpe (2013), provides valuable insights into the scientific and public debates about social and economic mobility.

A key conclusion that we draw about the UK debate, not least in light of the divergent US estimates of both levels and trends, is that much richer data than those provided by the NCDS and BCS cohort studies are needed in order to draw firm conclusions about the level and trend in UK income mobility.[76] It is also possible that class and income mobility are diverging because the processes that generate transitory errors are changing in ways that suggest intergenerational advantage is increasingly transmitted through deviations from the systematic components of income. In our view, the UK debate underlines the need for high-quality data to resolve what has turned out to be a question of great social concern.

Corak and Heisz (1999) provide Betas for both earnings and total market income for Canadian father-son pairs, using (at most) a five-year average of parental

[76]We note, in passing, that the dominance analysis conducted by us of the income quintile group transition matrices reported Blanden et al. (2010, Table 3) for the NCDS and BCS suggest that, except for the cells (5,3) and (5, 4) all BCS–NCDS differences in the cumulated matrices are positive (but those entries are negative). Thus, there is no dominance between the two cohorts.

income and a single year for sons' in 1995 at which point they are 29–32 years old. They find elasticities for earnings of 0.131 and for market income of 0.194. In addition to transition matrices, discussed below, they also estimate the conditional expectation, and its slope, of sons' earnings with respect to fathers', nonparametrically. They find that the elasticity varies substantially and quite nonmonotonically across the distribution of fathers' earnings.

Leigh (2007) estimates intergenerational earnings elasticity for Australian men using two-sample methods. For men born in 1949–1979, he estimates an elasticity of 0.181. This compares to a US elasticity for a similar cohort of sons, obtained using similar estimation methods, of 0.325. The difference is statistically insignificant, but still suggests Australian persistence is lower. His results for older cohorts, vary substantially, however. For men born in 1911–40 and 1919–43, the point estimates are 0.26, but for men born in 1933–62, the estimate is 0.413. Gibbons (2010) estimate intergenerational mobility for New Zealand father-son and father-daughter pairs of 0.25 and and 0.17, respectively,

Lefranc (2011) uses two-sample methods to estimate Betas for cohorts of men born between 1931 and 1975 in France. The estimates, which start at 0.626 for men born 1931–35 decline to 0.441 for cohorts born 1956–60 and increase thereafter, being 0.559 for cohorts born 1971–75. Estimates for Spain are provided e.g. by Cervini-Plá (2009) and for Italy by Mocetti (2007), both of which are high by interenational standards at about 0.4 and 0.5, respectively.

Pekkala and Lucas (2007) estimate intergenerational elasticities for Finnish cohorts born between 1930 and 1970, using census data on annual earnings for offspring and family income for parents. The intergenerational elasticities declined substantially; for sons from more than 0.30 to around 0.20, and for daughters from 0.25 to around 0.15 for cohorts born in 1930 to those born in 1950 and later. It may be of special interest to note that Pekkarinen et al. (2009) find comprehensive school reform, treated as a quasi-experiment, reduced the Finnish Beta by almost a third. The Norwegian trend studies have focused on the post-1950 cohorts. Bratberg et al. (2007) find a small decline in father–son and father–daughter elasticities from 1950 to 1965 cohorts. However, Hansen (2010) reports that this result does not hold when using the income of both parents. Instead, she finds a small increase in the elasticities for the 1955–70 cohorts. This difference suggests an increasing role for mothers, which has not been much explored in the literature. The Beta for Swedish father-son pairs is around 0.25 (see e.g. Björklund and Chadwick, 2003) but much higher at the top of the distribution (Björklund et al., 2012). Estimates from Denmark suggest quite low levels of persistence (e.g. Bonke et al., 2005)

Lefranc et al. (2013) estimate Betas for Japanese sons and daughters using two-step sample methods. The estimates for men are all quite close to 0.35. For daughters, estimates vary between 0.182 and 0.367. The evidence on whether or not the Betas increased for younger cohorts is mixed, at best. Ueda (2009) uses instrumental-variable techniques to estimate elasticities for men and women in Japan also, and find elasticities of around 0.411–0.458 for men, and 0.229–0.361 for women, depending on marital status and the use of family or individual income.

Non-linearities in the parent-child conditional income expectation were explored in a multi-country study by Bratsberg et al. (2007), who find the data for the USA, UK, Denmark, Norway and Finland all suggest the relationship is convex, with elasticities low at low levels of parental income, and increasing thereafter. At all quantiles of parental income, the elasticities are lower for the Nordic countries, than for the UK and the USA. Interpreted in terms of borrowing constraints on investments in child human capital, the results suggest capital market imperfections may be more of an issue not at the bottom more around the middle of the distribution of parental income.

Raaum et al. (2007) tackle another question in a multi-country study, namely how the mobility of daughters compares with that of sons across countries. Drawing on Chadwick and Solon (2002) and Björklund and Chadwick (2003), they find women's intergenerational income persistence is very similar across countries relying only on individual earnings. When family earnings are used for both men and women, the country ordering of intergenerational persistence for women looks very much like that for men. Using a framework that involves the intergenerational transmission of human capital endowments, assortative mating, and labour supply that responds to both own and spouse's wage, they infer that female labour supply is likely more (negatively) responsive to husband's earnings in the UK and especially the USA than in the Nordic countries.

We proceed to compare transition matrices across countries. To illustrate, consider the decile group transition matrices for the USA and Canada shown in Table 5 and derived from Mazumder (2005a) and Corak and Heisz (1999). Using the dominance approach discussed in Section 3.2, we can cumulate the transition matrices and take the USA-Canada difference. This leads to the results shown in Table 7. The vast majority of the cell entries are positive, suggesting Canada dominates the USA. However, given the two negative entries in cells (10, 1) and (10, 9), this result does not hold, strictly speaking.[77]

[77] We have not forced the rows or columns of either transition matrix to sum to 1, as they should

Table 7 Cumulated differences in intergenerational mobility tables across earnings decile groups for father-son pairs in Canada and the USA (USA-CAN)

	Son									
	1	2	3	4	5	6	7	8	9	10
Father										
1	6	10	9	8	9	11	8	6	1	1
2	2	9	11	13	12	14	11	6	3	2
3	1	6	8	16	18	18	15	8	4	2
4	8	11	13	21	16	20	15	10	4	2
5	10	12	15	19	17	19	15	9	7	4
6	9	12	11	15	14	19	17	11	5	4
7	8	9	12	15	15	18	22	18	10	3
8	8	9	11	17	17	17	21	21	13	2
9	4	5	7	9	10	12	12	9	9	2
10	−1	0	0	2	2	0	0	2	−1	0

Note: Cell entries are in percent. See notes to Table 5.
Source: Authors' derivations using transition matrices shown in Table 5 from Mazumder (2005a) and Corak and Heisz (1999).

Recall from Figure 16 that between the Betas and *r*s for disposable income, the USA and the UK were re-ranked, while Germany was least persistent in both. In Table 8, we illustrate again the use of the dominance approach, this time using quintile group transition matrices also from Eberharter (2013). The differences in the cumulated transition matrices suggest the Germany dominates both the UK and USA (all entries in the US–Germany and UK–Germany matrix are positive), but that the USA and UK cannot be ordered. Note, however, that there is only one strictly positive entry in cell (3,2), indicating the USA is close to dominating the UK.

5.4. Evidence on sibling correlations

In this subsection, we show evidence on sibling correlations, and relate them to intergenerational correlations. Why are sibling correlations of interest in the study of intergenerational income mobility? One way to motivate the interest in intergenerational mobility is to argue that it is related to equality of opportunity (see Section 2). A society in which a person's position is heavily dependent on the family he/she is born into is one where there is likely to be less equality of opportunity than one in which intergenerational persistence is very low.[78] But if we would like to understand how important family background is for the distribution of economic status, a focus on parent-child association captures only one part of the association. A fuller (but still incomplete) accounting of the importance of family background can be done by comparing the economic status of siblings. It

in a bistochastic matrix. The US matrix in particular fails this condition (most likely due to a smaller sample size and rounding error). If we do force the rows to sum to 1, the negative entries vanish and we have dominance.

[78]As we argued earlier, and to underline a point made repeatedly in the literature, the link between intergenerational mobility and equality of opportunity is far from straightforward.

Table 8 Cumulated differences in intergenerational transition matrices in disposable income among all persons for Germany, the UK and the USA

A. USA – Germany						B. USA – UK						C. UK – Germany					
	Offspring						**Offspring**						**Offspring**				
	1	2	3	4	5		1	2	3	4	5		1	2	3	4	5
Father						**Father**						**Father**					
1	3	5	5	1	0	1	−10	−1	−1	0	0	1	14	6	7	2	0
2	9	11	4	2	0	2	−11	−5	−2	−6	0	2	20	16	6	8	0
3	9	18	6	2	0	3	−11	1	−4	−9	0	3	20	18	11	11	0
4	9	18	9	9	0	4	−8	−3	−12	−10	−1	4	17	20	21	19	1
5	4	13	1	2	0	5	−10	−11	−21	−20	−1	5	15	24	22	23	1

Note: Cell entries are in percent. See notes to Figure 16.
Source: Authors' calculations from Eberharter (2013, Table 3).

turns out that the *sibling correlation* can be thought of as an R^2 of family background, capturing the importance of factors that siblings share in (most often) the variance of log income or earnings. While part of what siblings share is parental income, a large part is not. That is why sibling correlations are useful in assessing the importance of family background in the distribution of economic status.

To clarify the interpretation of a sibling correlation, we follow the exposition of Solon et al. (1991). Suppose that we observe annual income, assumed to equal long-run income plus transitory errors, assumed to be classical. The natural logarithm of income in year t, y_{ijt}, for sibling j in family i, for brevity, assumed to be measured as deviations from the population average, is modelled as

$$y_{ijt} = a_i + b_{ij} + v_{ijt}, \tag{24}$$

where a_i is a permanent component common to all siblings in family i, and b_{ij} is a permanent component unique to individual j, which captures individual deviations from the family component. The error term v_{ijt} picks up deviations of annual income from long-run income. The family and individual components are orthogonal by construction, so the long-run income variance is the sum of the family and individual component variances, $\sigma_a^2 + \sigma_b^2$. The share of the variance of long-run income that can be attributed to family background is

$$\rho = \frac{\sigma_a^2}{\sigma_a^2 + \sigma_b^2}. \tag{25}$$

This share coincides with the Pearson correlation in long-run income of randomly drawn pairs of brothers, which is why ρ is called a sibling correlation. Note that, as the conceptual model underlying the sibling correlation is defined in terms of variances, it can only vary between 0 and 1, i.e. negative correlations are ruled out.

A sibling correlation can be thought of as an omnibus measure of the importance of family and community effects. It includes anything shared by siblings – parental income and parental influences such as aspirations and cultural inheritance, as well as neighborhood influences such as from school, church and peers. Genetic traits not shared by siblings, differential treatment of siblings, time-dependent changes in neighborhoods and so on are captured by the individual component b_{ij}. The more important the effects that brothers share, the larger is the brother correlation.

Part of what siblings share in a is parental income. A useful analytical insight is that (assuming for ease of exposition marginal distributions are in 'steady state') the brother correlation in income can be thought of as the sum of the intergenerational income correlation squared and the correlation of other factors siblings share but that are orthogonal to income:

$$\rho = r^2 + \text{correlation of other shared factors.} \tag{26}$$

When the steady-state assumption is untrue, the first part of the sum on the right-hand side of equation 26 also involves the marginal distributions of income in the two generations. This decomposition allows us to apportion the overall importance of family background, as captured by ρ, onto that part accounted for by intergenerational persistence, measured by Beta or r, and the other factors siblings share that affect income.

Evidence about sibling correlations in earnings (and income) was surveyed by Solon (1999), Björklund and Jäntti (2009), and also by Schnitzlein (2013), who provides new estimates for Denmark, Germany, and the USA. We show in Table 9 evidence based mostly but not only on long-run earnings for several countries. The evidence is based on three main methods for estimating the variance components that constitute the sibling correlation: (unbalanced) ANOVA, restricted maximum likelihood estimates (REML), and generalized method of moments (GMM). As Björklund et al. (2009) report, whether or not the transitory errors are allowed to be autocorrelated has a big impact on the estimated sibling correlations. Allowing errors to be autocorrelated tends to reduce the individual variance, so increasing the estimated sibling correlation, so cross-country comparisons should be made across similarly defined models.

Although there are multiple estimates for several of the countries, we have sibling correlations in earnings or income for no more than seven countries for brothers, and six for sisters. The estimates for Nordic countries are low (and by far the lowest for Norway), highest for China, and of similar magnitude in Germany and the USA. For men, 43 per cent to 49 per cent of the variance in long-run earnings in Germany and the USA is accounted for by family background. This compares to 14 per cent in Norway, and 20 per cent to 25 per cent in the other Nordic countries. The ordering is similar but levels for women are lower across the board. Family background accounts for 30 per cent to 39 per cent of long-run earnings in Germany and the USA and between 11 per cent and 23 per cent in the Nordic countries.

The sibling correlation is a ratio of the variance of the family component in income to the variance of long-run income. In the spirit of the 'Great Gatsby' curve, shown in Figure 13, it is of interest to compare now another measure of persistence, the sibling correlation, with another measure of cross-sectional inequality, namely that of permanent earnings or income. We plot in Figure 17 the brother and sister correlations against the standard deviation of (the natural logarithm of) permanent earnings/income for those case listed in Table 9 where we have been

Table 9 Sibling correlations in earnings and income

Brothers					Sisters				
Denmark	0.23	1951–1968	ANOVA	Björklund et al. (2002)	Denmark	0.19	1958-1971	REML	Schnitzlein (2013)
Denmark	0.20	1958-1971	REML	Schnitzlein (2013)	Finland	0.13	1950–1960	ANOVA	Österbacka (2001)
China	0.57	Not reported	REML	Eriksson and Zhang (2012)	Finland	0.11	1955–1965	ANOVA	Björklund et al. (2004)
Finland	0.26	1953–1965	ANOVA	Björklund et al. (2002)	Germany	0.39	1958-1971	REML	Schnitzlein (2013)
Finland	0.26	1950–1960	ANOVA	Österbacka (2001)	Sweden	0.15	1951–1968	ANOVA	Björklund et al. (2004)
Finland	0.24	1955–1965	ANOVA	Björklund et al. (2004)	Sweden	0.23	1953	REML	Björklund et al. (2010)
Germany	0.43	1958-1971	REML	Schnitzlein (2013)	Norway	0.12	1953–1969	ANOVA	Björklund et al. (2004)
Norway	0.14	1950–1970	ANOVA	Björklund et al. (2002)	USA	0.34	1947–1955	REML	Mazumder (2008)
Norway	0.14	1953–1969	ANOVA	Björklund et al. (2004)	USA	0.28	1951–1958	ANOVA	Solon et al. (1991)
Sweden	0.37	1962–1968	GMM	Björklund et al. (2009)	USA	0.29	1958-1971	REML	Schnitzlein (2013)
Sweden	0.25	1953	REML	Björklund et al. (2010)					
Sweden	0.25	1948–1965	ANOVA	Björklund et al. (2002)					
Sweden	0.22	1962–1968	REML	Björklund et al. (2007a)					
Sweden	0.19	1951–1968	ANOVA	Björklund et al. (2004)					
USA	0.49	1947–1955	REML	Mazumder (2008)					
USA	0.45	1944–1952	REML	Levine and Mazumder (2007)					
USA	0.45	1951–1958	ANOVA	Solon et al. (1991)					
USA	0.43	1951–1967	ANOVA	Björklund et al. (2002)					
USA	0.45	1958-1971	REML	Schnitzlein (2013)					

Note: Estimates are all based on multi-year averages of earnings or income, adjusted for stage in lifecycle. We have relied in part on the compilation of evidence in Schnitzlein (2013) in constructing this table. Source: Schnitzlein (2013) and authors' compilation from sources listed in last column.

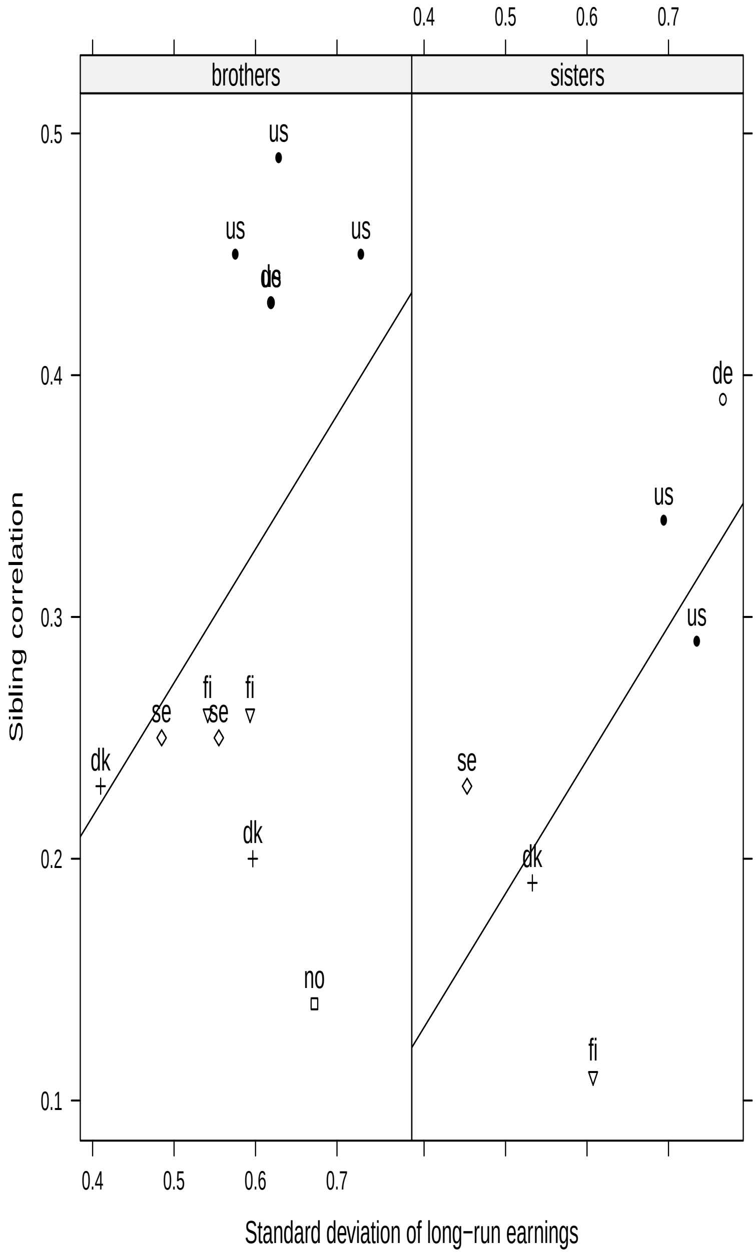
able to find all variance components.[79] In each panel, we have drawn the least squares regression line.

Due to the small number of countries included in the graph, it should be interpreted cautiously, but some insights can be gained. Among men, the estimated levels of permanent income inequality are consistent with very different degrees to which family background accounts for long-run earnings. Finland, Denmark and Norway have a standard deviation of log permanent earnings on either side of 0.4, as do Germany and the USA but, in the former group of countries, brother correlations are between 0.14 and 0.25. In the latter group, they are 0.43 and as large as 0.49. The regression line for men has all negative deviations for low brother correlations and all positive ones for high correlations, suggesting the least squares line gives a poor fit. Indeed, if we look at the two 'clusters' in each panel – the Nordic countries as one and Germany and the USA as the other – one conclusion may be simply that the Nordic countries differ form the USA and Germany. Thus, although the least squares line in both panels has a positive slope, it may be premature to talk of a 'Great Gatsby' curve for sibling correlations.

There is some evidence about changes across time in brother correlations both in the USA and Sweden. Levine and Mazumder (2007) examine brother correlations in earnings, family income, and hourly wages for two sets of cohorts: those born in 1942–1952 and those born in 1957–1965. The brother correlations in earnings increase from 0.263 to 0.452, in family income from 0.207 to 0.415, and those in hourly wages from 0.277 to 0.472. In no case is the change statistically significant at the conventional 5% level but, taken together, the estimates suggest

[79]Figure 13 has a single point for each country whereas we have included repeated observations for a country in Figure 17 for some cases.

Figure 17 Sibling correlation and long-run earnings inequality



Note: We have plotted on the horizontal axis the sum of the family and invividual components, which captures the variance of long-run earnings or income. The vertical axis shows the level of the estimated sibling correlation. Also shown in each panel is the least-squares regression line.
Source: See Table 9.

the importance of family background may have increased quite substantially. By contrast, Björklund et al. (2009) study change in brother correlations in Sweden starting with cohorts born 1932–38 and ending in 1962–1968, and find a decline in the importance of family background in the long-run income of men of roughly 13 percentage points. Although the authors are unable to pinpoint the reason for the decline, it coincides with the development of various welfare-state institutions.

We close by noting that, as with intergenerational associations, research on sibling associations should in the future provide more estimates for us to be able to draw robust conclusions about the importance of family background. Apart from the obvious question of why it is that siblings are so similar (what is it that families do?), we would like to see sibling correlations estimated (using the same methods and definitions) for a much wider group of countries than those seven for which we now have information.

We would also like to see rank correlations, not only Pearson correlations, to allow for a full standardization of the marginal distribution when comparing across time, countries, as well as estimates for both women as well as men. A minor point in that regard concerns estimation. Most of the estimates of sibling correlations in Table 9 rely on either unbalanced ANOVA or REML to estimate the variance components. While REML estimates could in principle be defined for data that follow an arbitrary distribution, in practice the likelihood is that of a normal distribution, as a and b are both modelled as conditionally normally distributed variables. While this may produce reasonably accurate estimates for the log of earnings or income, it is unlikely that REML would produce good estimates if applied to ranks, which are uniformly distributed by definition. Thus, the most feasible way of estimating sibling rank correlations would be to work with pairs of siblings rather than multi-level models.[80]

5.5. Other approaches to intergenerational mobility

In this subsection, we discuss three other approaches to intergenerational mobility. Two, based on occupation and on analysis of surnames, have recently been used to study very long-run trends in intergenerational mobility for which income information is not available. The third concerns an emerging literature on intergenerational links across more than two generations.

Economists have much to learn from sociologists when it comes to the study of intergenerational mobility. The study of the transmission of socio-economic advantage from generation to generation is one of the core issues in sociology. Empirical research has taken place for almost a hundred years and the theoretical discussion is also rich. Not surprisingly, the available data, the statistical techniques as well as the possibility to handle large data sets with statistical techniques have improved markedly in the last couple of decades. Hence, the prospects for comparative research based on reasonably comparable data have improved. Nonetheless, comparability is a major concern in the literature that we have come across.

One can distinguish between two strands of intergenerational research in modern sociology.[81] One of them focuses upon the relationship between status or prestige attainment of two generations, in general fathers and sons. Occupation is used as the basis to define status and alternative scales that attach status levels to occupations have been suggested in this literature. For example, the famous

[80]Note that the GMM-based estimates for Sweden reported in Table 9 also rely on pairs of brothers.

[81]Ganzeboom et al. (1991) provide an informative survey of this literature.

Duncan status index (Duncan, 1961) used the average education and income of each occupational category. Treiman (1977) has constructed prestige scales from survey data on the average prestige that people attach to various occupations.

The other strand of research defines socioeconomic status in terms of social class but emphasizes that social classes are intrinsically discrete and unordered. Hence, the analytical task is to measure mobility between these classes. The pros and cons of these two approaches to intergenerational mobility have been subject to a more than lively discussion within the sociological research community. Both approaches are prevalent and each has strong support.[82] The sociological literature on social mobility is far too vast to be reviewed here. One milestone is the monumental book by Erikson and Goldthorpe (1992b), discussed e.g. in Erikson and Goldthorpe (1992a), Hout and Hauser (1992) and Sorensen (1992). We note that this is a highly mature field that has generated enormous insight into intergenerational mobility.

Indeed, in order to study long-run changes in intergenerational mobility, class mobility may be the only option. Using census data on with names and occupational information, Ferrie (2005) and Long and Ferrie (2007, 2013a) identify father-son pairs by tracking the son of a given father in a later census in the USA and the UK. Ferrie (2005) studies long-run trends in occupational mobility in the United States, and Long and Ferrie (2007, 2013a) compare long-run trends in the USA and the UK. They find that the USA was more fluid in the late 19th century than either the UK or the USA in the 3rd quarter of the 20th century, a finding for which changes in agricultural occupations is central. Their paper generated

[82]For discussions see e.g. Ganzeboom et al. (1992, 3–7), Erikson and Goldthorpe (1992a), Hout and Hauser (1992), and Sorensen (1992).

two critical comments by prominent sociologists, Xie and Killewald (2013) and Hout and Guest (2013), to which they replied (Long and Ferrie, 2013b). Taken together, these papers provide a useful introduction to the use of historical census data to study intergenerational mobility across long periods of time.

Another emerging strand of literature relies on the fact that surnames convey information on social status. Gregory Clark and collaborators have researched social mobility using data about surnames in Sweden, the USA, England, Japan, India, and China.[83] Güell et al. (2007)and Collado et al. (2012) examine intergenerational mobility in Spain using surnames. This approach has great promise, but it would be more convincing if it could be validated using data that contain either occupation or income so mobility using names could be compared with other, more traditional methods.[84]

Finally, there are a handful of papers that examine intergenerational persistence across more than just two generations. The multigenerational view is lucidly discussed by Mare (2011). Income persistence across multiple generations are estimated at least by Marchon (2008) and Lindahl et al. (2012). In both of those papers, both the parents' and grandparents' income affects offspring income, suggesting that the simple 'AR(1)' model of intergenerational transmission is incomplete. These papers provide a perspective that most often goes unremarked on in the intergenerational mobility literature, namely that it relies on a 'dynastic' view of parent-child associations. Once grandparents are included in the analysis, care must be taken to distinguish between maternal and paternal grandparents.

[83] See e.g. Clark (2012), Clark (2010), Clark and Cummins (2012), Clark and Ishii (2012), Clark and Landes (2012), Clark et al. (2012), and Hao and Clark (2012).

[84] The UK and US data used by Long and Ferrie (2013a) would be ideal for validating the use of names because the father-son link was initially established using names.

5.6. Summary and conclusions

The large literature on intergenerational income mobility that has been surveyed in this section suggests that incomes are, indeed, persistent across generations. What has been learnt?

The main lesson is that differences in data (the three 'W's discussed in Section 5.1) may account for many of the differences in estimates. Put another way, because of the impact of the combination of lifecycle effects and and transitory variation in both parent and offspring generations, combined with other data issues, we know surprisingly little either about how income persistence varies across countries, or how it changes within countries over time. We also know very little about exchange mobility (fully standardizing for differences in marginal distributions).

Thus, despite the public prominence of the 'Great Gatsby' curve, very little is known about how intergenerational income persistence and mobility vary across countries and how this relates to cross-sectional inequality. More research, using comparable data for multiple countries across multiple cohorts of parents and offspring, is required. With a set of stylized facts about mobility differences and trends, we can then set out to try and explain them.

6. Conclusions

This chapter shows that substantial progress has been made in the analysis of income mobility over the last few decades, much of which has been stimulated by the increasing availability of suitable longitudinal data. For within-generation analysis, new household panel and administrative record data abound by comparison with the situation described by Atkinson et al. (1992). For between-generation analysis, the number of suitable data sets has also increased substantially, though not to the same extent (for obvious reasons), and issues of data quality remain relatively more important. Put another way, there has been a more general increase in the availability of good quality intragenerational income data sets across a relatively large number of rich countries. Good quality data for analysis of intergenerational income mobility is concentrated among a smaller number of countries. Most longitudinal data (in either context) refers to rich industrialized nations, and it would be interesting to examine the extent to which the patterns found also extend to middle- and low-income countries.

Although the availability of good data has increased substantially, many substantive issues of interest are not yet resolved. Our discussion of within-generation mobility revealed few clear cut conclusions about whether mobility has been increasing over time or decreasing in particular and whether mobility is greater in one country rather than another. The same can be said in regard to the evidence about income mobility between generations. In short, there remains much scope for systematic empirical analysis.

We have also shown that there has been a substantial increase in the number of mobility measures per se, but the literature has not yet matured in the same way as the measurement of (cross-sectional) income inequality has. Relatively underdeveloped are measures of individual income growth and, especially, of income risk. We would like to see empirical researchers making greater use of the descriptive methods that we have outlined – in order to show the data 'as they are' as far as possible – while also carefully selecting summary measures that reflect the mobility concept that is of particular interest. In the intergenerational mobility context, for instance, we have recommended greater use of measures of positional change and less reliance on Beta. More generally, transition matrices are under-used.

Our discussion of income mobility has focussed on mobility between two time points (with the exception of the discussion of mobility as longer-term inequality reduction). This simplifies the measurement task substantially, but does not remove the need for development of methods for describing individual income *trajectories* over multiple periods. In the intergenerational context, the interest is in not simply the similarities or differences between parents' and children's income, but also the prevalence of 'rags to riches and back in three generations' trajectories (for example) relative to other patterns. In the intragenerational context, we are interested not simply in each person's total lifetime income, but also the patterns of variation over calendar time and age, and how these patterns differ across individuals.

With multi-dimensional (multi-period) data, the natural reaction of most analysts is to fit models, with a small number of parameters summarizing the key differences between trajectory patterns. In the Introduction, we briefly cited literatures about the modelling of incomes within or between generations. One of the greatest challenges facing income mobility researchers is to develop tractable models of household income dynamics (not simply earnings dynamics for individuals) both within- and between-generations. Compared to the field of mobility measurement that we have reviewed in this chapter, mobility modelling is under-developed and deserves greater attention in future.

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