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Intergenerational mobility of socio-economic status in comparative perspective

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This paper reviews three strands of literature on socio-economic intergenerational mobility. The first is a mostly recent and rapidly growing economics literature that measures mobility in labour earnings and income. This approach is compared with two classical sociological approaches that measure the mobility in class and status. The United States seems to rank quite high in terms of class and status mobility, but low in terms of earnings and income mobility. This seemingly contradictory result can be accounted for by lower earnings mobility within occupations in the United States. JEL-codes: D1, D3, I62.

International comparisons of economic inequality suggest that the United States has the most unequal distribution of disposable income and earnings among OECD countries. Studies that use data from the Luxembourg Income Study, which has collected comparable microdata on incomes from many OECD countries, consistently rank the United States as having the most unequal distribution of disposable income, both in terms of relative inequality and of alternative definitions of relative poverty.[2] The same pattern emerges in studies that focus on hourly earnings or annual earnings of full-time workers.[3] Moreover, it seems that especially the lower half of the U.S. earnings distribution has particularly high dispersion.

Such findings, however, do not convince everyone that the United States is a more unequal society than other advanced countries. One common objection is that the time period over which incomes are measured – in

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2. A prominent example of such international comparisons is Atkinson et al. (1995).
3. Recent studies include Freeman & Katz (1995) and OECD (1996).

general one year – is too short. If U.S. workers' incomes vary relatively more over time, lifetime income is not necessarily more unequally distributed than in other countries. The high inequality of annual income and earnings in the United States would then be the consequence of high transitory variance and therefore an exceptionally high *intra*-generational mobility of income. Such transitory variance could in turn be a consequence of a highly competitive and "dynamic" economy. And, goes the story, it is the inequality of life-time, or permanent, income that is a matter of public concern, rather than short-term transitory fluctuations.

Comparative studies of intragenerational income and earnings mobility are – for obvious data-availability reasons – not as frequent as those that rely on annual data. However, thanks to the availability of new longitudinal data, a few comparative studies of intragenerational income mobility and income inequality over longer periods than a year have recently been done. These studies show, maybe to the surprise of some observers, that this type of mobility is not particularly high in the United States.[4] Long-run income is also relatively unequally distributed in the United States.

A second objection to the claim that the United States is an unequal society is that the country has been successful in achieving a more fundamental goal of equality, namely the equality of economic opportunity. Even though fundamental and frequently cited, this concept of equality is difficult to define in an operational way. One criterion for the equality of opportunity, which *can* be operationalized, is that the economic outcome of an individual is independent of her family background. The extent to which this holds can be gauged by studying the correlations of the economic outcomes of members of different generations of the same families – such as fathers and sons.[5] Low correlations would suggest high intergenerational mobility and a high degree of equality of opportunity. Those who advocate equality of opportunity would strive for a society in which economic outcomes are attributable only to individual effort and not to inherited status and wealth.

That such an "open society" is considered an ideal one among many Americans is obvious to anyone who has listened to the rhetoric of e.g. a U.S. presidential election campaign. Our impression is that quite many – both Americans and non-Americans – believe that the United States is and has been relatively successful in this respect.

The analysis of intergenerational mobility of economic and social status has a long tradition in sociology. The notion of high social mobility in the United States can also be found in the sociological literature. "American exceptionalism" is a commonly used expression of this notion. It goes far back in time to the writings of Tocqueville and Engels and Marx during the 19th century.[6] Engels and Marx and several of

4. Burkhauser & Poupore (1997) compare the United States and Germany and find quite similar magnitudes of earnings and income mobility in the two countries. Aaberge et al. (1998) compare the United States with Denmark, Norway and Sweden and find similar magnitudes of earnings and income mobility. OECD (1996, 1997) and Fritzell (1990) also present evidence on this issue. The *similarities* among countries are striking.
5. We are acutely aware of the fact that focusing on fathers and sons captures only a fraction of the possible intergenerational linkages. In this review paper, however, we decided to delimit comparisons to fathers and sons, because we think the comparability problems involved are slightly less serious than would be the case were we also considering mothers and daughters.
6. Erikson & Goldthorpe (1985) offer an introduction to the sociological discussion on American exceptionalism in social mobility.

their followers explained the weakness of organized labor in the United States by – a conjectured – high social mobility which prevented workers from becoming conscious of their class origins.

In this paper, we summarize and discuss the available empirical evidence on intergenerational mobility of economic and social status in a cross-country, comparative perspective. We think it is timely to write such a survey because there is a rapidly expanding new economics literature on intergenerational mobility of earnings and income that adds new insights to the existing sociological research on intergenerational mobility of social class and of social status attainment. We also believe that economists can learn from the rich tradition in sociology of studying social mobility.

We continue the paper in the next section with a survey of this new economics literature. The concepts of mobility that have been used in this research are defined and we summarize the available empirical results. The major result is that the United States (together with the United Kingdom) has the highest intergenerational earnings correlations – and hence lowest mobility – among the seven countries for which we have found empirical estimates. The following section offers a similar overview of the concepts and results from the sociological research on mobility of social class and of status attainment. In this literature the United States ranks quite highly in terms of social mobility (even though there is no ground for talking about "American exceptionalism"). These seemingly contradictory results lead us naturally to a methodological discussion of the approaches to the analysis of intergenerational mobility. We devote the following to discussing the differences between and the similarities of these strands of literature. In the last section we summarize and speculate about the future prospects of research on these issues.

The new economics literature on earnings and income mobility

Economists have shown some interest in intergenerational relationships for quite a long time. In the tradition of family economics, Becker & Tomes (1986) suggest a theoretical model of the intergenerational transmission of family status. What can be learned from such models over and above what is learned from 'mechanical' models of the same was called into question by Goldberger (1989), who argues there is little value added in at least the model of Becker & Tomes (1986). Recent theoretical contributions include Checchi et al. (1999), who examine the role of school financing in intergenerational mobility and Han & Mulligan (1997), who examine biases that result from misspecified population regression models (see also Mulligan1997). The earlier empirical literature was summarized by Becker & Tomes (1986), who showed results from nine studies covering five countries in which some measure of the income of sons was regressed on some measure of the income of fathers.

The interest in these questions was, however, markedly fueled by two papers that appeared in the same issue of the *American Economic Review* in 1992, by Gary Solon and David Zimmerman.[7] These papers made two major contributions. First, they pointed out some of the statistical problems involved in estimating the relationship between "long-

7. Altonji & Dunn (1991) was published at about the same time and contained similar – and in some respects additional – analyses but did not receive as much attention.

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 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 correlations between father's and son's long run incomes as high as 0.4 or 0.5, numbers which are much higher than those in the previous studies surveyed by Becker & Tomes (1986). Solon and Zimmerman obtained similar results using two different data sets, which lends additional credibility to their findings.

The parameters of interest

The aim in these studies has been to estimate the relationship between the long-run economic status of fathers and sons (or some other combination of representatives of two different generations). More specifically, let y_{si} and y_{fi} be the permanent components of the log of annual incomes of sons and fathers, respectively. The intergenerational relationship is specified as:

$$y_{si} = \alpha + \beta y_{fi} + \varepsilon_i \qquad (1)$$

The coefficient β is the elasticity of the son's income with respect to the father's income and is one candidate summary measure of intergenerational (im)mobility. It is straightforward to extend (1) by adding a quadratic term of y_{fi}, or otherwise specify the functional form in a more general way. For example, the impact of father's income might be higher in one part of the distribution than in another. If fathers' and sons' income have equal variances, β is also the intergenerational correlation coefficient, which we will denote by ρ. When the variances are unequal, the correlation coefficient is obtained by multiplying β by the ratio of the standard deviations of fathers' and sons' income.

The challenge is to estimate relationships like equation 1 from actual data. Income data of both fathers and sons are needed. Retrospective questions to adult sons about the incomes of their fathers a long way back in time would not provide reliable information.[8] Instead, Solon and Zimmerman used U.S. longitudinal data sets in which two generations of the same family had been interviewed in adult age. Further, in these data sets the fathers had been followed over a longer period so that long-run measures, rather than noisy short-run measures, of fathers incomes could be used. Solon and Zimmerman offered a variety of methods to estimate β (or ρ) from such data sets. Perhaps most importantly, the time-averaging of the explanatory variable in equation 1 reduces the errors-in-variables bias that measurement errors cause and thus yield an estimate of β that is closer to its true value.[9]

A transition matrix provides an alternative way to depict intergenerational mobility, and a number of summary mobility measures can be computed from the information in such a matrix (see e.g. Checchi et al.1999). The information in the matrix would also be

8. It is more likely that such questions yield reliable information about the occupation of the father in the period when the son grew up.
9. See the Appendix and Solon (1992).

able to tell us more about the kind and direction of mobility that is occurring. A problem with such an approach is that it requires long-run incomes of both sons and fathers.[10]

The evidence

In Table 1 we show intergenerational correlations of long-run incomes (and earnings) of fathers and sons that have been estimated in 15 studies covering 7 countries.[11] Solon identified 348 pairs of fathers and sons in the Panel Study of Income Dynamics (PSID) and used several years of earnings and income data of the fathers. A technique that used fathers' average earnings over five years (the TA-technique) gave an estimate of 0.41 for the intergenerational earnings correlation. Solon demonstrated that this TA-technique yields a downwards-biased estimate of ρ, but that the bias is smaller than if using single-year fathers' earnings. He further showed that under plausible assumptions, using education as an instrument most likely produces upwards-biased estimates of the same parameter. In the instrumental variable (IV) case, the upwards-biased estimate is 0.53, not too far above 0.41 for the downwards-biased TA estimate. The estimates for hourly wages and family incomes are of similar magnitudes.

Zimmerman used the National Longitudinal Survey (NLS), which also follows the children of families when they become adults. In addition to time averaging, he applied a few different techniques to purge fathers' earnings of its transitory component. Basically, he obtained estimates that are similar in magnitudes to Solon's.

Altonji & Dunn (1991) also employed the NLS. Their estimates are in the lower range of those obtained by Solon and Zimmerman. Taken together, these three studies suggest that the magnitude of the correlations are in the range 0.35-050, i.e. much higher than previously believed; Becker & Tomes (1986) concluded on the basis of earlier studies that the correlation does not exceed .20. Recent, so far unpublished, studies have looked more closely into the relationship between fathers' and sons' earnings. Reville (1995) started out by replicating the findings of Solon using 1984 as the year of sons' earnings. His more detailed analysis suggested, firstly, that the correlations rise markedly with the age at which sons' earnings are recorded,[12] and, secondly, that there seem to be period effects in the correlations for given age groups. Buron (1994) uses several years of both fathers' and sons' earnings and uses alternative methods to adjust for life-cycle effects. The estimate in his work that is most easily comparable to the other U.S. studies we cite is 0.39. Recent work by Lillard & Kilburn (1996) and Lillard & Reville (1996) suggests that, once the time-series structure of the transitory component of annual earnings is properly handled, the intergenerational correlation may be even higher than the numbers we show here.[13]

10. In estimating a setup such as that in equation (1), long-run incomes are required only for fathers, not necessarily for sons, as measurement errors in the independent variable bias the estimate, but measurement errors in the dependent variable are subsumed in the residual variance.
11. Results for other combinations of family members can be found in Altonji & Dunn (1991), Couch & Dunn (1997) and Dearden et al. (1996).
12. This result is also clear in Swedish data, see Björklund & Jäntti (1993).
13. The measurement errors ε that enter the measurement model for permanent earnings, $y_{jt} = y_j + \varepsilon_{jt}$, may themselves have an intertemporal covariance structure rather than being white noise. Efforts to take this into account seem to yield higher correlations than when these are assumed to be white noise.

Eide & Showalter (1999) study if the son's income elasticity wrt. to father's income varies across the distribution of son's income. They examine both the mean of son's earnings conditional on father's income (and, in some analyses, other explanatory variables) and the 10th, 50th and 90th percentile of son's income, conditional on father's. They use a specification similar to that in Solon (1992), except that they use three as opposed to five years of father's income and they include both the Survey of Economic Opportunity (SEO) and Survey Research Center (SRC) subsamples of the PSID.[14] The estimated coefficients in the OLS regressions are fairly similar (although somewhat lower in magnitude) than the results found in Solon (1992). While they find the coefficients in their quantile regressions to vary – the coefficient on father's income is highest at the 10th quantile – it would appear to be more interesting to systematically study whether the coefficient on father's income varies depending on where in the distribution of *father's* income a son comes from.

A further non-linearity concerns the "impact" of fathers with no earnings. Couch & Lillard (1998) argue that excluding fathers with missing five-year average earnings leads to upward-biased estimates of the correlation. Imputing one dollar to these fathers and including them leads to a large revision downward of the estimated coefficient. As a treatment of missing values, however, the imputation of a very low amount (ln 1 = 0) strikes us as daring, especially as least squares estimates are known to be sensitive to outliers.[15] A dollar a year does not strike us as a very convincing estimate of true economic status.

Many theoretical contributions to the economics literature view the distribution of (permanent) consumption as the more welfare relevant. Earnings are taken to be a worse but more readily available and therefore more often used measure of long-run income. Aughinbaugh (1996) estimates the correlation in the consumption of fathers and sons and finds this correlation to be even higher, on the order of .7.[16]

Interestingly, Bowles & Nelson (1974) studied the role of the inheritance of IQ in accounting for the intergenerational transmission of economic status. Their methods do not easily lend themselves to the framework we have used in constructing Table 1, since their measure of socio-economic background is a linear combination of parent's income, father's schooling and father's socio-economic status. However, their estimated correlations between that measure and son's income for different cohorts of men range from .324 to .464, fairly close to the later US estimates, using father's long-run earnings, listed in Table 1.

The rest of the table contains similar estimations for other countries. Most of these studies – except the early contribution by Atkinson (1981) – have been stimulated by the contributions by Solon and Zimmerman. Two of the studies are explicit comparisons between the United States and another country, using the same age limits for sons and fathers, the same methods, and as identical definitions of earnings (or other outcome measure) as possible. Couch & Dunn (1997) compare Germany and the United States

14. The PSID consists of two subsamples. The SEO subsample oversampled low-income households and the SRC again is a nationally representative cluster sample. See Hill (1993) for details.
15. See Solon (1998) for discussion.
16. See Han & Mulligan (1997) for alternative theoretical models that yield (partially) testable predictions about the magnitudes of the coefficients on consumption and earnings.

using the German Socio-economic Panel and the PSID. In order to achieve comparability between the two countries, they had to measure the outcome of sons at a younger age than Solon and Zimmerman. Therefore, it is not surprising that the level of their U.S. correlations is lower than the ones obtained by the former authors. Their German correlations are quite close in magnitude to the U.S. ones. Most German ones are lower, but some are higher. Thus, a conservative reading of their evidence suggests the correlations are more or less the same.[17]

The evidence from Germany is mixed, however. Wiegand (1997) uses more recent data from the GSOEP, which allows him to use a similar age-range for the sons as that in Solon. He also argues that the measure of annual earnings in Couch & Dunn is inferior to gross monthly earnings. His evidence suggests that the correlation is lower in Germany.

Neither Couch & Dunn nor Wiegand present formal statistical tests for the difference in correlations between the two countries. The standard errors of the regression coefficients suggest that, at conventional levels of significance, the estimates are not different. Wiegand, however, uses the IV method to obtain an upper bound on the German correlation. This upper bound is lower than the lower bounds presented by Solon, which adds to the evidence that suggests the U.S. is less mobile.

The other comparative study we are aware of compares Sweden with the United States (Björklund & Jäntti, 1993, 1997). Using identical methods and age limits for the two countries, we find higher correlations for the United States. However, a one-sided test of the difference only yields *p*-values around 0.20, so strong conclusions cannot be drawn. The rather low correlations for Sweden are lent some additional credibility by another study by Gustafsson (1994), who gives similar estimates although the data are not representative of the whole population.

The ordering of countries is our focus here, which puts the issue of excluding fathers with zero earnings in a slightly different perspective (see above). In Björklund & Jäntti (1997), we estimated the correlations using also unlogged incomes (and thus did not exclude zero-earnings fathers or sons). The result, that we could not statistically distinguish the Swedish and U.S. correlations remained intact and the point estimates still suggested lower mobility in the United States.

Among the other countries for which there are estimated correlations, two independent studies for the United Kingdom suggest similar magnitudes of intergenerational earnings mobility as in the United States. The study for Malaysia also reveals correlations close to the ones obtained for the United States. However, the estimates of correlations for Canada and Finland are lower than in the United States.

The overall impression of these studies is that the United States (and United Kingdom) tend to have quite high correlations and hence low mobility compared to other countries. The estimates are not in all cases very precise, so caution is called for in interpreting the differences.

The approaches in sociology

The study of the transmission of socio-eco-

17. It seems to have become conventional to present a "pyramid" of correlations, corresponding to different combinations of years over which father's earnings are averaged. If there are T=5 years of father's earnings, the "pyramid" contains $1/2\,T\,(T+1)$=15 correlations.

Table 1
Intergenerational correlations of long-run income or earnings between fathers and sons

Authors	Country and Data set	Measures of income	Sample size	Age of sons	Estimate of ρ	Comments
Solon (1992)	United States, PSID	1. Annual earnings 2. Hourly wage 3. Family income	348 father-son pairs	25–33	1: TA: 0.41 (0.09) 1: IV: 0.53 (0.14) 2: IV: 0.45 (0.10) 3: IV: 0.53 (0.12)	S. showed that the TA-technique provides a downwards biased estimate and the IV-technique an upwards bias.
Zimmerman (1992)	United States, NLS	1. Wage + salaries 2. Hourly wage 3. Duncan index of status	876 father-son pairs, but fewer in most estimations	29–39	1: TA: 0.54 (0.08) 1: MM: 0.41 (0.05) 2: TA: 0.39 (0.07) 2: MM: 0.38 (0.04) 3: TA: 0.33 (0.08)	The presented estimate are elasticities, which are close to the correlations that are reported in the paper. Z. also presented IV estimates that are close to those obtained by TA and MM techniques.
Altonji & Dunn (1991)	United States, NLS	1. Family income 2. Annual earnings 3. Hourly wage	675-735 father-son pairs	29–39	1: TA: 0.32 1: MM: 0.36 2: TA: 0.22 2: MM: 0.39 3: TA: 0.34 3: MM: 0.42	A & D argue that the MM-estimates are more reliable than the TA-estimates.
Buron (1994)	United States, PSID	Annual earning	253 father-son pairs	25–33	1: TA: .39 (.07) 2: GA: .47 (.07)	B. uses the over-time (1984–88) average of sons' ln earnings. The presented estimates are elasticities, but the correlations are very close.
Reville (1995)	United States, PSID	1. Annual earnings 1980 2. Annual earnings 1984 3. Annual earnings 1987	Around 300	25–38	1: TA: 0.26 2: TA: 0.45 (0.09) 3: TA: 0.32 (0.06)	
Eide & Showalter (1999)	United States, PSID	1. Annual earnings 1984 2. Annual earnings 1989 3. Annual earnings 1991	1459 2580 3635	25–41	1: TA: 0.33 (0.05) 2: TA: 0.36 (0.05) 3: TA: 0.34 (0.05)	Numbers shown are elasticities. E & D estimate elasticities in every year from 1984 to 1991. The lowest is .33 in 1984 and highest is .38 in 1985 and 86.
Atkinson (1981), Atkinson et al. (1983)	Britain, a sample from York	1. Weekly earnings 2. Hourly earnings	307 pairs of fathers and sons	Probably in their 40s.	1. 0.36 2. 0.43	Probably understated because of homogenous samples.
Dearden et al. (1996)	Britain, National Child Development Survey	Weekly wages	1565 pairs of fathers and sons	31	IV: 0.59 (0.07) TSIV: 0.39 (0.03)	The presented estimates are elasticities.
Corak & Heisz (1999)	Canada , register data	1. Annual earnings 2. Annual market income	≈ 350 000 pairs of fathers and sons	28–31	1. TA: 0.13 2. 0.19	The estimates are elasticities. Non-linearities implying greater mobility at the lower end of the distribution were found.
Jäntti & Österbacka (1995)	Finland, register data	Annual earnings	22 324 pairs of fathers and sons	Average age: 34.8	TA: 0.22	
Couch & Dunn (1997)	Germany, German Socio-Economic Panel	1. Annual earnings 2. Annual hours 3. Years of education	272 pairs of fathers and sons	Average age: 22.8	1: TA: 0.12 2: TA: 0.19 3: TA: 0.42	The sons are younger in this sample than in the other studies using US data which might explain why the earnings correlation is lower.
	United States, PSID	1. Annual earnings 2. Annual hours 3. Years of education	322 pairs of fathers and sons. Larger samples of education.	Average age: 24.9	1: TA: 0.17 2: TA: 0.17 3: TA: 0.24	
Wiegand (1997)	Germany, German Socio-Economic Panel	1. Monthly earnings 2. Annual earnings	1102 2130 pairs of fathers and sons	25–33	1: TA: 0.29 (0.07) IV: 0.36 (0.15) 2: TA: 0.10 (0.08) IV: 0.20 (0.27)	The presented estimates are elasticities.
Lillard & Kilburn (1995)	Malaysia, Family Life Survey	Annual earnings	343 pairs of fathers and sons	Around 25 on average	MM: 0.33–0.37	

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Authors	Country and Data set	Measures of income	Sample size	Age of sons	Estimate of ρ	Comments
Björklund & Jäntti (1997)	Sweden, Level of Living Surveys	1. Annual earnings 2. Market income (incl. income of capital)	400 sons, 500 fathers	29–38	1: TSIV: 0.23 (0.07) 2: TSIV: 0.29 (0.09)	The Swedish data set lacks information on fathers' age which is not controlled for in the estimations of neither the Swedish nor the US ρ. Adding such controls for the US raises the correlations by around 0.05.
	United States, PSID	Annual earnings	About the same as Solon	28–36	TSIV: 0.33 (0.10)	The method and data set differ slightly from the one used by Solon, but is the same for both countries.
Gustafsson (1994)	Sweden, 222 Stockholm boys.	Market income	185	25–41	OLS: 0.14 (0.07)	The estimates are elasticities. As emphasized by G., the estimates are downwards biased for two reasons: fathers' incomes are noisy because they only exist for one year, and the sample is homogenous. A correction for the former problem done by G. suggests that the coefficient should be in the range 0.17–0.20.

Note: Data sources: PSID: Panel Study of Income Dynamics. NLS: National Longitudinal Study. Econometric technique: TA: time averaging. MM: method of moment estimation. IV: instrumental variable technique. TSIV: two-sample IV (see Björklund & Jäntti (1997) and Dearden et al. (1996)). OLS: ordinary least squares using annual data for fathers. See Appendix.

nomic 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.[18] 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 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 by social class but emphasizes that social classes are intrinsically discrete and unordered. Hence, the analytical task is to measure mobility among 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. We will not take a stand in this discussion, but simply note that both approaches are prevalent and have strong positions in modern sociology.[19] We discuss each approach in turn.

Class mobility

The parameters of interest

The cross-national comparative perspective has a long tradition in the sociological research on class mobility. Leaving the early discussion by Weber and Marx aside, Lipset & Bendix (1959) presented the – much debated – hypothesis that mobility rates are the same in Western industrialized societies because of industrialization and the nuclear-family system. This hypothesis was formulated in terms of absolute mobility rates, i.e. deviations from the main diagonal in the class mobility matrix.

The traditional way to study class mobility tables, i.e., contingency tables with the origins – in this case, father's class – as one dimension and destinations – currently son's class – as the other was to compare the relative frequencies in each cell. These, however, depend on the marginal distributions of father-son pairs across origins and destinations. Structural shifts that occur across generations necessarily generate mobility, in the sense that it is arithmetically impossible for all observations to be found on the main diagonal. An important structural change in society, the timing of whose occurrence varies widely across countries, is the declining importance of the farming sector. Obviously, the sons of farmers, who thus have a common class origin, had to move to other

18. Ganzeboom et al. (1991) offer a most informative survey of this literature.
19. For discussions see e.g. Ganzeboom et al. (1992, pp. 3-7), Erikson & Goldthorpe (1992 a), Hout & Hauser (1992) and Sorensen (1992).

sectors of the economy, resulting in high rates of absolute mobility.

The response to this problem has been to focus on a concept of mobility, relative mobility, that is invariant with respect to such structural changes. In terms of the mobility matrix, mobility was to be compared in terms of a measure that was invariant to the marginal distributions of origins (fathers' classes) and destinations (sons' classes). Patterns of relative mobility can be estimated by fitting suitable log-linear models to the mobility tables (Andersen, 1996; Hout, 1983).

Relative mobility is defined in terms of log-odds ratios. The log-odds ratio in a 2×2 model is

$$1\mathrm{n}\,\frac{F_{11}\times F_{22}}{F_{12}\times F_{21}}, \qquad (2)$$

where F_{ij} is the frequency in cell (i,j) The collection of all 2×2 subtables define all combinations of odds ratios in a table. A model with k origin and destination classes has $(k^2-k)^2/4$ such odds ratios. These odds ratios embody 'the endogenous mobility regime' or 'the pattern of social fluidity' (see Erikson & Goldthorpe 1992*b*, p. 56). In the simple case of independence between origins and destinations, the log of each of these ratios equals zero.

The log-odds ratios can in turn be used to model the expected frequency F_{ij} in cell (i,j) of the mobility matrix:

$$\ln F_{ij} = \mu + \lambda_i^O + \lambda_j^D + \lambda_{ij}^{OD}, \qquad (3)$$

where μ is a scaling factor and the λs are parameters to be estimated from the grouped data that form the empirical mobility table. The parameter λ_i^O is an origin and λ_j^D a destination effect, and are proportional to the number of fathers in row i and sons in column j, respectively. The cell-interaction effects λ_{ij}^{OD} adjust the frequencies that are expected under independence to be higher or lower, depending on whether the (i,j) cell exhibits a higher or lower than expected frequency. Independence of origins and destinations implies that all cell-interaction effects λ_{ij}^{OD} in equation 3 are zero, i.e., cell frequencies depend on the marginal effects λ_i^O and λ_j^D only.

If *all* λ_{ij}^{OD} are estimated from the data, the model is said to be fully saturated. The trick in fitting log-linear models to mobility tables is to define and estimate models that give an acceptable fit to the empirical data without being saturated. Such specifications of equation 3 can involve, as we shall see below, inheritance effects (or immobility effects) within classes – implying less mobility than expected under independence, as well as particular patterns of higher or lower than expected mobility (under independence) across classes.

The evidence

The literature on class mobility is very rich indeed, both in volume and in complexity. In a review such as this, it is very difficult to adequately summarize the vast number of studies that have been carried out. One of the reasons such a summary is difficult is that the study of mobility tables rarely generates a single parameter or statistic that summarizes mobility in a particular country across two particular generations. Each study applies its own modeling strategy and the evidence of higher or lower mobility needs to be assessed within that particular strategy. There are good reasons for why, e.g., the study by Erikson & Goldthorpe is spread across more than 400 pages. We shall here in the main concentrate on the evidence in that book.

Cross-national comparability of the data

on classes has been a major concern in this literature. As the literature now stands, there are two major research projects that have recently devoted considerable resources to coding occupational information into class schema that are comparable across several countries. One of these programs is represented by the work by Erikson & Goldthorpe (1992*a*).[20] Under the auspices of the CASMIN (Comparative Analysis of Social Mobility in Industrial Nations) project, they have recoded occupational data for 12 countries. The resulting class information forms the basis of their empirical work. The output of the second project has been published by Ganzeboom et al. (1989). All in all, they have classified occupational data from 149 sources of information covering 35 countries.

These two projects start with a common conceptual class scheme, the EGP scheme, elaborated by Erikson et al. (1979). The EGP can be seen as a typology formed from four different job attributes: *sector* (non-manual, manual, farm workers), *employment* (self-employed vs. salaried), *skill level* and *supervisory status.* Erikson & Goldthorpe used the following collapsed seven-class version of the original scheme:

I + II Service class: professionals, administrators and managers; higher-grade technicians; supervisors of non-manual workers

III Routine non-manual workers: routine non-manual employees in administration and commerce; sales personnel; other rank-and-file service workers

IVa + b Petty bourgeoisie: small proprietors and artisans, etc., with and without employees

IVc Farmers: farmers and small holders and other self-empoyed workers in primary production

V + VI Skilled workers: lower-grade technicians; supervisors of manual workers; skilled manual workers

VIIa Non-skilled eorkers: semi- and unskilled manual workers (not in agriculture, etc.)

VIIb Agricultural labourers: agricultural and other workers in primary production

Ganzeboom et al. (1989) aggregate the last two classes into one and hence use six classes in their analysis.

The 7×7 national class mobility tables, using the above categories for samples of fathers and sons from different countries form the empirical basis of assessments of class mobility. We show first a measure of absolute mobility – the total mobility rate – which measures the percentage of men in each nation who are found off the main diagonal (Table 2). We note that Hungary, Australia, Sweden and the United States are in the high end of this ordering. Total mobility rates are of course affected by structural changes. We turn next to the evidence on relative mobility.

To study patterns of relative mobility, Erikson & Goldthorpe (1992 b) formulate a model of 'common social fluidity' (CmSF), defined on a collection of European mobility tables (adding a dimension, nation, to the log-linear model). Of the European nations they study (see Table 2) two, namely England and France, turn out to have the most 'central' fluidity patterns, as measured by the 'closeness' of the residuals of the fitted tables. The pattern of fluidity, as embodied in the predicted pattern of odds ratios from the

20. References to their earlier work can be found in Erikson & Goldthorpe (1992 b).

Table 2
Absolute mobility – total mobility rates in selected industrial nations

Nation	Total mobility rate
Australia	73
England	65
France	65
Federal Republic of Germany	62
Hungary	76
Ireland	58
Northern Ireland	63
Poland	60
Scotland	64
Sweden	73
United States	70

Note: The total mobility rate is the percentage of all men who are found off the main diagonal in the 7×7 mobility table.
Source: Erikson & Goldthorpe (1992 b), Tables 6.3, p. 195 and 9.4, p 330.

CmSF fitted to France and England, is taken as the data to which a 'core model of social fluidity' (CSF) is fitted.

The CSF model is of the general form

$$\ln F_{ij} = \mu + \lambda_i^O + \lambda_j^D + \lambda_{ij}^I, \qquad (4)$$

where, as before λ^O, λ^D and μ are the origin and destination main effects and the scaling factor, respectively. Rather than just empirically fit all of the interaction effects and fully saturate the model, they postulate a set of more specific terms, defined in terms of the whole mobility matrix, whose sum then gives the interaction effect λ^I, for each *(i,j)* cell. At this stage, macro-sociology enters the modelling, i.e., each cell interaction parameter is interpreted as being jointly composed of specific, sociologically-motivated associations between origin and destination class.

That is, the authors postulate that social fluidity patterns are subject to the relative desirability of different class positions, the relative advantage of each class origin and relative barriers individuals face in gaining access to different class positions. Erikson & Goldthorpe (1992*b* pp. 122-130) identify four classes of effects that are associated with the desirability, advantage and barriers of class positions, within which they identify specific patterns that affect the propensities for relative mobility: hierarchy (2 matrices of HI effects), inheritance (3 matrices, IN), sector (1 matrix, SE) and affinity (2 matrices, AF).

Each effect matrix consists of ones and twos, where cells that are theorized to be especially mobile (have positive coefficients) or immobile (have negative coefficients) are given a two whereas the base level has a one (corresponding to zero when logarithms are taken). For instance, the two hierarchy effects are obtained by noting that there are three hierarchical levels among the seven classes, with farmers (IVc), non-skilled workers (VIIa) and agricultural labourers (VIIb) being the lowest and the service class (I + II) being the highest. Movements from

one class to another are presumably more difficult if they entail crossing hierarchical levels and thus all such cells are "marked" in the matrix, allowing for these cells to have a lower propensity than that given by the marginals. The second hierarchy matrix again singles out those cells that involve movements across two hierarchical levels, presumably even more difficult than moving to the next higher level.

The inheritance effects capture the relatively high propensity to stay within one's origin class. The first matrix gives greater weight to all cells on the main diagonal, while the second and third matrices give a higher propensity to remain within the petty bourgeoisie and farming classes. The sector matrix in turn postulates that movements from the agricultural (IVc and VIIb) to non-agricultural classes are less likely than within-sector movements. Finally, the two affinity matrices capture additional forces that affect the likelihood of movements between particular origins and destinations, such as movements from agricultural classes to service classes and so on. The interaction effects resulting from these matrices are shown, in terms of the 7×7 mobility matrix, in Table 3. Depending on whether the postulated matrix captures barriers to or propensities for movements, the estimated parameters will be negative or positive, respectively.

The "core model of social fluidity" log-linear model shown in Table 3 can be written as

$$\ln F_{ij} = \mu + \lambda_i^O + \lambda_j^D + \lambda_{ij}^I \qquad (5)$$

$$= \mu + \lambda_i^O + \lambda_j^D + \lambda_{a(i,j)}^{HI1} + \lambda_{b(i,j)}^{HI2} + \lambda_{c(i,j)}^{IN1} + \lambda_{f(i,j)}^{IN2} + \lambda_{e(i,j)}^{IN3} + \lambda_{f(i,j)}^{SE} + \lambda_{g(i,j)}^{AF1} + \lambda_{h(i,j)}^{AF2}.$$

The log-linear model that is fitted to the data is thus not fully saturated (the cells filled with only a dash have no interaction effects) and the fitted cell interaction parameters have sociologically motivated interpretations in terms of the decomposition in terms of the four different forces, hierarchy, inheritance, affinity and sector that influence fluidity patterns.

This core model is found to fit data from each of the nations adequately when estimated jointly, although nationally varying parameters provide an even closer fit. Reproducing the parameter estimates for all nations here is unnecessary. Rather, we show the estimates from the core model (estimated jointly on the European nations listed above) along with those estimated for Sweden and the United States in Table 4.

The parameter estimates suggest relative mobility in Sweden is higher than that implied by the core model, in the sense that it is closer to the pattern of neutral fluidity, which is in turn the fluidity pattern implied by the marginal effects alone. The fit of the model for the United States is substantially improved by adding (after inspection of the residuals) an additional affinity effect (AFX), which allows for additional mobility between the service class (I+II) and both routine non-manual workers (III) and non-skilled workers. As slightly different models are accepted for Sweden and the United States, it is instructive to examine the patterns of social fluidity in the full table for these two countries. These patterns are displayed in Table 5.

The numbers shown measure deviations of nationally estimated fluidity patterns from core model fluidity patterns. Specifically, they measure the percentage difference in cell interaction parameters between the relevant national variant model (estimates of equation 5 which are shown in Table 4) and the neutral-fluidity level *compared* to the corresponding difference under the 'core model'.

Table 3
The core model of social fluidity – the decomposition of cell interaction parameters

Origin	Destination						
	Service class (i+II)	Routine non-manual workers (III)	Petty bourgeoise (IVa+b)	Farmers (IVc)	Skilled workers (V+VI)	Non-skilled workers (VIIa)	Agricultural labourers (VIIb)
Service class (1+II)	1N1 + IN2	HI1 + AF2	HI1 + AF2	HI1 + SE	HI1	HI1 + HI2	HI1+HI2 + SE + AF1
Routine non-manual workers (III)	HI1 + AF2	IN1	–	SE	–	HI1	HI1 + SE
Petty bourgeoise (IVa + b)	HI1 + AF2	–	IN1 + IN2	SE + AF2	–	HI1	HI1 + SE
Farmers (IVc)	HI1 + HI2 + SE	HI1 + SE	HI1 + SE + AF2	HI1 + IN1 + IN2 + IN3	HI1 + SE	SE + AF2	–
Skilled workers (I + VI)	HI1	–	–	SE	IN1	HI1 + AF2	HI1 + SE
Non-skilled workers (VIIa)	HI1 + HI2	HI1	HI1	HI1 + SE	HI1 + AF2	IN1	SE
Agricultural labourers (VIIb)	HI1 + HI2 + SE + AF1	HI1 + SE	HI1 + SE	HI1	HI1 + SE	SE + AF2	IN1

Note: The interaction parameters in each cell consist of a linear combination of specific effects, to be estimated from the relative mobility table, that are associated with postulated fluidity patterns between classes that derive from: hierarchy – that movements between classes at three different hierarchical levels are less likely, (2 matrics if HI effects), inheritance – that staying in the orgin class may be more likely because of inherited knowledge etc, (3 matrics, IN), sector – that movements between agricultural and non-agricultural sectors is less likely (1 matrix, SE) and affinity – movements between certain classes separated by other factors may be less likely (2 matrics, AF). Once parameters attaching to each of the different forces have been estimated, the predicted interaction parameters can be calculated for each call as shown in the Table.
Source: Erikson & Goldthorpe (1992 *b*), Tables 4.3, p. 133.

Table 4
Estimated specific fluidity parameters and model fit – core model of social fluidity

Nation	Goodness of fit $G^2(S)$	Estimated interaction effects: Hierarchy HI1	HI2	Inheritance IN1	IN2	IN3	Sector SE	Affinity AF1	AF2	AFX
CORE		-.22	-.42	.43	.81	.96	-1.03	-.77	.46	-
Sweden	33.7	-.16	-.45	.28	.65	.78	-.62	ns	.37	-
United States	42	-.13	-.49	.41	.35	.89	-.68	-.74	.28	.20

Note: See Table 3 for the implied cell interaction parameters. The measure of goodness of fit is the standard G^2-statistic standardized to be more comparable across sample sizes. Values of around 40 for this statistic are taken to represent an acceptable fit.
Source: Erikson & Goldthorpe (1992 b), Tables 5.3, p. 147 and 9.2, p. 319.

Table 5
Estimated social fluidity in Sweden and the United States – differences in estimated patterns of fluidity in accepted national models and national neutral fluidity as a percentage of the difference between the core model of social fluidity relative to core neutral fluidity

Origin		Destination						
		Service class (i+II)	Routine non-manual workers (III)	Petty bourgeoise (IVa+b)	Farmers (IVc)	Skilled workers (V+VI)	Non-skilled workers (VIIa)	Agricultural labourers (VIIb)
Service class (1+II)	Sweden	73	97	97	63	94	97	30
	US	62	111	91	65	92	80	67
Routine non-manual workers (III)	Sweden	97	96	100	67	100	94	63
	US	111	98	100	71	100	92	65
Petty bourgeoise (IVa + b)	Sweden	97	100	73	73	100	94	63
	US	91	100	62	85	100	92	65
Farmers (IVc)	Sweden	65	63	68	65	63	73	100
	US	69	65	78	63	49	85	100
Skilled workers (I + VI)	Sweden	94	100	100	67	86	97	63
	US	92	100	100	71	98	91	65
Non-skilled workers (VIIa)	Sweden	97	94	94	63	97	86	67
	US	80	92	92	65	91	98	71
Agricultural labourers (VIIb)	Sweden	30	63	63	94	63	73	86
	US	67	65	65	82	49	85	98

Note: Numbers are calculated as shown in equation 6
Source: Erikson & Goldthorpe (1992 *b*), Tables A5.1, pp. 185–187.

The numbers in Table 5 are calculated as

$$100\exp[sgn(I_{cij} - I_{c0})\ \{(I_{vij} - I_{v0}) - (I_{cij} - I_{c0})\}], \qquad (6)$$

where I_{cij} is the interaction parameter in the "core model" and I_{c0} is the neutral fluidity parameter under that model, whereas I_{vij} is the national variant interaction parameter in cell (i,j) and I_{v0} is the parameter for neutral fluidity (Erikson & Goldthorpe 1992*b*, Table A5.1). A value of 100 in this table suggests that for the country and cell concerned, the national variant model is as distant from the national neutral fluidity estimate as the core model estimate for this country is from the core neutral fluidity estimate.

The deviations of Swedish and U.S. fluidity patterns tend to be in the same direction and suggest relatively small deviations from patterns of neutral social fluidity. Erikson & Goldthorpe (1992 a) conclude that Sweden and the United States appear to be at "the upper end" of relative class mobility among the countries that were investigated but neither is exceptionally mobile.[21] In their con-

21. See also Erikson & Goldthorpe (1992 b, ch. 11) for estimates of overall social fluidity rates.

cluding discussion, Erikson & Goldthorpe express concern over the comparability of the basic U.S. data with those from other countries. It appears that the set of occupational codes in the 1960 U.S. census that forms the basis of the data that were available to them was not very well suited to recoding along the criteria of the EGP schema. Cross-national comparability of data sources would seem to be an unresolved issue also in this literature.[22]

Ganzeboom et al. (1989) apply other variants of log-linear models to their cross-country data that contain multiple surveys and years for each country. Their emphasis is on changes in fluidity patterns over time rather than on cross-country fluidity differences in their results. However, the estimated immobility coefficients (see Ganzeboom et al. 1989, Table 5) suggest that Sweden and the United States are among the most mobile countries in this study as well. With respect to the mobility ordering of countries, their results lend support to the findings of Erikson and Goldthorpe.

Mobility of status or prestige attainment

The parameters of interest

Much work in this field has been directed to estimating the parameters of the causal ("structural") model of status attainment suggested by Blau & Duncan (1967). This model contains direct as well as indirect effects of family background on occupational status. The occupational status of the respondent (at the peak of his career) is in this model a linear function of the status of his first job, the occupational status of his father, and his own educational level. The status of the first job is (also linearly) determined by the occupational status of the father and own education. Finally, own education is a linear function of the occupation of the father. In such a framework it is possible to distinguish between direct effects of family background on status achievement as well as indirect effects via the education of the respondent.

Even though the simple correlations between father's and son's occupational status is one output of such a model, this research approach has the more ambitious goal of finding the causal mechanisms via which family background affects socioeconomic outcomes. The parameter of interest to us here is the correlation of the (current) socio-economic status of fathers (S_{fi}) and sons (S_{si}):

$$S_{si} = \rho S_{fi} + \eta_i. \qquad (7)$$

The evidence

The relatively few existing comparative studies focus on the parameters in structural models such as that in Blau & Duncan (1967). However, in order to achieve comparability simpler versions have been estimated. For example, Treiman & Yip (1989) examine the importance of fathers' occupation, which reflects ascription, and own education, which reflects achievement. They find that the latter is most important, at least in industrialized countries. Treiman & Ganzeboom (1990) present similar conclusions in a survey of studies from a large number of countries. That achievement is more important than ascription can be interpreted as one kind of openness of the society.

For our purposes, it is unfortunate that it has not been customary to present the simple correlations (or regression coefficients) between the occupational status of fathers and

22. See also the special issue of the European Sociological Review in 1992 (volume 8, no. 3) devoted to Erikson & Goldthorpe (1992 b).

Table 6
Intergenerational correlations of occupational status between fathers and sons

Country	Sample size	Estimated correlation
Austria	452	0.500
Finland	388	0.342
Germany	2964	0.419
Ireland	1793	0.491
Italy,	372	0.372
Netherlands	2246	0.404
Northern Ireland	2250	0.410
Switzerland	380	0.474
United Kingdom	6895	0.351
United States	20490	0.343

Note: The sons are 21-64 years old, and must work at least 30 hours a week to be included in the sample. The ISEI status scale is used as the measure of status. Source: Ganzeboom and Treiman, (unpublished computations made available to us).

sons. Fortunately though, we have obtained estimated correlations from a new data set developed by Ganzeboom et al. (1992) particularly for cross-country comparative research purposes.[23] More specifically, they have designed a status scale (International Socio-Economic Index of occupational status, ISEI) for the 271 occupations in the International Standard Classifications of Occupations (ISCO). ILO released this classification of occupations in 1988.

This new status scale is derived in the spirit of the Duncan index so that a weighted average of the educational and income levels of occupations determine their status. 31 data sets covering 16 nations were stacked into one data file. Next, the status scale was constructed so that the direct effect of education on income was maximized (and the indirect effect via occupation minimized) in a basic status attainment model estimated using the stacked data. The status index thus achieved is positively correlated with the Treiman prestige index and the EGP class scheme (10 classes scaled with ISEI means): 0.76 for the prestige index and 0.90 for the EGP.

In Table 6 we show the estimated correlations for 10 countries. The one for the United States is 0.341. It is interesting to note that Zimmerman obtained an almost identical estimate when he used the Duncan status index (Table 1) on the NLS data.[24] Our main conclusion from reading this table,

23. See also Ganzeboom & Treiman (1996).

24. Two further technical points are noteworthy. In contrast to the results for earnings, Zimmerman's estimated correlation for the Duncan index with the TA-technique did not rise when the number of years was increased, suggesting that the occupation of the father in a single year was not a particularly noisy measure in the 1960s. Further, the correlation obtained in Table 6 with retrospective information on fathers' occupation is about the same as the one obtained by Zimmerman using information reported by the fathers themselves. Some results in Björklund & Jäntti (1997) also suggest that using sons' retrospective information about fathers' occupation and education yields more or less the same results as when fathers' own information is used. Cf. Hauser &Warren (1997).

though, is that the United States ranks quite highly in terms of mobility of occupational status. Only Finland and the United Kingdom have as low correlations. (Standard errors are, of course, necessary in order to draw strong conclusions about the differences among countries.)

A comparison with results for earnings correlations reveal two contradictory results. First, Finland and the United States have about the same correlations of status attainment, whereas Finland seemed to have a lower correlation of earnings. Second, status attainment is more strongly correlated among fathers and sons in Germany than in the United States, whereas the opposite pattern was found for earnings correlations. A consistent result from the two approaches to intergenerational socio-economic mobility is that the United States and the United Kingdom have about the same correlations in both types of analysis.

Comparing approaches to intergenerational socio-economic mobility

How should we view these three approaches to the mobility of social and economic status? Is one "better" than the other? Or are these research approaches complements rather than substitutes? Further, is it possible that a country gets a high rank in terms of mobility for one outcome measure but a low one for another measure?

From a purely conceptual point of view, status (prestige), class and income (earnings) are obviously different aspects of a person's position in society. Hence, we ought to treat these three branches of mobility analysis as complementary; they will offer different information about the nature of intergenerational socio-economic mobility. On the other hand, there are strong reasons to believe that the mechanisms that generate mobility of status, class and earnings are quite similar. We should therefore expect quite similar results. The literature in all three fields of intergenerational socio-economic mobility has also shown what a difficult task it is to achieve comparability among countries. Therefore, the possibilities to solve the pure practical problems to achieve comparability among countries must be given high priority in any attempt to find out which approach to analysis of intergenerational socio-economic mobility that has produced the most credible findings. What outcome variable is most easy to compare among countries: status, class or income?

Nonetheless, the results also indicate that countries tend to rank differently depending on outcome measure, so we must also ask whether it is possible to reconcile the results from the three fields of research. We continue to discuss these issues under the headings "practical issues" and "reconciliation of the results".

Practical issues

In the first place, practical problems arise when measuring current income and occupation. If the researcher is confined to collecting this information by means of surveys, income is definitely the more problematic variable to measure accurately. As stressed by Hauser & Warren (1997), people are much less willing to reveal their income than their occupation to an interviewer. On the other hand, occupational data place more of a burden on the collector and the coder of the data. In order to achieve a high degree of cross-country comparability, the coding procedures must not differ too much between countries. Indeed, part of the controversy between Erikson & Goldthorpe

(1992 a) and Hout & Hauser (1992) had to do with the occupational coding procedures in the United States compared to other countries. We believe that the corresponding problems in interpreting income data are less, even though there could be differences among countries in e.g. the inclusion of the monetary value of work-related benefits in reported income.

The problems in measuring income in surveys are even more severe when it comes to parents' income. Whereas it is feasible to get reliable answers to questions about parents' (main) occupation, it is more or less impossible to get accurate information about parents' income by means of retrospective interview questions.

When data from administrative records are available, the situation changes to the advantage of income data. Today, income data covering (at least) two generations are available for research purposes in several countries.[25] Not only are non-response and recall errors eliminated in this way, but, in addition, sample sizes are often large making statistical inference easier and estimation of more elaborate relationships feasible.

A second practical problem arises because of intragenerational mobility of income and occupation. There is of course intragenerational mobility in both these variables, but most likely income is the more mobile one. Since income is a continuous variable, there are – as we have seen – now quite simple techniques to examine the extent of intragenerational mobility and correct for the bias that it causes. Studies that use income and are based on panel data sets like the PSID can solve problems that are due to intragenerational mobility by using incomes measured over several years. The studies that use status and class, however, are based on the notion that the father had one major occupation and that the same holds for the son. This assumption might have been reasonably valid in the past, but might be less valid in the modern labor market that we believe is characterized by higher mobility of workers between sectors of the economy. Modeling the impact of intragenerational occupational mobility on observed measures of mobility is less analytically tractable and has to the best of our knowledge been studied only rarely (see Breen & Jonsson, (1997) and Hauser & Warren, (1997) for recent treatments of these issues).

A counterargument could be that income and earnings are more difficult to compare across countries than is commonly believed. Problematic groups in this respect are farmers and self-employed persons. In the sociological literature, the problems associated with treating these groups have been discussed at length. Maybe these problems are more severe in the analysis of income mobility. This issue has not yet (to the best of our knowledge) been addressed among the economists who work in this field. A related problem is that occupations differ in terms of working conditions. Economists would characterize this as a problem of taking compensating wage differentials into account. Current earnings could therefore be an incomplete indicator of the attractiveness of an occupation. One advantage of using the status of an occupation instead of current earnings could be that status captures the value of working conditions.[26] The fact that it

25. The Canadian and Finnish studies in Table 1 are based on incomes from such registers.
26. It is common to compare university professors (with low earnings but quite high status) and car salesmen (with higher earnings but lower status).

often is hard to get significant coefficients on working conditions variables in micro data wage equations (see Duncan & Holmlund, 1983) could be taken as an indication that this is not a severe flaw of earnings as outcome measure in intergenerational studies, but we believe that the problem deserves some attention in future research.

One factor that may confound the assessment of mobility in incomes and earnings, but not in occupation-based outcomes, is cross-sectional price variation. The same income in, say, a large urban center may generate less well-being than in a rural setting. Thus, a higher income for the urban son than that enjoyed by his rural father may in fact represent no mobility at all once cost-of-living differences are taken into account. Urban areas do, on the other hand, offer services that may only be participated in at great expense in rural areas, such as e.g. theater and opera. The scope for price variation across areas is likely in part a function of the size of a country. It is difficult to determine, with no empirical analyses, if and in what direction regional price variations generate biases. It is even more difficult to assess how the handling of such a complication would affect the mobility ranking of countries.

Finally, we note that an important advantage of the analysis of class mobility using methods for the analysis of frequency tables is that no artificial restrictions on the form of the relationship – such as symmetry – are imposed on the observed patterns. A weakness in the present economics literature is that it has mostly been conducted in terms of correlation (or regression) coefficients rather than in terms of more flexible measures of association (directly estimated mobility tables being an exception, of course). There are, a priori, no good reasons to believe that the association between fathers' and sons' incomes is the same throughout, over e.g. the income range of fathers. The Canadian study by Corak & Heisz (1999) is an exception to this point. They estimate a more general functional form and do, indeed, find that the association between fathers' and sons' incomes varies over the distribution of fathers' income. Their analysis of a more flexible functional form is much facilitated by a large sample size that is due to the data's stemming from registers. An obvious direction in which to move also the comparative literature is to estimate more flexible, preferably non-parametric measures of association, such as correlation curves (see Bjerve & Doksum1993).

Reconciliation of the results

The three major ways of measuring socio-economic intergenerational mobility rank the United States differently. Comparative studies of occupational and class mobility as well as of mobility of status attainment tend to rank the United States fairly high in the mobility ranking. Studies that measure the correlation of long-run income (or earnings) tend to place the United States quite low in the mobility ranking. On the face of it, this might appear to be contradictory.

However, one can obtain higher intergenerational earnings mobility in country A than in country B even if B has higher occupational mobility or mobility of status attainment. The vehicle for this re-ordering is the extent to which differences in the part of earnings that depends on occupation and actual earnings is distributed and, importantly, how highly correlated these deviations are across generations. The greater the role of fathers' unobserved influences in earnings is on the economic position of sons, the more scope there is for re-ranking in moving from occupational mobility to earnings mobility.

Class and earnings mobility

Is it possible that a country with higher occupational mobility has lower earnings mobility? We believe it is. Consider a simple model of earnings or income, where the long-run earnings or income, y, of both the father f and son s in the ith pair is determined by observed variables, X, with associated coefficients, β, and an unobserved effect, a:

$$y_{fi} = X_{fi}\,\beta_f + a_{fi}, \tag{8}$$

$$y_{si} = X_{si}\,\beta_s + a_{si}. \tag{9}$$

The observable variables, X, consist of a set of indicator variables of class position. The unobserved effect a captures the variation within classes, but can also capture intergenerational influences that are not captured by the measured association between classes. The coefficient vector β consists of within-occupation income averages (or the "returns to occupation"). We abstract here from problems associated with *estimating* β and need not assume anything about the correlation of X and a. We can write the covariance of fathers' and sons' incomes as

$$\sigma_{yf,ys} = \beta'_f\,\overline{X'_f X_s}\,\beta_s + \sigma_{af,as}, \tag{10}$$

where we have, for expositional simplicity, ignored the fact the the unobserved parts of both fathers' and sons' are likely to be related to the observed parts. The matrix $\overline{X'_f X_s} = 1/n\ \sum_i X'_{fi}\, X_{si}$ gives the relative frequency of each combination of fathers' and sons' occupation, i.e., the cross-tabulation of fathers' and sons' occupation divided by the number of father-son pairs, n.

Suppose that you are comparing earnings and class mobility in two countries, A and B. There are, in this setup, several ways in which higher class mobility in A compared to B might translate into lower mobility of earnings in A than in B. First, the β-coefficients might be such that for country A, although it in some sense has more mobility, it ends up having a larger covariance. Second, the covariance of the unobserved terms might be such that although $\beta'_f\,\overline{X'_f X_s}\,\beta_s$ is equal across countries, the correlation is higher in A. If we write the difference in the earnings covariance in terms of the decomposition in equation 10, we can see how this can happen:

$$\underbrace{\sigma^A_{yf,ys} - \sigma^B_{yf,ys}}_{I} = \beta^{A\prime}_f\,\underbrace{\overline{X'_f X_s}^{A}\beta^A_s - \beta^{B\prime}_f\,\overline{X'_f X_s}^{B}}_{II}\beta^B_s$$

$$+ \underbrace{\sigma^A_{af,as} - \sigma^B_{af,as}}_{III} \tag{11}$$

In terms of equation 11, we have lower mobility of earnings in A, $I>0$, if both the difference of the observed parts is positive, $II>0$, and the difference of the unobserved parts is positive, $III>0$, or as long as either of them is large and positive enough, $II+III>0$.

Status attainment and earnings mobility

What about status attainment and earnings mobility? We can write out the model that links socio-economic status to its determinants:

$$S_{i,j} = X_{i,j}\delta, \quad j = f,s \tag{12}$$

where X is a vector of occupational indicators, δ is the vector of weights (the same for father and son) used to calculate the index of status attainment and S is the value of the index itself. Doing some injustice to the idea of how occupational prestige can be attained (but to be able to compare it with the earnings mobility – literature) we write:

$$y_{i,j} = X_{i,j}\delta\gamma + e_j, \quad j = f,s \tag{13}$$
$$= S_{i,j}\gamma + e_j.$$

The coefficient γ links status to earnings, with an error term e, allowing for differences between the prestige of an occupation and where it places a person in the distribution of long-run income. We can then write the intergenerational earnings correlation (again, for simplicity ignoring possible correlation between the e's and S's) as

$$\sigma_{yf,ys} = \gamma^2\, \sigma_{Sf,Ss} + \sigma_{ef,es}. \tag{14}$$

In comparing two different countries in terms of the three elements above, it is clear that we could have lower intergenerational earnings mobility in country A even if there was lower mobility of status attainment in B, as long as the correlation of disturbances was sufficiently lower in B – low enough to offset the higher correlation of attained status in B.

$$\underbrace{\sigma^A_{yf,ys} - \sigma^B_{yf,ys}}_{I} = \underbrace{\gamma^{A2}\sigma^A_{Sf,Ss} - \gamma^{B2}\sigma^B_{Sf,Ss}}_{II} + \underbrace{\sigma^A_{ef,es} - \sigma^B_{ef,es}}_{III} \tag{15}$$

Again, $I>0$ if both $II>0$ and $III>0$, or as long as $II+III>0$.

Note that equations (11) and (15) suggest ways in which class and status mobility differ from income or earnings mobility. Namely, factors that are unassociated with class or occupation (the a's and e's) generate within occupation or class dispersion of earnings, but may also be correlated across generations, in which case they enter the income correlation.

Conclusions

One major conclusion arises from the three broad strands of literature that we survey, as well as the literatures on *intra*generational economic mobility and cross-sectional inequality: There is no simple relationship between cross-sectional and longer run inequality, whether intra-, or intergenerational. General stories of conflicts between equality of opportunity and of outcome, for instance, fail to account for the evidence.

In terms of the broad question of "American exceptionalism", it is hard to conclude anything more than this: The United States is exceptional only in so far as it has extra-ordinarily high relative income differences. These are not accompanied by exceptionally high rates of intergenerational socio-economic mobility – at least not across the three dimensions scrutinized here.

We believe that intergenerational socio-economic mobility is a question that rightly is quite central to both economics and sociology. Efforts to collect new, and to continue to collect old, longitudinal household surveys that address all three types of mobility discussed in this paper should be given as much support as possible. The micro data should also be made available to individual researchers in order to facilitate the overcoming of the fairly severe comparability problems that are present. Also, in the absence of observed data on several generations, more effort can and should go into studying the accuracy and reliability of recall data, i.e., the accounts of a current generation about the activities of their parents.

We want to distinguish between two different tasks for future research, for both of which it is fruitful to incorporate income in the analysis. First, (reasonably) simple summary measures of intergenerational associations are needed to compare the overall

magnitude of mobility among countries and over time. It is, after all, an interesting social statistic in its own right, the magnitude of which has been subject to much speculation. The prospects of getting reliable estimates of such measures based on income have improved considerably in recent years, and will probably increase in the near future as well.

Even though it would be to go too far to say that the problem of estimating correlations of long-run (or permanent) income between fathers and sons has been settled, estimation techniques suggested by Solon and Zimmerman, perhaps augmented by more complex error processes, applied to data sets such as the PSID make it possible to find a fairly narrow bracket within which the true parameter lies. Further, in several countries, access to income data from administrative records eliminates the need for surveys to collect the data that are needed. No doubt, studies based on occupational status or class are more reliable when data collection is restricted to surveys, and in particular when information on parents' must be based on retrospective questions. However, we foresee several studies of intergenerational income mobility based on large administrative data sets in the near future. The second task is to learn more about the mechanisms that generate the strong intergenerational connections in socio-economic outcomes, and especially how various public policies affect the relationships. Most likely, parents' education and occupational prestige are important per se in this process. Parents' education offers useful information and knowledge, and their social status provides networks that facilitate the labor market career. However, we would be most surprised if income did not matter in this process. Some investments in children must directly or indirectly be paid by the parents and therefore money cannot be neglected in this analysis.

Our final, rather obvious, observation is that economists and sociologists have much to both learn from and to teach each other in bringing this research forward.

Appendix

Techniques to estimate the intergenerational correlation in income

Let y_{si} be the log of long-run economic status, or permanent income of the son and y_{fi} be the corresponding variable for the father of the ith father-son pair. Let ρ represent the population correlation between sons' and fathers' economic status, our measure of - intergenerational income mobility. If the y's were directly observed, ρ could be estimated by applying ordinary least squares (OLS) to

$$y_{si} = \rho y_{fi} + \varepsilon_i \tag{16}$$

where y_{fi} and y_{si} are measured as deviations from means and are assumed to have equal variances.[27]

The time-averaging (TA) technique

Long-run economic status is not directly observed. Rather, we observe annual incomes, assumed to equal true long-run status and random transitory fluctuation:

$$y_{fit} = y_{fi} + v_{fit}, \tag{17}$$

27. The assumption of equal variances is adopted here to facilitate the presentation of the technique we employ. The assumption used here for expositional purposes.

and

$$y_{sit} = y_{si} + v_{sit}, \qquad (18)$$

where the v's represent random transitory fluctuation, assumed to have variance σ^2_v and to be uncorrelated with each other or true status. An OLS estimate of the correlation, using annual measures, yields a downwards inconsistent estimate of ρ (Solon, 1989,1992):

$$\operatorname*{plim}_{N\to\infty} \hat{\rho}_{OLS} = \frac{\text{Cov}(y_{fi}, y_{si})}{\text{Var}(y_{fi})+\text{Var}(v_{fit})} \approx \frac{\sigma^2_y}{\sigma^2_y+\sigma^2_v}\rho < \rho. \qquad (19)$$

If several observations of the annual income of fathers are available, less biased estimates of the correlation can be obtained. While we still can't observe y_{fi} directly, we use

$$\bar{y}_{fi} = \frac{1}{T}\sum_{t=1}^{T} y_{fit} = \frac{1}{T}\sum_{t=1}^{T}(y_{fit}+v_{fit}) = y_{fi} + \bar{v}_{fi}, \qquad (20)$$

where $\bar{z}_i = \frac{1}{T}\sum_{t=1}^{T} z_{it}$ for any variable z. Then the probability limit of the correlation is

$$\operatorname*{plim}_{N\to\infty} \hat{\rho}_1 = \frac{\text{Cov}(y_{fi}, y_{si})}{\text{Var}(y_{fi})+\text{Var}(\bar{v}_{fi})} = \frac{\sigma^2_y}{\sigma^2_y+\frac{\sigma^2_v}{T}}\rho<\rho. \qquad (21)$$

The inconsistency of this estimator diminishes with the number of years over which incomes are averaged. Using this approach, Solon (1992) showed that previous estimates of the intergenerational correlation were seriously downward biased.

The two sample instrumental variable (TSIV) technique

In Björklund & Jäntti (1997) we have no actual Swedish father-son pairs. What we do have, however, are data on the fathers of young men. Specifically, the Swedish Level of Living Survey (SLLS) asks respondents what their parent's education and occupation were. By estimating earnings equations for a sample of older men, our set of \`synthetic fathers', we construct two different estimators of the intergenerational correlation in order to gauge the magnitudes of these parameters in Sweden. We are able to use data from the PSID to mimic these estimators for the United States, enabling us to compare estimates obtained using similar methods in the two countries.

What the technique does is to use information from two samples, estimating coefficients from a sample of adult men with sons with which to predict father's income for a sample of sons, who have reported their father's education and occupation. Dearden et al. (1996) call this the prediction technique and in the statistical literature it is known as two sample instrumental variables (TSIV) estimation, see Arellano & Meghir (1992) and Angrist & Krueger (1992).

Assume that a father's permanent income can be written as

$$y_f = X_f\, \beta_f + \eta_f, \qquad (22)$$

where X_f and β_f are vectors of explanatory variables and coefficients and η_f is an unobserved term affecting permanent income,

assumed to be independent of X_f (suppressing the index i. The corresponding equation for the son is

$$y_s = X_s\beta_s + \eta_s. \qquad (23)$$

The assumption that the unobserved component is orthogonal to the observed component in permanent income allows us to estimate the β_s-coefficients by use of OLS. Observe that we do not assume that the sons' and fathers' unobserved components are uncorrelated. The estimator of the correlation between fathers' and sons' actual incomes is, using the assumption that $Cov(X_f \eta_f)=0$,

$$\rho_1 = \frac{\beta'_f\, Cov(X_f, X_s)\beta_s + \beta'_f\, Cov(X_f,\eta_s)}{\beta'_f\, Var(X_f)\beta_f + Var(\eta_f)} + \frac{Cov(\eta_f,X_s)\beta_s + Cov(\eta_f,\eta_s)}{\beta'_f\, Var(X_f)\beta_f + Var(\eta_f)} \qquad (24)$$

the correlation between fathers' observed characteristics (predicted income) and sons' observed income,

$$\rho_2 = \frac{\beta'_f\, Cov(X_f, X_s)\beta_s + \beta'_f\, Cov(X_f,\eta_s)}{\beta'_f\, Var(X_f)\beta_f}, \qquad (25)$$

is useful. As we argue in Björklund & Jäntti (1997), under reasonable assumptions this estimator provides an upper bound on the true intergenerational correlation, ρ. On the other hand, the correlation between fathers' observed characteristics (predicted income) and sons' observed characteristics (predicted income) is

$$\rho_2 = \frac{\beta'_f\, Cov(X_f, X_s)\beta_s}{\beta'_f\, Var(X_f)\beta_f}, \qquad (26)$$

This estimator measures the strength of the association between the observable component in the permanent income of the fathers and sons.

The denominators of ρ_2 and ρ_3 are smaller than that of ρ_1, because the variance of the unobserved component only enters in ρ_1. The estimators ρ_2 and ρ_3 will differ to the extent that the unexplained part of sons' income is correlated with fathers' observed characteristics [$\beta'_f\, Cov(X_f, \eta_s)$]. The larger the covariance of sons' and fathers' observed characteristics, the closer will the estimators ρ_2 and ρ_3 be. The smaller the variance and covariance terms involving unobserved parts are, the closer will all three estimators be. At the extreme, if observable characteristics account for all of permanent income, all three estimators will be equal.

The education-as-instrument (IV) technique

In addition to the estimator ρ_1, Solon (1992) estimates the correlation by use of an instrumental variables (IV) method, arguing that this produces an *upwards* inconsistent estimate of ρ. The argument, in brief, is as follows. Assume that the sons' long-run status is determined by

$$y_{si} = \gamma_1\, y_{fi} + \gamma_2 E_i + \varepsilon_i, \qquad (27)$$

where E_i is the father's education. Estimating the parameter we are interested in, namely, the projection of y_{si} on y_{fi} leads to, using the standard omitted variable formula,

$$\operatorname*{plim}_{N\to\infty} \hat{\rho}_1 = \gamma_1 + \gamma_2\frac{\lambda\sigma_E}{\sigma_y}\rho, \qquad (28)$$

where λ is the correlation between E and y_{fi}. An IV estimate of ρ, using father's education as the instrument, has the probability limit

$$\operatorname*{plim}_{N\to\infty} \hat{\rho}_4 = \frac{\mathrm{Cov}\,(y_{si,}\ E_i)}{\mathrm{Cov}\,(y_{fi,}\ E_i)}$$

$$= \frac{\mathrm{Cov}\,[(\gamma_1 y_{fi} + \gamma_2 E_i + \varepsilon_i),\ E_i]}{\mathrm{Cov}\,(y_{fi,}\ E_i)}$$

$$= \gamma_1 + \gamma_2 \frac{\sigma_E}{\lambda\sigma_y} = \rho + \gamma_2\sigma_E \frac{1-\lambda^2}{\lambda\sigma_y}. \quad (29)$$

If education has a positive effect on income, i.e. $\lambda>0$ and $\gamma_2>0$, this estimator will be upwards biased. The closer to one λ is, the smaller will the upwards bias be, as there is less variation in income that isn't captured by education.

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