Review of Income and Wealth
Series 46, Number 4, December 2000

INTERGENERATIONAL INCOME MOBILITY IN SWEDEN: WHAT DO TAX-DATA SHOW?

BY TORUN ÖSTERBERG

University of Goteborg

The aim of this paper is to investigate intergenerational income mobility in Sweden by means of a representative sample drawn from tax-data files. Longitudinal data on actual parent-child pairs spanning 1978–92 are employed. Regression and correlation coefficients are analyzed and transition matrices calculated in order to investigate income mobility over generations. The results achieved show high intergenerational income mobility in Sweden between fathers and sons in comparison to estimations performed in most other countries and more especially compared to the U.S. This indicates that Sweden does not only have lower cross-sectional income inequality, but also higher intergenerational income mobility than those countries. The mother's earnings influence children's earnings less than the father's. However, the mother's earnings correlate more strongly with a daughter's earnings than they do with that of a son. The major indication of immobility across generations is found in the upper income deciles.

INTRODUCTION

The question whether economic status is transferred from one generation to another has gained increased attention in recent years. Consequently there is a substantial body of literature where the correlation in permanent earnings is estimated across generations. The aim of this paper is to investigate intergenerational income mobility in Sweden using a representative sample drawn from tax-data files. To study intergenerational transmission of earnings status is of interest when studying the character of inequality in a specific society. If there is perfect mobility across generations, it will not matter if your parents belonged to the 90th income percentile or the 10th. The child is independent of its parent's income position. If on the other hand there is no mobility, the child will end up in exactly the same income position as its parents. Most societies are somewhere in-between these extremes and in this study the position of Sweden's intergenerational income mobility will be investigated.

Various attempts have been made to estimate the intergenerational income correlation in Sweden. One example is Gustafsson (1994), who uses a homogenous sample where the fathers' incomes were observed for one year only.[1] However, most studies on intergenerational income mobility have been generated from U.S. data (Behrman and Taubman (1985), Peters (1992), Solon (1992) and Zimmerman (1992)), though some studies of other countries exist as well.[2] Recently,

Note: Financial support from Swedish Council for Social Research (SFR) is gratefully acknowledged. The author would like to thank Björn Gustafsson, Katarina Katz, Pernilla Dahlin and an anonymous referee for helpful comments and suggestions.

[1]The data consists of sons born between 1939 to 1946 living in Stockholm in the middle of the 1950s.

[2]Examples of this are United Kingdom [Atkinson (1981) and Dearden *et al.* (1997)], Sweden [Gustafsson (1994) and Björklund and Jäntti (1997)], Canada [Corak and Heiz (1996)] and for Malaysia [Lillard and Kilbrun (1995)].

income mobility has been compared cross countries; examples are Couch and Dunn (1996) who perform a cross-country study of Germany and the United States, and Björklund and Jäntti (1997) who compare Sweden with the United States.

When analyzing economic status across generations in a cross-country perspective, data availability is one of the major problems. Björklund and Jäntti, for instance, cannot observe actual incomes on dependent generations when estimating intergenerational income mobility in Sweden. They handle this difficulty by using a new technique. They estimate intergenerational income correlations for independent fathers and sons in order to obtain the Swedish estimates. Their data consist of each son reporting his father's education and occupation, and from this information they have predicted the father's earnings, by means of using an independent sample of fathers. The results of Björklund and Jäntti point in the direction of higher intergenerational income mobility in Sweden than in the United States. Their evidence is, however, not entirely conclusive.

This paper analyses whether their conclusion is also valid when examining earnings information on dependent fathers and sons in Sweden from the tax-authorities. The analysis will also provide estimates of the intergenerational income correlations for father-daughter, mother-son and mother-daughter pairs, using actual earnings for both generations. The difference here from earlier estimates of Swedish intergenerational income mobility is not only the use of tax-data, but also the use of a representative sample which is quite large. This enables the detection of where in the income distribution mobility is greatest.

In the analysis both regression and correlation coefficients on income over generations will be estimated, in order to compare the results with other studies. Transition matrices will also be calculated. The fact that Sweden represents a country with a fairly low level of cross-sectional income inequality is quite well known. But is Sweden also a country with high intergenerational mobility?

Data and How They Are Analyzed

The data used in this analysis are drawn from The Swedish Income Panel (SWIP). The panel consists of a representative 1-percent sample drawn from the register of the total population (RTB) of Swedish-born individuals. The sample was first drawn for people residing in Sweden in 1978. Then, the individuals were followed each year until 1992. Income information is provided by the Swedish tax-register, which also includes those who do not pay any income tax. An advantage of the panel is that it provides earnings information on parents for a subset of the sample. The subset consists of sampled persons born in 1941 or later whose mothers and fathers could be identified.

The identification was successful in most but not all cases; it was possible to identify at least one parent in 82 percent of the cases.[3] Parents are identified and

[3]Out of 29,469 Swedes born between 1941 and 1965, I have found parents to 24,275 of them. The inability to find parents for all individuals could be caused by emigration from the country, but the primary reason is probably the parents' deaths. When excluding those parents that were 65 years or older in 1978, there are 19,455 sons and daughters left for the analysis. All children are Swedish-born, while 6.7 percent of the fathers and 2.3 percent of the mothers are foreign-born.

matched to children by means of central registers. These include information for each individual on who has the formal guardianship of the child, i.e. usually the biological parents unless the child is adopted. As soon as a child is born, a formal guardianship is registered by the authorities. The matching of parents and children was executed by Statistics Sweden in the early 1990s. If the mother is unmarried, she can choose to be the sole guardian of the child, which is why slightly more mothers than fathers are found in the registers.[4] Despite this, the registers are considered to be of high quality and accuracy.[5]

The longitudinal composition of the data provides a more reliable measure of long-run status than a cross-section would do, as transitory fluctuations in earnings can be avoided in the former. The problem posed by homogenous samples discussed by Solon (1992) should not cause any difficulty here, since the data are based on a representative national sample from register data, which results in no attrition (other than from death and migration) and no sample non-response. Estimation based on hourly earnings would be informative. However, one disadvantage of the data is that it does not provide any information on working hours or wage per hour, so one has to rely on annual earnings alone.

The analysis is based on pairs, in which the individuals of the second generation were 25 years or older in 1990 and the parents were 64 years or younger in 1978. These restrictions exclude most students and retired individuals. The following numbers of pairs are obtained; 8,336 father-son pairs, 8,053 father-daughter pairs, 8,753 mother-son pairs, and 8,387 mother-daughter pairs.[6] Income of both parents and children is observed for a 3-year period; the parents from 1978 to 1980 and the children from 1990 to 1992, which means one and a half decades between the intervals of measurement. This design is the result of a trade-off between the desirability of observing income in each generation for many years and that of observing the two generations at the same stage of the life cycle. To include more years in the observation period would give a better approximation of long-run economic status, but with each year included in the observation period observations would be lost due to the age restrictions. Previous results indicate stabilization of the estimates using a three-year average [Zimmerman (1992) and Solon (1992)]. Compared to most data used to estimate intergenerational mobility the sample size in SWIP is quite large, which enables analysis of sensitivity of the estimates and the possibility of discovering non-linearities.

The two generations are not observed at identical stages in their life cycles. The father's mean age in 1978 was 52 and the mother's 49, while the children's mean age was 37 in 1992. This indicates an average age difference of about 12–15 years between the generations, at the point of observation. When Björklund and Jäntti (from this point on referred to as B&J) estimate the intergenerational

[4] 8,336 father-son pairs can be found compared to 8,753 mother-son pairs. Corresponding numbers for daughters are 8,053 and 8,387.

[5] The definition implies that in some cases the child is not living with its parent. This is the case when, for example, the child lives with a stepfather and the biological father is registered as the guardian. One can perhaps argue that the formal guardian has a greater impact on the child's development than the stepfather, but these issues will not be further dealt with here.

[6] Siblings could be included in the sample as it is a 1 percent sample out of the Swedish population. However, the probability that siblings would be included in a 1 percent sample must be considered very low and should not cause any major problems.

correlation, the analyzed sons are born between 1952 and 1961 and the mean ages of the sons and the fathers are 34 and 43 respectively. This gives on average a 10-year difference between the generations in B&J's Swedish sample. To estimate intergenerational income correlations, B&J use earnings and market income where capital income is added. The corresponding annual incomes will be used in the estimations performed in this analysis.[7]

Various panels in Sweden and the United States permit analyses of intergenerational mobility. The data used here, SWIP, consist of a large number of observations where actual income is observed for both generations. The data used by B&J are the Swedish Level of Living Surveys (SLLS). B&J use 550 fathers to estimate earnings and 327 independent sons. The published U.S. estimates are mostly based on two panels; the Panel Study of Income Dynamics (PSID) (used by among others, Solon (1992) and B&J, and the National Longitudinal Survey (NLS) used for example by Zimmerman (1992). However, in their studies, sample sizes are quite small; the PSID sample used by Solon consists of about 350 father-son pairs and NLS of about 900 father-son pairs.

METHOD

To estimate intergenerational mobility a simple Markov process is used.

$$Y_{i,c} = \beta Y_{i,p} + \varepsilon_i \tag{1}$$

where $Y_{i,c}$ is the long-run economic status or permanent income of generation c (the child) and $Y_{i,p}$ is the corresponding variable for generation p (the parent). When ordinary least square is applied to equation (1) the estimate of β will be downward inconsistent, as permanent income cannot be observed, only annual income can. The annual incomes that are observed can be assumed to consist of permanent income plus a random transitory fluctuation.

$$Y_{i,c,t} = Y_{i,c} + \upsilon_{i,c,t} \tag{2}$$

where $\upsilon_{i,c,t}$ is the random transitory fluctuation for individual i in generation c at time t.[8] The downward bias would be of the magnitude $plim\,\hat{\beta} = \sigma_y^2/\sigma_y^2 + \sigma_\upsilon^2)\beta < \beta$ when estimating β by OLS (Solon, 1992). However, if the annual income is averaged over several years the bias in the estimate of β can be reduced to $plim\,\hat{\beta} = \sigma_y^2/(\sigma_y^2 + \sigma_\upsilon^2/T)\beta < \beta$. The inconsistency will not entirely vanish, only be diminished, as the transitory components of annual income can be serially correlated. As mentioned above, a three-year average will be used in this analysis. The coefficient β corresponds to the elasticity of the son's income with respect to the father's income, when $Y_{i,c}$ and $Y_{i,p}$ are measured in logarithms. In the case of equal variances across generations, β will also be the intergenerational correlation coefficient denoted by ρ. In the case of differing variances, ρ can be estimated by $\beta(\hat{\sigma}_p/\sigma_c)$.

[7] The earnings consist of income from labour and income from business. The market income consists of earnings plus income from capital, deductions of deficits and income from real estate. As deductions are included, the variable can take on negative values.

[8] A corresponding equation is assumed for generation p; the parents.

In the estimations, only those individuals who were identified all three years will be included.[9] Three different ways of measuring annual income will be used: Income in levels for both earnings and market income and the logarithm of earnings (not of market income as it contains negative values). Last, to be able to compare with other reported results, the logarithm of earnings for those with a discontinuous work-history and those that worked all years will be separated in the estimations. For those with a zero-income in some years the average over years will be based only on the years that an income was earned.[10]

To obtain income measures corrected for difference in life-cycles between generations, the income for both generations is regressed on age and age square according to

$$I_{i,t}(c) = \gamma_{1,t}(c) + \gamma_{2,t}(c)age_i + \gamma_{3,t}(c)age_i^2 + v_{i,t}(c). \tag{3}$$

where $I_{i,t}(c)$ is the income for individual i in year t for generation c and $v_{i,t}(c)$ is the individual specific residual. The model is estimated for each generation, sex, year, earnings in levels, market income in levels and logarithm of earnings, by ordinary least squares. The residuals are saved from these 24 regressions and age adjusted income is then calculated as expected income evaluated at the mean age for that generation. The individual specific residual is then added.

$$\text{Adj}\, I_{i,t}(c) = \hat{\gamma}_{1,t}(c) + \hat{\gamma}_{2,t}(c)\overline{age_i} + \hat{\gamma}_{3,t}(c)\overline{age_i^2} + \hat{v}_{i,t}(c). \tag{4}$$

Hats represent the estimated coefficients and residuals and $\text{Adj}\, I_{i,t}(c)$ is the life-cycle adjusted income for individual i in year t for generation c. The approximation of long-run economic status or permanent income $Y_{i,c}$ in equation (1) is then calculated as

$$Y_{i,c} = \sum_{T} \frac{\text{Adj}\, I_{i,t}(c)}{T_i} \tag{5}$$

where T_i is the number of years with an income for individual i. What is used here is the average of logarithm of income and not the logarithm of averaged income. $Y_{i,c}$ is the measure of permanent income that will be used in the Markov process estimations (equation 1) and also in mobility matrices and figures.

Estimation Results

Table 1 displays the regression and correlation coefficients. The estimated regression and correlation coefficients based on SWIP for pairs of fathers and sons are generally lower than those of B&J for Sweden, based on the SLLS, both using annual earnings. This could be caused by the fact that B&J predicted the father's earnings based on his education and occupation. The prediction method could therefore result in incorrect estimated earnings for individuals with earnings atypical of their education and occupation, thereby decreasing the variability in the sample. This conjecture is also supported by B&J's own evidence from the

[9] Even though an individual did not earn an income that year he/she will be observed.

[10] I.e. an individual that had an income for one year only will have this year as the approximation of permanent income in the estimations with zero-earners included.

TABLE 1

REGRESSION AND CORRELATION COEFFICIENTS

Variable	Regression coefficients: Earnings Corrected for Age	Regression coefficients: Market Income Corrected for Age	Correlation Coefficients: Earnings Corrected for Age	Correlation Coefficients: Market Income Corrected for Age
	For Son and Fathers			
Actual incomes	0.139	0.185	0.182	0.181
$n = 8{,}336$	(0.008)	(0.011)	(0.011)	(0.01)
Logarithm of actual income	0.125		0.114	
$n = 8{,}336$	(0.012)		(0.011)	
Logarithm of actual income all those with zero income	0.129		0.131	
excluded[1] $n = 7{,}711$	(0.011)		(0.011)	
	For Daughters and Fathers			
Actual incomes	0.062	0.083	0.134	0.140
$n = 8{,}053$	(0.005)	(0.006)	(0.011)	(0.011)
Logarithm of actual income	0.076		0.069	
$n = 8{,}053$	(0.012)		(0.011)	
Logarithm of actual income all those with zero income	0.071		0.071	
excluded[1] $n = 7{,}412$	(0.010)		(0.010)	
	For Sons and Mothers			
Actual incomes	0.104	0.172	0.068	0.092
$n = 8{,}753$	(0.016)	(0.020)	(0.011)	(0.011)
Logarithm of actual income	0.018		0.026	
$n = 8753$	(0.007)		(0.011)	
Logarithm of actual income all those with zero income	0.022		0.030	
excluded[1] $n = 7{,}361$	(0.009)		(0.012)	
	For Daughters and Mothers			
Actual incomes	0.110	0.139	0.121	0.123
$n = 8{,}387$	(0.010)	(0.012)	(0.011)	(0.011)
Logarithm of actual income	0.025		0.036	
$n = 8387$	(0.007)		(0.011)	
Logarithm of actual income all those with zero income	0.036		0.053	
excluded[1] $n = 7{,}143$	(0.008)		(0.012)	
	(Standard errors in parenthesis)			

[1]If either generation has a discontinuous work history the observation is deleted.

PSID, which shows larger elasticity estimates when the father's income is imputed than when the father's observed income is used. As mentioned above, B&J were not able to estimate the regression and correlation coefficients based on actual income because of lack of information on the father's income in SLLS. The estimates from SWIP are based on actual earnings or actual market income, for both fathers and sons. These estimates make it possible to contrast the Swedish intergenerational mobility with the American estimates based on actual earnings from PSID and reported by B&J.

The estimates provided in this paper can thereby be seen as candidates for the omitted estimates for SLLS in B&J's, i.e. those based on actual income in both generations. B&J find the American elasticity estimates to be 0.392 (std:

0.082) when corrected for age differences. When measured in logarithms and when including those with zero income for a given year, the result for the Swedish sample is lower and estimated to 0.125.

B&J estimate the correlation coefficients to 0.309 (std: 0.069) for the United States and the correlation coefficient based on SWIP is 0.114 based.

In this study, market income gives higher regression and correlation coefficients compared to earnings, which is consistent with the B&J results. The highest regression coefficient for the fathers and sons is found for the market income corrected for age (0.185). The differences between the Swedish estimates based on SWIP and B&J's estimates for the U.S. are significant and clearly indicate, when tested for, that intergenerational income mobility is higher in Sweden than in the United States.[11]

I now focus on the additional estimates of intergenerational mobility in the United States, based on the PSID and NLS. Do these results differ from the estimates found in B&J and if so, what might cause the difference? Zimmerman (1992) and Solon (1992) both estimate the intergenerational income mobility correlation in the United States to be around 0.4 after correction for several sources of bias. The result seems robust as the authors use different data. However, estimates based on the PSID seem to be sensitive to different sample selection criteria. This can be illustrated by results reported in Couch and Dunn (1996), who also use the PSID. Their estimation of the correlation in earnings of 0.168 for fathers and sons in the United States is considerably lower compared to both the Solon and B&J estimates on the PSID. Couch and Dunn argue that these differences are caused by different sample selections.

One reason for the discrepancy in estimates, is that B&J have selected their sample by observing the fathers and sons at approximately the same stage of their life cycles. Thus in the B&J study, the average age-difference for the generations when observed is only 10 years, while the equivalent average in the Couch and Dunn study is 20 years. The average age-difference in this study is 12 years between generations, and is thereby closer to the B&J selection. The effect of age differences between generations at the point of observation on the estimated coefficients is examined for these data, and the results do not seem overly sensitive to this restriction.[12]

Other explanations for the lower estimate in the Couch and Dunn estimation on the PSID data is due to their inclusion of sons aged 18 or older, while Solon (1992) uses a stronger restriction of an age 25 cut-off.[13] Finally the lower estimates for Couch and Dunn could be caused by the fact that they allow yearly observations of zero earnings to enter into their yearly average before taking the logarithm. Solon excludes individuals with discontinuous work-history in his analysis. Zimmerman (1992), on the other hand, analyzed only fully employed father and

[11] A test is performed to investigate whether $\rho^{US} > \rho^{SW}$ and the test-statistic is 3.22 for the difference between the regression coefficient and 2.79 for the correlation coefficients. With p-values < 0.001 and 0.0026 respectively. The test-statistic is asymptotically normally distributed.

[12] When excluding all those with more than 10 years of age difference when observed; for example the father's and son's earnings uncorrected for age rise from 0.119 to 0.127.

[13] The higher age 25 cut-off is also used here.

sons, which is not possible in this analysis due to lack of information. B&J compare the estimates attained with and without the zero-earners, and their results are not sensitive to this restriction. According to the estimation results found in Table 1, the same conclusion can be drawn for the data used here.[14] The estimates obtained by the PSID appear to be sensitive to the use of different restrictions; a fact that does not hold for the estimates based on SWIP.

The SWIP data used in this analysis appear to be affected primarily whether levels or logarithms of earnings are used. When the earnings are logarithmically transformed, the coefficient estimates decrease. This holds especially for the mother-children pairs, where the coefficients approach zero, i.e. there is no correlation between earnings of different generations. When using actual earnings, the regression and correlation coefficients vary for the mother-daughter pair between 0.110-0.139 and between 0.068-0.172 for the mother-son pair in Table 1.

The lower estimates for the models using logarithm of earnings can stem from the fact that the relationship between incomes of different generations is stronger at the upper end of the income distribution. This can be seen in Figure 1a-1d where earnings adjusted for life-cycle differences have been computed in deciles for both generations. Given the parent's income decile, the probability for children to end up in a certain decile is reported in these figures.[15]

If earnings mobility were independent across generations, no probabilities would differ greatly from 0.10. Most probabilities reported in Figure 1 are neither distinguishable from 0.10 except for the lowest and highest deciles. If the father belongs to the highest decile there is a 27 percent chance that the son will also end up in the highest decile. The equivalent figure for the mother-son pair is 17 percent, father-daughter pair 22 percent and 20 percent for the mother-daughter pair. In the case of father and son, lower mobility in the bottom decile can also be seen; there is a 19 percent chance for the son to end up in the lowest decile if the father had done so.

However, this has no parallel for the mother and children pairs. Given that the mother belongs to the lowest decile, the children have an almost equal possibility of ending up in any other decile. The reasons for mothers belonging to the lowest income decile may differ. First she could be constrained in the labor market and not have the possibility to supplying labor as she wishes. Another possibility is that she could have a husband with high earnings to reduce her labor supply. The fact of whether the small labor supply is voluntary or not will probably have different implications for the effect on the children.

In sum these data show clear non-linearities in the relation between income of subsequent generations. These results are in line with Johnson and Reed (1996), who show in a sample from the U.K. National Child Development Survey that the immobility of income over generations is most present at the extreme ends of the income distribution. Also Corak and Heiz (1996), using a giant tax-data set of more than 440,000 observations, display clear non-linearities in the intergenerational income mobility for Canadian fathers and sons. They find greater immobility at the bottom and top income deciles for fathers and sons, though the

[14]3.8 percent of the children had zero income for at least one year. 5 percent of the fathers had zero income for at least one year between 1978–80; the equivalent figure for the mothers is 18 percent.

[15]As variable of analysis earnings corrected for age is used.

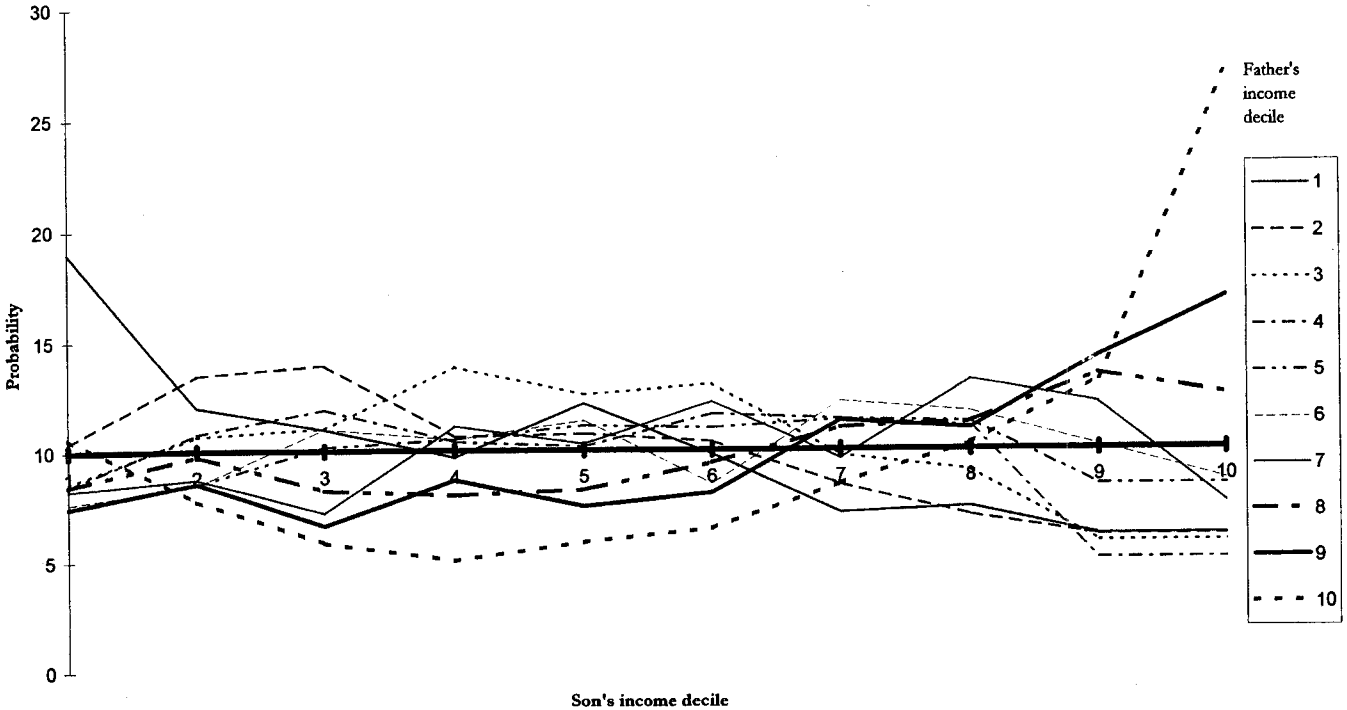

Figure 1a. Probability for son to end up in a certain decile given the father's decile

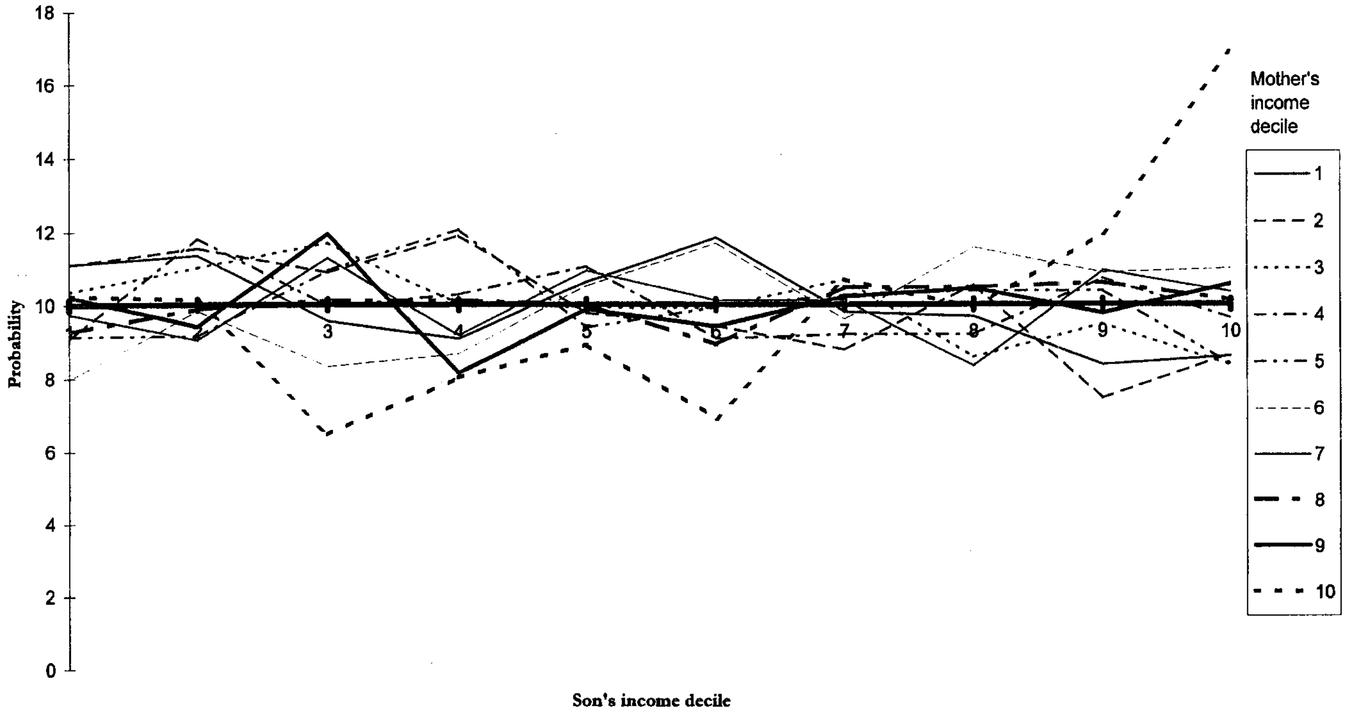

Figure 1b. Probability for son to end up in a certain decile given the mother's decile

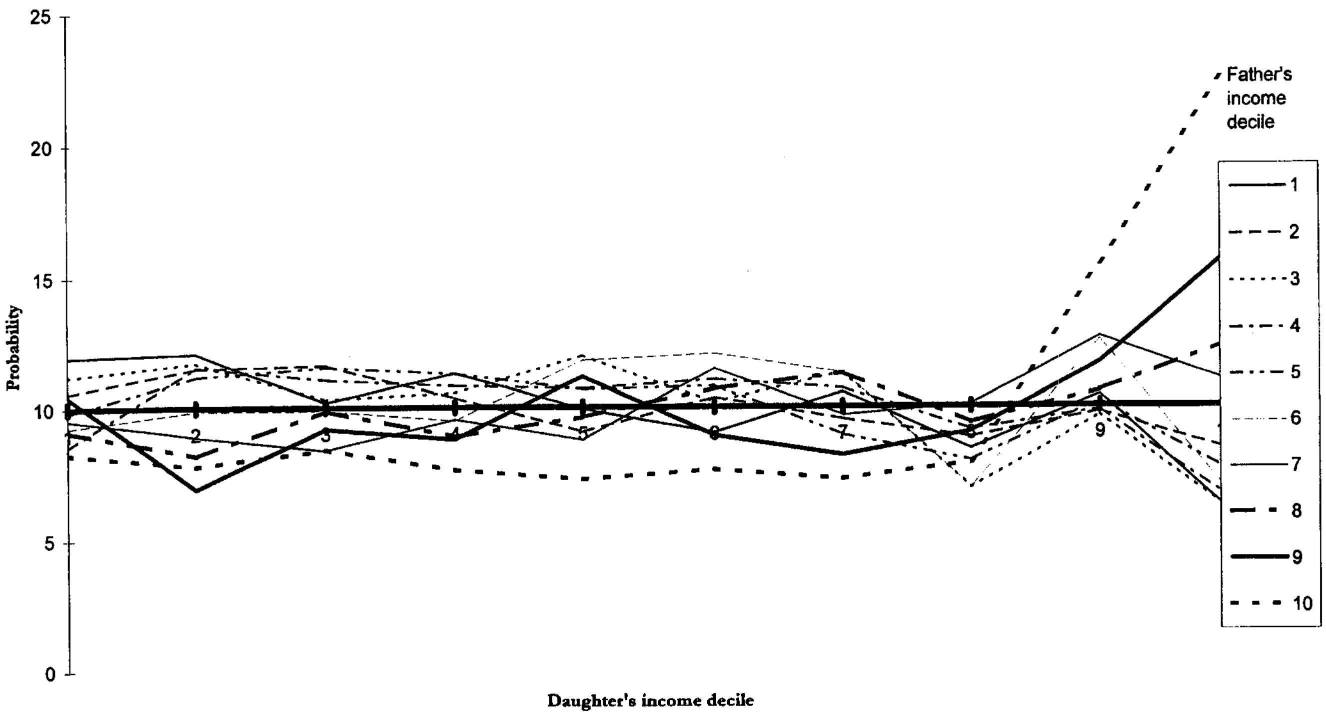

Figure 1c. Probability for daughter to end up in a certain decile given the father's decile

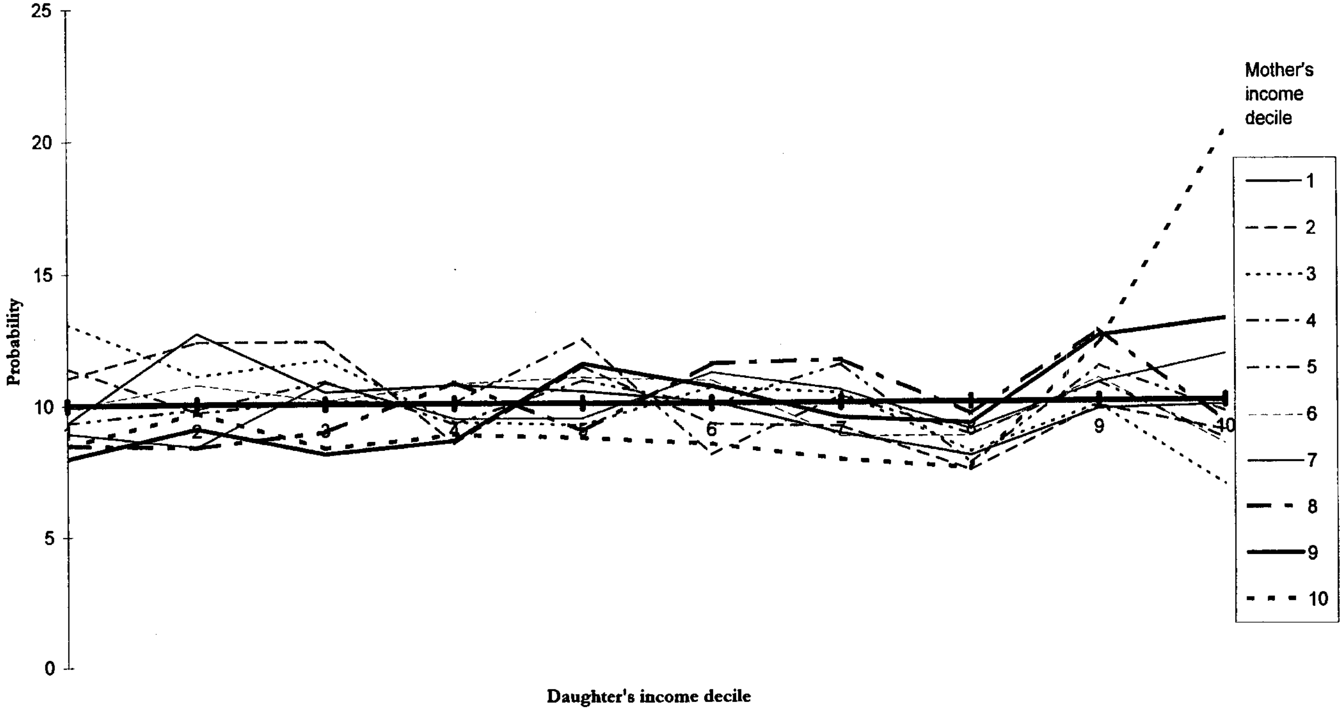

Figure 1d. Probability for daughter to end up in a certain decile given the mother's decile

mobility is larger than the one found in this analysis at the highest and lowest deciles.

I now turn more explicitly to estimates for child-parent pairs, other than that of father-son, from Table 1. The results show that the regression coefficients for fathers and daughters are about half the magnitude of the regression coefficients for the pairs of father and sons. The elasticity estimates for the father-daughter pairs are 0.062–0.083, depending on different sample selections, while the father-son elasticity estimates are in the range 0.125–0.185.

Peters (1992), Lillard and Kilburn (1995) and Dearden *et al.* (1997), who also estimated daughters' intergenerational income mobility, have shown on the other hand a greater mobility for sons with respect to their parent's income, in comparison to daughters. However, the Couch and Dunn results support the findings in this analysis. They estimate correlations not only for the father-son pair but also for the mother-daughter pair. The mother-daughter pair tends to have greater mobility compared to the father-son pair in both the German data (GSOEP) and the American (PSID). For the mother-daughter pair in GSOEP the correlation is insignificant and approaches zero, a result also valid in this analysis. Here the correlation between fathers and daughters is also lower than the father-son correlation, and varies between 0.069–0.140 compared to 0.114–0.182 for the father-son pair. The impact of the mother's income on the children's income is weaker than the father's. The regression and correlation coefficient varies between 0.018–0.0172 in relation to the son and 0.025–0.139 in relation to the daughter.

Finally mobility matrices are shown in Table 2. Sons, daughters, fathers and mothers are divided into four different income classes to examine the intergenerational income mobility in mobility matrices. The limits of the income classes are the same as applied by B&J: **poor**—those with earnings below 50 percent of sample median earnings; **lower middle income**—those with earnings from 50 percent of the median up to the median; **higher middle income**—from the median up to 1.5 times the median and **well-to-do**—earnings above 1.5 times the median earnings.

B&J point toward larger inequalities in earnings in the United States than in Sweden, and their conclusion is further strengthened by the results presented in this paper. Only 10 percent of the fathers can be defined as poor in this analysis, compared to 15 percent in B&J based on SLLS and 20 percent for the U.S. based on PSID. This indicates that Sweden has a more equal earnings distribution than the United States.

The conditional probabilities reported here are also consistent with the B&J results. The most dominant signs of immobility are apparent in the middle income classes. Given that the father can be found in the lower middle income class (2), the son has a 45 percent chance of ending up in the same income class. The immobility in the highest income classes, shown in Figure 1a–1d, is, however, not possible to detect when the income classes are defined in this broad manner. More than 60 percent of the children have moved to an income class other than that of their father's. In relation to the mother the mobility is even higher: about 70 percent of the children are in an income class that is different from their mother.

TABLE 2

MOBILITY MATRICES

Mobility matrices for children and parents.
The probability conditional on parent's income class.

Father's Income Class	Son's Income Class 1	2	3	4	Daughter's Income Class 1	2	3	4
1	17.4	44.7	30.5	7.3	9.2	45.5	38.3	7.0
2	8.9	44.8	37.3	9.0	7.9	45.4	39.4	7.3
3	8.4	38.4	41.4	11.7	7.7	40.4	42.4	9.6
4	12.9	26.3	35.5	25.3	8.3	35.0	38.3	18.7
		$n = 8{,}336$				$n = 8{,}053$		

Mother's Income Class	Son's Income Class 1	2	3	4	Daughter's Income Class 1	2	3	4
1	11.2	40.8	36.2	11.8	8.3	43.3	37.8	10.6
2	10.1	41.9	36.2	11.7	9.0	42.2	40.6	8.2
3	9.4	40.4	38.7	11.6	7.2	43.0	40.4	9.4
4	11.0	35.3	38.6	15.0	7.0	38.0	40.5	14.5
		$n = 8{,}753$				$n = 8{,}387$		

Mobility matrices for Swedish children and parents.
Unconditional bivariate probabilities.

Father's Income Class	Son's Income Class 1	2	3	4	Daughter's Income Class 1	2	3	4
1	1.64	4.24	2.89	0.70	0.90	4.46	3.76	0.69
2	3.62	18.17	15.14	3.65	3.19	18.21	15.83	2.92
3	2.86	13.02	14.04	3.97	2,66	13.90	14.59	3.31
4	2.08	4.22	5.70	4.06	1.29	5.45	5.93	2.91
		$n = 8{,}336$				$n = 8{,}053$		

Mother's Income Class	Son's Income Class 1	2	3	4	Daughter's Income Class 1	2	3	4
1	2.05	7.46	6.62	2.15	1.43	7.48	6.52	1.83
2	3.24	13.46	11.64	3.77	2.90	13.61	13.09	2.63
3	2.57	11.07	10.61	3.18	2.08	12.44	11.70	2.70
4	2.43	7.84	8.57	3.34	1.51	8.21	8.74	3.14
		$n = 8{,}753$				$n = 8{,}387$		

Income classes are defined according to: 1—Poor < 50 percent of median; 2—Lower middle 50 percent of mediant to median; 3—Higher middle median to 150 percent of median; 4—Well-to-do ≥ 150 percent of median.

CONCLUSIONS

The results reported in this analysis based on observed income on dependent generations, indicate a high intergenerational income mobility in Sweden between fathers and sons, compared to estimations from most other countries. The correlation between father's and son's income varies between 0.11 and 0.18, depending on different restrictions. The estimation results appear stable and are not seriously affected by the inclusion of those with a discontinuous work-history. Neither are the estimation results altered when pairs with more than 10 years of age difference at point of measurement are excluded. The main differences in estimation results arise when logarithmic values of income are used.

More conclusive evidence than the Björklund and Jäntti (1997) results has been provided showing that the intergenerational correlation in earnings appear to be higher in the United States than in Sweden. With the tests given, the hypothesis that Sweden does not only have lower cross-sectional income inequality, but also higher intergenerational income mobility can be accepted. The estimate of the correlation in earnings across generations is 0.114 in Sweden, when based on SWIP and corrected for age; this in comparison to 0.309 which is Björklund and Jäntti's estimate for the United States. The estimations of Zimmerman (1992) and Solon (1992) of around 0.4 for the United States do not contradict my results.

This analysis also provides estimates for the other generational pairs. The father's earnings seem to correlate more weakly with the daughter than with the son. The influence of the mother's earnings on her children's earnings is weaker compared to the influence of the father's earnings. The mother's earnings correlation with her children is however stronger for the daughter compared to the son; a sign of a role-model effect.

Earlier studies argue that it is essential to observe individuals in both generations for several years and approximately at the same stage of the life-cycle, and not to use homogenous samples for the reduction of bias in the estimates of intergenerational mobility. It should also be added that the use of large sample and non-linear methods could also be important. In this analysis the highest income classes tend to be the least mobile. The fact that the father belongs to the tenth income decile triples the chances for the son to end up in the equivalent decile. In the case of mother and son, the chance is nearly doubled. To detect these signs of immobility, large samples are needed to arrive at a good description of intergenerational income mobility.

References

Atkinson, A. B. , On Intergenerational Income Mobility in Britain, *Journal of Post-Keynesian Economics*, 3, 2, 194–217, 1981.

Behrman, J. and P. Taubman, Intergenerational Earnings Mobility in the United States: Some Estimates and a Test of Beckers' Intergenerational Endowments Model, *The Review of Economics and Statistics*, 67, 144–51, February 1985.

Björklund, A. and M. Jäntti, Intergenerational Income Mobility in Sweden Compared to the United States, *American Economic Review*, 87, 5, December 1997.

Couch, K. A. and T. A. Dunn, Intergenerational Correlations in Labor Market Status—A Comparison of the United States and Germany, *The Journal of Human Resources*, XXXII, 1, 1996.

Corak, M. and A. Heiz, The Intergenerational Income Mobility of Canadian Men, *Analytical Studies Branch Research Paper Series*, 89, Statistics Canada, 1996.

Dearden, L., S. Machin, and H. Reed, Intergenerational Mobility in Britain, *The Economic Journal*, 107, 47–66, January 1997.

Gustafsson, B., The Degree and Pattern of Income Immobility in Sweden, *Review of Income and Wealth*, 40, 1, 67–86, March 1994.

Johnson, P. and H. Reed, Intergenerational Mobility Among the Rich and Poor: Results from the National Child Development Survey, *Oxford Review of Economic Policy* 12, 1, 1996.

Lillard, L. A. and M. R. Kilburn, Intergenerational Earnings Links: Sons and Daughters, RAND: Santa Monica, *Labor and Population Program, Working Paper, Series* 95–17, 1995.

Peters, H. Elisabeth, Patterns of Intergenerational Mobility in Income and Earnings, *Review of Economics and Statistics*, 74, 456–66, 1992.

Solon, G., Biases in Estimation of Inter-generational Earnings Correlations, *The Review of Economics and Statistics*, 71, 172–4, 1989.

Solon, G., Intergenerational Income Mobility in the United States, *American Economic Review*, 82, 3, 393–408, June 1992.
Zimmerman, D. J., Regression Toward Mediocrity in Economic Stature, *American Economic Review*, 83, 409–29, June 1992.