DISCUSSION PAPER SERIES

IZA DP No. 797

Assessing Changes in Intergenerational Earnings Mobility

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June 2003

Forschungsinstitut
zur Zukunft der Arbeit
Institute for the Study
of Labor

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Discussion Paper No. 797
June 2003

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IZA Discussion Paper No. 797
June 2003

ABSTRACT

Assessing Changes in Intergenerational Earnings Mobility*

Previous research on changes in intergenerational mobility suggests that mobility is decreasing over time. One explanation for this pattern is increased cross-sectional income inequality. In contrast to most other OECD countries, income inequality in Norway has been remarkably stable through large parts of the 1980s and the 1990s, not least due to a compression of the earnings distribution during the same period. Using longitudinal data for Norwegian children born in 1950, 1955, 1960, and 1965, we find a relatively high degree of earnings mobility. Furthermore, there is no tendency to increasing inequality along this dimension. This finding supports the hypothesis that intergenerational mobility is positively correlated with a compressed income distribution. Quartile father-child earnings transition matrices, together with non-parametric regressions, indicate quite high mobility in the middle of the distribution and somewhat more persistence at the top and bottom. This approach also reveals increased mobility over time for sons, but a less clear picture for daughters.

JEL Classification: J62, C23

Keywords: intergenerational mobility, intertemporal change

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* The data were provided by the Norwegian Social Science Data Services, who share no responsibility for the authors' conclusions. We would like to thank The Norwegian Research Council for financial support, grant no 140127/530 ("Velferdsprogrammet"). The paper has benefitted from comments and suggestions at presentations at the ESPE 2002 meeting in Bilbao, Zentrum für Europäische Wirtschaftsforschung, Mannheim, and the HERC Centre, University of York. The paper was partly written while Bratberg was visiting the University of York, whose hospitality is gratefully acknowledged.

1 Introduction

Although it is hard to imagine a society without inequalities, equality of opportunity seems a commonly agreed-upon ideal. Equal opportunities do not necessarily imply an equal income distribution, but do seem less compatible with a society which is immobile in the sense that families remain in the same position in the income distribution year after year and generation after generation. Indeed, one interpretation of the modern welfare state is that it attempts to reduce the importance of family background for an individual's economic failure or success. In particular, there is a concern that parents' poverty might be passed on to their offspring, implying persisting income inequalities at any point in time. Many institutions and regulations, like progressive taxation, subsidies for education, etc., are (partly) motivated by society's attempts to increase intergenerational mobility.

Against this background we have witnessed a growing body of research on intergenerational correlations of social and economic status, see Solon (1999) for a recent review. Of course, the discussion of parental heritage is a broad one, including income and wealth, genetic endowment, cultural values, social skills, etc.[1] In this paper, however, the focus is on the correlation of earnings across generations. The seminal papers within this branch of the literature are Becker and Tomes (1979, 1986). Based on their theoretical model where utility-maximising families invest part of parents' earning in the human capital of the children, they

[1] The closely related sociological literature on the mobility of social classes and reproduction of inequality is vast, even in Norway, and predates much of the research performed by economists, see Ganzeboom et al. (1991) and Björklund and Jäntti (2000).

performed empirical analyses that concluded with a relatively high degree of mobility. More recent studies, notably Solon (1992) and Zimmerman (1992), have reached other results. By means of improved data it has been possible to replace single-year earnings with long-run income measures, and to avoid the homogeneity of the samples used in many of the early studies. Homogeneity and short-run instead of long-run income both have the same effect: it leads to downward bias in the estimated mobility parameter. Solon (1999) concludes that "... 0.4 or a bit higher [...] seems a reasonable guess of the intergenerational elasticity in long-run earnings for men in the United States." (p.1784), and this is about twice the magnitude of the elasticity estimated by Becker and Tomes (1986).[2]

Even though most of the literature refers to US longitudinal surveys, several studies based on data from European countries are available. For the UK, Atkinson et al. (1983) and Dearden et al. (1997) are comparable to Solon (1992) in terms of data and methodology. They report intergenerational elasticities between fathers and sons even slightly higher than what is found in the US. Jäntti and Osterbacka (1999) and Österbacka (2001) report an intergenerational elasticity in Finland of about 0.2. Gustafsson (1994), Björklund and Jäntti (1997) and Österberg (2000) report elasticities based on intergenerational data from Sweden. They all find that the elasticity is considerably lower and Sweden therefore seems to be a more mobile society than the US and the UK. Even though Björklund and Jäntti (1997) apply a different estimation method, their conclusion remains unaltered after using the same method in US surveys. The indication of higher

[2] Lillard (1998) argues, however, that there is no consensus in the literature, and that the main reason for the diversity is the variation in sample selection rules.

mobility is robust or, if anything, even strengthened in Österberg (2000), whose data and method are comparable to Solon (1992).[3]

The cross-sectional income distribution in the US (and, for that matter, the UK) on the one hand and in the Scandinavian countries on the other is strikingly different. Accordingly, several authors have suggested that the relatively low cross-sectional income inequality in Scandinavia is associated with higher intergenerational mobility. As noted by Björklund and Jäntti, the question as to whether equality of opportunity and equality of outcomes are independent of each other can be illuminated by further international comparisons of intergenerational income mobility. For Norway, a country with income inequalities about the same range as in Sweden, low intergenerational elasticities would be a support for the alleged positive correlation between cross-sectional inequality and cross-generational mobility.

There is, however, another way to shed light on this issue. For instance, changes in the cross-sectional income distribution from one period to another should be reflected in changes in the intergenerational mobility during the same period of time. Differences in policy, institutions, etc. from one period to another within one country might be less than across country-differences in the same factors. If so, we should expect less noise in within-country than in across-country analysis.

The assessment of changes in intergenerational mobility is difficult not least due to extremely strong data requirements. Mayer and Lopoo (2001), Chadwick

[3] There are no previous studies of intergenerational earnings mobility in Norway. However, Raaum et al. (2001) consider sibling and neighbourhood correlations.

(2002), and Blanden et al. (2002) are among the very few studies that address this issue. The former two are based on US data and the latter on data from the UK. Both the US and the UK are examples of countries that have experienced a relatively steep increase in income inequality since the late 1970s. Blanden et al. (2002) find that mobility appears to have fallen when they compare the 1958 with the 1970 birth cohort. Chadwick (2002), investigating the period 1976 to 1996, finds that intergenerational mobility has decreased during this period for daughters, and possibly for sons. Mayer and Lopoo (2001) report the opposite when they compare the 1949-1952 and the 1960-1962 cohorts, but suggest that this might be a consequence of the increased state investment in education over the period in question.

Norway is, as mentioned, a country with small cross-sectional income inequalities compared to other OECD countries, see e.g., Atkinson et al. (1995). Moreover, in the preceding 2-3 decades, when most other comparable countries have been experiencing increased inequality, wage distribution in Norway has been relatively stable, and has even become more compressed at the bottom (Kahn, 1998). Income inequalities also remained quite stable through the part of the 1980s and 1990s when unemployment increased (Aaberge et al (2000)). Armed with data for the 1950, 1955, 1960, and 1965 cohorts and their families, we estimate separately the intergenerational mobility for each cohort. Hence, the cohort comparison constitutes an interesting experiment along two dimensions. On the one hand, we are able to test whether the relatively stable income distribution has a parallel in stable mobility parameters. On the other hand, we can do crosscountry comparisons. For example, to the degree that reduced mobility in the UK is explained by increasing income inequality, we have no reason to expect similar movements in Norway.

The data used in our analysis are based on census data that are linked to public register data on earnings, education, marital status, and migration. The database contains information on the full population of children born every fifth year between 1950 and 1990. Each child may be linked to parents and grandparents. Importantly, our data contain annual gross taxable earnings for the period 1967-1995 for the generations that are linked together. We focus on today's middle-aged: the cohorts born from 1950 to 1965. For the two oldest of these cohorts we are able to track their earnings trajectories at least until the age of 35, and earnings at 30 are available for them all. The problems of homogeneous sample and/or short-term earnings measure do not, therefore apply, neither do the problems of small samples (Mayer and Lopoo, 2001) or different data bases for the different cohorts (Blanden et al., 2002).

2 Estimating earnings correlations

The effect of parents' economic status on the status of a child may be formulated simply as

$$y_i^c = \alpha + \beta y_i^p + \epsilon_i, \tag{1}$$

where c and p denote "child" and "parent", respectively, y is a measure of "lifetime" or "permanent" income in logs, β is the slope coefficient, and ϵ_i is a random error term. This reduced form may be motivated by utility-maximising families investing a part of parents' earnings in the human capital of children, cf. Becker and Tomes (1979, 1986), Solon (1999). The closer to zero β is, the higher is intergenerational mobility. Often the focus is on the parent-child earnings correlation coefficient, ρ. If $var(y_i^c) = var(y_i^p)$, ρ equals the OLS estimator of β. Otherwise, $\rho = \widehat{\beta}^{OLS} \frac{SD(y_i^p)}{SD(y_i^c)}$.

Simple as it may be, a major obstacle to estimating equation (1) is obtaining estimates of permanent income. Typically, only a single or a few observations for each generation is available. This leads to attenuation bias: If one uses period t earnings as a proxy for permanent income, we may write this as $y_{it}^g = y_i^g + \eta_{it}^g, g = c, p$, where η_{it}^g is a transitory component consisting of observed and unobserved factors. Due to unobservables, inserting y_{it}^g into (1) and estimating the equation by OLS leads to classical errors in variable problem even after controlling for observables. Thus the estimate of β will be biased towards zero.

We may decompose the relation between observed and permanent status further. Assuming linearity in the parameters, the observed income for generation g in period t is

$$y_{it}^g = y_i^g + \gamma_g' x_{it}^g + u_{it}^g, \tag{2}$$

where x_{it}^g is a vector of time-varying characteristics, γ_g is a vector of coefficients, and u_{it}^g is a random term. Typically $\gamma_g' x_{it}^g$ consists of an age profile. With a long panel available, an estimator of y_i^g would be a panel fixed effect. Typically that is not feasible due to the limited time-span of the data. Given an estimate of γ_g, however, an available estimator of y_i^g is the residual from a regression of equation (2),

$$\widehat{y}_i^{g,R} = y_{it}^g - \widehat{\gamma}_g' x_{it}^g = y_i^g + w_{it}^g, \tag{3}$$

where $w_{it}^g = (\gamma_g - \widehat{\gamma}_g)' x_{it}^g + u_{it}^g$ is an error and we let the superscript R denote "residual". If we allow that $\widehat{\gamma}_p \neq \widehat{\gamma}_c$, it is unnecessary to estimate the model in the two steps suggested by (3). Instead one could solve the two equations given by (2) for $y_i^g, g = c, p$, and substitute these into (1), to obtain the estimable equation

$$y_{it}^c = \alpha + \beta y_{it}^p + \gamma_c' x_{it}^c - \beta \gamma_p' x_{it}^p + u_{it}^c - \beta u_{it}^p + \epsilon_i. \tag{4}$$

Using the residual estimator in (1) still gives a downwards bias due to errors in variables. It is straightforward to show that if w_{it}^g and y_i^g are uncorrelated, the probability limit of this estimator is $\beta \frac{var(y^p)}{var(y^p)+var(w^p)}$. However, the bias may be reduced by averaging values of y_{it}^g and x_{it}^g for several years. If w_{it}^p is white noise, a T-period average is biased by a factor

$$\frac{var(y^p)}{var(y^p) + \frac{var(w^p)}{T}} > \frac{var(y^p)}{var(y^p) + var(w^p)}.$$

If the white noise assumption does not hold, the expression is more complicated, cf. Zimmerman (1992) for details. Alternatively, one may use instruments for parental income. If the instruments also belong in a structural equation for children's income and have positive effects, the IV estimator is biased upwards, cf. Solon (1992), Zimmerman (1992). In that case, OLS and IV estimates represent lower and upper bounds of β, respectively.

An alternative to the regression approach described above is to compute parent-child quantile transition matrices instead of estimating β or ρ. One simply divides the income distributions of children and parents into q percentiles and compute transition matrices for the fraction of children that belong to each percentile j, given that the parent belongs to percentile k for all (k, j) pairs. In a perfectly mobile society the fraction in each cell will be $\frac{1}{q}$, whereas in a perfectly immobile society, all the diagonal elements will be 1 with 0s elsewhere. In this paper we follow most of the received literature and report quartile transition matrices. An advantage of this method is that it provides insights into which parts of the income distribution the intergenerational mobility, or lack of such, is largest. The issues related to measurement problems discussed above clearly relate to the income measures used in transition matrices, too, and one should use as good a measure of permanent income as available for both generations.

3 Data and sample

The data in this study are extracted from the Norwegian Database of Generations (DBG), which contains information from several public registers, merged by Statistics Norway. This database includes the full cohorts born every fifth year in the period from 1950 to 1990, merged with data on their parents and grandparents. Thus the observational unit is at the individual level (the child). For each generation there is information about family characteristics, education and variables describing the labor market attachment of the individuals, together with information about income. The data sources are:

- Censuses 1960-90 (10-yearly), with information such as family type, housing type, area of residence, parents' year of birth, occupation, education and labour market participation. There is also some information on grandparents.
- Income registers (based on tax reports) with yearly earnings from 1967 to 1995 for parents and children.
- Other public registers that provide information on childbirths, migration, and changes in marital status and education, for children as well as their parents. These are recorded in the database as monthly events.

Our sample consists of the four cohorts born from 1950 to 1965. The main focus will be on the 1950 and 1960 cohorts, whose 10-year spacing is comparable to the research by Blanden et al. (2002), and who also have a reasonable range of adult earnings available (until the age of 35 for the 1960 cohort).

We have limited our study to individuals whose father was younger than 40 when the child was born. That is because our earnings series starts in 1967, when someone born in 1910 was 57 years old. We use parents' earnings until the age of 65, when retirement becomes significant. (The official retirement age is 67, but many wage earners have the opportunity to retire at 65.) We also exclude individuals with missing information on certain variables that are potential instruments: fathers' education, area of residence, housing type, and house ownership. Census data is taken from the 1960 censuses for the fathers of the 1950 and 1955 cohorts, and from 1970 for the 1960 and 1965 cohorts.

The following (log) earnings measures are used for sons and daughters: Averages 1980-84 for the 1950-cohort and 1990-94, along with earnings at age 29 and 30 for all cohorts except those born in 1955, whose 1985 earnings are not used. This exclusion is due to information from the data providers that the earnings data from 1985 are of low quality. Moreover, inspection of the earnings data from 1986 and 1987 has given rise to suspicion of problems also for these years, and have been excluded from the analysis, too. The years 1970 and 1971 have turned out to contain an excessive number of zero earnings.[4] We therefore exclude those years in the analysis to be reported. In all averages, years with zero earnings are excluded, but we do not exclude individuals without a complete series. E.g., if one out of three years is zero when computing the three-year average, we use the average of (the log of) the two remaining years. This is in line with previous research using similar data, e.g. Österberg (2000).[5]

(Table 1 about here)

Table 1 shows descriptive statistics for sons and daughters, respectively, along with their fathers.[6] There are fewer observations for the 1950-cohort than for the younger ones. Moreover, there are only half as many daughters as sons in this

[4]The data coding makes it impossible to distinguish true zeroes from missing income information.

[5]We have checked the sensitivity of the results to conditioning on complete series of years with positive income in the averages.

[6]The sample sizes are defined by 1. non-missing information on the background information mentioned in the text, 2. at least one year of non-zero earnings for the relevant three years for fathers, 3. non-zero earnings either at age 29 or 30 for children. In some of the regressions the sample sizes will be slightly larger due to using longer averages.

cohort. This is related to the way the data were constructed: the initial linking of children to parents is based on the 1970 census, identifying the parents of children who were still living with their parents at that time. Obviously the tendency to leave the parental home before the age of 20 was larger for young women than men in 1970. Neither can we exclude the possibility that young women undertaking education are over-represented in the 1950 sample. These problems are discussed in the analysis. As to the statistics, we note that the average age of fathers of 17-year-olds has decreased by about two years from the 1950 to the 1965 cohort, whereas the mean and variance of earnings have increased, as has the educational level. For the younger generation, the education level also has increased, and in the 1960 cohort more than 90 per cent have education above primary school level.

4 Results and discussion

The analysis proceeds as follows. First we compare son-father correlations for the 1950 and 1960 cohorts, where five-year averages of sons' earnings (age 30-34) are available. Regressions of earnings at ages 29 and 30 are then carried out for all the cohorts (except age 30 for the 1955-cohort), and the results are interpreted in the light of the results where longer averages for the sons are available. The same analyses are then performed for daughters and fathers. Quartile transition matrices are then reported for the 1950 and 1960-cohorts. Finally, to address potential nonlinearity in the transmission coefficients, we perform non-parametric regressions of children's earnings on fathers' earnings. In all the regressions, fathers' earnings are adjusted by a quadratic age profile, i.e. in equation (2),

$x_{it}^f = (age_{it}^f, (age_{it}^f)^2)$. As there can be no age variation within each cohort of sons or daughters, their earnings are unadjusted.

(Table 2 about here.)

Table 2 serves a dual purpose: it compares the 1950 and 1960 cohorts, and it addresses the measurement issue discussed in Section 2 by comparing results using a three-year average of fathers' earnings results using single years and two-year averages. Sons' earnings are averages 1980-84 and 1990-94, respectively. We report regression coefficients (β), as well as correlations (ρ). For the 1950 cohort, we find that for single-year measures of fathers' earnings, the regression coefficients are quite sensitive to which year is chosen, but the correlations are more stable. For the three-year average 1967-69, we find that $\beta = 0.165$ and $\rho = 0.145$. In the 1960 cohort there is less variation in the results. Using the three-year average, $\beta = 0.123$ and $\rho = 0.109$.[7] Thus the association between fathers' and sons' earnings appears to have been reduced, i.e. intergenenerational mobility has increased.

(Table 3 about here)

In Table 3 we investigate further the revealed pattern of increasing intergenerational mobility, found in Table 2, by regressing sons' earnings at age 29 and 30 on three-year averages of fathers' earnings.[8] It is reasonable to expect that

[7] As noted in Section 3, there are excessive number of missing in 1970 and 1971, and we have therefore abstained from using averages involving those years. Regressions on these years deviated from those reported here by being much smaller. On the other hand, regressions for the 1960-cohort involving 1980 and 1981 were quite similar to those in the fourth column of Table 3, and a five-year average gave almost identical results to the three-year average.

[8] As noted before, we only report age-29 results for the 1955 cohort because of unreliable 1985 earnings data.

a trend in the single-year correlations would carry over to the 30-34-correlation. Moreover, Mayer and Lopoo (2001) and Blanden et al. (2002) in their similar comparisons also use earnings at about 30 (respectively 30, and 30 for the older and 33 for the younger cohort). For sons, we find a reduction in β as well as ρ at both ages. For a 30-year-old in 1980, the elasticity was 0.100 and the correlation 0.096, reduced to 0.068 and 0.075 for a young man at the same age in 1995. Thus the trend apparent in Table 2 is verified in Table 3, and it seems reasonable to claim that the earnings correlation between young men and their fathers has actually decreased from the early eighties to the mid-nineties.

(Table 4 and 5 about here)

Tables 4 and 5 report results according to tables 2 and 3 for women and their fathers. The sample selection problems discussed in the previous section may cast doubts on the results for the 1950 cohort. With this qualification in mind, in column three of Table 4 we observe a quite high elasticity (0.222) for the regression on the three-year average 1967-69, but a moderate correlation coefficient of 0.091. The difference between the two measures, not found for sons, reflects the large variance in daughters' earnings. If some selection actually is on education, again indicating that the fathers invest in daughters' human capital to a larger degree, we would expect an upward bias in the association between daughters' and fathers' earnings. This notion is consistent with a regression coefficient that seems too high compared to sons, as well as the received literature, which quite unambiguously finds lower associations between daughters and fathers than sons and fathers. Turning to the 1960 cohort, we find $\beta = 0.137$ and $\rho = 0.087$ using the three-year average for fathers' earnings. A conservative judgement is that earnings mobility for women has not decreased from the 1950 to the 1960 cohort. Turning to Table 5, this seems to warrant a conclusion that there is actually a trend towards an even smaller daughter-father earnings association: for 29-year-olds, ρ has decreased from 0.083 for the 1955-cohort to 0.054 for the 1965-cohort, and β from 0.147 to 0.065. The numbers at age 30 are not available for the 1955 cohort, but for the 1960 and 1965 cohorts they are almost identical.

In addition to the reported results, we have estimated several instrumental variable (IV) regressions on single year earnings. As noted in Section 2, IV may reduce the bias from using single (or a few) years earnings to measure permanent earnings, but may also be biased upwards. We have used as instruments groups of dummy variables indicating fathers' educational level, housing type, house ownership and whether living in an urban area. However, the IV results are very sensitive to what instruments are included and in some cases give elasticities that seem unrealistically high. Moreover, with very few exceptions the exclusion restrictions are rejected by Sargan tests. We have therefore decided not to report them. On the other hand, when checking the sensitivity of the results to extending fathers' earnings from a three-year to a five-year average (available for the 1960 cohort but not the 1950 cohort), the results were quite similar to those reported.

(Table 6 about here)

Elasticities and correlations as reported above provide nice summary measures of intergenerational mobility but build on the implicit assumption that the mobility is independent of position in the earnings distribution. Therefore, we have also computed transition matrices for the 1950 and 1960 cohorts, reported in Table 6. Fathers' earnings are measured by three-year age-adjusted averages, and children's by the average at age 30-34. The following observations are conspicuous: 1: There is a larger tendency to persistence in the upper and lower parts of the earnings distribution. 2: This tendency is larger in the upper than in the lower quartile. 3: The second point is more accentuated for women than for men, i.e. they tend to have larger upward mobility from the bottom quartile. Comparing the cohorts is not straight-forward, as each matrix is made up of 24 elements. The results for the 1950 cohort of women may also be hampered by non-random missing observations. With this caveat, comparisons may be facilitated by using an "immobility index" computed as the sum of the leading diagonal and adjacent cells, following Dearden et al. (1997) and Blanden et al. (2002). "Perfect mobility" implies that the probability in each cell is 0.25 with an index of 2.5, whereas "perfect immobility" would imply an index of 4. Judged by this criterion, mobility has increased for men, with an index reduced from 2.858 to 2.799. The index for women has actually increased, but we are skeptical of the result for the 1950 cohort. However, we also observe that for the 1960 cohort, the immobility index is lower (2.725) for women than for men.

(Figure 1-4 about here.)

The lack of symmetry in the transition matrices may raise doubt about the linear restriction imposed on the coefficients in the regression approach. To supplement the regression results hitherto reported, therefore, we have carried out non-parametric regressions of sons' and daughters' earnings on their fathers, for the 1950 and 1960-cohorts. The earnings measures still are three-year averages 1967-69 and 1977-79 for fathers, and average at age 30-34 for sons and daughters. We apply the lowess (locally weighted regression) estimator, first suggested by Cleveland (1979).[9] This estimator essentially draws a smooth curve through a bivariate scatterplot by regressing each observation of the dependent variable on the neighbouring observations of the independent, with weights that are a decreasing function of distance.[10] Figures 1-4 show the results together with the linear prediction, with the vertical lines indicating fathers' 10th, 25th, 50th, 75th, and 90th centiles. It is difficult to compare the cohorts based on these graphs, but a similar picture emerges from them all: children's earnings are U-shaped functions of fathers' earnings, and the pattern seems somewhat more accentuated for women. It deserves mention that the slope (elasticity) is lower in the lower end of the distribution, even though the tendency to remain in the fathers' earnings quartile is larger here than closer to the median. On the other hand, the steeper slope in the other end of the distribution is consistent with the tendency to "stay on top". It must be noted that for the 1950-cohort of sons, the slope flattens at the very top of the distribution. The lesson learned from this exercise, along with the transition matrices, is that even though β (and ρ) provide relevant parameters describing the intergenerational earnings correlation, relying on them alone may conceal important characteristics of the dynamic relationship.

[9] A similar approach is used by Corak and Heisz (1999).

[10] Each y_i is regressed on the $\frac{1}{2}kN$ neighbouring values x-values on each side of x_i, where N is sample size and k is a smoothing parameter, and each x_j is weighted by a decreasing function (the tri-cubic) of the distance from x_i. The predictions $\widehat{y}_i$ are then used to draw a line through $(\widehat{y}_i, x_i), i = 1, ..., N$. The larger k is, the larger the degree of smoothing. We used $k = 0.8$, but have also experimented with lower values. The software Stata was employed.

We conclude this section with a short comparison of our results with some previous findings. Österberg's (2000) results from Sweden apply similar earnings data, where parents are observed in 1978-80 and children 1990-92. Thus her time window is comparable to the one we use for the 1960 cohort, but the second generation in her sample on average are five years older (37 in 1992). Her estimates of ρ, 0.114 for sons and 0.069 for daughters are quite similar to those found in our tables 2 and 4. For Finland, Österbacka (2001) find somewhat higher correlations, 0.156 and 0.121 for sons and daughters, respectively. However, the young generation in her sample is born between 1950 and 1960, with earnings measured in 1985, 1990, and 1995, so we must take into consideration that her sample is on average older.

As regards other assessments of changes over time, it is notable that our results differ from those of Blanden et al. (2002) in the UK and Chadwick (2002) in the US in that they find an opposite tendency: intergenerational mobility seems to have decreased. Even though there are some differences in the age and income measures of children and parents, that is hardly enough to explain the reverse tendency. Rather, it seems likely that the diverging results are due to differences in the countries' cross-sectional income inequalities. Blanden et al. and Chadwick both explain the decreasing mobility across cohorts with the increase in the intragenerational inequalities in the respective countries. Our findings suggest, in accordance with the UK and US results, that the compression of the earnings distribution in Norway is associated with increasing generational mobility.

In contrast to Chadwick, Mayer and Lopoo (2001) report increasing income mobility for the US, and suggest that this might be a consequence of increased state investment in education. The large expansion in educational achievement during the previous 2-3 decades is, however, common to the UK and Norway. But as noted by Blanden et al., this development is not necessarily associated with increasing mobility. If children from lower income families were relatively more benefited by the expansion than rich ones, then the mobility would rise. According to Mayer and Lopoo, this is what has been taking place in the US. In the UK, on the other hand, Blanden et al. argue that the educational upgrading of the population for the most part occurred for people from richer parents, and that this fact has reinforced the decrease in mobility generated by increasing income inequalities.

As for Norway, the increased educational attainment has, to a large extent, been a result of several school reforms, where a central aim has been to enhance equality of opportunity along the socioeconomic dimension. Aakvik et al. (2002) provide empirical evidence that there has been a particularly steep increase in attainment among the disadvantaged groups that were the main target of the educational reforms. The younger the cohorts in our sample, the more they have benefited from the reforms. Although evidence is modest, the alleged re-distributive changes in the Norwegian educational system lend support to the tendency of increasing income mobility across generations. The framework and the data in the present paper do not allow a decomposition of increased educational attainment and compression of the income distribution as sources to increased generational mobility. This is, however, on the agenda for future research.

5 Concluding remarks

The purpose of this paper is to assess potential changes in intergenerational mobility in Norway for cohorts born in 1950-1965, using the Database of Generations. We estimate parent-child earnings correlations and quartile transition matrices for more than 160 000 father-child pairs, where the second generations are born in 1950, 1955, 1960, and 1965. For sons, the OLS elasticity of sons' average earnings at age 30-34 w.r.t. a three-year average of fathers' earnings is 0.166 for the 1950-cohort and 0.123 for the 1960 cohort, with correlation coefficients of 0.147 and 0.108. Daughters have a higher elasticity than sons in the 1950 cohort (0.222), but more similar in the 1960 cohort (0.137). The correlation coefficients are lower. Comparisons at ages 29 and 30 for all four cohorts show the same tendency to increasing mobility, but the coefficients are lower. Quartile father-child earnings transition matrices indicate quite high mobility in the second and third quartiles, but some persistence at the top of the distribution. This pattern is the same for both cohorts, clearest for sons. The results show that Norway is characterised by a high degree of earnings mobility, and that there is no tendency to increasing inequality along this dimension. Our findings also add to previous evidence indicating that countries with low levels of intragenerational inequality have a high degree of intergenerational mobility.

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Table 1. Descriptive statistics for children and their fathers by the children's birth cohort

	1950		1955		1960		1965	
	Mean/%	Std. Dev.	Mean/%	Std. Dev.	Mean/%	Std. Dev.	Mean/%	Std. Dev.
Fathers of sons								
3-year earnings average	11.94	0.54	12.07	0.68	12.24	0.68	12.09	0.86
Age	48.16	4.91	47.66	4.96	47.22	5.30	46.26	5.28
Sons								
Earnings at 29	12.13	0.60	12.06	0.82	12.11	0.69	12.06	0.83
Earnings at 30	12.17	0.55	-	-	12.13	0.73	12.15	0.74
Av. earnings 30-34	12.12	0.61			12.13	0.77		
Education								
-9	23.08		15.58		12.47		9.09	
10-12	45.57		54.22		60.98		63.33	
13-16	21.26		21.74		19.15		20.24	
17+	10.09		8.46		7.41		7.35	
N	18920		22447		23873		25503	
Fathers of daughters								
3-year earnings average	12.02	0.51	12.08	0.68	12.24	0.69	12.11	0.85
Age	48.24	4.81	47.72	4.94	47.19	5.35	46.23	5.28
Daughters								
Earnings at 29	11.19	1.24	11.21	1.20	11.44	1.07	11.62	1.03
Earnings at 30	11.17	1.22	-	-	11.47	1.08	11.71	0.93
Av. earnings 30-34	10.98	1.24			11.41	1.08		
Education								
-9	23.82		14.47		10.19			6.75
10-12	39.45		51.74		58.66			58.38
13-16	31.94		30.59		27.42			30.10
17+	4.79		3.20		3.73			4.78
N	9352		17721		21965			22270

Notes: Earnings (1995 NOK) in logs
3-year averages: 1967-69, 1972-74, 1977-79, 1982-84
Fathers' age: in 1967/72/77/82

Table 2. Intergenerational earnings correlations for sons, various series length of fathers' earnings.

		Fathers' earnings measure					
		1950-cohort			1960-cohort		
Year of fathers' earnings (1950 cohort / 1960 cohort)		Single year measure	Two-year average	Three-year average	Single year measure	Two-year average	Three-year average
1967/1977	β	0.204 (0.010)			0.121 (0.008)		
	ρ	0.154			0.104		
	N	18343			23650		
	β		0.172 (0.008)			0.123 (0.008)	
	ρ		0.147			0.108	
	N		18814			23815	
1968/1978	β	0.142 (0.008)		0.166 (0.008)	0.134 (0.008)		0.123 (0.007)
	ρ	0.134		0.145	0.111		0.109
	N	18500		18920	23461		23873
	β		0.143 (0.008)			0.129	
	ρ		0.136			0.110	
	N		18825			23615	
1969/1979	β	0.135 (0.007)			0.123 (0.008)		
	ρ	0.133			0.102		
	N	18409			23243		

Notes: The sons are from the 1950 and 1960 cohorts.
The log of sons' earnings are regressed on fathers' earnings together with fathers' age and age squared.
Sons' log earnings measure: Average age 30-34 (1980-84 for 1950-cohort, 1990-94 for 1960-cohort).
The number of observations, N, varies from one cell to another due to the fact that only years with earnings > 0 are included.
Standard errors in parentheses.

Table 3. Intergenerational earnings correlations for sons at ages 29 and 30, four various cohorts of sons

Sons' age		1950	1955[a)]	1960	1965
Age 29	β	0.080 (0.008)	0.085 (0.008)	0.078 (0.007)	0.066 (0.006)
	ρ	0.071	0.071	0.074	0.069
	N	18650	22447	23503	25003
Age 30	β	0.100 (0.008)	-	0.090 (0.007)	0.068 (0.006)
	ρ	0.096		0.083	0.075
	N	18644		23310	25017

Notes: The sons are from the 1950, 1955, 1960 and 1965 cohorts.
The log of sons' earnings are regressed on fathers' earnings together with fathers' age and age squared.
Fathers' earnings measures: three-year age adjusted averages 1967-69, 1972-74, 1977-79, 1982-84
Standard errors in parentheses
[a)] We report only age-29 results of the 1955 cohort because of unreliable 1985 earnings data.

Table 4. Intergenerational earnings correlations for daughters, various series length of fathers' earnings.

		Fathers' earnings measure					
		1950-cohort			1960-cohort		
Year of fathers' earnings (1950 cohort / 1960 cohort)		Single year measure	Two-year Average	Three-year Average	Single year measure	Two-year average	Three-year average
1967/1977	β	0.259 (0.029)			0.138 (0.011)		
	ρ	0.092			0.083		
	N	9079			21764		
	β		0.221 (0.026)			0.142 (0.011)	
	ρ		0.088			0.088	
	N		9297			21906	
1968/1978	β	0.176 (0.023)		0.222 (0.025)	0.148 (0.011)		0.137 (0.011)
	ρ	0.080		0.090	0.088		0.087
	N	9133		9352	21605		21965
	β		0.187 (0.023)			0.138 (0.011)	
	ρ		0.083			0.086	
	N		9321			21742	
1969/1979	β	0.173 (0.023)			0.131 (0.011)		
	ρ	0.077			0.080		
	N	9101			21403		

Notes: The daughters are from the 1950 and 1960-cohorts.
The log of daughters' earnings are regressed on fathers' earnings together with fathers' age and age squared.
Daughters' log earnings measure: Average age 30-34 (1980-84 for 1950 cohort, 1990-94 for 1960 cohort).
The number of observations, N, vary from one cell to another due to the fact that only years with earnings > 0 are included.
Standard errors in parentheses.

Table 5. Intergenerational earnings correlations for daughters at ages 29 and 30, four various cohorts of daughters

		1950	1955[a)]	1960	1965
Age 29	β	0.218 (0.029)	0.147 (0.013)	0.126 (0.011)	0.065 (0.008)
	ρ	0.086	0.083	0.078	0.054
	N	7877	17721	20070	22270
Age 30	β	0.179 (0.027)	-	0.126 (0.011)	0.069 (0.008)
	ρ	0.074		0.080	0.058
	N	7924		20102	22299

Notes: The daughters are from the 1950, 1955, 1960 and 1960-cohorts.
The log of daughters' earnings are regressed on fathers' earnings together with fathers' age and age squared.
Fathers' earnings measures: three-year age adjusted averages 1967-69, 1972-74. 1977-79, 1982-84
Standard errors in parentheses
[a)] We report only age-29 results of the 1955 cohort because of unreliable 1985 earnings data.

Table 6. Quartile transition matrices, 1950- and 1960-cohorts

	1950-cohort							
	Sons				*Daughters*			
Fathers	Bottom	2nd	3rd	Top	Bottom	2nd	3rd	Top
Bottom	0.351	0.275	0.214	0.160	0.297	0.257	0.243	0.202
2nd	0.249	0.295	0.260	0.195	0.254	0.270	0.258	0.218
3rd	0.207	0.257	0.284	0.253	0.250	0.255	0.249	0.246
Top	0.193	0.173	0.242	0.392	0.199	0.218	0.25	0.334
Inequality index [a)]		2.858				2.670		
	1960-cohort							
Bottom	0.325	0.291	0.226	0.159	0.295	0.271	0.243	0.191
2nd	0.246	0.285	0.264	0.206	0.269	0.270	0.263	0.198
3rd	0.228	0.243	0.275	0.254	0.233	0.252	0.262	0.254
Top	0.202	0.181	0.235	0.381	0.204	0.207	0.232	0.357
Inequality index [a)]		2.799				2.725		

Notes: [a)] Inequality index: sum of leading diagonal and adjacent cells
Earnings measures for fathers: 1967-69 and 1977-79 (age adjusted)
Earnings measures for children: 1980-84 and 1990-94

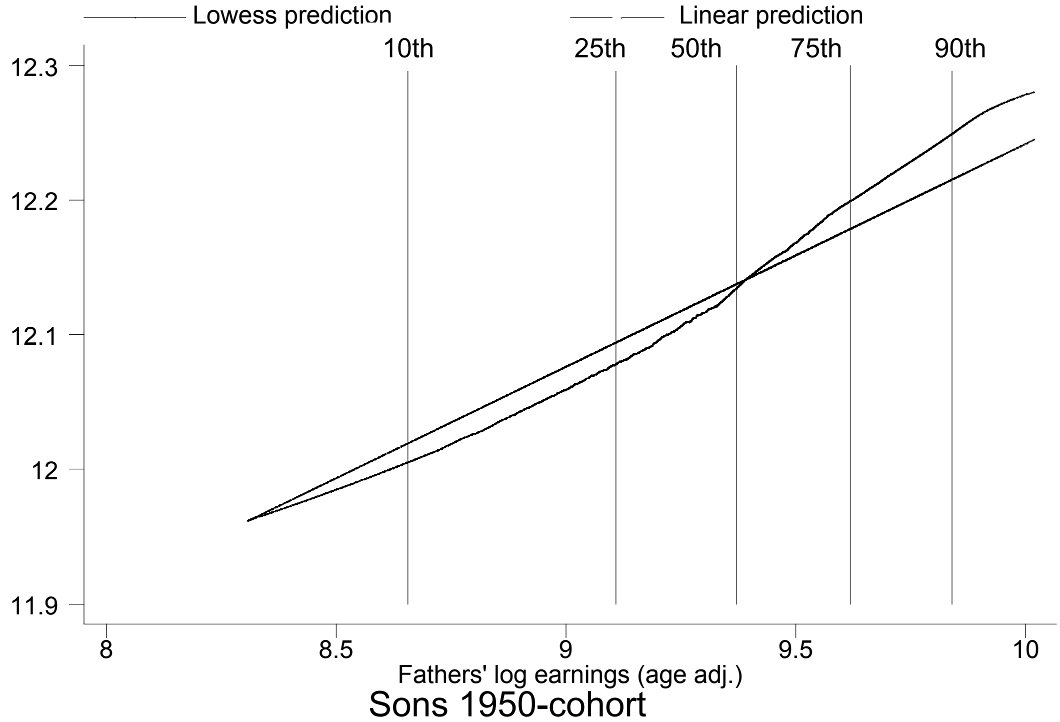

Figure 1

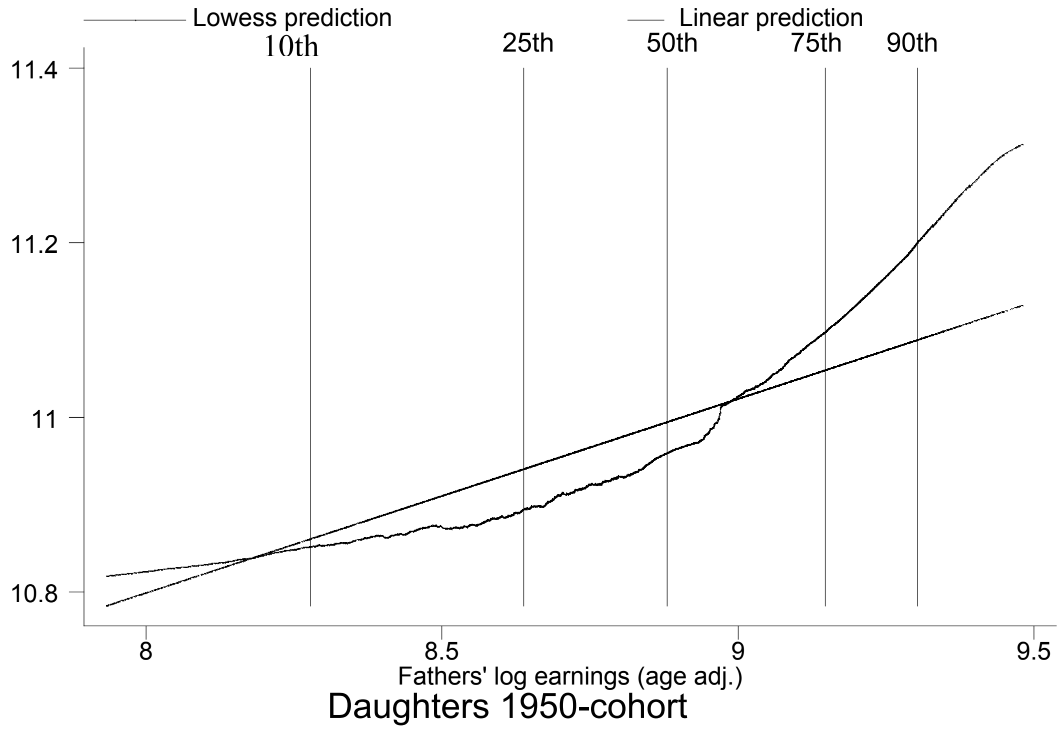

Figure 2

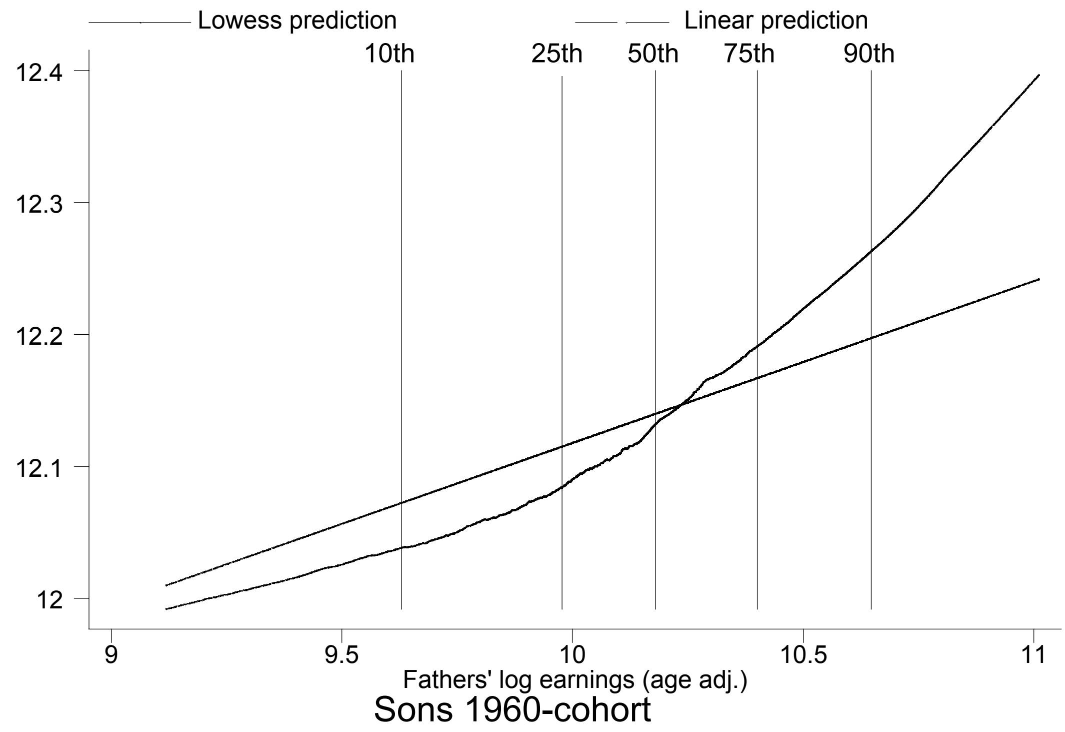

Figure 3

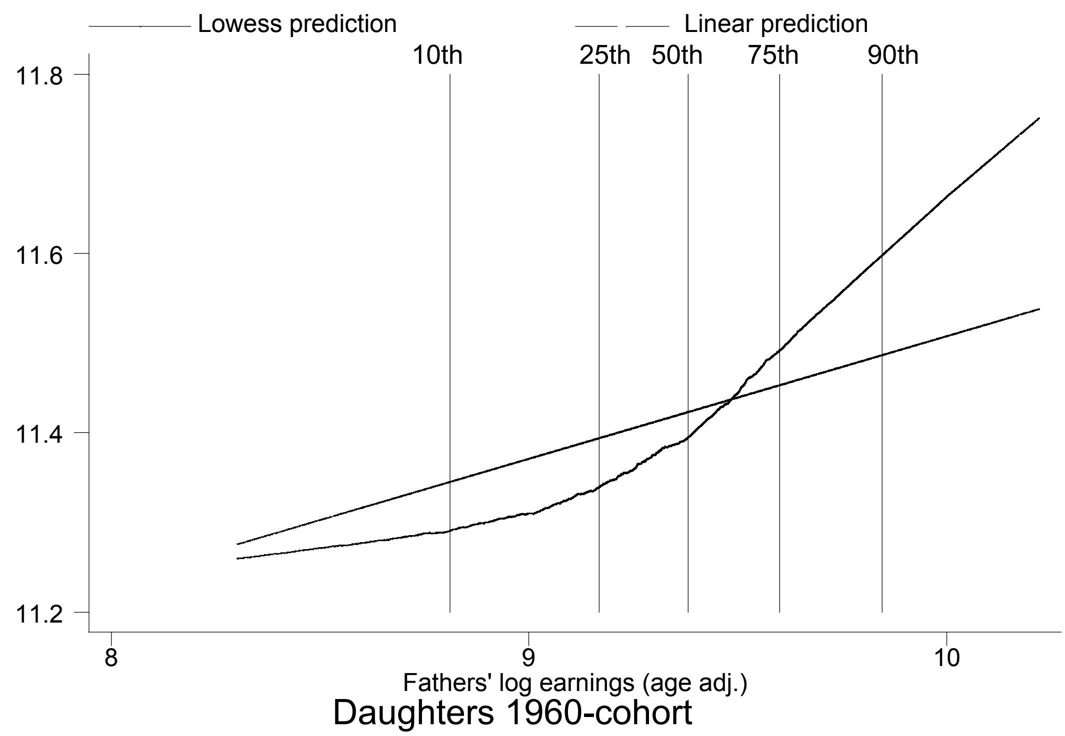

Figure 4

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