Trends in brother correlations in class and incomes in Finland:

A comparison of cohorts born in 1932-62

Jani Erola Turku School of Economics, jani.erola@tse.fi

Juho Härkönen Yale University, juho.harkonen@yale.edu

Markus Jäntti Åbo Akademi, markus.jantti@iki.fi

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Abstract

Recent assessments of cross-national and cross-cohort comparisons of intergenerational mobility have found intriguing differences in income and class mobility. For example, the United States seems to be an open society when it comes to class mobility but a rigid one in terms of income mobility. Furthermore, recent analyses have shown differences in trends in income and class mobility in Britain. In this paper, we compare trends in brother resemblance in class and incomes in Finland across cohorts born in 1932-62. We use population register data from the Finnish Census Panel to calculate brother correlations for class and incomes. We also compare brothers using loglinear models. We find diverging trends in brother resemblance in class and incomes, suggesting that intergenerational mobility in and family background effects on incomes and class may be governed by different processes.

Introduction

The analysis of the effects of family background on socioeconomic success is a central field of research in sociology and economics alike. The most common approach in this tradition is a comparison of the incomes or class positions of parents and their children (e.g., Blau and Duncan 1967; Erikson and Goldthorpe 1992; Solon 1992; Björklund and Jäntti 1997; 2000; forthcoming; Breen 2004; Corak 2004). This research has found surprising robustness, but also international and cross-temporal variation in socioeconomic mobility between parents and children.

While measures of the association between parents' and their children's socioeconomic positions appeal to our common-sense understandings of socioeconomic mobility and the role of family background, they provide a limited picture of the role of family background, as parents' socioeconomic background is only one measure of the multiple family background factors that affect adulthood outcomes. To better account for these factors, several studies have used correlations between siblings' socioeconomic statuses to estimate the "total" effect of family background and other factors that are shared by siblings (such as neighborhood conditions and schools) (e.g., Hauser and Mossel 1985; Ganzeboom et al 1992; Sieben and DeGraaf 2001; Sieben et al. 2001; Björklund et al. 2002). Not surprisingly, these correlations are generally higher than those between parents and their children (on the relationship between intergenerational correlations and sibling correlations, see Solon (1999)).

While questions of intergenerational mobility and family background have been of interest to both sociologists and economists, there has been surprisingly little overlap between the two disciplines (for a recent exception, Morgan et al. 2006). However, a comparison of the estimates of intergenerational mobility from economics and sociology can produce a very different picture of the extent of mobility in societies. For instance, the United States appears as a highly mobile society when class mobility is concerned, but more rigid when one looks at intergeneration mobility in incomes (Björklund and Jäntti 2000) (although according to other studies the US does not seem exceptional even in class mobility, e.g. Erikson & Goldthorpe 1992; Beller and Hout 2006). Measures of class and income mobility can also produce a very different picture of trends in intergenerational mobility within countries. A recent example concerns mobility trends in England (Blanden et al. 2004; Goldthorpe and Jackson 2007; Erikson and Goldthorpe 2008). According to measures of income mobility, intergenerational inheritance seems to have strengthened considerably between two cohorts born just 12 years apart (Blanden et al. 2004). At the same time, Goldthorpe and Jackson (2007) found notable stability in class mobility using the same data. Erikson and Goldthorpe (2008) also found that associations between fathers' and childrens' class were generally stronger than those between family incomes and childrens' earnings in adulthood. This is the only explicit test that has systematically compared the measures that we are aware of.

These differences in the perceptions of intergenerational inheritance and openness are intriguing. One would certainly wish and expect consensus between different measures on this important social issue. However, one can think of reasons for as to why the results differ.

First, social class and income are complementary, although related, measures of social position. Therefore, it may simply be that mobility in one is stronger than in the other, and countries differ with respect to their relative strength. We can, for example, think of a society in which parents are more interested in ensuring their children's social class or status than in maximizing their incomes, and the other way around. Parental class and income can also capture a different range of factors that affect socioeconomic outcomes. For example, class can be more strongly correlated with cultural capital than income. Furthermore, cultural and social capital may be a more important determinant of socioeconomic success in one society, whereas parental wealth and incomes open doors in another society, due to for instance differences in education systems (cf. Checchi et al. 1999). Björklund and Jäntti (2000, 24-6) offered a more formalised version of a related argument, according to which countries' (and periods') relative ranking of intergenerational inheritance may vary across mobility measures depending on the relationship between class-independent factors that affect the incomes of both parents and children.

The differences in estimates of the degree of mobility can, second, result from methodological differences across the disciplines or different methodological and data problems facing researchers on income and class mobility. Sociologists have often analysed intergenerational mobility with loglinear models in the case of social classes and correlations and regression analysis in the case of socioeconomic indeces, while economists establish intergenerational correlations and elasticities (by regressing childrens' log incomes on parents' or families' log incomes). These intergenerational measures based on correlations and elasticities on the one hand, and parameters from log-linear models on the other are obviously not directly comparable, in part because most former measures do not take account of possible nonlinearities (there are exceptions, however, e.g. Bratsberg et al. 2007). Thus, the comparability of the association measures poses an obvious challenge to comparisons of income and class mobility. Income measures are also more prone to measurement error than questions about respondents' occupation, which form the basis for measures of social class and socioeconomic status. Typical analyses of income mobility also focus on earnings, which often excludes incomes from self-employment and farming (e.g. Björklund and Jäntti 2000). Recent results have also shown that the picture of income mobility is sensitive to the number of years over which (parental) incomes are averaged in order to create a proxy for permanent incomes (e.g., Solon 1992; Mazumder 2005). The estimates are also sensitive to the stage of the life-course in which permanent (both of parents and children) incomes are measured (Böhlmark and Lindquist 2006; Haider and Solon 2006). All these issues increase the demands for suitable data for studying income mobility. Class mobility, on the other hand, can be more readily analysed with a wider range of data. As mentioned, respondents' answers on their (and their parents) occupations are generally more reliable. Classes and occupations also arguably show more year-to-year stability, which is why the above issues have not received similar focus in standard analyses (Ganzeboom ym. 1991), and the early examples of the application of multiple class observations on the analysis of class mobility of Bielby et al (1977) and Hauser et al (1983) have not inspired much replication. However, the downside of using class as the measure of status is that it is not always clear what distinguishes the classes and how to determine class membership. Due to this class membership is likely to include more measurent error than for example individual's earnings, which are more straightforward to define (assuming no recall or other measurent error as in the case of register data).

In this paper, we estimate brother correlations in class and earnings across five birth cohorts (1932-38, 1939-44, 1945-51, 1952-56, and 1957-62) in Finland using register data for a total of 33,767 men from 20,715 families from the Finnish Census Panel. We have two main questions, one substantive and one methodological. Regarding the former, we ask whether the brother correlations--and thus, whether the effect of factors shared by brothers--have changed across the three decades. The five cohorts grew up in very different socioeconomic circumstances. While poverty was widespread during the earlier periods and the older cohorts grew up during the hardship of the second world war and its direct aftermath, younger cohorts experienced more favorable conditions as education and the welfare state expanded and living conditions improved. We would thus expect that family background mattered more for the older cohorts than the younger ones. Our methodological question concerns the (dis)similarity in the results gained from different measures of social position, namely, class and incomes. We focus on brothers due to lower stability in class and earnings patterns of women (e.g., Böhlmark and Lindquist 2006), which may affect our estimates and also compromise comparability of income and class measures.

This descriptive paper makes three contributions. First, our paper is, to our knowledge, the first to compare brother resemblance in income and class (as a categorical variable) using the same data. In fact, our paper is to our knowledge also the first to provide estimates of sibling resemblance in class (instead of socioeconomic or prestige indeces). Second, we use multilevel modeling to provide comparable estimates of brother resemblance in earnings and class. We use identical cohorts, class and income measures across the same stages in the life-course, and similar models. These improve the comparability of the estimates across income and class. We also use high-quality register data, which help avoid common problems of measurement error and attrition in income measurement in particular. To further facilitate comparability, we also use multiple measurements both for income and for class for the same phase in the life-course. Third, we provide evidence of changes and stability in brother resemblance across five Finnish birth cohorts, who came to age under very different socioeconomic conditions. Our results suggest that the picture of the total family effect on socioeconomic attainment and its trends may indeed depend on whether one looks at class attainment or at incomes. Furthermore, our results for both brother resemblance in class and in incomes are somewhat sensitive to the choice of method (variance decomposition using multilevel models versus loglinear models).

The structure of the paper is following. First we provide a brief overview of previous related research in Finland and elsewhere. After that we describe the data and methods. The empirical section begins with the description of the changes in class structure in Finland for the analysed cohorts. We then analyse sibling resemblance with multilevel ordered probit (class) and multilevel linear regression (income) models. Then we compare the difference of the background effect applying the "traditional" loglinear models for the father-brother mobility in class for the studied cohorts. The last section concludes.

Previous research

Occupational class mobility has been studied fairly little in Finland, if compared to many other European countries. The studies conducted in the 1970s and 1980s showed that a large share of the white collar and manual labourer classes had their origins among the self-employed farmers and farm workers. This was explained with the rapid pace of industrialization that took place during two decades after the Second World War. The studies suggested that social mobility in Finland was very similar to that in Sweden and Norway, two other Nordic welfare states with high level of equality of opportunity. (Pöntinen, Alestalo & Uusitalo, 1982; Pöntinen, 1983; Erikson & Pöntinen, 1985).

After the eighties, there was a long gap in mobility studies. In 2002, Erola and Moisio analysed occupational class mobility of the 31-35 years old in 1990 and 1995, also using the Census Data. Although the analysed period was very short and only two cohorts were used, the study suggested that Finland was still a very equal society in terms of social mobility. (Erola & Moisio 2002, 2005.) Moisio (2006) analysed the change in social mobility by comparing mobility differences between different cohorts, using a single cross-sectional survey data set collected in 2005. The result indicated that social mobility in Finland had increased since the World War II. This was also suggested by a study of social mobility between three consequent generations by Erola and Moisio (2007). When social fluidity between two consequent generations were studied, it signalled the weakening of the social inheritance of status. It also showed that the effect of grandparents' class on the class status of the grandchildren was practically non-existing in Finland, when the association between two consequent generations were controlled for.1

Income mobility in Finland has been quite frequently studied in the past few years. Jäntti and Östebacka (1995) report father-son correlation of 0.22 in incomes according to five year average incomes. Österbacka (2001) reports correlations in earnings between fathers and children that are as low as 0.12-0.16.Bratsberg et al (2007) report intergenerational elasticities of about 0.20 for the cohort born 1956-60, but find strong non-linearities across the distribution. The association is very weak in the lowest income decile, and much stonger in the highest. Generally, the association is approximately at the same level as in other Nordic countries, but much weaker than in US or UK.

Importantly for the current study, there are also sibling association models for income available for Finland. Österbacka (2001) reports a correlation of 0.26 for brothers and 0.11 for sisters. Björklund et al (2002) report borther income correlation of 0.25 for Finnish 25-42 years old brothers. However, when controlling for autocorrelated errors between years of observations (1985, 1990, 1995), Finnish estimate increases to 0.52, which is on par with the results from US rather than with other Nordic countries.

Although not studied in Finland, sibling associations of index-based measures of occupational class have been previously elsewhere. One of the early examples by Hauser and Mossel (1985) estimated that family background explained about one third of the sibiling resemblance in SEI index between brothers in US. Sieben and De Graaf (2001) compared brother resemblance according to ISEI in England, Hungary, the Netherlands, Scotland, Spain, and the USA, and found that shared family background contributed about 36.4 percent of the variance in occupation. Conley and Glauber (2004, 2007) found the sibling correlation in SEI of 0.418. This was clearly lower than in education (0.576) but fairly comparable to income (0.458). As mentioned above, to our knowledge there are no similar tests applying class classifications, such as EGP.

Due to obvious data restrictions, there are few studies looking at trends in sibling correlations (however, Kuo and Hauser 1995; Mazumder and Levine 2003; Björklund et al. 2007). In an analysis of males born between 1932-1968 in Sweden, Björklund et al (2007) report a dramatic decline in the brother correlation in income from 0.34 to 0.23 between cohorts born in 1932-38 and 1944-50, respectively, after which the association remained rather stable. The authors of this study associated this decline to the effects of family background on education between these cohorts.For the United States, Mazumder and Levine (2003) found increasing brother correlations in earnings across time despite stable brother correlation in education.

Models and estimation

Sibling correlations in income and class

Our modeling strategy begins with a simple description of socioeconomic attainment, where we assume that we have a measure of long-run socioeconomic attainment of sibling *j* in family *i*

$$y_{ij} = \mu + \varepsilon_{ij} \qquad (1),$$

where μ is the population mean and ε_{ij} is an individual specific component with variance σ_ε^2. This component shows the individual's relative long-term socioeconomic attainment. This can be decomposed into a family specific component of permanent attainment (a_i) that is common to all siblings in a family and an individual specific component of permanent attainment b_{ij}:

$$\varepsilon_{ij} = a_i + b_{ij} \qquad (2).$$

The former captures the family-specific deviations from the population mean whereas the latter denotes individual deviations from the family component average. These components are also independent by construction. The variance, σ_ε^2, of the individual specific component in the population model is similarly a sum of the variances of the common family component and the individual deviations from the family mean

$$\sigma_\varepsilon^2 = \sigma_a^2 + \sigma_b^2 \qquad (3).$$

We can calculate the share of the variance in long-term socioeconomic attainment that stems from family background with

$$\rho = \frac{\sigma_a^2}{\sigma_b^2 + \sigma_b^2} \qquad (4),$$

which equals the correlation between two randomly drawn pairs of brothers. The bigger the *rho*, the stronger the influence of shared effects shared by the brothers (including family background, neighborhoods, schools, and genetic traits).

To estimate this brother correlation, we need to estimate the variance of the family component and the variance of the individual component, respectively.

In the case of (logged) income, we have multiple observations of annual income so that

$$y_{ijt} = \mu_t + a_i + b_{ij} + v_{ijt} \qquad (5),$$

where y_{ijt} is logged annual income for sibling *j* in family *i* in year *t*, μ_t is the population mean of logged incomes in year *t*, a_i is a family specific component of permanent income, b_{ij} is an individual specific component of permanent income, and v_{ijt} denotes possible measurement error and annual transitory income shocks. We estimated this two-level model the nlme package in R (R Development Core Team 2008).

Models for estimating the variance components for social class face additional difficulties due to the categorical nature of commonly used class variables, including the one (eight class Erikson-Goldthorpe class variable) used here. The optimal way of estimating the variance components for classes is through multilevel multinomial logit or probit models. To estimate the variances for the (long-term) individual and family components, one needs to specify both separate random effects at each level for each outcome (excluding the reference category) *and*, at each level, a full set of covariances between these random effects. This exercise proved computationally very burdensome and has to wait for future versions of the paper.

Therefore, we instead estimate ordered probit models. The model is written as

$$y^*_{ijt} = a^*_i + b^*_{ij} + v^*_{ijt}, \text{ where } y_{ijt} = \begin{Bmatrix} 0 \text{ if} & y^*_{ijt} < \tau_1 \\ 1 \text{ if} & \tau_1 \leq y^*_{ijt} < \tau_2 \\ & \vdots \\ n \text{ if} & \tau_n \leq y^*_{ijt} \end{Bmatrix} \quad (6).$$

Here y^*_{ijt} is the latent propensity of class attainment of individual *j* in family *i* in year *t*, a^*_i is a latent family specific component of permanent class attainment, b^*_{ij} is a latent individual component of class attainment, and v^*_{ijt} denotes a transitory component in class attainment in year *t*. The individual is in the lowest class (here, 0) if his latent class attainment propensity falls below the threshold τ_1, in the second lowest class if his latent class attainment propensity is between τ_1 and τ_2, and so on. The individual is in the highest class if his class attainment propensity is above τ_n.

This operationalisation of class assumes a hierarchical class structure. Although social class is in everyday parlance commonly conceptualised in such terms, assuming a hierarchy between classes as defined by the Erikson-Goldthorpe schema (see below) is in no ways straightforward. Here we simply rank the classes in the order as defined below, which also tracks rather closely the differences in incomes between classes in Finland. We estimated the variance components of this three-level ordered probit model with the statistical package aML 2.0 (Lillard and Panis 2003). Standard errors and confidence intervals for the sibling correlation were calculated using the

Maximum Likelihood method as described in Donner and Wells (1986, 405).

Log-linear analyses of sibling resemblance

As secondary means to examine trends in sibling resemblance in income and class, we use loglinear models, which are commonly used in sociological analyses of intergenerational mobility. Typically, the class position of a parent and a child form the dimensions of the mobility table; in sibling models the statuses of two siblings are used instead. We construct mobility tables for brothers according to EGP classes and income octiles. In addition, we match the brothers to their fathers, so that we can apply the usual parent-child setting to the data as well. We apply four rather commonplace mobility models: the independence model of no association between variables; the model assuming changing marginals but no association between class or income variables; the constant social mobility (fluidity) model, allowing interaction between positions of two brothers or fathers and sons in addition to changing marginals; and finally, the unidifference (or log-multiplicative layer effect) model for the sibling or father-son association according to cohort, testing the significance of the change in the strength of status association (Erikson & Golthorpe 1992; Xie 1992). As in studies applying the mobility tables approach, the models are fitted to one observation only, in this case to the first observation of class or incomes of brothers. Loglinear models were estimated using gnm-package in R (Turner & Firth 2006, R Development Core Team 2008).

Data

We use data on 33,767 Finnish men born in 1932-1962 from the Finnish Census Panel (FCP). The FCP is a register-based panel database compiled and provided by Statistics Finland. Since the data are collected from administrative registers, it suffers less from missing data, measurement error, and attrition problems than common survey data. Therefore, it promises a good source for comparing results for family background effects using income and class. We classified our sample into five mutually exclusive birth cohorts, born in 1932-38, 1939-44, 1945-51, 1952-56, and 1957-62. Table 1 shows the number of individuals (males), families, and observations in each cohort.

Table 1

The dataset includes 20 waves: 1950, 1970, 1975, 1980, 1985, and annual observations from 1987 to 2001. From this information we matched brothers and compiled follow-up data for our five cohorts. We matched brothers by, first, defining males as children if they were age 18 or less, and second, if they resided in the same family during our observations. Less than 10 percent of the individuals lived in independent households at age 18 in each cohort. Brothers were defined as belonging to the same cohort if they were born during the same range of years: brothers belonging to different cohorts were excluded from the cohort-specific analyses. Therefore, we calculate correlations for brothers from the same family who belong precisely to the same birth cohort.

Research on intergenerational income mobility and sibling correlations in income have paid much attention to attenuation bias that arises from the use of (parental) income observations from a single or a limited number of years (Solon 1992; Zimmerman 1992; Björklund and Jäntti 2000; Mazumder 2005; Björklund et al. 2007). This arises because of fluctuations in annual earnings, which produce measurement error to measures of permanent earnings, and thus bias intergenerational correlations downwards.

Another source of bias concerns lifecycle bias. Haider and Solon (2006) and Böhlmark and

Lindquist (2006) examined this issue in the United States and Sweden, respectively, by analysing the correspondence between short-term and longer-term earnings across different stages of the life-course. They concluded that for men, the correspondence is the highest during the early/mid-30s to early 40s. Since life-cycle bias can lead to biased estimates of the intergenerational or sibling correlations, incomes should be measured during this phase where the correspondence between short- and long-term incomes is the highest (see appendix in Björklund et al. 2007 for discussion of life-cycle bias in analyses of sibling correlations in income).

Building on these results, we restrict our observations from age 33 to age 43, both for incomes and class. Issues of biases stemming from poor proxies of permanent status or life-cycle bias have not received much attention in sociological studies (the latter may be why, for example, Hauser et al. (1999) and Conley and Glauber (2007) find strengthening sibling correlations across the life course, a result in line with Haider and Solon's (2006) findings of a weakening link between short-term and permanent incomes at later ages). However, although we are here not able to examine these issues further, we believe that they are of possible concern also to sociologists.

To reduce bias in estimating long-term incomes (and class), we would optimally include observations for each year during this age range (cf. Mazumder 2005). However, our data enable annual observations in this age range only for cohorts born 1954 or later, and for the eldest cohorts, we only have observations for every fifth year. Therefore, to maximize comparability, we observe incomes and class every five years for all birth cohorts. This leads to a three-level data structure, with observations at the lowest level, males at the middle, and families at the highest level.

Table 2

Table 2 describes our observation window. We have observations from eight years: first, from 1950 and then from 1970 to 2000 for every fifth year (x-axis). Each row shows the age of each annual birth cohort during these years. The ages shaded in grey represent the ages during which the males were defined as children (0-18 years): males residing in the same family during childhood were further defined as brothers. The ages shaded in black represent the ages during which we collected information on their incomes and class (33-43 years). For example, we observe the oldest males (born in 1932) first in 1950, when they were 18 years old, and thus were defined as children and eligible to be included in our analyses. They were next observed in 1970, when they were 38 years old, and then in 1975, when they were 43 years old. For this cohort, we have income/class data for these two years, and for some cohorts, we have data for three years.

Variables

Our class variable is based on a eight-class version of the Erikson-Goldthorpe class schema. The classification coding was based on the coding key by Ganzeboom and Treiman (2001). Our class schema and its distributions are described in Table 3. Our income variable is deflated and includes all personal taxable gross incomes during a year. These include earnings, social benefits, capital incomes, and other incomes. In the multilevel analysis, we used a logged transformation of incomes, while controlling for age and the year of income observation. For loglinear models, we instead use income octiles in order to enhance comparability with the class schema.

Table 3

Changes in class structure, absolute career mobility and income distribution

Table 2 shows the class structure of the analysed cohorts at the age of 33-37 using EGP class schema (see above). It can be seen that contrary to the expectations of some, the growth of service classes with the expense of the size of the manual labour classes has been very modest, almost non-existing since the second cohort.

Table 4

Although the class structure remains largely the same over time, there is quite much variation in the class positions of individuals between observations not only between the first and second, but also the second and third observation; whereas one fourth of the cases change occupations between two first observations, around one fifth of the cases do so also between second and third observation. This applies also to vertical career mobility between observations (this refers to the mobility across three hierarchy levels of the classes, as proposed by Erikson and Goldthorpe (1992), according to which the classes I, II and IVa+b are most advantageous and the classes VIIa and VIIb least advantageous, the other classes remaining between them.) The amount of mobility is about one third lower than according to absolute career mobility, but nonetheless far from being non-existent. There is little variation between the cohorts. The tables are reported in Appendix Table 1 and 2).

Figure 1

Figure 1 shows boxplots of the distribution of income in the five cohorts, measured in each case in using the second income observation we use. The boxplots show the median (the vertical line where the box is tallest), the mean (the black dot) and the percentiles 5, 12.5, 25, 37.5, 62.5, 75, 87.5 and 95, indicated by changes in the height of each box. The graphs suggest that average incomes have grown quite substantially across these cohorts. The average among those born between 1932 and 1938 was about 19000 euro, while for the last cohort born between 1957 and 1962 it was about 30000 euro. The next to last cohort, born between 1952-1956, is shown here with incomes in 1995, just after the deep 1990s recession and they are the only cohort to have a decline in mean income relative to their preceding cohort. The dispersion of income appears also to have increased, as suggested by the increased difference between the mean and the median. If we redraw the same picture using log income to reveal relative income differences (not shown here), the increase in relative dispersion turns out to be more modest.

Trends in sibling correlations in class and income

Table 5 shows the variances and 95 % confidence intervals of the family and individual components and the sibling correlations for permanent class and income. The changes in sibling correlations are further illustrated in Figure 2.

Table 5

Figure 2

The individual components are higher in each case, leading to sibling correlations below 0.5. The sibling correlations are higher for class than for income throughout the five cohorts. However, the gap between the sibling correlations with the two measures varies across cohorts. Likewise, the trends in family effects on socioeconomic attainment are different depending on the measure at hand.

The total estimated effect of family background on income decreased somewhat between the first and second birth cohorts (born in 1932-38 and 1939-44, respectively), after which it increased, decreased again between the third (1945-50) and fourth (1952-1956) cohort, and then increasing somewhat again for the youngest cohort (1957-62). The trend seems to be towards increasing family background effect. Overall, the estimates are in line with those reported by Österbacka (2001) and Björklund et al (2002). On the other hand, the effect of family background on social class *increased* clearly between the first two birth cohorts, then declined rather linearly from the second birth cohort to the fourth, after which it again increased slightly for the youngest cohort. We checked the results against ordered logit measures, and found a similar pattern, although the estimated brother correlations were generally somewhat stronger (except in the oldest cohort) and more stable across cohorts.2 Therefore, we cannot yet find a definite answer to the surprising increase in the family effect between the first two birth cohorts. However, all cohorts the family background effect on class appears stronger clearly stronger than for incomes, although the gap seems to be closing after the second cohort.

Especially to the question of whether we find diverging trends in the total family background effect on class and income, we next turn to results from loglinear models on the association between brothers' class position and their income octiles.

Loglinear models for brother resemblance and father-son mobility

As a further investigation for our results, in this section we model brother resemblance in class and income (octiles) using loglinear models. We also compare these results to models of father-son mobility tables. As discussed above, an important difference between these and the multilevel models shown above—apart from the obvious difference of different models—is that here we resort to single observations (namely, the first) per individual.

In Table 6 we provide the loglinear models for brother association using eight-class EGP and income octiles when observed at the age of 33-37 years. In the lowest part of the panel we show the similar models for father-son occupational class mobility. Models A are independence models (baseline models). Models B that only marginals change and there is no association between either the positions of bothers or positions of fathers and sons. Here the dissimilarity index is very similar in the case of class-relate models, about one third lower according to income. This suggests that marginal changes play a bigger role altogether in the case of income than in the case of class. Models C test the constant association either between classes or incomes. Now the model fit statistics change so that in the case of brother associations the fit is in practice equal in the case of class and income. This means that for the overall model fit the association between incomes play a smaller role than between classes. Thus like in the case of multilevel models we observe a stronger background effect for class than for incomes. Further, model fit is much better in the case of father-son class associations; the dissimilarity index is about the half of the brother association models. This reflects the general motivation for sibling models; inequalities according to socio-economic background are stronger when it comes to sibling associations than when measured directly between parents and children.

Table 6

Thus far the results are in line with the multilevel models; when modeling the change they depart. In all cases, assuming change between cohorts in the strength of association (unidifference models; Models D) does not improve the model fit. In the case of class measures we could linearize the effect, and get a statistically significant model fit improvement according to conventional levels of significance; the magnitude of improvement according to dissimilarity index would not, however, change. Linearization would not make a difference in the case of income.

In Figure 3 we show the unidifference coefficients according to cohorts from Models D. It can be seen that although the change over time is insignificant, the variation in coefficients is far greater for income associations. Like already pointed out above, the strength of background effect seems stronger according to class. The pattern of change seems very different to that from the multilevel models, almost the opposite. In the case of multilevel models for class we observed a reducing background effect for the second to the fourth cohort; in the case of loglinear models the significance of linearly reducing background effect are due to the lower association in the last cohort. In the case of income we observe no pattern of change whatsoever. It is likely that the results in the case of income are much more effected by the usage of only single observation; it may be questioned whether the approach is at all applicable to the income associations.

Figure 3

Conclusions

In this paper we have compared brother resemblance models for incomes and class for five different cohorts in Finland, using Finnish Census Panel data. The cohorts analysed were born in 1932-38, 1939-44, 1945-51, 1952-56, and 1957-62. The main question in our analyses was whether similar conclusions of brother resemblance in socioeconomic attainment could be reached when analysing class and income. This question relates to the recently increasing interest in cross-fertilization in research on intergenerational mobility between sociology and economics.

We used multilevel models (ordered probit for class and OLS for (logged) income) with up to three observations on class and income for each member of the data; at the ages of 32-37; 37-42 and 43. We compared these results with loglinear models for class and income octiles each pair of brothers using the first observation. Further, we also matched the brothers with their fathers and applied loglinear modelling also for these more conventionally used intergenerational assocations.

According to our findings, the brother correlation is approximately 0.4 and above for class and around 0.3 for income. The class effect is approximately at the same level as reported elsewhere accoding to socio-economic status indeces. The income effect was approximately at the same level as reported previously elsewhere. The finding of a stronger effect according to class than income was also confirmed by the loglinear models, and is in line with the previous finding elsewhere.

Overall, these results suggest that the picture of the effects of family background on socioeconomic attainment can vary rather considerably depending on whether one analyses class or incomes. In this sense, the diverging results on intergenerational class and income mobility reported by Goldthorpe and Jackson (2007) and Blanden and associates (2004), respectively, can possibly be real, that is, not methodological artefacts.

According to the brother correlations estimated from the variance decomposition using multilevel models, the family background effect on income decreased somewhat between the first and second birth cohorts, then increased a bit between the second and third cohort, decreased between the third and fourth cohort, and then increased somewhat again for the youngest cohort. In the case of income, the influence of the total family effect seems to be growing. The results for the changing family background effect according to class were more mixed, however. On the other hand, the effect of family background on social class *increased* clearly between the first two birth cohorts, then declined rather linearly from the second birth cohort to the fourth, after which it again increased slightly for the youngest cohort. The results indicate that the gap between the two socioeconomic measures may be closing. In the case of loglinear models for class, we observed rather clear stability with some decrease in brother resemblance in the youngest cohort. For income the results from loglinear models were clearly more mixed.

However, one should treat our results with caution given that in this version of the paper, we had to resort to ordered models of class attainment instead of more appropriate multinomial methods. The issue thus clearly warrants further investigation, and we seek to address it in the future. Neither—due to the essentially descriptive nature of this paper—do we here seek to explain the differences in the strength of the family background effect on class and income.

In addition to the methodological improvements, one of the important extensions of the multilevel analyses conduced here is to model class and income together. We also modeled each cohort separately. Albeit providing less accurate parameter estimates, the dissimilarity index from the loglinear model for class suggest that the changes in background effects may have played -- although statistically significant if linearized -- a very small overall role. The real magnitude of the changes over time could be more accurately estimated if the cohorts were modeled together.

Our analysis is--to our knowledge--the first one to examine sibling correlations in class (as a categorical variable). Despite the rather considerable challenges in estimation (which could not be fully resolved for this version of the paper), we argue that using categorical class measures in addition to socioeconomic indeces promises to link sibling analyses better to the tradition in sociological research on intergenerational mobility, which has dominated the scene especially in Europe.

Notes

1 According to our knowledge, all studies on class mobility in Finland have applied categorical status measures and loglinear models of mobility tables. However, Björklund and Jäntti (2000) report ISEI correlation between fathers and sons of 0.346 in the case of Finland, and Erola and Moisio (2007) ISEI elasticity of 0.28 between fathers and sons and 0.22 between mothers and daughters.

2 The rhos for the ordered logit analysis are reported in Appendix Table 3. If the equivalent analysis is done applying logged ISEI-index, the RHOs more or less overlap with the values from the probit-analysis (see Appendix Table 4).

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TABLES AND FIGURES

Table 1. Basic descriptives of the data

	N families	N siblings	N observations
1932-38	3739	4585	9654
1939-44	3684	4396	9279
1945-51	5197	6426	13822
1952-56	5747	6810	14681
1957-62	7590	9147	19690

Table 2. Observation window of the cohorts.

	1950	1970	1975	1980	1985	1990	1995	2000
1932	18	38	43	-	-	-	-	-
1933	17	37	42	-	-	-	-	-
1934	16	36	41	-	-	-	-	-
1935	15	35	40	-	-	-	-	-
1936	14	34	39	-	-	-	-	-
1937	13	33	38	43	-	-	-	-
1938	12	32	37	42	-	-	-	-
1939	11	31	36	41	-	-	-	-
1940	10	30	35	40	-	-	-	-
1941	9	29	34	39	-	-	-	-
1942	8	28	33	38	43	-	-	-
1943	7	27	32	37	42	-	-	-
1944	6	26	31	36	41	-	-	-
1945	5	25	30	35	40	-	-	-
1946	4	24	29	34	39	-	-	-
1947	3	23	28	33	38	43	-	-
1948	2	22	27	32	37	42	-	-
1949	1	21	26	31	36	41	-	-
1950	0	20	25	30	35	40	-	-
1951	-	19	24	29	34	39	-	-
1952	-	18	23	28	33	38	43	-
1953	-	17	22	27	32	37	42	-
1954	-	16	21	26	31	36	41	-
1955	-	15	20	25	30	35	40	-
1956	-	14	19	24	29	34	39	-
1957	-	13	18	23	28	33	38	43
1958	-	12	17	22	27	32	37	42
1959	-	11	16	21	26	31	36	41
1960	-	10	15	20	25	30	35	40
1961	-	9	14	19	24	29	34	39
1962	-	8	13	18	23	28	33	3

Table 3. The eight class Erikson-Goldthorpe class schema

Class		% (observations)
I	Upper service class (Upper-grade professionals, administrators, and officials; managers in large establishments; large proprietors)	10.93
II	Lower service class (Lower-grade professionals, administrators, and officials; higher-grade technicians; managers in small establishments; supervisors of non-manual employees)	19.91
III	Routine non-manual employees	5.30
IVa+b	Self-employed (non-farming: small proprietors etc., with and without employees)	8.18
IVc	Farmers	6.39
V+VI	Skilled manual workers (lower-grade technicians, supervisors of manual workers, skilled manual workers)	25.90
VIIa	Semi- and unskilled manual workers (not in agriculture)	19.94
VIIb	Agricultural workers	3.45

Table 4. Class structure by cohort, first observation.

	1932-38	1940-44	1945-50	1952-56	1957-62
I	6.99	11.70	11.49	10.91	11.61
II	15.90	20.43	18.91	19.43	21.27
III	4.84	5.26	4.98	5.55	6.38
IVa+b	5.99	5.03	6.67	8.44	8.62
IVc	12.12	7.03	5.89	5.63	4.42
V-VI	25.66	26.22	28.93	27.99	24.61
VIIa	21.19	21.35	19.81	19.09	20.20
VIIb	7.30	2.99	3.32	2.96	2.89
Total	4,521	4,352	6,387	6,789	9,122

Table 5. Point estimates and 95 % confidence intervals for family and individual components and brother correlations for permanent social class (Columns 1-3) and income (Columns 1-3) by cohort.

	Class			Income		
	Family component (σ^2_a)	Individual component (σ^2_b)	Rho (ρ)	Family component (σ^2_a)	Individual component (σ^2_b)	Rho (ρ)
	(1)	(2)	(3)	(4)	(5)	(6)
1932-1938	2.085	3.486	0.374	0.111	0.315	0.260
	{1.706; 2.503}	{3.098; 3.898}	{0.319; 0.429}	{0.078; 0.157}	{0.279; 0.357}	{0.256; 0.265}
1939-1944	2.608	2.931	0.471	0.084	0.265	0.240
	{2.237; 3.008}	{2.604; 3.277}	{0.419; 0.523}	{0.057; 0.139}	{0.233; 0.300}	{0.235; 0.244}
1945-1951	2.292	3.215	0.416	0.099	0.224	0.306
	{1.997; 2.606}	{2.935; 3.507}	{0.373; 0.459}	{0.078; 0.125}	{0.202; 0.249}	{0.303; 0.308}
1952-1956	2.403	3.944	0.379	0.071	0.184	0.279
	{2.081; 2.748}	{3.625; 4.275}	{0.330; 0.428}	{0.053; 0.095}	{0.163; 0.206}	{0.276; 0.282}
1957-1962	2.512	3.694	0.405	0.107	0.196	0.353
	{2.227; 2.815}	{3.435; 3.958}	{0.366; 0.444}	{0.090; 0.127}	{0.178; 0.216}	{0.351; 0.354}

Table 6. Fit statistics of loglinear models for brother and father-son associations according to first observation.

Brothers, Class						
	L2	df	Delta	BIC	AIC	Var
A. Coh + Egp1 + Egp2	1887.5	301	0.205	-729.3	1285.5	1005.7
B. Coh:Egp1 + Coh:Egp2	1538.2	245	0.189	-591.8	1048.2	853.5
C. C+Egp1:Egp2	350.4	196	0.073	-1353.6	-41.6	127.1
D. Unidiff (Egp1:Egp2,Coh)	343.5	192	0.072	-1325.7	-40.5	124.8
Brothers, Income						
	L2	df	Delta	BIC	AIC	Var
A. Coh + Inc1 + Inc2	1762.9	301	0.208	-830.4	1160.9	952.9
B. Coh:Inc1 + Coh:Inc2	643.6	245	0.121	-1467.2	153.6	325.2
C. C+Inc1:Inc2	331.5	196	0.072	-1357.2	-60.5	115.6
D. Unidiff (Inc1:Inc2,Coh)	324.7	192	0.070	-1329.5	-59.3	108
Fathers and Sons, class						
	L2	df	Delta	BIC	AIC	Var
A. Coh + FEgp + Egp	7620.6	301	0.2	4580.2	7018.6	3907.6
B. Coh:FEgp + Coh:Egp	6389.4	245	0.177	3914.6	5899.4	3074

C. B+C	303.9	196	0.032	-1675.9	-88.1	103.1
D. Unidiff (FEgp:Egp,Coh)	294.5	192	0.032	-1644.9	-89.5	103

Figure 1. The marginal distribution of income by cohort (in year 3)

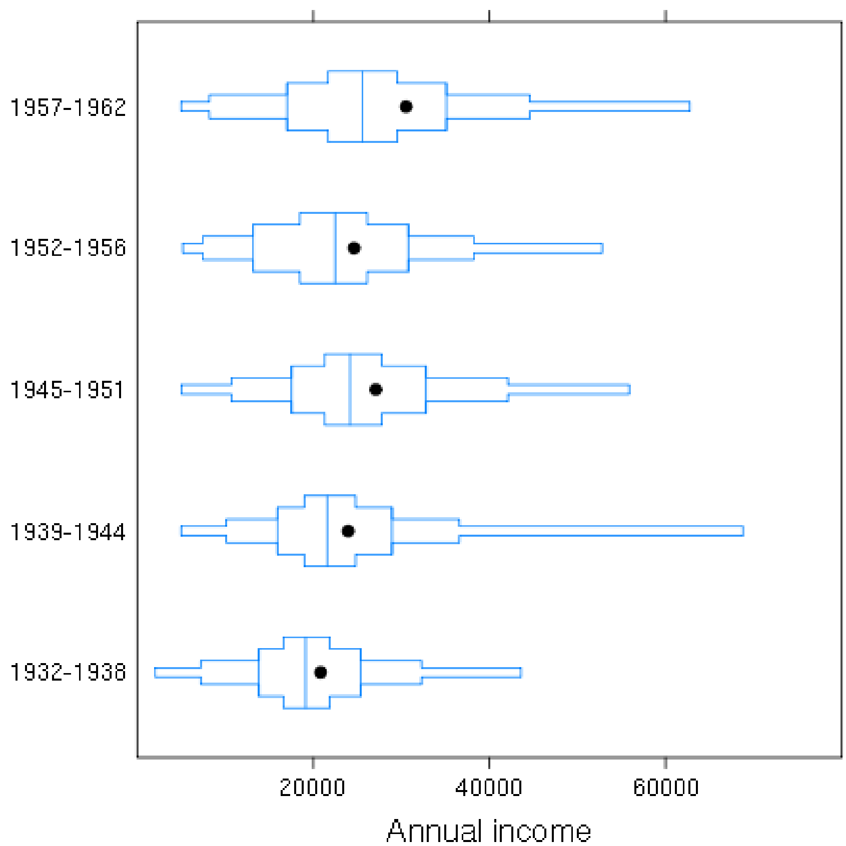


Figure 2. Trends in brother correlations in class and income

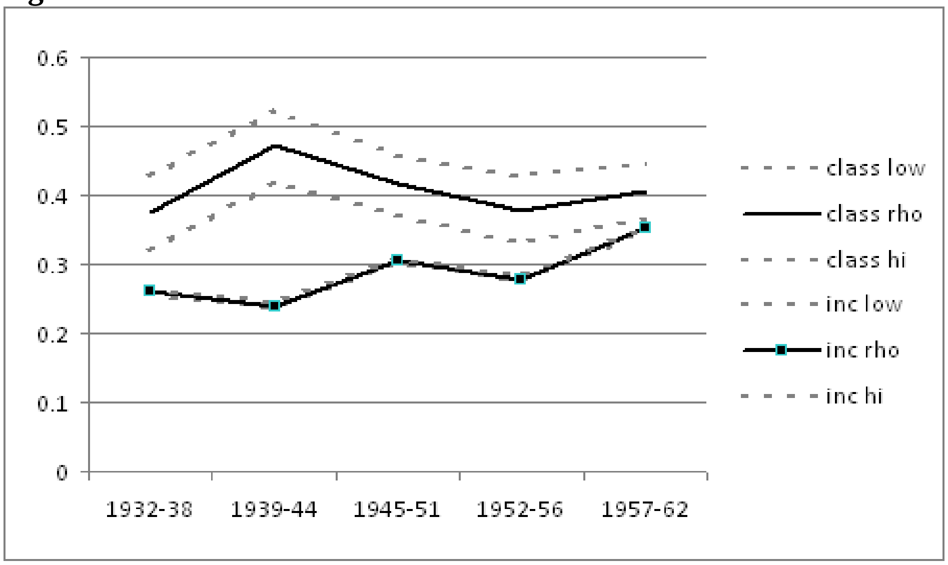
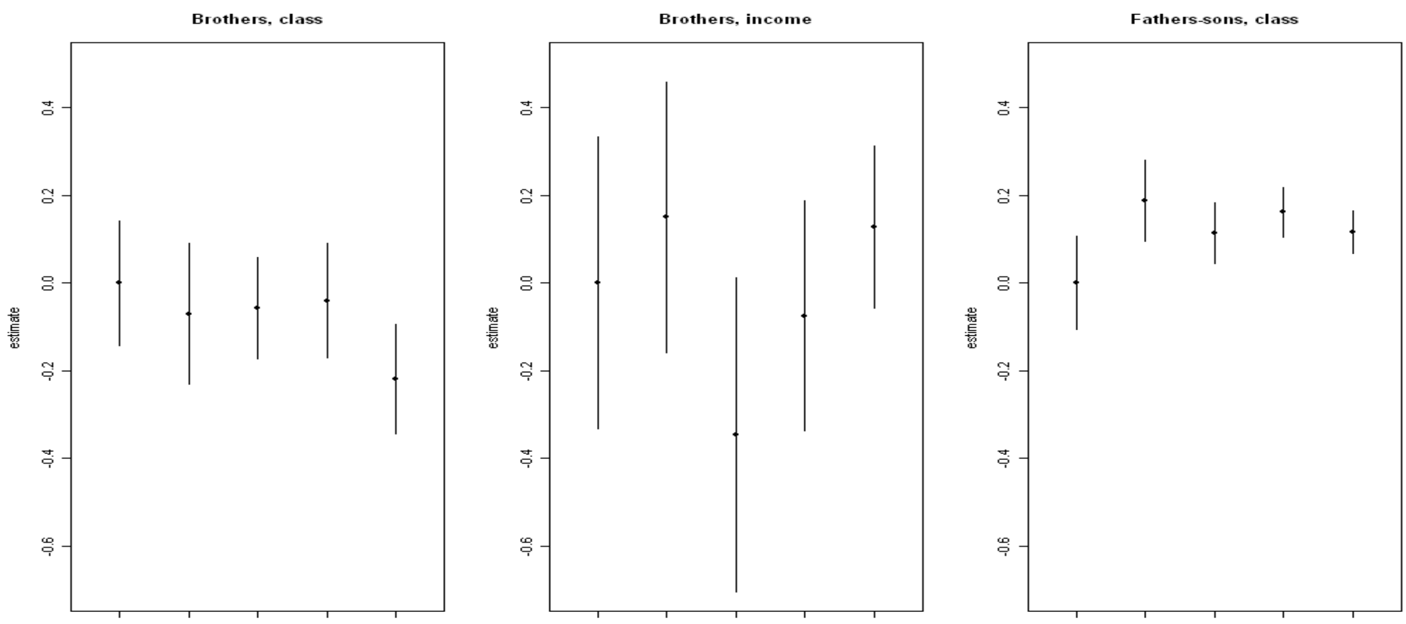


Figure 3. Coefficients for unidiff-association per cohort for pairs of brothers and fathers and sons according to first class observation. Coefficients and their quasi-standard errors.



Appendix

Appendix Table 1. Percentage of career mobility between observations by cohort.

	1932-38	1940-44	1945-50	1952-56	1957-62	Total
1 – 2	24.84	25.38	22.72	22.03	23.98	23.62
2 – 3	23.16	19.90	20.16	20.14	21.07	20.77
1 – 3	33.94	29.49	32.18	31.34	32.03	31.83

Appendix Table 2. Percentage of vertical career mobility between observations by cohort.

	1932-38	1940-44	1945-50	1952-56	1957-62	Total
1 – 2	18.73	18.28	15.42	14.37	16.36	16.35
2 – 3	16.77	13.38	13.47	11.72	14.73	13.84
1 – 3	25.70	18.98	21.79	20.58	21.02	21.42

Vertical levels: 1. I, II, IVa+b; 2. III, IVc, V-VI; 3. VIIa, VIIb (see Erikson & Goldthorpe, 1992)

Appendix Table 3. Rhos for the ordered logit multilevel models with EGP.

	rho low	rho	rho high
1932-38	0.235	0.297	0.359
1939-44	0.440	0.490	0.540
1945-51	0.417	0.457	0.497
1952-56	0.417	0.460	0.503
1957-62	0.448	0.482	0.516

Appendix Table 4. Rhos for the multilevel models with logged ISEI-index.

ISEI

	rho low	rho	rho hi
1932-38	0.341	0.370	0.400
1939-44	0.397	0.430	0.462
1945-51	0.404	0.436	0.468
1952-56	0.356	0.380	0.403
1957-62	0.365	0.386	0.407