

## **Educational Expansion and Social Mobility in the Twentieth Century**

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## **Educational Expansion and Social Mobility in the Twentieth Century**

### *Abstract*

Sociologists have long considered inequality in educational attainment to be a major cause of inequality between people in their chances of occupying a more advantageous class position. However, there is some dispute as to whether or not educational inequality according to social class background declined during the 20<sup>th</sup> century. What is not in doubt is the marked expansion of education in the advanced economies during this time, and several authors have pointed to educational expansion as a mechanism by which inequalities in social mobility chances may be reduced. I measure the magnitude of such an effect and compare it with the impact of educational equalization on social mobility in Britain, Sweden and Germany during the twentieth century. I find that in all three countries educational expansion has had a positive effect in promoting greater social mobility.

## **Educational Expansion and Social Mobility in the Twentieth Century**

Sociologists of stratification and inequality have a long standing interest in the role of the educational system in helping to determine the occupations that individuals come to hold. The underlying assumption is that a weakening of the relationship between social origins and educational attainment will lead to a weakening of the relationship between origins and destinations. In the log-linear tradition of social mobility research this has often been asserted, but the link between educational and occupational mobility has never been adequately demonstrated. The first goal of this paper, therefore, is to use simulations, based on log-linear models, to gauge the impact of increasing educational equalization on the origin-destination association. The second goal is to illustrate the impact that the expansion of education, aside from any equalization, may have on the relationship between social origins and destinations. Under certain conditions, educational expansion alone may lead to a weakening of this relationship. I calibrate the size of this effect and compare it with the impact of educational equalization in Britain, Germany and Sweden.

The paper begins with a review of recent research into the evolution of the relationships between social origins and, on the one hand educational attainment and, on the other, class destination. I then turn to the two mechanisms – equalization and a compositional effect – by which changes in the educational system bring about changes in social fluidity and I illustrate how these might work via a simple fictitious example. Using data from Britain, Sweden and Germany I quantify the impact of both of these mechanisms on the trend in social fluidity

during the 20<sup>th</sup> century. The paper concludes with a discussion of the roles of these two mechanisms and the conditions under which they might be expected to operate. Various technical aspects of the method developed and used here have been relegated to a technical appendix.

### *Educational equality*

One of the most debated questions in recent research in social stratification concerns whether, over the course of the twentieth century, inequalities in educational attainment according to social background diminished, or remained more or less unchanged. Until the late 1980s it was widely supposed that a decline in educational inequality would be an inevitable consequence of modernization, and, particularly, of industrialization. Treiman (1970: 217) argued that not only should 'the direct influence of education on occupational attainment ... be stronger in more industrialized societies' but 'in more industrialized societies parental status should play a less important role in educational attainment than in less industrialized places'. Yet, in a study published in 1993, Shavit and Blossfeld found stability in educational inequality among industrialized countries during the course of the twentieth century. Notwithstanding substantial expansion of educational systems during the century, particularly at the lower secondary level, in only two out of thirteen countries studied - Sweden and the Netherlands – was any reduction found in the strength of the associations between social origins and educational attainment. The editors concluded (page 19) that, with the two exceptions referred to, 'there has been *little change in socioeconomic inequality of educational opportunity*' (italics in original), and thus 'the impact of educational reforms on

changes in educational stratification seems to be negligible' (page 21). In other words, although educational attainment levels have everywhere increased, the relative chances of continuing to further levels of the educational system among people of different social origins have remained generally unchanged.

Shavit and Blossfeld's results echoed, to a large extent, earlier findings from the USA, France, the Netherlands, Britain and elsewhere.<sup>1</sup> In a recent review, however, Breen and Jonsson (2005: 226) reported that

Subsequent analyses have clearly shown equalization in the case of Germany (Henz & Maas 1995, Jonsson et al. 1996, Müller & Haun 1994), France (Vallet 2004a), Italy (Shavit & Westerbeek 1998), and probably Norway (Lindbekk 1998), while the results for Sweden (Jonsson & Erikson 2000) and the Netherlands (Sieben et al. 2001) have been corroborated.

As well as studies of single nations, there have been at least two comparative analyses of educational inequality with respect to social origins: these are Rijken's (1999) PhD thesis, and the recent papers by Breen, Luijkx, Müller and Pollak (2009a, 2009b). These authors find a tendency for social origins to have a diminishing influence on educational attainment over cohorts born during the first three-quarters of the twentieth century though, in the case of the latter papers, some important cross-national variations in this trend are found.

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<sup>1</sup> For the USA, Featherman and Hauser (1978); for France, Garnier and Raffalovich (1984); for the Netherlands, Dronkers (1983) and for Britain, Halsey, Heath and Ridge (1980) and Heath and Clifford (1990).

It is not immediately obvious why there should be such conflicting findings about temporal trends. It is certainly true that different studies use different measures of social origins (for example, measures of occupational status or prestige or categorical measures of social class membership and measures of parental education) and employ different statistical models, but it is nevertheless somewhat surprising that the basic pattern – whatever it is - is not robust to these kinds of variations. All studies are, however, agreed that there was a massive increase in formal levels of educational attainment in the course of the twentieth century.

### *Social Fluidity*

A common approach to the study of social reproduction is to examine the relationship between the social class position held by an individual and the social class that his or her parents occupied when the respondent was growing up. There are various ways that this relationship could be modeled, but much emphasis has been given to the ratio of the odds, among respondents born into one class origin compared with those born into another, of coming to occupy one social class destination rather than another. Odds ratios formed in this way are sometimes said to capture 'relative mobility chances' or 'social fluidity'. For analysts of social mobility, education is one of the major channels through which intergenerational class reproduction occurs (Ishida, Müller and Ridge 1995) and so educational inequality is chiefly important because of how it might affect inequalities in social mobility chances, or, in other words, how social fluidity might be driven by educational fluidity.

Perhaps unsurprisingly, disagreements about temporal trends in social fluidity parallel those about educational inequality. For example, Erikson and Goldthorpe (1992: 367) argued for 'a high degree of temporal stability' in social fluidity, yet their analyses were based on the proxying of birth cohorts by the use of age-groups drawn from a single survey per country. Indeed, few studies have analyzed trends in fluidity using a series of mobility tables collected over time (when a series is defined as consisting of more than two over-time observations),<sup>2</sup> and even fewer have drawn comparisons of trends between countries. The best known exceptions to this are Ganzeboom, Luijkx and Treiman (1989: 47) who claimed to demonstrate 'substantial cross-national and cross-temporal differences in the extent of mobility', and Breen and Luijkx (2004) who document significant variation in social fluidity among the countries of Europe and a fairly widespread temporal trend towards greater fluidity in the closing decades of the 20<sup>th</sup> century.

Most temporal comparisons have adopted a period perspective, but a small number of single country studies have drawn comparisons over birth cohorts: these include Hout (1988, USA) and, more recently, Vallet (2004a, France), Müller and Pollak (2004, Germany), and Breen and Jonsson (2007, Sweden). In all these cases, repeated cross-sectional surveys are used, so allowing for the possibility of comparing period and cohort change in a way that cannot be done with a single

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<sup>2</sup> Exceptions are Hout's (1988) work on the General Social Survey (see also DiPrete and Grusky 1990; Hauser and Huang 1997), the work of Luijkx and Ganzeboom (1989) on Dutch data, Jonsson's and Mills' (1993) analysis of Swedish and English data, and Vallet's (1999) comparisons using the French FQP surveys.

survey. These cohort based analyses show a gradual decline in the association between class origins and destinations during the twentieth century.<sup>3</sup>

*How education might influence social mobility: (1) an example*

Following Breen and Jonsson (2007), there are two ways through which education might cause social fluidity to change. The first, 'equalization', is a decline in the association between class origins and educational attainment: this is what has mainly concerned analysts of social mobility. The second process is 'compositional': if the origin–destination association is weaker at higher levels of education, and if educational expansion results in increasing shares of each cohort reaching higher educational levels, then this compositional change will lead to a reduction in the gross association between origins and destinations. It is clear that the compositional effect requires both educational expansion and the correct sort of three-way interaction between class origins, educational qualifications and class destinations. I shall use the term 'differential association' to refer to the case in which the origin – destination association tends to be weaker at higher levels of education. Such differential association may be present when, for example, higher qualifications are a powerful signal for employers that leaves little leeway for social network effects, or when the job markets in which degree-holders operate are particularly meritocratic. Hout (1988) was the first to draw attention to this phenomenon and its possible impact on fluidity and, subsequently, a weaker origin–destination association at

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<sup>3</sup> There are good reasons for supposing that change in fluidity is better captured as a cohort than a period phenomenon. For most people, educational attainment is the most important single determinant of their class position, and education is attained during childhood, youth and early adulthood (or, as Müller and Pollak 2004: 110 succinctly state, 'Educational participation and class inequality in education is a cohort phenomenon'). Thus, changes in the distribution of educational attainment affect specific birth cohorts and not the entire population (Breen and Lujckx 2007).

higher levels of education has been reported for France (Vallet 2004b), Sweden (Erikson and Jonsson 1998) and Germany (Breen and Luijkx 2007). It is therefore possible that, in these countries, an expansion of higher education across cohorts led to increasing fluidity. If this is the case, then it follows that social mobility analysts' concern with only educational equalization is too restrictive, because educational expansion itself – about the occurrence of which there is no controversy – may also have the capacity to weaken the association between class origins and destinations.

These possibilities can be illustrated by means of the hypothetical 2 by 2 by 2 table shown in Table 1. The situation at time=1 is given on the left side of the table. The top rows, labeled 'OE' show the relationship between class origins (middle class or working class) and educational level (low or high) and the odds ratio (OR) capturing the association between these. Below that are the tables showing the origin – destination distribution and odds ratios at each level of education (OD|E) and, below them, the overall OD distribution, which is formed by adding the OD|E=low and OD|E=high tables. By design the odds ratio is lower at the high level of education to capture differential association. The right-hand part of the table, labeled Time = 2, shows the effects of different kinds of change in the origin – education relationship. Case (a) deals with equalization alone, which in this case takes the form of a reduction in the origin – education odds ratio from its time=1 value of 1.5 to 1.31, while preserving the marginal distributions of both origins and education. Keeping the two partial origin – destination associations unchanged (as I do throughout), causes the overall origin – destination relationship to decline from 4.14 to 4.03. This is the effect of equalization. The next column, (b), shows the impact of expansion

when differential association holds. This time the origin – education odds ratio and the origin distribution remain the same but the marginal distribution of education shifts towards a larger share of people at the higher level. This causes the overall odds ratio to decline to 3.87. Finally case (c) shows the impact of both processes. The origin – education table now preserves only the marginal distribution of origins and is subject to exactly the same equalization as in case (a) and the same expansion as in case (b). The origin – destination association odds ratio in this case is 3.77.

[TABLE 1 HERE]

*How education might influence social mobility: (2) formal derivation*

The outcomes shown in Table 1 can be derived formally using log-linear models and some elementary results about conditional probabilities. Consider a four way cross-tabulation of class origins (O, indexed by  $j$ ) by educational attainment (E, indexed by  $k$ ) by current class position or destination (D, indexed by  $l$ ) by cohort (C, indexed by  $i$ ), with all the variables treated as categorical. A log-linear modeling approach would consider two tables: the EOC table, which permits the analysis of the distribution of E, conditional on O and C, (E|OC), and the EOCD table, for the analysis of D conditional on O, E and C, (D|OEC). Any pair of models for E|OC and D|OEC implies a set of estimated frequencies in the OEDC table and thus, by collapsing over the E margin, a set of frequencies for the ODC table. In this way one can assess the effect of different models for E|OC and/or D|OEC on the trend in OD over C.

We can write the joint probability distribution of these four variables as the product of conditional probabilities that capture the dependencies among them:

$$p_{ijkl} \equiv pr(C = i, O = j, E = k, C = l) = pr(D = l | O = j, C = i, E = k) pr(E = k | O = j, C = i) pr(O = j, C = i) \quad (1)$$

We always treat the joint distribution of origins and cohorts,  $pr(O = j, C = i)$ , as exogenous since the OC margin is fitted exactly in all the models considered here.

The three-way probability distribution of cohort by origins by destinations is arrived at by summing over the levels of E:

$$p_{ij.l} = \sum_k pr(D = l | O = j, C = i, E = k) pr(E = k | O = j, C = i) pr(O = j, C = i) \quad (2)$$

Specific models of the exact nature of the dependencies in (1) will thus give rise to different joint distributions of ODC via (2). It is intuitively clear that equalizing the origin – education relationship will reduce the origin – destination association, but the consequences of the compositional effect are less immediately obvious. A log-linear model for this effect can be written:

$$f_{ijk.} = \mu \gamma_j^O \gamma_k^E \gamma_{il}^C \gamma_{jk}^{OE} \gamma_{ij}^{CO} \gamma_{ik}^{CE} \quad (3a)$$

$$f_{ijkl} = \alpha \beta_j^O \beta_l^D \beta_k^E \beta_i^C \beta_{jk}^{OE} \beta_{jl}^{OD} \beta_{kl}^{ED} \beta_{ij}^{CO} \beta_{il}^{CD} \beta_{ik}^{CE} \beta_{ijk}^{COE} \beta_{jkl}^{OED} \quad (3b)$$

Here  $f$  denotes the expected frequencies under the model. The effect of origins on education does not vary over cohorts and so there is no equalization: this is captured by the absence of the term  $\gamma_{ijk}^{COE}$  from (3a). However, E has an additive dependence on O and C (so the distribution of education is allowed to change over cohorts). D depends on C and also on the joint EO distribution (this is the differential association effect). In this case:

$$\sum_k pr(D | O, E) pr(E | C) pr(O, C) = pr(D | OC) pr(OC) = pr(DOC) = pr(OD | C) pr(C)$$

The same argument can be applied to log-linear models for educational equalization and for the combined equalization and compositional effects: in both cases they imply a dependence of the origin – destination association on cohort.<sup>4</sup>

I use the foregoing results as the basis of simulations of the effect, on the change in the Origin - Destination odds ratios (i.e. social fluidity), of educational equalization and of the compositional effect. The analysis is carried out using data on men in Great Britain, Germany and Sweden, for birth cohorts covering the first two-thirds of the twentieth century.

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<sup>4</sup> These results can be obtained on request from the author. In case a skeptical reader thinks that all models give rise to the three-way ODC distribution, consider the case in which E depends additively on C and D depends additively on E. Then

$$\sum_k pr(D | E) pr(E | C) pr(O, C) = pr(D | C) pr(OC) = pr(D | O),$$

and now the OD relationship does not vary with C.

### *Data*

The sources of data for the three countries are shown in Table 1. For Britain the data come from the General Household Survey (GHS) for the years 1973, 1975-6, 1979-1984, 1987-1992. Although the GHS has been fielded annually from 1971 onwards, information on the employment of the respondent's father (or head of the family) was not collected in 1977 or 1978 or after 1992, while information on the last job held by those not currently working was not gathered in 1985 and 1986. I omit the 1974 data because Goldthorpe and Mills (2004: 202), who have recently used the GHS for mobility analysis (and who kindly provided the data), report that 'some serious problems' exist with the 1974 data.

The German data, which cover the period 1976 to 2002, come mainly from the Allbus survey, but are supplemented by other surveys. Müller and Pollak (2004: 85), who kindly provided the data, write that these surveys 'all use similar questionnaires and procedures to collect the data' and they express their confidence that 'cumulating these databases does not introduce systematic inconsistencies'. To preserve comparability over time, they refer only to Germans resident in the former West Germany.

The Swedish data set is a compilation of the annual surveys of living conditions (ULF) 1976-1999, conducted by Statistics Sweden (Vogel et al. 1998). Each survey is representative of the adult Swedish population, aged 15-75 (Breen and Jonsson 2007 for more detail).

[TABLE 2 HERE]

I analyze data for men (the reason for this is explained in the technical appendix) and, as is conventional in mobility analysis, I confine attention to the working age population. For Germany and Sweden this is men aged 25 to 64 at each survey. In the British data the age range is narrower: 25 to 49, because, for most years, respondents to the GHS aged 50 or over were not asked the occupation of their father.

Class origins (O, for short) and destinations (D) are defined using the EGP class schema (Erikson and Goldthorpe 1992: ch. 2) as follows:

I Upper service class - higher grade professionals, administrative and managerial workers;

II Lower service class - lower grade professionals, administrative and managerial workers;

III Routine non-manual workers;

IVab Self employed and small employers,

IVc Farmers,

V+VI Skilled manual workers, technicians and supervisors, and

VIIab+III Semi- and unskilled manual, agricultural, and lower grade routine non-manual workers.

These seven classes are used in the German case. Unfortunately the limitations of the GHS force the use of a six, rather than seven, class classification for Britain, because it is not possible to distinguish between classes I and II or between these and class IVa (that is, small employers): accordingly the first class in Britain is I+II+IVa while the third is IVb (the self-employed without employees). In Sweden, on the other hand, the last two classes cannot be distinguished, so here too there are six, rather than seven, classes.

Respondents' education is measured by their highest level of educational attainment (education (E), for short) categorized using the CASMIN educational schema. Categories 1a, b and c, and also 2a and b of the original scheme (Müller, Lüttinger, König, and Karle 1989) have been amalgamated, giving five educational categories:

- 1abc (compulsory education only),
- 2ab (secondary intermediate education, vocational and general),
- 2c (full secondary education),
- 3a (lower tertiary education), and
- 3b (higher tertiary).<sup>5</sup>

The exception here is Sweden where there are six educational categories because the distinction has been made between 1ab and 1c.

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<sup>5</sup> Higher tertiary education, 3b, means the successful completion (with examination) of a traditional, academically-oriented university education. Lower tertiary education, 3a, is usually characterized by a shorter length of study and more practically oriented study programs (e.g. technical college diplomas, social worker or, non-university teaching certificates).

Birth cohorts are defined differently because of the different survey dates and age ranges covered. So for Germany the cohorts are born 1912-21, 1922-30, 1931-39, 1940-48, 1949-57, 1958-66 and 1967-77 and for Britain 1924-27, 1928-34, 1935-41, 1942-48, 1949-55, 1956-62 and 1963-67. Thus the German data span a slightly longer period than the British. Omitting cases with missing information on one or more variables gives sample sizes of 18961 for Germany and 58125.5 for Great Britain. For Germany the data are unweighted data but for Britain they are weighted to correct for over-sampling of Scottish respondents. For Sweden there are 24 annual surveys and they are grouped into six, 4-yearly periods (1976-79; 80-83; ... 96-99). Within each period I identify ten four-year age groups (25-28; 29-32... 61-64) which allow me to define 15 overlapping age cohorts. The oldest cohort was born in 1912-18, the youngest in 1968-74. The sample size is 33281.

In all three data sets not every cohort can be observed in every survey: the oldest and youngest cohorts are observed least often, while the cohorts born around the middle of the century are observed most frequently. This means that there is more data in the middle of the series and rather little at the extremes – something which should be kept in mind when interpreting the results.<sup>6</sup>

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<sup>6</sup> The focus of this paper is change in social fluidity across birth cohorts but the data comprise samples from each birth cohort drawn in one or more periods and this raises the question of whether there should be controls for period in the models fitted here. I have not done this because the need to further cross-classify the data by period would have led to very sparse tables, especially in the German case. But ignoring period effects can be justified by reference to the results of earlier research which found that, once cohort effects are included in a model of fluidity, period effects become largely insignificant. For Britain and Germany, Breen and Luijkx (2007:112) write: '(i)n both cases, differences between periods are in fact the result of cohort replacement' and Breen and Jonsson (2007: 1805) write similarly: 'period change disappears when we control for differences between birth cohorts.'

### *Models and results*

The existence of an educational equalization effect requires that the education – class origin association (EO) weakens over cohorts, while the compositional effect requires both that the distribution of education has shifted upwards over successive birth cohorts – expansion – and that there exists a three-way interaction between origins, education and current class position, or destination, such that the OD association is weaker at higher levels of education (differential association).

Figure 1 shows the trend over cohorts in class differences in educational attainment using the results of ordered logit models. In the figure each line refers to a class origin and shows how the coefficients for that class evolve over cohorts, with class I always acting as the reference category and having a coefficient of zero. The overall impression is of a decline in class inequalities in Germany and Sweden but a somewhat less clear picture in Great Britain. On the other hand, it appears that among the older cohorts class differences were less marked in Britain and even among the youngest cohort they are still less than in Germany.

Figure 2 shows the strength of the association between origins and destinations at each educational level (educational categories 1ab and c for Sweden were amalgamated for this figure). The parameters come from a log-multiplicative layer effect model (Xie 1992) which gives a parsimonious account of the association between origins and destinations across different educational levels. The associations are measured relative to that in the lowest educational level, where it is fixed at 1. In Sweden and Germany there is a tendency for the association to weaken

at higher levels of education (though in Germany it is weakest at the lower rather than the upper tertiary level). In Britain too there is variation across educational levels but the differences are a little less marked.

[FIGURE 2 HERE]

Figure 3 shows the expansion of education among men over the twentieth century by comparing the educational distributions of the second oldest and second youngest cohorts in each country. It is clear that all three countries have shown an expansion, particularly marked in Britain where the oldest cohort was very heavily clustered at the lowest educational level.

[FIGURE 3 HERE]

Figure 4 shows the trend in the gross association between class origins and class destinations in the three countries.<sup>7</sup> For the purposes of the figure, the Swedish birth cohorts have been collapsed into seven to aid comparability. It is immediately evident that the three countries display different patterns of change. In Sweden there has been a roughly constant decline in the origin – destination association over successive birth cohorts: this replicates findings of earlier studies such as Breen and Jonsson (2007). In Germany, an initial strengthening was followed by a steep decline, commencing with the third birth cohort (born 1931-39). This latter pattern is already well known from earlier studies of German social mobility (Müller and Pollak 2004).

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<sup>7</sup> Here the trend is captured using the log-multiplicative layer effect model (Xie, 1992). In none of the countries does this fit the data using the usual  $p < .05$  criterion (deviance statistics are 276 on 210 df for Germany; 392 on 336 for Sweden and 205 on 144 for Britain) but the trends discovered using less parsimonious models do not differ substantively from those shown in figure 4.

The British case also shows a strengthening of the origin – destination association, followed by a very modest decline commencing with the cohort born 1942-48.

### *Simulations*

So far I have concentrated on changes in social fluidity that may be induced by educational equalization and through the workings of the compositional effect. But there are two other factors that can also play a role. On the one hand, there may be changes over cohorts in the residual, direct effect of origins on destinations. Not all of the origin – destination association is mediated via education and, if those non-educational factors which are related to class origins and which constitute this direct effect change, or their effects change, then so will the unconditional origin – destination association. This would be captured in the  $\beta_{ijl}^{COD}$  term of a log-linear model. On the other hand, the impact of education on class destination may also change. This is another way of saying that the returns, in terms of class position, to educational attainment may change over cohorts. Formally:

$$\sum_k pr(D | EC)pr(E | O, C)pr(O, C) = pr(D | OC)pr(OC) = pr(DOC) = pr(DO | C)pr(C)$$

This can itself lead to cohort changes in social fluidity. The effect is captured in the  $\beta_{ikl}^{CED}$  term of a log-linear model. In assessing the impact of the equalization and compositional effects I would like to take account of these other two means by which the origin – destination relationship might change. Thus I generate a set of counterfactual ODC tables by successively removing terms from an initial log-linear model that includes all four avenues by which the unconditional origin – destination

relationship can change over cohorts. So I begin with (1) a model that includes all effects in the E|OC table and includes all effects in the D|OEC table except for the four way interaction,  $\beta_{ijkl}^{COED}$ . This model is:

$$f_{ijk.} = \mu \gamma_j^O \gamma_k^E \gamma_l^C \gamma_{il}^{OE} \gamma_{jk}^{CO} \gamma_{ij}^{CE} \gamma_{ik}^{COE} \gamma_{ijk}^{COE} \quad (4a)$$

for the EOC table; and

$$f_{ijkl} = \alpha \beta_j^O \beta_l^D \beta_k^E \beta_i^C \beta_{jk}^{OE} \beta_{jl}^{OD} \beta_{kl}^{ED} \beta_{ij}^{CO} \beta_{il}^{CD} \beta_{ik}^{CE} \beta_{ijl}^{COD} \beta_{ijk}^{COE} \beta_{ikl}^{CED} \beta_{jkl}^{OED} \quad (4b)$$

for the DOEC table. Given the fitted values from this model the table is computed by collapsing over levels of E.

I then, (2), remove the term  $\beta_{ijl}^{COD}$  from the second equation of the previous model and, once again, generate a hypothetical ODC table. Any log-linear model that includes the  $\beta_{ijl}^{COD}$  term will always yield an implied ODC table that reproduces the observed ODC distribution exactly: thus this term must be the first one removed from the model. The difference between the ODC distributions implied by models (1) and (2) tells us the effect of changes over cohorts in the partial effect of origins on destinations.

Next, (3), I remove the  $\beta_{ikl}^{CED}$  term from the second equation. The difference between the ODC distribution to which this gives rise and the previous one is a

measure of the consequences, for social fluidity, of changes in the returns to education, given that the impact of changes in the residual effect of origins on destinations has already been removed. Then, (4), I remove the  $\gamma_{ijk}^{COE}$  term from the first equation of the model. This yields the model already shown in equations 2a and 2b. Comparing the ODC distribution generated from this model and that generated from model (3) is a measure of the impact of educational equalization, given the removal of the two earlier effects.

Lastly, (5),  $\gamma_{ik}^{CE}$  is left out of the first equation, so removing educational expansion, and the  $\beta_{jkl}^{OED}$  term is left out of the second equation, removing differential association. Comparing the ODC distribution this model implies and that implied by model (4) allows one to gauge the compositional effect.

In this way I generate five counterfactual ODC tables for each of the three countries.<sup>8</sup> The goal is to compare across the set of five tables to assess the relative importance of the four effects listed above – and particularly educational equalization and the compositional effect – in contributing to a reduction in the overall association between origins and destinations. I do this by using the log-multiplicative layer effect model (Xie 1992) to estimate the trend in the origin – destination relationship in each of the counterfactual ODC tables. In its more usual implementation this model would set the log odds ratios in each cohort equal to the corresponding log odds ratio in the oldest cohort, multiplied by a cohort-specific

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<sup>8</sup> The reasons for choosing to remove the various effects in this particular order are explained in the technical appendix. The appendix also contains the syntax to fit these models in Lem (Vermunt 1997).

parameter:  $\ln(\theta_{jkl}^{ODC}) = \varphi_l \ln(\theta_{jkl}^{ODC})$  for  $l = 2, \dots, L$ , where  $\varphi_l$  is the cohort-specific parameter, conventionally set equal to one for the oldest cohort. The model used here, however, replaces this set of  $\varphi_l$  coefficients with a linear trend over cohorts:  $\ln(\theta_{jkl}^{ODC}) = (1 + \varphi \times [l - 1]) \ln(\theta_{jkl}^{ODC})$ . So there is a single slope coefficient,  $\varphi$ , which is multiplied by  $(l-1)$  to estimate the log odds ratios for the  $l^{\text{th}}$  table. Applying this approach yields a set of five values of  $\varphi$  each derived from one of the simulated ODC tables. By comparing these, the relative importance of each of the four factors can be measured.<sup>9</sup>

The results are shown in Table 4 and the Swedish case can be used to explain the method further. The first row ('Observed' model) relates to the log-multiplicative layer effect model just described applied to counterfactual table (1). The estimate of  $\varphi$  (multiplied by 100 for presentational purposes) in this case is -0.96, reflecting the decline in the origin – destination association across successive birth cohorts. A test that the 'true' slope is zero can be rejected at  $p = .027$ .<sup>10</sup> The second row reports the results of the same model applied to the second counterfactual table. Here the slope is more strongly negative, showing that, once the partial effect of origins on

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<sup>9</sup> The attraction of this approach is that, in each case, a single  $\varphi$  coefficient is generated and this makes comparisons straightforward. But the model assumes a linear trend in the origin – destination association and, as Figure 4 shows, at least for the observed data, this is not an accurate depiction of change in Britain or Germany. On the other hand, this linear specification can be thought of as a summary measure of the overall change between the oldest and youngest cohorts. I investigated other ways of making comparisons across the models: for example, fitting a quadratic trend and then computing the derivative of change in fluidity wrt cohort. In this case the derivative itself depends on cohort, and so I computed the average of this derivative across cohorts. This method gave a result very similar to the approach reported in the text.

<sup>10</sup> All the tests reported in Table 4 have one degree of freedom. Of course since the data to which the model is fitted are themselves generated from a model, this is not a valid test of statistical significance.

destinations is held constant over cohorts, inequality counterfactually declines more steeply than it is observed to. From this it follows that changes in the partial effect of origins on destinations have had a disequalizing effect, causing odds ratios to be generally higher than they would have been otherwise. The magnitude of that effect is shown in the final column: removing the ODC effect makes the slope more negative by a factor of 1.69. The negative value means that this effect acts to reduce social fluidity. When the impact of changes in the ED effect are removed inequality declines slightly more, showing that this too has a disequalizing effect, though of a much smaller magnitude (a factor of 1.14). However, removing the change in the origins – education association over cohorts shifts the slope in the opposite direction. It becomes much flatter, showing that, without equalization, the origin – destination association would not have declined (social fluidity would not have increased) as much as it has been observed to. Finally, applying the same logic, the compositional effect also contributes to greater social fluidity, though not quite to the same extent as educational equalization. Once the compositional effect is removed the remaining slope is not significantly different from zero in any country. This is because this model contains no pathways through which OD depends on C. In other words, the true slope in this case is zero, and any non-zero slope must be due to sampling variability.

Turning to Germany, the estimates of the slope of the origin – destination relationship are much more strongly negative than in Sweden.<sup>11</sup> In this case, all

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<sup>11</sup> The units in which the slope is measured are cohorts and since there are 15 cohorts in Sweden and seven in Germany and Britain comparisons of slopes between countries should be adjusted to take this into account.

effects, except EDC, contribute to increasing social fluidity. Whereas in Sweden the residual OD effect changed in such a way as to reduce social fluidity, in Germany it had the opposite effect. Once again, however, the equalization and compositional effect have played the greatest role in accounting for the overall increase in German social fluidity, with the latter being somewhat more important than the former.

Lastly, in Britain, it can be seen that in only one of the simulated distributions is the slope significant. Once the ODC effect is taken away the slope is essentially flat across cohorts. The ODC effect itself has operated in the direction of reducing social fluidity. There is also a paradoxical result that 'equalization' (but which should, more accurately, be termed changes in the origin – education association) has actually been slightly disequalizing. The only source that acted to increase social fluidity is the compositional effect, though it fails to reach statistical significance. It is notable that, although it was necessary to include, in the simulations for all countries, variation over cohorts in the relationship between education and class destination, this was nowhere an important source of change in social fluidity.

### *Conclusions*

In this paper I have sought to shed some light on a set of counterfactuals. These have been of the form: what would have happened had the *effect* of, say, class origins on educational attainment been constant over time compared with what actually occurred. The findings cannot support strong causal claims. As far as the equalization effect goes, it may be, for example, that left-leaning governments introduced policies that simultaneously made labor markets function more equitably and

opened the educational system to more working class children. In this case social fluidity and equalization would be correlated, and, while it is implausible to assume no causal effect, its impact may not have been correctly assessed. Similarly for the compositional effect, which might be observed if there was some characteristic of individuals that led them to have higher levels of education and made their class origins less relevant to the class position they came to occupy. An obvious possibility is that more able students reach higher levels of education, and, for able adults, class origins do not matter so much to the employers who hire and promote them. On the other hand, one argument against this interpretation is that there is no evidence of the compositional effect weakening over successive birth cohorts among which the selectivity of higher levels of education has steadily diminished.

In both Sweden and Germany the equalizing of educational chances had a consistent effect on increasing social fluidity. In Sweden equalizing educational outcomes has involved such things as the removal of fees for secondary and tertiary education, the introduction of free school meals and health care in schools, and free books and teaching aids in primary and secondary school. All these policies made educational opportunities less dependent on economic resources in the family of origin. Educational reforms, such as the comprehensive school reform in the 1950s, have aimed at increasing opportunities and educational attainment, particularly among those from less well off families (e.g., Erikson and Jonsson 1996). At the same time, policies were implemented to reduce the variation in the consequences of class position, and, according to Erikson (1996), this led to more equal educational attainments between children of different origins. In particular, he points to the

importance of equality of condition with respect to income and the risk of prolonged unemployment. Political means towards this include progressive taxation and redistribution mainly through the welfare state, for example by the introduction of the child allowance (Korpi and Palme 2004) and through Sweden's labor market policies.

On this account, educational equalization is the result both of specific policies within education and of an increasing equality of condition in society as a whole. If this is true (and crucially, if this is the only way to attain equalization) then Germany should have followed a similar path and Britain not. Jonsson, Mills and Müller (1996), who compare educational inequality in the same three countries, argue that Germany shares several relevant features with Sweden. They note that a high degree of economic security (prosperity and job stability) for all social classes was secured during the period of the *Wirtschaftswunder* while 'strong trade unions have, much like their Swedish counterparts, managed ... to secure for their members advantageous working conditions, pension schemes, and other arrangements...' (Jonsson *et al* 1996: 202). As far as specific educational policies are concerned, they point to the increase in the period of compulsory education which reduced the costs between *Hauptschule* and *Realschule*, and the increase in the level of general education needed to acquire an apprenticeship. This meant that although many working class children acquired higher levels of education in order to take up apprenticeships, they in fact equipped themselves 'with greater opportunities for continuing on to higher levels of education' (Jonsson *et al* 1996: 203).

Britain enacted many of the same types of policy changes in education as Sweden, yet, apparently, without the same effect. The 1944 Education Act in Britain, although it made secondary education free, preserved a highly differentiated system of schools, including a small private sector: differentiation within the state sector was much reduced after the late 1960s but has never quite disappeared. Access to tertiary education was quite stringently restricted at least until the 1970s. Within such a selective system (one of sponsored mobility, as it is sometimes called) educational expansion seems to have occurred without the same degree of equalization as in Germany and Sweden. In other words, working class children were not disproportionately the beneficiaries of educational reforms. Certainly post-compulsory levels of education were open to working class children, but they were quite stringently selected on the basis of examination performance. Similarly, the welfare state in Britain has never played the same kind of redistributive role seen in Sweden, even though income inequality was modest by international standards during the period between the end of the Second World War and the 1980s. Britain never attempted to introduce Swedish-type labor market policies and, unlike most of continental Europe, never developed any corporatist institutions. The persistence of a labor market for unskilled or semi-skilled labor may also have been an important factor in maintaining class differences in educational attainment, most obviously manifested through the high proportion of working class children who left school at age 15.

The motivation for this paper was to quantify the impact of educational equalization and the compositional effect on the trend in social fluidity. In all three countries the

compositional effect appears to have had a positive impact on social fluidity. In Germany its effect was stronger than that of equalization while in Sweden, where equalization was the main cause of increasing social fluidity, its impact was correspondingly less. In Britain it was the only effect that pointed in the direction of greater fluidity. Thus, to return to the initial discussion, sociologists of education and of stratification have largely overlooked the potential for educational expansion to promote greater equality in mobility chances even in cases where equalization did not occur.

Whether educational expansion is observed to be correlated with equalization in social fluidity will depend on whether it is accompanied by educational equalization, educational disequalization, or no change in educational fluidity. In the examples considered here there are cases in which expansion and equalization reinforce each other (Germany and Sweden) and one case in which expansion occurred with little or no change in educational fluidity (Britain). But one might also find cases in which the equalizing impact of expansion on social fluidity was offset by the manner in which equalization affected educational equalization; in particular, if the beneficiaries of expansion were mainly young people from more advantaged class origins this could lead to more educational inequality whose impact on social fluidity would run counter to that of equalization itself (Jonsson and Erikson 2007). Educational expansion alone is thus a somewhat risky policy if one's goal is to increase social fluidity, and, indeed, in many European countries, educational expansion *per se* was not an important strategy. On the other hand, and provided that the conditions hold for the compositional effect, then promoting educational equalization is likely to be

a very effective strategy because it also advances compositional change. This is because equalization almost certainly implies expansion of the middle and/ or higher levels of education, given that it is unlikely to occur through a reduction in educational participation by the middle classes (as noted by, among others, Raftery and Hout 1993: 60-1; and Breen and Jonsson 2007: 1805-6).

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## Technical Appendix

### Why I analyze only men

In analyses involving class position it is necessary to restrict attention to men and women who are currently participating in the labour force: failure to do this would mean, for example, the inclusion of some men and women whose class position is one that they occupied in the distant past or which is derived from that of their spouse, depending on the conventions followed in the particular survey. But, given a focus on those in the labor force, temporal comparisons will be compromised if rates of labor force participation change within birth cohorts as they age. In general terms this will not be a problem where men are concerned, but in Britain and Germany not only is there a historical trend towards greater female labor force participation but there are also pronounced life-cycle patterns of participation and non-participation. In Sweden this is much less of a problem because here the pattern of women's labor force participation varies less over the life course and is much more similar to men's. So if the focus was confined to Sweden it would be unproblematic to include women and men in the analysis (as in Breen and Jonsson 2007), but the inclusion of the other two countries makes this not feasible.

### Order of the decompositions

The method described in the text tests the impact of the four components by successively removing them from the model that generates the observed ODC distribution (shown as the log linear model of equations 4a and 4b). The components are removed in this order:  $\beta_{ijl}^{COD}$ ,  $\beta_{ikl}^{CED}$ ,  $\gamma_{ijk}^{COE}$  and, lastly,  $\gamma_{ik}^{CE}$  and  $\beta_{jkl}^{OED}$  together. There is little discretion involved in this choice. As long as  $\beta_{ijl}^{COD}$  remains in the log linear

model the implied ODC margin will always be identical to the observed ODC margin, and so it must be removed first. In addition,  $\beta_{jkl}^{OED}$  cannot be removed before  $\gamma_{ijk}^{COE}$  because the compositional effect requires two terms in the model, namely  $\beta_{jkl}^{OED}$  and  $\gamma_{ik}^{CE}$ , the latter capturing shifts in the educational distribution over cohorts. Thus to remove the compositional effect would require the removal of both of these; but in that case, the model which seeks to capture equalization would be non-hierarchical, including  $\gamma_{ijk}^{COE}$  but not  $\gamma_{ik}^{CE}$ .

### Comparing slopes

The effect of different influences on cohort change in the OD relationship is measured through comparisons of slopes from different models. While this approach is parsimonious it also deals with a difficulty that arises in making such comparisons. When a particular term, such as  $\beta_{ijl}^{COD}$  is removed from the log linear model, the result is that the partial association between origins and destinations is set at some fixed value over all cohorts. Unfortunately this is not its value in the oldest cohort but, rather, a complicated 'average' value taken over all cohorts. If the  $\beta_{ijl}^{COD}$  effect is such that the partial OD association weakens over cohorts, removing it will tend to reduce the association in older cohorts and increase it in younger ones. This general problem means that straightforward comparisons of fluidity in a given cohort are not a feasible way of measuring the impact of effects such as  $\beta_{ijl}^{COD}$  because the model that omits such a term does not represent a situation in which the partial OD association remains fixed at its value in the oldest cohort. Comparing slopes circumvents this problem because the slope is invariant to rescaling the associations

in every cohort according to the value of  $\beta_{jt}^{OD}$  - in other words, the slope is unaffected by whether the partial OD association is set at its value in the oldest cohort or at some average (across cohorts) value.

#### Fitting the models

The models were fit using the Lem program (Vermunt 1997). The 'Observed' model is fitted using the 'mod' command:

```
mod E|CO {EOC}
```

```
D|OCE {DOC DEC OED}
```

together with the write margin command:

```
wma ODC allchange.txt
```

which saves the fitted ODC margin to the file 'allchange.txt'. The other four models were fitted analogously.

Models were fit to the saved marginal distributions using the 'mod' and 'des' commands like this:

```
mod {CO CD spe(OD,1a,C,b,-2)}
```

```
des [1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1
```

```
0 1 2 3 4 5 6 7 8 9 10 11 12 13 14]
```

This example is for the Swedish data, fitting a continuous version of the log-multiplicative layer effect model over the 14 cohorts defined there.



Table 1: Effects of Hypothetical Changes in Education

		Time = 1			Time = 2								
					(a) Equalization only			(b) Expansion only			(c) Equalization and expansion		
	Educational level	Low	High	OR	Low	High	OR	Low	High	OR	Low	High	OR
<b>OE</b>	<i>Middle class</i>	50	50	1.50	52	48	1.31	40	60	1.50	42	58	1.31
	<i>Working class</i>	90	60		88	62		75	75		73	77	
		<i>Middle</i>	<i>Working</i>	<i>OR</i>	<i>Middle</i>	<i>Working</i>	<i>OR</i>	<i>Middle</i>	<i>Working</i>	<i>OR</i>	<i>Middle</i>	<i>Working</i>	<i>OR</i>
<b>OD   E=low</b>	<i>Middle class</i>	35	15	5.4	36.4	15.6	5.4	28	12	5.4	29.4	12.6	5.4
	<i>Working class</i>	27	63		26.4	61.6		22.5	52.5		21.9	51.1	
		<i>Middle</i>	<i>Working</i>	<i>OR</i>	<i>Middle</i>	<i>Working</i>	<i>OR</i>	<i>Middle</i>	<i>Working</i>	<i>OR</i>	<i>Middle</i>	<i>Working</i>	<i>OR</i>
<b>OD   E=high</b>	<i>Middle class</i>	40	10	2.7	38.4	9.6	2.7	48	12	2.7	46.4	11.6	2.7
	<i>Working class</i>	36	24		37.2	24.8		45	30		46.2	30.8	
		<i>Middle</i>	<i>Working</i>	<i>OR</i>	<i>Middle</i>	<i>Working</i>	<i>OR</i>	<i>Middle</i>	<i>Working</i>	<i>OR</i>	<i>Middle</i>	<i>Working</i>	<i>OR</i>
<b>OD overall</b>	<i>Middle class</i>	75	25	<b>4.14</b>	74.8	25.2	<b>4.03</b>	76	24	<b>3.87</b>	75.8	24.2	<b>3.77</b>
	<i>Working class</i>	63	87		63.6	86.4		67.5	82.5		68.1	81.9	



Table 2: Samples from each country

<b>Country</b>	<b># tables</b>	<b>Sources of data</b>	<b>Years for which data are included</b>
Germany	30	Zumabus	1976-77 1979(2) 1980 1982
		Allbus	1980 1982 1984 1986 1988 1990-92 1994 1996 1998 2000 2002
		Politik in der BRD	1978 1980
		Wohlfahrtssurvey	1978
		German Life History Study	I (1981-1983) II (1985-1988) III (1988-1989)
		German socio-economic panel	1986 1999 2000
Great Britain	15	General household survey	1973 1975-76 1979-1984 1987-1992
Sweden	24	Annual surveys of living conditions (ULF)	1976-1999

Table 3: Estimated change in the Origin-Destination relationship in different simulations

(a) Sweden

Model	Slope (*100)	Test of slope		Contributions %
		Chi-square	p	
<i>Observed</i>	-0.96	4.891	0.0270	-
<i>No_ODC</i>	-1.62	16.1304	0.0001	-68.75
<i>No_EDC</i>	-1.75	19.3875	0.0000	-13.54
<i>No_equalization</i>	-0.76	3.0995	0.0783	103.13
<i>No_compositional</i>	0.00	0.6305	0.4272	79.17

(b) Germany

Model	Slope (*100)	Test of slope		Contributions %
		Chi-square	p	
<i>Observed</i>	-4.71	30.2204	0.0000	
<i>No_ODC</i>	-4.08	21.7865	0.0000	13.32
<i>No_EDC</i>	-4.2	23.4263	0.0000	-2.54
<i>No_equalization</i>	-2.43	7.0039	0.0081	37.42
<i>No_compositional</i>	0.02	0.9664	0.3256	51.80

(c) Gt. Britain

Model	Slope (*100)	Test of slope		Contributions %
		Chi-square	p	
<i>Observed</i>	1.98	6.2643	0.0123	
<i>No_ODC</i>	-0.77	1.1257	0.2887	-154.49
<i>No_EDC</i>	-0.86	1.3998	0.2368	-5.06
<i>No_equalization</i>	-1.2	2.7741	0.0958	-19.10
<i>No_compositional</i>	0.2	0.0713	0.7895	78.65



Figure 1: OE association over birth cohorts

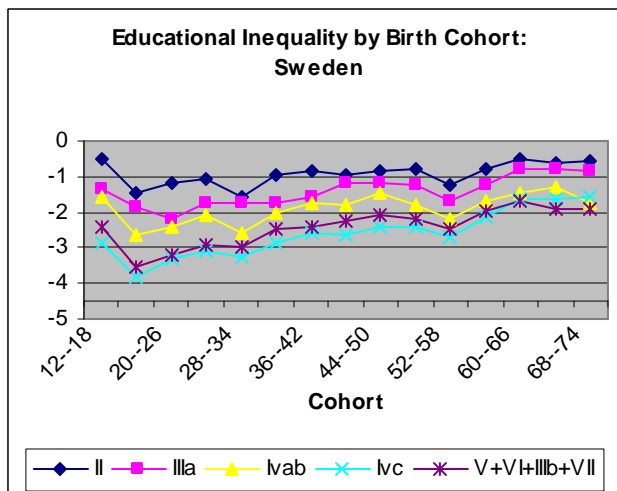
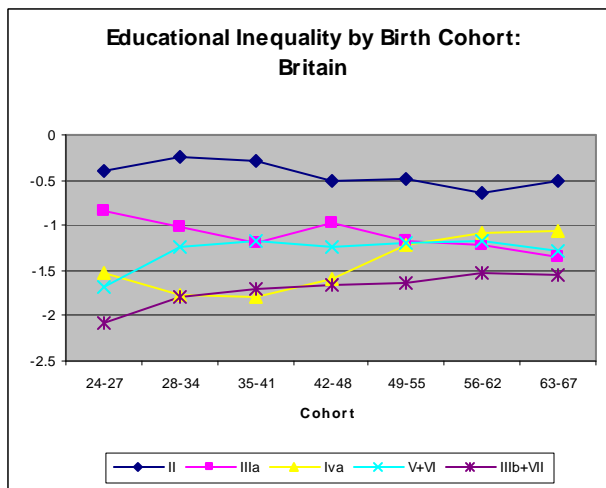
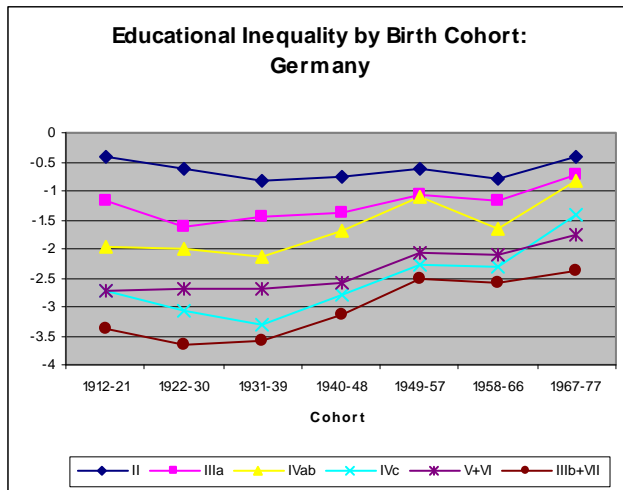


Figure 2: Association between origins and destinations at each level of education

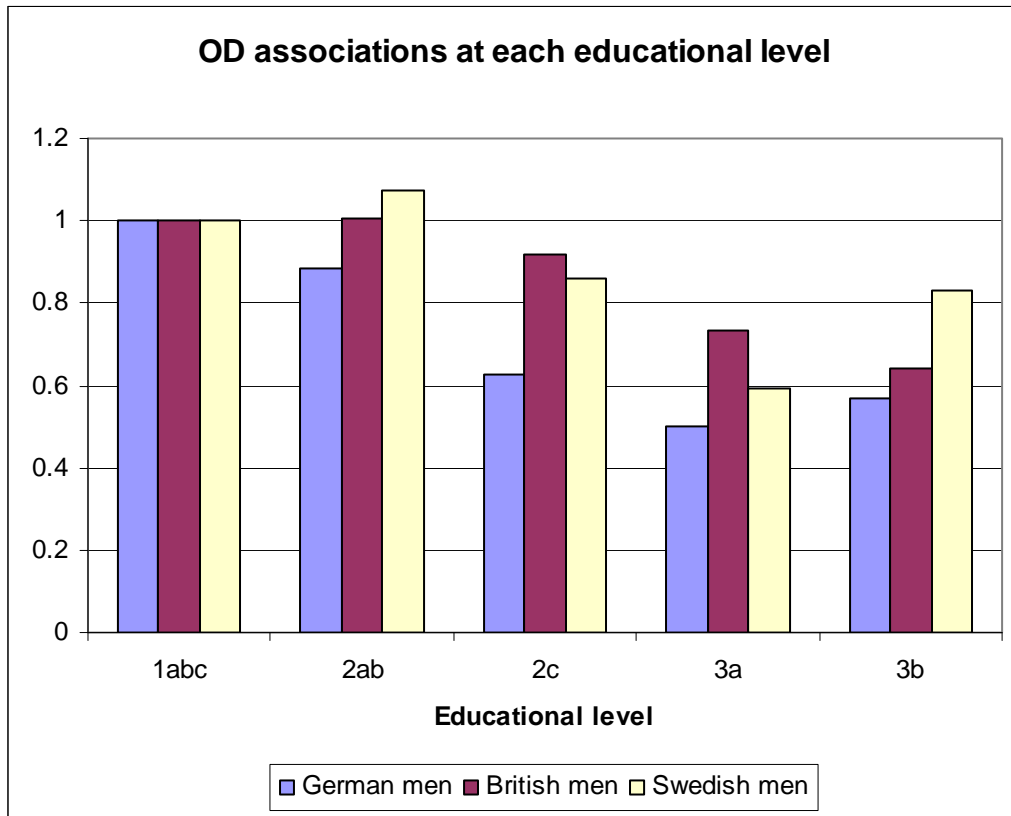


Figure 3: Educational distributions in oldest and youngest cohorts

