INTERGENERATIONAL EDUCATIONAL MOBILITY IN THE NETHERLANDS FOR BIRTH COHORTS FROM 1891 THROUGH 1960*

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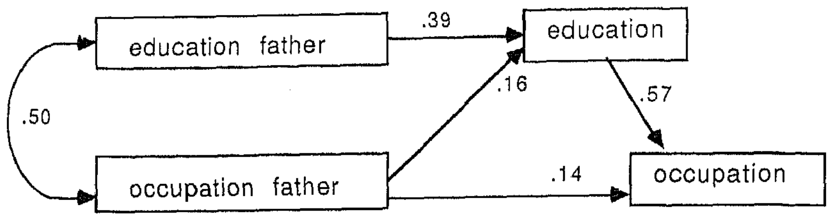
Introduction – The Analysis of Educational Mobility

Occupational achievement is the most important issue for the sociological study of intergenerational transmission of social inequality. Modelling the association between occupational positions of fathers and sons has received most attention; mothers and, to a lesser extent, daughters are often neglected. From Sorokin's landmark monograph on social stratification (1927) to the more recent work of Blau & Duncan (1967), Featherman & Hauser (1978) and Goldthorpe (1980), there has been an unbroken series of empirical and theoretical investigations on this form of intergenerational mobility. In the Netherlands, a similar stress on occupational status of men as the main dimension of social inequality can be observed in the work of the 'Leyden School' (Van Heek, 1945; Van Tulder, 1962; Ultee, 1984) and, more recently, in a number of studies of historical trends (Ganzeboom & De Graaf, 1984; Ganzeboom *et al.*, 1987; Luijkx & Ganzeboom, 1989).

Concentration on the intergenerational transmission of occupation rests upon a number of considerations, some of which are of dubious merit. First, researchers are often interested in occupational achievement because it is thought to be the 'single best dimension of social inequality' (Blau & Duncan, 1967: pp. 6–7). These authors argue that most inequalities within a society are reflected in the occupational distribution. A second, related argument is that occupational achievement is the principal criterium for social interaction patterns. The prestige or status provided by an occupation is thought to be the most important determinant of general esteem and, as such, it is supposed to regulate the attribution of respect or contempt. A third argument is that occupation, especially when categorised by class, is best equipped to predict political dispositions and behaviour. Finally, the one-sided concentration in occupational mobility research on the positions of fathers and sons is supported by the conventional view that husband's occupation determines the social status of the entire family, and that the wife's status is more or less derivative (Goldthorpe, 1983).

Figure 1 The status attainment model for the Netherlands in 1977, men aged 25-64



Source: De Graaf, 1987, p. 5; based on the CBS Life Situation Survey 1977

In the research reported here, we will focus on educational mobility between generations rather than on occupational mobility. We think that there are a number of arguments which strongly support this approach.[1]

First, it is empirically doubtful whether the central dimension of social stratification is in fact occupation. There is much evidence to the contrary that it is not an individual's position in the occupational structure, but an individual's educational attainment which is the appropriate central dimension. Values and norms, choice of partner and friends, cognitive abilities, cultural activities, leisure time behaviour in general, and children's school success, all are more closely associated with educational attainment than with occupational position (cf. literature reviewed by De Graaf, Ganzeboom & Kalmijn, 1989).

A second and related argument in favour of the analysis of educational mobility data is in the very structure of the status attainment model. Figure 1 gives an estimation of the classical status attainment model for the Netherlands in 1977 (De Graaf, 1987). Like most models of this type, the model shows that the transmission of educational positions from one generation to the next is *the* vehicle of intergenerational mobility. There is a strong connection between the educational positions of the two generations, which largely determines the relation between the occupational statuses of two generations. These processes are hidden in the occupational mobility table. The other side of this argument is in the observation that the relations in the educational mobility table reflect direct processes. Relations in the occupational mobility table are a composite of two processes: one indirect, via educational attainment, and one direct, via the straightforward transmission of occupational position from parents to children. To summarise: intergenerational transmission of social status should be studied through the transference of educational attainment, because it is the most important component in the status attainment process.[2]

The third argument in favour of the analysis of educational mobility is in the absence of a gender bias in the study of educational processes. Occupation is an unsatisfactory indicator of social status among women, especially for older cohorts, since they are less often employed than men. On the other hand, all women have some level of education, as a matter of course, and this implies that education supplies mobility data for men and women with an equal measurement procedure. In the Netherlands this advantage is extremely important due to the traditional dearth of female participation in the labour market. Gender bias caused by an elevated number of unemployed women is not the only point of neglect in the occupational mobility table: the same is true of all other groups outside the labour market, such as the unemployed, the retired, and the disabled. A fourth advantage is that educational attainment is a fairly fixed characteristic over a person's life course. At a certain age, which for present purposes we will assume to be twenty-five years, the highest level of education attained is no longer subject to change. In other words, age effects are absent. The consequence is that the processes observed in an educational mobility table are fixed in historical time.

Events within an educational career can therefore be attributed with reasonable accuracy to a certain year, on the basis of the respondent's age. It is consequently possible to assess historical changes by cohort analysis, without confusions with age or period effects. Events underlying educational mobility can also be allocated to a fixed historical period and are thus ready to be related to external historical conditions, such as legislation and economic development. This brings a time series design for the analysis of intergenerational transmission of inequality within reach. Cohort analysis is problematic in the case of occupational mobility, because occupation varies during the life cycle. Only with occupational data which are fixed at the beginning of the occupational career (first occupation), or with data on full occupational careers and exact dates of job changes, does satisfactory cohort comparison become possible (Blossfeld, 1987).

Fifthly, the study of educational mobility is facilitated by a relatively large amount of existing data. Historical comparison of social mobility data are necessarily based on the analysis of repeated cross-sectional surveys. Occupational mobility tables for the whole population are compared over survey years. The disadvantage here is not only that the data within these occupational mobility tables reflect events which cannot be linked to specific external variables, but also that these data are scarce and do not go far back in time. The present analysis allows us to report on educational mobility for nearly 23,000 men and women born in the period between 1891 and 1960, where the same data sets would supply information on occupational mobility data of only some 10,000 men and women, distributed over fourteen population surveys with known measurement times, but with unknown 'event' times.

In summary, there are several good arguments for giving the analysis of educational mobility a more central position on the sociological agenda. This article is intended as a contribution to this aim. We are not, of course, attacking the assumption that occupational mobility is a sociologically relevant phenomenon. Occupation is and remains an individual characteristic that is decisive for a wide range of life chances in our society, especially for (and via) income attainment. Occupational mobility has aspects that have no parallel in educational mobility: this is especially the case for direct material transmission of occupation from parents to children via the inheritance of possessions, and for the analysis of parental influences during the rest of the life cycle. Nevertheless, greater attention to the transmission of educational mobility seems to be justified. This article is limited to a descriptive analysis of intergenerational educational mobility of the Dutch population over a long historical period. We aim to merge as many data as possible and to test the presence or absence of shifts in the association between parents' and children's educational attainments over successive birth cohorts. Our research questions are: How can the relative educational opportunities of men and women from different family backgrounds (indicated by father's educational attainment) be described, and have these opportunities changed during the twentieth century? An explanatory analysis, in which changes in educational mobility are related to external changes, is postponed: for the moment we simply aim to establish the presence or absence of social change in educational opportunities in the Netherlands.

Research on the intergenerational association of educational positions of men and women in the Netherlands has been dominated by research on specific transitions in educational careers (cf. Dronkers, 1983; Vrooman & Dronkers, 1986). The main focus largely has been the transition from primary to secondary education, when parents and children (of twelve years) must choose between different levels of secondary education. Comparison of cohorts crossing this major transition point in the Dutch educational career at different periods, from the 'forties to the 'eighties, led to the conclusion that socioeconomic background is as important for younger cohorts as it was for older ones.[3]

Mare (1981) has argued that cohort change (or stability) in family effects on school continuation probabilities does not necessarily mean that the relationship between family background and total educational attainment has also changed (or been stable). Using data from the United States, he was able to demonstrate that increased family effects on school continuation decisions are in fact coupled with a stable relationship between family background and total years of schooling. This paradox probably originates firstly in the loss of control by parents over decisions about schooling which they take when their children are young, but which later on the children take for themselves. Thus, family influence decreases over the course of the educational career. Secondly, the general process of educational expansion means that more and more children survive the early hurdles in their educational careers. As a consequence later decision points in educational careers, which are less dependent upon family background, have grown in importance. The convergence between increasing family influence on school continuation decisions seen across cohorts, the decrease in family effects over the educational career, and the general increase in school enrolment apparently produced a stable overall relationship between family background and educational attainment in the United States. This is of course not necessarily the case for other countries. For the Philippines, Smith and Cheung (1986) found stable background effects on school continuation probabilities and decreasing overall associations, and logically the same could be the case for the Netherlands. Only further historical comparative research, on both school continuation decisions other than the transition from primary to secondary education, and on the overall association would provide an answer.

Empirical studies by Peschar on the overall association between family background and educational attainment have concluded that there has been hardly any change (Peschar & Popping, 1986; Peschar, 1987). Peschar used the same analytical strategy of analysis that we ourselves employ. He analysed educational mobility tables for four successive birth cohorts and, applying log-linear analysis, he found no change in association at all. Our analysis differs from that of Peschar basically only in the amount of data used for comparison, and hence by the statistical *power* of the research design.[4] Apart from the national labour market survey (1982), which Peschar used, we have thirteen other data-sets (1970 through 1986) at our disposal. There are two main advantages in doing this. Firstly, by employing older population surveys, we are able to extend the period investigated much further back in history. Secondly, our much larger total sample provides more statistical power with which we can look at the processes involved. Whereas Peschar's analysis is based on 840 men and 946 women, born between 1925 and 1965, we will analyse data on 11,892 men and 10,895 women born between 1891 and 1960. As we shall see, this permits us to uncover significant and relevant trends which were logically unavailable to Peschar.

Research design and data

For the analysis of stability or change in educational mobility for a maximum time span and a maximum number of individuals, we have merged all available national Dutch cross-sectional data-sets with information on the educational attainment of the respondents and their fathers, together with respondent's gender and year of birth. The data-sets are described in Table 1; they were all collected between 1970 and 1986. We have limited our analyses to respondents with complete information on education, father's education, age and gender, who were at least 25 years of age at the time of the interview. Without this latter limitation we would have created a selection on the dependent variable 'educational attainment', because the longest educational careers last until approximately 25 years of age. This selection means that our analysis is limited to cohorts born before 1960. On the other hand, we

Table 1. Description of the data-sets used

Title of the data-set (researcher, no. Steinmetzarchives)	Year	Number of cases analysed
Progressiveness and Conservatism (Middendorp, P0079)	1970	1114 (4.9%)
National Election Survey 1970 (Stouthardt et al., P0136)	1970	972 (4.3%)
Political Participation Survey (Research Group National Election Survey, P0355)	1971	731 (3.2%)
Life Situation Survey 1974 (Central Bureau of Statistics, P0210)	1974	3515 (15.4%)
Income Satisfaction 1976 (Hermkens & Van Wijngaarden, P0653)	1976	683 (3.0%)
Life Situation Survey 1977 (Central Bureau of Statistics, P0328)	1977	3067 (13.5%)
National Election Survey 1977 (Research Group National Election Survey, P0354)	1977	1467 (6.4%)
Political Action, second survey (Barnes & Kaase, P0322)	1979	1083 (4.8%)
Prestige and Occupational Mobility (Ultee & Sixma, P0839)	1982	565 (2.5%)
National Labour Market Survey Program (Heinen & Maas, P0748)	1982	1593 (7.0%)
Life Situation Survey 1983 (Central Bureau of Statistics, P0761)	1983	3098 (13.6%)
Organisation Strategic Labour Market Research (OSA, not in Steinmetzarchives)	1985	2991 (13.1%)
National Election Survey 1986 (Van der Eijk et al., P0866A)	1986	1234 (5.4%)
Income Satisfaction 1987 (Hermkens & Van Wijngaarden, not in Steinmetzarchives)	1987	674 (3.0%)
Total		22787 (100%)

Table 2. Numbers of cases, according to birth cohort and sex

Birth cohort	Men	Women
1891–1895	89	96
1896–1900	184	202
1901–1905	323	382
1906–1910	429	513
1911–1915	601	598
1916–1920	749	708
1921–1925	1017	915
1926–1930	1100	1037
1931–1935	1273	1020
1936–1940	1330	1110
1941–1945	1548	1322
1946–1950	1626	1437
1951–1955	1037	984
1956–1960	586	571
Total	11892	10895

have little information on birth cohorts born before 1891, and so our historical comparison deals with birth cohorts between 1891 and 1960.

To facilitate the analysis we have categorised all cases into fourteen birth cohorts, each of five years; the oldest cohort was born between 1891 and 1895, and the youngest one between 1956 and 1960. Table 2 shows the numbers of men and women in each birth cohort.

The oldest cohorts are those born in the first half of the 1890s; the first persons in this cohort to complete their educational careers did so around the year 1900. The youngest persons in our analysis were born in 1960 and have only recently finished their educational careers. This implies that our analysis covers processes between about 1900 and 1980 but, since both the oldest and youngest cohorts are sparsely represented, our conclusions primarily reflect events in the period between 1910 and 1975.

Different coding schemes for educational attainment were used in the fourteen population surveys. These sometimes varied between respondents and their fathers within a single survey. We sought a 'common denominator', an overarching categorisation to render the codings comparable, and this we found in the following four level scheme:[5]

1. Less than primary education or primary education (ISCED levels 0 and I);
2. Lower secondary general education or lower secondary vocational training (ISCED level II.1);
3. Higher secondary general education or middle level vocational training (ISCED level II.2);
4. Higher vocational training or university level (ISCED level III).

Respondents and their fathers were recoded within this four level categorisation without any problem, using the detailed educational codings of each individual survey. We next constructed four-to-four mobility tables for each birth cohort, separately for men and women. We shall thus analyse a total of two times fourteen tables to establish trends in intergenerational educational mobility in the Netherlands. The mobility tables are listed in the Appendix.

Analysis: Models

We shall analyse educational mobility, analogously with the analysis of occupational mobility tables, by way of log-linear analysis (cf. Ganzeboom *et al.*, 1987). The model we shall apply here is Goodman's (1979) scaled association model II. The four educational categories in this model are scaled in such a way that the total association in a table can be expressed with a single parameter, similar to a correlation or regression coefficient. Next to this general association parameter, a second set of parameters is estimated for the relative densities on the main diagonal of the table, where fathers and respondents are in identical educational categories:

$$\ln (F_{ij}) = O_i + D_j + U^*U_i^*U_j + DIA_i$$

The model contains the following parameters:
- F_{ij} are the expected frequencies.
- O_i and D_j are sets of parameters to reconstruct the exact distributions in the marginals of the tables.
- the $U^*U_i^*U_j$ parameter set consists of a general association parameter U, which generates that the expected F_{ij} are smaller for cells which represent larger mobility. This type of model is usually denoted as the (quasi-)uniform scaled association model, since the pattern of association in the table (with exception of the main diagonal) is described in terms of one uniform parameter, U, which describes the expected frequencies in the table as a whole.
- the U_i and U_j parameters stand for a scaling of the categories such that the model optimally reproduces the observed frequencies. It is conventional to assume that rows and columns, i.e. educational categories of respondents and their fathers, have equal scaling (U_i=U_j).
- the DIA_i parameter takes care of a special treatment of the diagonal cells in the mobility table. The densities in these cells tend to be larger than might be expected from the other components in the model and, especially for occupational mobility, tend to vary across categories. If the DIA_i coefficients do not differ between categories, we will refer to them as INH.

The differences between the models will be evaluated with the Bayesian Information Coefficient BIC (Raftery, 1986), which takes into account not only the number of degrees of freedom and the likelihood ratio statistic, but also the (large) number of cases. Models with a (more) negative BIC are preferred.

Analysis: Results

The analysis will be divided into two parts, both reported in Table 3. Panel A of Table 3 serves to establish a satisfying pattern for the association in the tables, whereas Panel B displays models which survey differences in association over cohorts.

The models in Panel A are related to two mobility tables (one for men and one for women), aggregated over all fourteen birth cohorts, as given in the bottom part of the Appendix. We estimate two models for each step: both are

Table 3. Log-linear models of intergenerational educational mobility in the Netherlands, birth cohorts 1891–1960

Panel A: models for patterns of association

				equal scaling:			optimal scaling:		
				NDF	L^2	BIC	NDF	L^2	BIC
(A.1)	(O+D)*S + DIA*S		+U*S	8	61.0	–19.3	6	58.9	– 1.2
(A.2)	(O+D)*S + DIA	+INH*S	+U*S	11	68.7	–41.7	9	66.0	–23.7
(A.3)	(O+D)*S + DIA		+U*S	12	70.2	–50.3	10	68.6	–31.8
(A.4)	(O+D)*S	+INH*S	+U*S	14	167.5	27.0	12	83.0	–37.4
(A.5)	(O+D)*S	+INH	+U*S	15	167.9	17.4	13	85.8	–44.7
(A.6)	(O+D)*S	+INH	+U	16	168.1	7.6	14	86.0	–54.5
(A.7)	(O+D)*S		+U	17	373.8	203.2	15	220.6	70.1

Panel B: models for differences in association between birth cohorts

			NDF	L^2	BIC	L^2	BIC
				men		women	
(B.1)	(O+D)*T + INH*T	+U*T	96	162.4	–738.2	136.4	–756.0
(B.2)	(O+D)*T + INH*Y	+U*T	108	170.2	–843.2	147.0	–847.0
(B.3)	(O+D)*T + INH	+U*T	109	171.2	–851.6	156.1	–857.2
(B.4)	(O+D)*T + INH	+U*Y	121	176.9	–958.5	176.4	–948.4
(B.5)	(O+D)*T + INH	+U	122	210.7	–934.1	190.5	–943.7

O=origins; D=destinations; S=sex; DIA=immobility parameters for separate diagonal cells; INH=general immobility parameter; T=cohort categorical; U=association parameter; Y=cohort linear; NDF=number of degrees of freedom; L^2 =likelihood ratio (chi squared distributed); BIC: Bayesian Information Coefficient (Raftery, 1986). The models of panel B are estimated with fixed (optimal) category-scalings.

(quasi-) uniform association models. We use *equal scaling* between the educational categories for the models in the left column, while *optimal scaling* (U_i=U_j) of the categories was applied in the right column. The differences between the successive steps lie in the different treatment of the diagonal cells and the potentially dissimilar association patterns for men and women.

Model A.1 estimates the simple quasi-uniform association model for each table, in which both the association coefficient U and the immobility parameters DIA_i are estimated differently for men and women. Model A.2 is simplified with regard to the densities on the main diagonal: these are split up in a parameter (DIA) which makes distinctions between the four educational categories in an equal way for both sexes, and one general coefficient (INH) that denotes a general difference on the main diagonal between men and women. This assumption improves the fit of the model. The assumption that there are no differences between men and women with respect to immobility is tested in Model A.3: this assumption also improves the model. Until this point no differences occur between the models with equal and with optimal scaling respectively. The advantage of optimal scaling, however, is evident in model A.4, where the assumption that there is unequal immobility over the four educational categories is removed. The model with equal scaling shows no improvement but, as the decreasing BIC statistic clearly demonstrates, the model with optimal scaling does.

Models A.5 and A.6 test the significance of the sex specific differences in the general immobility parameter (INH) and the general association parameter (U), respectively. Neither of these models weakens the model fit. In model A.7 the general immobility parameter INH is removed, but this hypothesis worsens the fit significantly, so that A.6 is our preferred model.

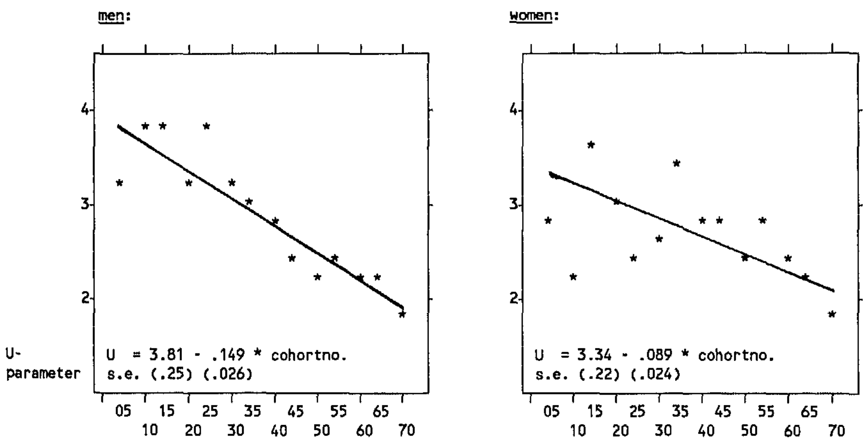
Thus, *three components* are necessary to describe the pattern of association in intergenerational educational mobility in the Netherlands: (a) the optimal scaling of the categories, (b) a general association parameter U, and (c) a general immobility parameter INH which indicates that respondents with the same educational attainment as their father are over-represented, given the general association pattern. The estimated optimal scaling of the four educational categories in model A.6 is –0.717, –0.156, 0.238 and 0.636. The differences as compared to equal scaling are not very large: they show that the distance between primary education and the three other categories is relatively large[6]. The value of the general association parameter in model A.6 is 2.713, and the value of the general immobility parameter is 0.250. These positive values imply that the probability of mobility decreases with the distance but that, given this structure, the probability for a son or daughter to be in exactly the same educational category as his or her father is even greater.

The models in Panel B of Table 3 analyse the presence or absence of trends in the association patterns over cohorts. This analysis reports on fourteen tables, and is performed separately for men and women.

Model B.1 is similar to model A.6, but is now estimated for each birth cohort, with a separate association parameter U and a separate immobility parameter INH for each of the fourteen cohorts. In the models which follow we attempt to improve and simplify this first one, summarising the differences in these two parameters by adding either a linear trend or an equality constraint. Model B.2 estimates a linear trend for the immobility parameter INH. Model B.3 removes this trend and constrains INH to be equal over all cohorts. Neither of these steps weakens the model fit.[7] Models B.4 and B.5 perform the same steps for the association parameter U. A linear trend for the association considerably improves the model for both men and women, whereas the equality constraint impairs the model fit once again. The model shows therefore that there has been significant social change for both men and women. According to the model fit this trend is stronger for men than for women.

The direction of this social change can be read from Figure 2, in which the (unrestricted) association parameter U from model B.3 is given for the year in which the cohort was twelve years old. The linear trend of preferred model B.4 is also displayed. There is a sharp decrease in the association coefficient for both sexes, which implies that educational mobility in the Netherlands

Figure 2 Intergenerational educational mobility in the Netherlands, birth cohorts 1891-1960; association parameters U of model B.3 and linear trend of model B.4



The cohorts are scaled according to the year in which their average age was twelve years has greatly increased during the twentieth century. The standard errors of the estimated equations, which are included in the graphics, indicate that the decrease in association differs significantly between the sexes and has been less steep for women than for men. In sum: there is a clear trend towards more openness, and this is stronger for men than it is for women.

Conclusions and Discussion

Our analysis of intergenerational educational mobility in the Netherlands, for birth cohorts from 1891 to 1960, has resulted in the following two conclusions:

a. The pattern of educational mobility can be described with a *simple* log-linear model, certainly when compared with the models for occupational mobility. The pattern can be summarised with an optimal scaling of the categories, a general association coefficient and a general immobility coefficient. Differences between birth cohorts can even be summarised by differences in one coefficient, namely the general association coefficient.
b. There is a *clear trend towards more mobility*, and this trend is linear. This conclusion, that Dutch society has become more open in the 20th century, is in line with empirical results on changes in occupational mobility (Ganzeboom & De Graaf, 1984; Ganzeboom et al., 1987), and in marriage patterns (Sixma & Ultee, 1984).

In this discussion, we would like to stress some of the strong points of our methods of data collection and analysis. We have brought together data from fourteen different national population surveys. Since the trend we observe is over birth cohorts – and not surveys – it is very unlikely that it can be reduced to the peculiarities which are undoubtedly inherent to the separate surveys. These peculiarities are made into random errors by our design. We would also like to emphasise the large number of individuals (22,787) we brought together in one data-set. The trend observed would certainly not have been so clear with a smaller number of cases. This, in our opinion, is why our conclusions differ from those of Peschar (Peschar, 1987; Peschar & Popping, 1986), who found no change in educational mobility over birth cohorts. Peschar used only one of the fourteen population surveys we have brought together here, and he has therefore analysed only 7% of our number of cases (Table 1). Additional calculations suggest that the rather strong decrease in association we observe, requires sample sizes of at least 1351 men and 2960 women to be statistically significant.[8] Peschar had data on only 840 men and 946 women, and thus was not able to find significant trends. Our material does have some disadvantages as well. The main disadvantage is that, although one can look relatively far in the past, it is much more difficult to observe recent developments. In order to have finished educational careers, a birth cohort must be at least twenty-five years old. If one supposes that the important decisions in educational careers are made when a student is about ten to fifteen years old, then a researcher lags at least fifteen years behind the facts. This period will be even longer in practice, since the age limit of twenty-five is rather low for certainty that everyone has finished his or her educational career, and because a researcher has to wait for the newest surveys. What this means for our analysis is that most of the processes which are included in our tables took place before 1975 and that recent changes in Dutch educational policies (during the 'seventies), which aimed to reduce the influence of family background, cannot be evaluated with these data. It may be more appropriate to focus on school continuation decisions, for example by following cohorts through the educational system, for the analysis of recent developments (Vrooman & Dronkers, 1986). However, we stress that such an analysis does not study the highest level attained and can therefore only provide a partial picture. Those wishing to study long-term developments in educational mobility, or to evaluate the total impact of educational reforms, must have patience.

NOTES

* A Dutch version of this article was published in *Sociale Wetenschappen*, 32 (1989) pp. 263–277. An earlier version of the English paper was presented at the meetings of the ISA Research Committee on Social Stratification and Mobility in Utrecht, April 26–29, 1989. We would like to thank the persons and organisations who permitted us to use their data-sets, and the Steinmetz Archives in Amsterdam for their assistance.

1. In this we partly follow Peschar (1987; Peschar & Popping, 1986), who compared educational and not occupational mobility for Dutch men and women with Polish and Hungarian data.
2. The relevance of this argument has also been recognised in government practice. Authorities attempting to banish intergenerational transmission of social inequality, have mainly focussed their efforts on trying to remove background influences in educational careers.
3. The analysis did however observe differences in the mechanism of transmission, where children's academic achievements have replaced parents' social status as the main selection criterion. The result of no change in final outcome is produced by the capacity of high status parents to promote the academic achievement of their children.
4. We also employ a slightly different log-linear model and more parsimonious restrictions, but this is of secondary importance as compared with the difference in the amount of data.
5. This scheme is identical to the one used by Peschar (1987) and has much in common with the first digit of the Dutch S.O.I. 1978 (Standard Educational Categorisation of the Central Bureau of Statistics) (CBS, 1986). The first digit of the S.O.I. has six levels, in which we have joined the upper two and the lower two. The first digit of the Dutch S.O.I. is equal to the level categories of the International Standard Classification of Education (ISCED).
6. This implies that the association between father's and respondent's education is stronger in the early parts of the educational career (cf. Mare, 1981).
7. There is however a slight trend in parameter INH. Although the coefficient is insignificant, its sign is negative. Its direction is therefore similar to the trend observed in parameter U and, because of this, it is well covered in the trend model for this parameter (B.4).

8. These additional calculations are based on the difference between models B.4 and B.5 of Table 3, which evaluates the significance of the trend. For men the difference in likelihood is 33.8, with one degree of freedom. Because the critical value of the chi2 distribution with one degree of freedom is 3.84 ($p<.05$), and because the likelihood is linear with the number of cases analysed, we have 33.8/3.8=8.8 times more cases than necessary to find a significant trend. Dividing our sample size of men (11,892) by 8.8 we find 1351. For women the difference in likelihood between models B.4 and B.5 is only 14.1, implying that here we have 3.67 as many cases as necessary. Dividing the sample size of women (10,895) by 3.67 we find 2969 as the necessary sample size to evaluate the trend to be significant for women.

Appendix Educational mobility tables, according to birth cohort and sex; respondents in rows, fathers in columns

men:

1891-1895:

64	0	1	0
11	1	1	0
6	0	1	1
1	0	1	1

1896-1900:

126	1	1	0
31	1	2	0
8	2	4	1
3	1	1	2

1901-1905:

179	2	0	1
64	8	1	0
30	7	6	5
8	6	3	3

1906-1910:

205	6	1	0
119	11	2	1
34	8	7	3
16	8	4	4

1911-1915:

262	5	2	0
154	26	3	1
65	17	15	3
20	5	7	16

1916-1920:

282	9	4	0
206	34	10	3
83	22	19	6
27	9	20	15

1921-1925:

293	17	6	1
282	51	13	6
146	44	38	5
37	18	28	32

1926-1930:

258	17	4	1
302	49	13	4
178	54	45	7
68	44	14	42

1931-1935:

296	17	9	1
310	63	23	8
215	64	47	17
81	50	35	37

1936-1940:

211	15	10	1
334	98	25	10
234	89	57	18
80	71	32	45

1941-1945:

205	24	9	4
372	130	26	5
269	120	78	21
101	67	54	63

1946-1950:

159	17	16	1
334	143	33	11
294	155	114	25
96	84	76	68

1951-1955:

96	15	7	3
164	72	25	5
179	120	88	29
57	75	51	51

1956-1960:

26	8	3	1
80	46	22	7
91	78	48	24
35	40	34	43

women:

1891-1895:

71	6	2	0
9	3	0	0
1	0	1	1
1	0	0	1

1896-1900:

149	5	3	0
19	4	2	0
9	2	2	0
5	0	1	1

1901-1905:

255	10	1	4
41	17	3	3
12	8	6	12
2	3	0	5

1906-1910:

328	8	4	1
69	16	5	7
21	15	9	7
11	2	5	5

1911-1915:

354	22	8	5
88	21	7	5
30	9	18	7
8	3	6	7

1916-1920:

383	17	4	5
137	29	16	3
47	12	15	9
12	6	6	7

1921-1925:

439	27	8	0
202	50	22	7
58	30	26	14
9	2	6	15

1926-1930:

440	27	9	1
251	48	24	16
77	35	34	19
19	8	10	19

1931-1935:

328	29	13	0
278	80	23	16
97	34	39	16
13	11	16	27

1936-1940:

249	34	9	3
329	101	40	13
117	56	49	18
21	23	25	23

1941-1945:

208	27	4	3
453	131	39	15
135	80	68	32
32	33	29	33

1946-1950:

191	37	8	3
436	137	44	19
178	101	94	32
44	39	33	41

1951-1955:

97	24	8	1
168	80	38	7
155	112	72	38
39	52	47	46

1956-1960:

40	12	5	2
75	57	14	8
104	67	52	20
29	30	29	27

accumulated tables over all birth cohorts, according to sex:

men:

2662	153	73	14
2763	733	199	61
1832	780	567	165
630	478	360	422

women:

3532	285	86	28
2555	774	277	119
1041	561	485	225
245	212	213	257

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