

Bivariate and multivariate scaled association models. An application to homogamy of social origin and education in Hungary between 1930 and 1979

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Abstract. In this article we compare bivariate and multivariate models for homogamy of social origin and education to test whether bivariate models of homogamy lead to biased results. We use data on Hungarian couples married between 1930 and 1979 and loglinear models of scaled association. The results indicate some differences between bivariate and multivariate analyses. At each point of time bivariate models overestimate homogamy, both with respect to education and social origin. However, results on trends in time do not differ much between the two analyses. The exception is the period 1940–1959, in which bivariate analysis showed decreasing educational homogamy, and multivariate analysis showed an increasing trend. The latter finding can be explained by declining homogamy of social origin, as well as the weaker reproduction and cross-effects in this period.

1. Introduction

Research on marital selection has generally shown great resemblance between spouses. Most people marry persons with similar characteristics. This tendency towards homogamy applies to wide range of characteristics, both in physical and social traits (for a recent overview see Surra, 1990), but is particularly strong with respect to status characteristics like education and occupation. For stratification analysts, homogamy with respect to social status has long been an important object of study (cf. Sorokin, 1927; Berent, 1954). Status homogamy namely directly indicates the extent to which members of different social groups accept each other as equal in the social hierarchy. This paper continues the line of research on status homogamy by studying homogamy of social origin (father's and father-in-law's occupation) and educational homogamy (husband's and wife's education).

The point of departure for the analysis reported here is the finding that in most industrializing countries after World War II the association between

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spouses' educations has declined (Ultee & Luijkx, 1990). That is, in most of the world's marriage markets people increasingly wind up marrying a person of dissimilar educational level. Exceptions to this world wide pattern of decreasing educational homogamy are Belgium (Ultee & Luijkx, 1990) and the United States (Kalmijn, 1991a; Mare, 1991).

From a theoretical point of view, the finding of an overall declining trend in educational homogamy is rather unexpected. This is firstly so, since from modernization theory (Davis & Moore 1945; Kerr *et al.*, 1960) one would expect that in modernizing societies education has replaced origin status as the main asset in the distribution of societal rewards. For marital selection this would imply that people should pay more, and not less attention to the education of the potential spouse. Secondly, due to the increasing labor market participation of women and the educational convergence between men and women, marital selection processes are more symmetric than they used to be (cf. Oppenheimer, 1977). Marriages of poorly educated, wealthy men with highly educated, poor women are nowadays less attractive for both men and women. This decreased popularity of the traditional trade-off in the marriage market should increase the likelihood of marriages that are homogamous with respect to education. Thirdly, as Mare (1991) points out, due to increasing length of education the time interval between leaving school and entering wedlock has narrowed. If one holds that friendships from school last some years, this smaller time interval will increase the chances of educational homogamy.

In this paper we propose that the solution to the contradiction between the empirical finding of declining educational homogamy and the theoretical expectation of increasing educational homogamy does not lie in revision of marriage market theory,¹ but in revision of the methodology applied. We assume that the methods used to model homogamy in previous research may have been too simple. In particular, in the large scale internationally comparative analysis of Ultee & Luijkx (1990), educational homogamy is assessed by comparing husband's education with wife's education. The problem of this bivariate approach to modelling homogamy is that educational homogamy is (implicitly) assumed to be an outcome of people choosing on only one characteristic (their educational attainment). However, the selection of a spouse may be a more complex process. In the marriage market people do not match on a single trait, but consider multiple traits such as age, religion, race, ethnicity, and social origin (e.g. Kalmijn, 1991b).

This complex structure of the spousal selection process must have consequences for the models needed to assess homogamy. The degree of homogamy observed in bivariate models may be a spurious *by-product* (Blau & Duncan, 1967: 358) of people's tendency to choose each other on related

characteristics. In this paper we argue that these by-product effects occur with respect to both educational homogamy and homogamy of social origin. Educational homogamy may be a by-product in the following way: if two people of high social origin choose each other as spouses, this homogamous marriage is likely to also be a marriage between two highly educated people, since people of high social origin likely have high educational attainment. This educationally homogamous marriage may even have occurred without preferences to match on education. Conversely, homogamy of social origin may be a by-product of people's tendency to match on education (cf. Warren, 1966; Blau & Duncan, 1967).

Given the potential by-product effects, it can be expected that studies using bivariate models for homogamy yield biased results. First, the propensity to marry someone of similar educational level may be lower than the results of a bivariate analysis would indicate, as is the tendency to match on social origin. A second, more serious problem is that trends in homogamy can be distorted as well. In particular, the world-wide pattern of decreasing bivariate educational homogamy observed by Ultee & Luijkx (1990), may be a by-product of a decreased tendency to match on social origin. To test such by-product effects, *multivariate* models of homogamy are needed in which the associations between spouses' educations and spouses' social origins are simultaneously analyzed. This paper's primary aim then is to further explain and test the by-product effect for the trend in educational homogamy by comparing results from bivariate and multivariate analyses. In the same manner it will be tested to what extent the bivariate trend in homogamy of social origin is a by-product of people's tendency to match on education.

2. The by-product explanation and research questions

In order to clarify the way in which a match on social origin influences a match on education, let us assume a marriage market in which besides partner's education, the occupational statuses of the father and the father-in-law are important in the decision whom to marry. In this particular situation, spouses' similarity with respect to education is a result of four relationships: (a) a comparison of one's own education with that of the education of the potential mate, (b) a comparison of one's own education with the social origin of the potential mate (i.e., father-in-law's occupation), (c) a comparison of one's own social origin (i.e., father's occupation) with the social origin of the potential mate (i.e., father-in-law's occupation), and (d) the relation

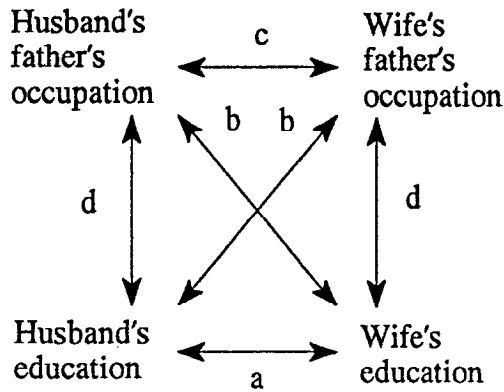


Fig. 1. Multivariate model for educational homogamy (arrow a), homogamy of social origin (arrow c), intergenerational reproduction (arrows d) and cross-effects (arrows b).

between father's occupation and one's own educational attainment. These relations are shown in Figure 1.

Relations (a) and (c) in Figure 1 are straightforward. They respectively indicate matching on education and matching on social origin. Relations (b) and (d) are of a more complex nature. Relations (b) in Figure 1 link husband's social origin to wife's education and wife's social origin to husband's education. They represent the (partial) association between a person's education and his or her father-in-law's occupation, over and above the association expected on the basis of educational homogamy and homogamy of social origin only. Substantively, these relations make sense because a match on the marriage market does not only involve a comparison of one's own trait with the same trait for another person, but also involves a process of exchange between certain characteristics. One could for example assume that a marriage between a man of high social origin and a woman with a high level of education forms a more attractive combination on the marriage market than a marriage between a man with a high level of education and a woman of high social origin (cf. Murstein, 1986). This stems from traditional sex-specific roles within marriage, in which the man earns the money and the woman raises the offspring. The first role requires a good occupation or high social origin, whereas for the last role cognitive skills are more important. In the sequel, such processes of exchange between traits will be referred to as 'cross-effects'.

Relations (d) in Figure 1 represent the association between one's social origin and one's educational attainment for both husbands and wives. Most likely, these association will be positive, as social status is inherited through the educational system. We may speak here of 'intergenerational reproduc-

tion'. In the status attainment literature this intergenerational reproduction is represented by a causal relationship, a causal influence of father's occupation on child's educational attainment. However, in marriage markets this causality assumption seems odd and the relationship needs no causal interpretation. The matching processes in these markets do not operate in a life-course perspective, but occur at single points in time. Kalmijn (1991a, p. 510) states this as follows: "In the process of marriage selection, individuals demonstrate a set of attributes to others and evaluate a similar set of attributes in their potential spouses. While social origins and education are causally related, it is doubtful whether the mechanisms of demonstration and evaluation have an underlying causal order. People look at each other's origins, given their current destinations, and they look at their destinations, given their origins".

If we return to Figure 1 and assume that the intergenerational reproduction (d) and cross-effects (b) both have positive parameters, it follows that a positive association between spouses' social origins (c) produces a positive association between spouses' educations. In other words: spouses are similar with respect to educational level, because they have matched homogeneously on social origin and because origin and education relate positively. This educational similarity may even arise when the tendency to marry someone of similar educational level (relation a in Figure 1) is absent. In that particular instance, bivariate educational homogeneity is completely due to other processes involved in marital selection.

When different points of time are compared (as we will do in the sequel by the analysis of cohorts), multivariate analysis of homogeneity of social origin and education may show that distinct and even opposing trends underlie changes in the observed homogeneity pattern. In particular, multivariate analysis may show that a trend towards less bivariate educational homogeneity conceals a stable or even upward trend in (partial) educational homogeneity. This would in particular be the case if intergenerational reproduction (d) and homogeneity of social origin (b) have positive parameters and these parameters decline over time. If these assumptions hold, one may – all other things being equal – expect a decline in bivariate educational homogeneity. Net of these processes, the tendency to match homogeneously on education (a), may however have remained stable. It may even be so, that the tendency to match on education shows an increase, but not so strong as to undo the expected decline in educational homogeneity due to the other variables at work. If the latter situation occurs, our initial contradiction between empirically found declining educational homogeneity and theoretically expected increasing educational homogeneity is resolved: the observed bivariate decline hides a multivariate increase in educational homogeneity.

To test these by-product explanations, we need a multivariate model for homogamy of social and education. In this article we develop such a multivariate model for cohorts in Hungary married between 1930 and 1979. Hungary was chosen because it has three highly comparable, large-scale household surveys with the relevant data (spouses' social origins and educations). By combining these data we are able to use a database of considerable sample size (total $N = 21,164$ couples). For a loglinear analysis such a large database is required because the cross-classification of spouse's social origins and educations produces many cells. Substantively, Hungary is of interest because it underwent strong economic and political developments in the period of investigation (cf. Ferge, 1979; Kolosi & Róbert, 1985). One may expect that the relations in the multivariate model have changed to a fairly large extent. This in particular makes a test of the by-product thesis for trends relevant.

In sum, we address the following questions:

- (1) *To what extent did the association between spouses' educational levels in Hungary change between 1930 and 1979?*
- (2) *To what extent did the association between spouses' social origins in Hungary change between 1930 and 1979?*
- (3) *To what extent do the bivariate trends (in questions 1 and 2) differ from trends in homogamy of social origin and education obtained with multivariate models?*

By answering these questions we improve upon earlier research on marital selection. Although in some previous studies both bivariate (e.g. Ultee & Luijkx, 1990) and multivariate models for homogamy were estimated (e.g. Kalmijn, 1991a; Hendrickx, 1994; Tsai, 1994), no comparison of both analyses is currently available, nor was it ever assessed to what extent trends observed with bivariate models conceal trends in educational homogamy.

3. Data

The data sets that we use in this paper are:

- (1) 'Social Mobility and Occupational Changes in Hungary', conducted in 1973. This survey originates from a Microcensus carried out in 1973 by the Central Statistical Office in Budapest under the direction of Rudolf Andorka (Andorka, 1973). For this Microcensus a stratified sample (by districts) of 0.5 percent of the total population over 14 years of age was taken. The questions in the 1973 survey were addressed to 40,426 mem-

- bers of different households, among them 7,661 couples who reported on their education, their father's occupation and their year of marriage.
- (2) 'A Model of Stratification Survey', conducted in 1981–1982 by the Institute for Social Sciences in Budapest and the Hungarian Central Statistical Office under the general direction of Tamas Kolosi (Kolosi, 1982). Its questionnaires were administered independently to all adult (18 years and older) members of approximately 9,000 households, among them 4,781 households containing two surviving marriage partners and reporting on the educational level and occupational status of the father and their year of marriage.
 - (3) 'Hungarian Social Mobility and Life History Survey', conducted by Kulcsar & Harcsa (1983). This is a household survey which addresses questions to each member of a household. In total 32,301 household members were interviewed. The file contains 8,722 couples that were married and reported on their educational levels and fathers' occupations and their year of marriage.

By pooling the three surveys we have a total sample size of 21,164 couples available. Trends in homogamy are assessed by assigning the couples from the pooled file to five cohorts married between 1930 and 1979 (1930–1939; 1940–1949; 1950–1959; 1960–1969; 1970–1979). Comparison of these cohorts will yield an estimate as to what extent homogamy has decreased, increased or remained stable over time.²

Father's occupation refers to the situation when the spouses were between 14 and 18 years old. The occupational categories used are: (1) farmers, (2) lower manuals, (3) higher manuals, (4) lower non-manuals, and (5) higher non-manuals. This classification covers the distinction between white-collar and blue-collar workers, makes further distinction in skill-levels and employs a farm category. Earlier research has proven these to be important distinctions in terms of social mobility chances (Ganzeboom *et al.*, 1989).

The educational levels of both partners were recoded into: (1) 0–5 classes of elementary school, (2) 6–7 classes of elementary school, (3) 8 classes of elementary school, (4) secondary education (academic, vocational and technical), and (5) post-secondary education (including college and university). Compared to conventional classifications of education this classification differs by distinguishing three categories of elementary school. This is relevant for the Hungarian case because selection of pupils into the labor market or secondary schools occurred at these different levels of primary school (Róbert, 1991). In addition, in our data there is a large group of persons with elementary education only. This makes a sub-distinction necessary.

Unfortunately, the data do not allow a distinction between first and second or later marriages. This can be a problem because people who remarry are – due to a more limited pool of spouses to choose from – less educationally homogamous in their latest marriage than people who marry once. Jacobs & Furstenberg (1986) observed this for two age cohorts of married American women. However, their findings also indicated stability of this remarriage effect over time. If one also takes into account that remarriage is fairly stable in the 1930–1979 period,³ the remarriage effect cannot confound the observed trends in homogamy in a substantial way.

4. Models

Loglinear models are applied to assess (trends in) homogamy. Recent studies of homogamy have applied these models with almost no exception (e.g. Hout, 1982; Jones, 1987; Ultee & Luijkx, 1990; Kalmijn, 1991a; Hendrickx, 1994). The advantages of this technique over other covariance models (cf. Warren, 1966; Blau & Duncan, 1967) are that loglinear models allow one to disentangle effects of marginal distributions (i.e. structural homogamy) and relative chances of marital association (i.e. relative homogamy). This is important when structural opportunities for marriages and preferences of partners would result in contradictory trends in homogamy. The second advantage of loglinear modelling is that one can define multiple parameters to represent the relation between two (or more) variables. In the analysis of intergenerational occupational mobility chances this is often done by specifying separate diagonal parameters for ‘excessive inheritance’ (e.g. for farmers), and the same can be applied to tables cross-classifying spouses’ characteristics (cf. Hout, 1982).

To model homogamy, several loglinear specifications are available. Here, we choose for a model of (scaled) association. More specifically, we use the log-multiplicative *Quasi Row and Column Effects Model II (RCII)* as proposed by Goodman (1979). This RCII model has three attractive properties. First, the RCII model yields a measure for the social distances between groups. Substantively, this is of interest since people’s preferences on the marriage market may be regarded as a function of these distances (cf. Bogardus, 1925a, 1925b). Second, the RCII model does not assume uniform, but scaled (uniform) association. For a marriage market the latter assumption seems to be appropriate, since some groups are closer to each other on the intermarriage dimensions than others. Third, the RCII model uses only one association parameter for homogamy outside the diagonal cells of a marriage table. This makes it easier to assess bias in the association parameter and

consequently facilitates the estimation of by-products over other loglinear models.⁴

For the multivariate table cross-classifying spouses' social origins by spouses' educations by time, the 'bivariate' RCII model that assumes educational homogamy only, can be specified as:

$$\ln(F_{ijklc}) = HO_{ic} + WO_{jc} + HE_{kc} + WE_{lc} + UE_c U_k U_l + DE_k, \tag{1}$$

where F_{ijklc} are the expected frequencies for the multivariate table (subscript i refers to husband's social origin, j to wife's social origin, k to husband's education, l to wife's education, and c to cohort). HO_{ic} , WO_{jc} , HE_{kc} , and WE_{lc} are the main effects of respectively husband's and wife's social origins and husband's and wife's education per cohort. UE_c is the association parameter for educational homogamy per cohort, and U_k and U_l are the scaled categories of respectively husband's and wife' education with (identifying) constraints $\sum_k U_k = \sum_l U_l = 0$, and $\sum_k U_k^2 = \sum_l U_l^2 = 1$. Finally, DE_k are diagonal parameters for each educational category.⁵

Ignoring the effects of cohort and social origin, the association parameter for educational homogamy is related to expected odds ratios in the following way (cf. Ganzeboom *et al.*, 1989):

$$\ln \frac{F_{kl} * F_{k'l'}}{F_{k'l} * F_{kl'}} = UE * (U_k - U_{k'}) * (U_l - U_{l'}), \tag{2}$$

where notations from Equation (1) apply and where k and k' are adjacent categories, likewise for l and l' . This means that the association parameter (UE) is equivalent to the log odds-ratio of the expected frequencies, but scaled by the distance between category scores. Hence, it is a model of scaled association. Note that if the intervals between adjacent educational categories are unity, that is $(U_k - U_{k'}) = (U_l - U_{l'}) = 1$ for all j and l , the scaled association model of Equation (2) becomes the model of uniform association in which all adjacent odds-ratios have identical values.

The 'multivariate' RCII model that specifies next to educational homogamy also homogamy of social origin and interactions between origins and educations, is the following model:

$$\begin{aligned} \ln(F_{ijklc}) = & HO_{ic} + WO_{jc} + HE_{kc} + WE_{lc} + UE'_c U_k U_l + DE_k \\ & + UO_m U_i U_j + DO_i + HOHE_{ikc} + WOVE_{jlc} + HOWE_{ilc} \tag{3} \\ & + WOHE_{jkc} \end{aligned}$$

where restrictions and notations from Equation (1) apply, and UO and DO

refer to respectively the association parameter for homogamy of social origin and its diagonal parameters. In this multivariate model, UE' refers to the association between spouses' educations net of the other relations in the model. Comparison of this association parameter with the bivariate parameter UE from Equation (1) yields an estimate of the by-product effect on educational homogamy. Similarly, we may compare the bivariate parameter for homogamy of social origin with its multivariate counterpart to obtain the by-product effect for this kind of homogamy. For reasons of sparsity, we do not write down these equations here.

In the analysis additional constraints are imposed upon the models of Equations (1) and (3) to make them better interpretable. One obvious restriction is to restrict U_k and U_l to be equal. This implies that the educational categories of husband and wife are equally scaled. In that case, the categories define an 'intermarriage dimension' along which categories are ordered according to their propensity to intermarry (relative to their absolute sizes). The same can be applied to the categories of social origin, U_i and U_j . Another restriction is to model the reproduction association ($HOHE_{ik}$ and $WOWE_{jl}$) and cross-effects ($HOWE_{il}$, and $WOHE_{jk}$) as uniform associations, given the category scalings found for the educational and social origin categories. This makes it easier to assess the sign of these effects and the consequences for both educational homogamy and homogamy of social origin.

For practical reasons, the RCII models are estimated using a two-step approach (see for this procedure also Ganzeboom *et al.*, 1989). First, scale values for the categories of spouses' educations and spouses' social origins are found, using the program AssocPc (Luijkx, 1988).⁶ Then, in the second step, the scale values are applied as fixed scores in loglinear models with the GLIM program (McCullagh & Nelder, 1983). To select a proper model, the *bic* statistic (Bayesian Information Criterion; Raftery, 1986) is used. This measure is preferred over the Log-Likelihood Ratio (L^2), because L^2 has the disadvantage that any small discrepancy between observed and expected frequencies turns out to be significant when the sample size is large. The *bic* measure adjusts for sample size as follows: $bic = L^2 - df * \ln(N)$, where *df* are the degrees of freedom and N is the sample size. If the *bic* measure is negative, the alternative model is more likely than the saturated model. The appropriate selection criterion then, is to look for a model with the most negative *bic*.

5. Results

In this section we describe results of bivariate and multivariate analyses of homogamy of social origin and education. First, we shed light on the strength

Table 1. Loglinear models for the association between spouses' educations and spouses' social origins for five marriage cohorts

	df	L ²	bic
1. <i>Marginals</i>	3040	22886	-7392
2. <i>Educational homogamy</i>			
a. 1 + one diagonal parameter	3039	17706	-12562
b. 1 + category-specific diagonal parameters	3035	16847	-13382
c. 2b + uniform association	3034	13367	-16852
d. 2b + scaled uniform association	3032	13131	-17068
3. <i>Homogamy of social origin</i>			
a. 2d + one diagonal parameter	3031	10130	-20059
b. 2d + category-specific diagonal parameters	3027	9641	-20508
c. 3b + uniform association	3026	9016	-21123
d. 3b + scaled uniform association	3024	8989	-21130
4. <i>Bivariate trend in educational homogamy</i>			
a. 3d + fluctuation in association	3020	8923	-21156
b. 3d + linear trend in association	3023	8958	-21151
c. 3d + curvilinear term in association	3022	8926	-21173
5. <i>Bivariate trend in homogamy of social origin</i>			
a. 4a + fluctuation in association	3016	8879	-21160
b. 4a + linear trend in association	3019	8892	-21177
c. 5b + curvilinear term in association	3018	8880	-21179
6. <i>Multivariate trend models</i>			
a. 5a + uniform reproduction- and cross-effects	3012	3391	-26608
b. 6a + fluctuation in reproduction- and cross-effects	2996	3331	-26509

of both kinds of homogamy and on the selection of a well fitting, parsimonious model. Then, we discuss (bivariate and multivariate) trend models and assess the by-product effects on homogamy.

To begin, Table 1 presents the loglinear models for the multiway table. It starts off with a model of null association, the 'marginals' model (Model 1). This model allows the educations and social origins of spouses to vary between cohorts, but assumes no further interactions between spouses' educations and social origins. The bic statistic of the marginals model shows a negative value (bic = - 7392) which implies that the no association model is to be preferred over the saturated model. This may lead one to accept the hypothesis of null association. However, comparing it to the other (association) models in Table 1, the marginals model fits badly. This indicates that spouses' educations and spouses' social origins are both associated, as could be expected.

Models 2a to 2d in Table 1 put restrictions on the association between spouses' educations. The first of these models - Model 2a - uses a single parameter for the diagonal cells. The diagonal cells are the cells in which

spouses have equal educational levels. The bic measure of this diagonal model shows a much more negative value ($\text{bic} = -12562$) than the marginals model, which indicates that the diagonal cells are disproportionately under- or overrepresented. Conform results from earlier research (cf. Ultee & Luijckx, 1990; Kalmijn, 1991a; Tsai, 1994), the diagonal parameter is positive (1.13; not shown in a table). This means that people have a general tendency to marry someone of identical educational level. The next model of Table 1 – Model 2b – specifies this inmarriage tendency to be dependent on the educational level by allowing category-specific diagonal parameters. Model 2b reduces the L^2 substantially at the cost of 5 degrees of freedom and is to be preferred to Model 2a according to the bic criterion ($\text{bic} = -13382$). Apparently, the tendency to marry within one's own educational group varies between educational groups. The diagonal parameters (from low to high respectively 1.96, 1.05, 0.69, 0.96, and 2.83) show that inmarriage is highest at the extremes of the educational hierarchy, and demonstrate that no such inmarriage tendency exists in the middle educational ranks.

Model 2c of Table 1 adds to the previous model (Model 2b) a uniform association parameter for the relation between spouses' educations outside the diagonal. The bic value of this uniform association model again shows a larger negative value compared to the previous models ($\text{bic} = -16852$). Since the uniform association parameter is positive (0.76), one may conclude that spouses do not only have a tendency for inmarriage, but also tend to associate with people near in educational status, and avoid relations at a large social distance. The next model, Model 2d, tests whether this association is dependent on the intervals between educational groups. This model of scaled association performs indeed better than the model of uniform association, Model 2c. With two more degrees of freedom for the scaling procedure, the scaled uniform association model has a bic of -17068 . The association in this model has a value of 7.39. The scale values for the five educational groups – which proved to be equal for husbands and wives – are respectively -0.46 , -0.45 , -0.08 , 0.28 , and 0.71 . These results suggest that on average people tend to associate with (educationally) likes more than with dislikes, but some educational groups (e.g. people with almost no education and people with some education) are closer to one another on the 'intermarriage' dimension than others (e.g. the highest educated versus the one-but highest educated).

To assess whether next to educational homogamy people prefer marrying someone of equal social origin, we add in Models 3a to 3d of Table 1 associations between spouses' social origins. These models yield approximately the same results as in the analyses of educational homogamy. Firstly, they indicate a tendency for inmarriage, because the diagonal model (Model

3a) improves upon our best fitting model so far (Model 2d). Secondly, the inmarriage tendency depends on specific categories of social origin. For the five groups the diagonal parameters are (from low to high) 1.38, 0.28, 0.60, 0.87 and 1.98. This again shows excessive inmarriage at the extremes of the social hierarchy, notably the farmers and higher non-manuals. Thirdly, next to the inmarriage tendency, people prefer to associate with persons near in origin status (cf. Model 3c). Fourthly, the association depends on the intervals between classes of origin because the model of scaled uniform association (Model 3d) fits better than the unscaled model. Furthermore, the scaling of the classes of origin (respectively -0.45 , -0.48 , -0.05 , 0.28 , and 0.70) highly correspond to the scaling of the educational categories. Although the association between spouses' social origins (2.34) is lower than the association between spouses' educations (7.39), it again indicates a tendency to marry likes and to avoid relations with dislikes.

Bivariate trends

Having found acceptable and parsimonious baseline models for homogamy of education and social origin, the central questions of this paper can now be addressed. The first two of these questions pertain to trends in respectively bivariate educational homogamy and bivariate homogamy of social origin. To what extent have these associations changed in Hungary between 1930 and 1979? To answer this question, Models 4a to 4c in Table 1 describe trends in educational homogamy, and Models 5a to 5c describe trends in homogamy of social origin. In the first of these models, Model 4a, educational homogamy is assumed to fluctuate between cohorts in a discrete way. That is, we add to Model 3d cohort-specific association parameters. According to the bic criterion this fluctuation model improves upon the static models addressed before (bic = -21156). Apparently, educational homogamy has not been stable in the 1930–1979 period in Hungary.

In panel A of Table 2 the exact values for the association between spouses' educations are shown by cohort, and in Figure 2 the corresponding trend line is drawn graphically. From these data one can observe that bivariate educational homogamy declined in the period 1930–1959, but increased thereafter from 1960 to 1979. The latter increase is stronger than the initial decline, which makes the association end up higher than it started off. Overall then, the trend in educational homogamy seems to be U-shaped.

This finding of a U-shaped trend in educational homogamy is substantiated by Models 4b and 4c in Table 1. Model 4b assumes a linear trend in educational homogamy. Not surprisingly, the fit of the linear trend model (bic = -21151) is worse than the fluctuation model. Note here that the linear trend

Table 2. Parameters of bivariate and multivariate trend models for educational homogamy and homogamy of social origin; selected models of Table 1

	Cohort				
	30-39	40-49	50-59	60-69	70-79
A. Bivariate trends (model 5a)					
1. Educational homogamy	7.85	6.77	6.63	8.08	9.51
2. Homogamy of social origin	3.85	3.15	2.37	1.99	2.13
B. Multivariate trends (model 6b)					
1. Educational homogamy	6.39	5.63	5.98	7.36	8.53
2. Homogamy of social origin	2.31	1.91	1.54	1.28	1.52
3. Reproduction men	0.30	0.22	0.25	0.27	0.32
4. Reproduction women	0.33	0.32	0.26	0.36	0.35
5. Cross-effect men	0.14	0.16	0.11	0.16	0.16
6. Cross-effect women	0.15	0.13	0.08	0.07	0.10

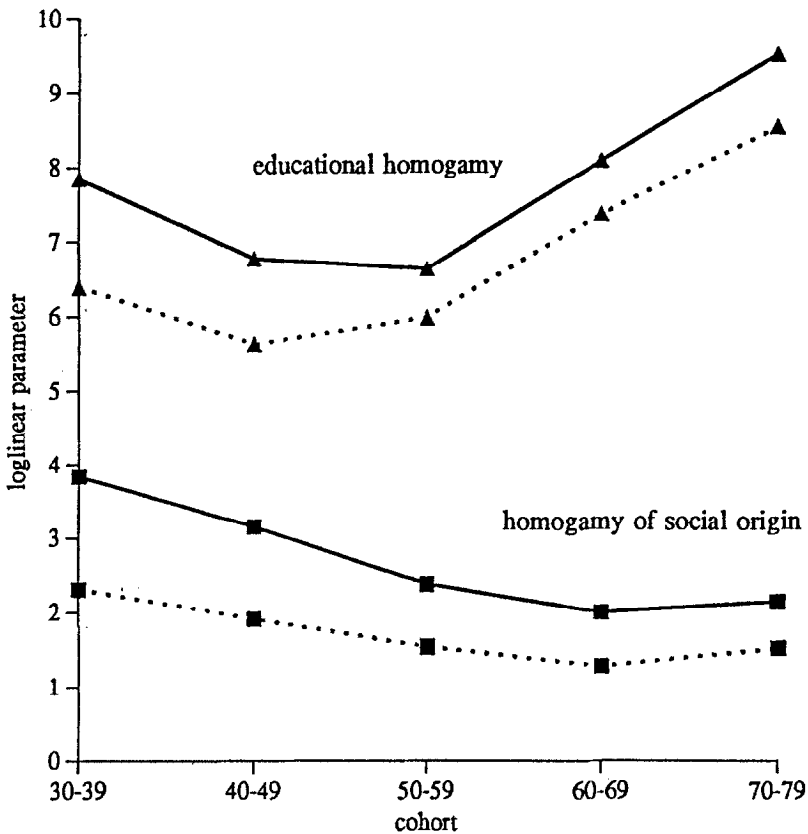


Fig. 2. Trends in bivariate (straight lines) and multivariate (dotted lines) homogamy of social origin and education; Hungary 1930-1979.

parameter is positive (0.61), which is in line with the finding that in our data educational homogamy is stronger for the younger than the older married. In Model 4c a quadratic term is added to the linear trend model. It states a curvilinear slope in educational homogamy. As may have been expected, this model fits the data better ($bic = -21173$) than the previous models. Hence, our conclusions from visual inspection stand firm. Educational homogamy shows a U-curved trend: it was strong in the beginning (the thirties), weaker during the war-period and the fifties, but rapidly increasing thereafter.

Our next question is whether homogamy of social origin declined in Hungary during the period of investigation (1930–1979). Models 5a to 5c in Table 1 provide an answer to this question. In the fluctuation model – Model 5a – Model 4a is taken as a baseline and cohort-specific association parameters for the relation between spouses' social origins are added. The model fit improves ($bic = -21160$), which shows that homogamy of social origin also fluctuates significantly between cohorts. In panel A of Table 2 and in Figure 2 we show the corresponding scaled association parameters per cohort. It can be seen that homogamy of social origin steadily declined from 1930 to 1969. However, in the last cohort (1970–1979) this decrease levels off and the association even shows a slight increase. Nonetheless, the overall decline in the association is large: in the last cohort (1970–1979) the parameter for homogamy of social origin (2.13) is almost half the size of the parameter for the first cohort (1930–1939). Models 5b and 5c specify this trend as respectively a linear and curvilinear trend. Not surprisingly, the bic statistics show that not a linear trend specification ($bic = -21177$), but a curvilinear trend specification fits best ($bic = -21179$). These statistics are however very close to one another, and do not alter the conclusion of strongly decreasing homogamy of social origin.

Multivariate trends

From the previous results one may conclude a shift in Hungarian marriage patterns from selection on the basis of social origin to selection on the basis of educational level. However, as was pointed out in the first two paragraphs of this paper, bivariate analyses of homogamy may yield trends that are distorted by other processes involved in marital selection. To assess whether this is true for our data, multivariate models for homogamy are applied. In these multivariate models we do not only assume that people match on origins and destinations (e.g. as in Model 5a), but also take into account the links between origin and destination (e.g. as in Equation 3). Model 6a and 6b of Table 1 contain these additional associations. In these models,

reproduction and cross-effects are added to the previous models. We take Model 5a, in which both educational homogamy and homogamy of social origin fluctuate between cohorts, as our baseline. The reason for doing so is that this allows us to find by-product effects on homogamy for cohorts separately.

In the first multivariate model of Table 1 – Model 6a – invariant reproduction and cross-effects are modelled. These effects – that have the form of scaled uniform associations (see also the models section) – improve the model fit greatly according to the bic criterion ($\text{bic} = -26608$). Hence, reproduction and cross-effects must be significant. Inspection of the corresponding parameters (not shown here) shows these effects to be positive. This implies that the by-product effect may indeed be at work. Given the positive associations between origins and destinations on the marriage market, homogamy on one dimension may cause homogamy on another dimension.

To see what consequences these reproduction and cross-effects have for trends in homogamy of social origin and education, the second multivariate model of Table 1 – Model 6b – allows these reproduction and cross-effects to vary between cohorts. The corresponding multivariate trend model fits worse ($\text{bic} = -24509$) than the previous one. This shows us that at least some, or maybe even all of the reproduction and cross-effects are invariant. Table 2 informs more specifically which of these effects changed and which not. The parameters in this table demonstrate that the intergenerational reproduction for both men and women and the cross-effect for women changed in a curvilinear fashion, while the cross-effects for men do not show a trend. Modelling these changes proves the U-shaped trends in reproduction to be significant, both for men ($\text{bic} = -26608$) and women ($\text{bic} = -26613$). The cross-effects for men do not show a particular trend, however, and are assumed to remain constant over time.

From Table 2 and Figure 2 one may conclude that the changes in the multivariate pattern of homogamy – including the reproduction and cross-effects – leave the trends in educational homogamy and homogamy of social origin almost unaffected. Although at each cohort bivariate homogamy is substantially lower than its multivariate counterpart, the trend slopes only slightly differ. This means that by-product effects on homogamy do occur, but these effects are more or less stable in time. Hence, multivariate homogamy analysis does not make for other trends in homogamy than bivariate homogamy analysis. This conclusion holds both for homogamy social origin and educational homogamy.

However, a more detailed cohort-specific inspection of bivariate and multivariate homogamy reveals some interesting differences between the two types of analysis. The largest of these differences occurs with respect to the change

in educational homogamy between the second (1940–1949) and third cohort (1950–1959). In the bivariate case, educational homogamy slightly *decreases* (from 6.77 to 6.63), whereas in the multivariate case it *increases* (from 5.63 to 5.98). Bivariate analysis distorts the ‘real’ trend here. The reason for this bias in the trend can be understood from the parameters confounding bivariate educational homogamy, notably homogamy of social origin, intergenerational reproduction and the cross-effects. As may be observed from Table 2, these ‘distorting’ factors decrease in size. Since we know from Model 6a that origins and education associate positively, such a decrease leads us – all other things being equal – to expect a decrease in educational homogamy. Given the relatively stable trend in bivariate educational homogamy during this period, the ‘residual’ homogamy (arrow d in Figure 1) must have increased to compensate the expected decrease in homogamy. Hence, the discrepancy between bivariate and multivariate homogamy analyses arises.

6. Conclusions

The analyses of marital patterns in Hungary between 1930 and 1979 in this paper have shown the following results. First, it was demonstrated that in Hungary between 1930 and 1979 homogamy of social origin decreased considerably. That is, those who marry pay less and less attention to the social origin of the prospective spouse, and increasingly prefer marrying someone of dissimilar origin status. This result is in line with expectations from modernization theory that hold that just like in labor markets, people’s choices in marriage markets shift from ‘ascriptive’ to ‘achievement’ values. The decline in homogamy of social origin may also be explained by the decreasing possibilities of parents to interfere in marital decisions of their offspring. Due to this development, prospective partners feel less pressure to marry within the social class they stem from. The observed decline in homogamy of social origin is not uniform throughout the 1930–1979, however. In the last marriage cohort studied (1970–1979) the decrease in homogamy of social origin has leveled off, and homogamy showed an increase. This was not expected. It may however be explained by the Hungarian experience with socialism. During the 1970s, socialism was at its weakest in Hungary, and private property gained in importance. Since property was still distributed unequally along lines of social origin, this may have increased the tendency to match on origin status.

Similarly to homogamy of social origin, educational homogamy showed a U-shaped trend. Until the 1960s educational homogamy decreased, while after this period educational homogamy increased. Contrary to the trend in

homogamy of social origin, the trend in educational homogamy did not end up lower, but considerably higher than it started off. Hence, taken over the full 1930–1979 period, educational homogamy became stronger. This result again confirms expectations from modernization theory, according to which achievement values – choosing someone of high educational level – become increasingly important in selecting an attractive partner. However, because the increase in educational homogamy was not linear, other factors must have influenced it. One of these factors may be the socialist regime. In the fifties for example, when educational homogamy was weakest, Hungarian socialism was in its most extreme form. During this period, quota recruitment was successfully maintained in educational selection. People from low social origin were given priority in the selection for higher education. This created opportunities for interaction between people that will finally achieve different educational levels. If friendships from schools hold some time, the quota system may consequently have led to a decrease in educational homogamy.

However, the chief purpose of this paper was not to replicate existing research on homogamy, but to find out whether trends in homogamy obtained with bivariate models distort ‘real’ trends in homogamy. The results did not indicate any major distortions of this kind. Although at each point of time bivariate models overestimate homogamy, both with respect to homogamy of social origin and education, multivariate analyses did not lead to substantially other trends than separate analyses. A relatively minor exception to these results is the 1940–1959 period. Between the two marriage cohorts for this period, bivariate educational homogamy decreased, but multivariate educational homogamy increased.

The reason why the multivariate approach to modelling marital homogamy only worked for the 1940–1959 period, lies in the fact that during this period the factors influencing educational homogamy (i.e. homogamy of social origin, intergenerational reproduction and cross-matching) changed to a considerable extent and also in the same direction, namely less association. For the other periods no such uniform and strong changes were observed. Consequently, these factors do not distort trends in homogamy to a large extent.

Notwithstanding this modest result, it would be wrong to conclude that multivariate homogamy models are of little use. In fact, we have specified the *conditions* under which multivariate trend analysis leads to different findings than bivariate trend analysis. In our data these conditions only applied to one period, and only to educational homogamy. However, they may very well apply to other countries or periods in time, or to other strongly related characteristics of spouses.

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Notes

1. One such revision could be Goode's version of modernization theory (Goode, 1964). He suggests that while in working life achievement values replaces ascriptive values, in other parts of life – including the choice of a partner – people have gained more freedom. Although this explanation is appealing, it does not lead to new predictions and does therefore not suffice as an alternative explanation.
2. We have tested whether between the three surveys cohorts change with respect to homogamy of social origin and education. This turned out not to be the case. Therefore our data do not suffer from the 'duration of marriage' problem (cf. Kalmijn, 1991a). This duration problem may cause bias in homogamy estimates when homogamous marriages are more stable and have lower chances of divorce than heterogamous marriages (Bumpass and Sweet, 1972).
3. In 1930 the percentage of couples in which one or both of the spouses remarried was 18.4%, while in 1980 it was 29.3% (Source: Demographic Yearbooks, 1920–1990, Central Statistical Office, Budapest).
4. We also tested other loglinear models such as the crossings parameter model (Hout, 1982). These models do not fit better than the RCII models.
5. Note that we do not assume interactions of cohort with the diagonal parameters (DE) and with the scale parameters (U_k and U_l). In the analysis these interactions proved to be non-significant. This is a nice outcome, because it makes it easier to compare the intrinsic association and diagonal parameters between tables.
6. To estimate the Equal Row and Column Effects Model II, Ruud Luijkx rewrote the program ANOASC (Shockey & Clogg, 1983). This adaption is known under the name AssocPc. The program uses an iterative proportional fitting algorithm to provide estimates of the parameters.

References

- Andorka, Rudolf (1973). *Social Mobility and Occupational Changes in Hungary, 1973* [MRDF], Budapest: Central Statistical Office.
- Berent, Jerzy (1954). Social mobility and marriage: a study of trends in England and Wales,

- pp. 321–338, in: D. V. Glass (ed.), *Social mobility in Britain*, London: Routledge/Kegan Paul.
- Bogardus, Emory S. (1925a). Social distance and its origins, *Journal of Applied Sociology* 9: 216–226.
- Bogardus, Emory S. (1925b). Measuring social distance, *Journal of Applied Sociology* 9: 299–308.
- Blau, Peter M. and Duncan Otis D. (1967). *The American Occupational Structure*, New York: John Wiley.
- Bumpass, Larry L. & Sweet, James A. (1972). Differentials in marital instability: 1970, *American Sociological Review* 37: 754–766.
- Burgess, Ernest W. & Wallin, Paul (1943). Homogamy in social characteristics, *American Journal of Sociology* 44: 109–124.
- Davis, Kingsley and Moore, Wilbert E. (1945). Some principles of stratification, *American Sociological Review* 10: 242–249.
- Ferge, Zsuzsa (1979). *A society in the making. Hungarian social and societal policies, 1945–1975*, Harmondsworth: Penguin.
- Ganzeboom, Harry, B. G., Luijckx, Ruud & Treiman, Donald (1989). Intergenerational class mobility in comparative perspective, *Research in Social Stratification and Mobility* 8: 3–79.
- Goode, William (1964). *The Family*, Englewood Cliffs: Prentice-Hall.
- Goodman, Leo A. (1979). Simple models for the analysis of association in cross-classifications having ordered categories, *Journal of the American Statistical Association* 74: 537–552.
- Hendrickx, John (1994). *The Analysis of Religious Assortative Marriage*, Ph.D thesis, Nijmegen: Nijmegen University.
- Hout, Michael (1982). The association between husbands' and wives' occupations in two-earner families, *American Journal of Sociology* 88: 397–409.
- Jacobs, Jerry J. & Furstenberg, Frank F., Jr. (1986). Changing places: conjugal careers and women's marital mobility, *Social Forces* 64: 714–732.
- Jones, Frank L. (1987). Marriage patterns and the stratification system: trends in educational homogamy since the 1930s, *Australian and New Zealand Journal of Sociology* 23: 185–198.
- Kalmijn, Matthijs (1991a). Status homogamy in the United States, *American Journal of Sociology* 97: 496–523.
- Kalmijn, Matthijs (1991b). Shifting boundaries: trends in religious and educational homogamy, *American Sociological Review* 56: 786–800.
- Kerr, Clark, Dunlop, John T., Harbison, Frederick H. & Harmondsworth, Charles (1960). *Industrialism and Industrial Man*, Cambridge, Massachusetts: Harvard University Press.
- Kolosi, Tamas (1982). *A Stratification Model Study, Hungary 1981–1982* [MRDF], Budapest: Social Research Informatics Institute (TARKI).
- Kolosi, Tamas & Róbert, Péter (1985). Die Rolle des Abends- und Fernstudiums bei der Gesellschaftlichen Mobilität, in: M. Kaiser, R. Nuthman & H. Stegman (eds.), *Berufliche Verbleibsforschung in der Diskussion*, Nürnberg: Institut für Arbeitsmarkt und Bildungsforschung.
- Kulcsar, Rosza & Istvan Harcsa (1983). *Hungarian Social Mobility and Life History Survey, 1983* [MRDF], Budapest: Central Statistical Office.
- Luijckx, Ruud (1988). *AssocPC. A Program to Estimate Goodman's Scaled Association Models*, Tilburg: Department of Sociology.
- Mare, Robert D. (1991). Five decades of educational assortative mating, *American Sociological Review* 56: 15–32.
- Marvin, Donald M. (1918). Occupational propinquity as a factor in marriage selection; *Quarterly Publications of the American Statistical Association* 16: 131–150.
- McCullagh, P. & Nelder, J. (1983). *Generalised Linear Models*, London: Chapman and Hall.
- Murstein, Bernard I. (1986). *Paths to Marriage*, Beverly Hills: Sage publications.
- Oppenheimer, Valerie K. (1977). The sociology of women's economic role in the family, *American Sociological Review* 42: 387–406.

- Raftery, Adrian E. (1986). Choosing models for cross-classifications, *American Sociological Review* 51: 145–146.
- Róbert, Péter (1991). Educational transition in Hungary from the post-war period to the end of the 1980s, *European Sociological Review* 7: 213–236.
- Rockwell, Richard R. (1976). Historical trends and variations in educational homogamy, *Journal of Marriage and the Family* 38: 83–95.
- Shockey, James W. & Clogg, Clifford C. (1983). *ANOASC: A Computer Program for the Analysis of Association in a Set of K I-by-J Contingency Tables*, University Park, PA: Pennsylvania State University, Population Issues Research Center.
- Sorokin, Pitirim A. (1959 [1927]). *Social and cultural mobility*, Reprint, New York: Free Press.
- Surra, Catherine A. (1990). Research and theory on mate selection and premarital relationships in the 1980s, *Journal of Marriage and the Family* 52: 844–865.
- Tsai, Shu-Ling (1994). Stratification of marriages in Taiwan: the relative importance of ethnicity and education, Paper presented at the meeting of the Research Committee 28 of the ISA during the World Congress of Sociology, Bielefeld, July 18–23.
- Ultee, Wout C. and Luijkx, Ruud (1990). Educational heterogamy and father-to-son occupational mobility in 23 industrial nations: general societal openness or compensatory strategies of reproduction?, *European Sociological Review* 6: 1–25.
- Warren, Bruce L. (1966). A multiple variable approach to the assortative mating phenomenon, *Eugenics Quarterly* 13: 285–290.