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No Decline in Assortative Mating for Educational Level

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Education data from the Norwegian twin panel reveal no decline in the marital correlation for educational attainment for the past 35 years. Comparable marital correlations are found for British and American samples. A higher marital correlation is obtained for the parents of the Norwegian twins and the parents of their spouses, but this is an artifact. A twin's recall of his/her parents' educational levels is shown, by model fitting, to be biased by his/her own education level. Allowing for this bias reduces our estimate of the parental marital correlation and reduces estimates of the broad heritability of educational attainment from 74–81 to 49–58%. Other, unrelated factors may also be biasing estimates of the similarity of their parents' educational levels by the twins and their spouses.

KEY WORDS: assortative mating; education; no secular changes.

INTRODUCTION

The contribution of assortative mating to social inequality has received little empirical study. Positive assortative mating for a trait which is

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transmitted within families, whether through genetic inheritance, environmental transmission, or a mixture of both, will increase the variance of that trait, compared with random mating (Fisher, 1918; Rice *et al.*, 1978). Positive assortative mating for a complex variable such as educational attainment, which is determined by a variety of cognitive and noncognitive traits (Lavin, 1965; Bennett *et al.*, 1966; Oden, 1968; Rutter *et al.*, 1970), will generate correlations between these variables in the next generation (Thompson, 1966; Heath, 1983; Eaves *et al.*, 1984). Thus recurrent assortative mating for educational level or other characters related to social status will ensure that individuals in a population who are disadvantaged in one respect will tend to be disadvantaged in many other respects too.

A secular decrease in the intensity of assortative mating will produce a decrease in the variability of a character in the next generation and, in the case of variables such as educational attainment, will also produce a change in the covariance structure of the determinants of that variable. Just such a decline in assortative mating has been suggested for cognitive ability (Johnson *et al.*, 1980) and for educational attainment (Ahern *et al.*, 1983). Unfortunately conclusions about cognitive ability were based on a comparison of a variety of different studies using different tests and different sampling procedures. The "secular change" in assortative mating for ability may rather reflect a decline in psychometric standards, recent studies using less reliable group tests (e.g., DeFries *et al.*, 1979), or samples with severe range restriction (e.g., Horn *et al.*, 1979). Conclusions about assortative mating for educational attainment were justified by comparing the correlation between the years of education reported by pairs of spouses to the marital correlations between the years of education which they reported for their parents. The generational difference could therefore reflect a tendency to overestimate the educational similarity of one's parents.

Data from the Norwegian twin register, a population-based panel of like-sex twins born throughout the period 1915–1960 (Magnus *et al.*, 1983), permit a powerful test of the hypothesis that assortative mating for educational level is declining. As part of a three-generation cross-cultural study of the inheritance of cardiovascular risk factors, twins and their spouses were asked, by mailed questionnaire, to report the years of education completed by themselves and by their parents. We can thus compare the marital correlations for educational attainment obtained for the twins and their spouses and for their parents, broken down by decade of birth of the twins. Since each member of a twin pair reported separately the years of education completed by his or her parents, a powerful test of the validity of such retrospective data is available.

Table I. Structure of the Sample

	Twin group			
	Male MZ	Female MZ	Male DZ	Female DZ
Twins	2433	3065	3356	3855
Spouses	1207	1637	1563	1958
Mothers	1445	1688	2127	2292
Fathers	1492	1662	2139	2249
Mothers-in-law	1115	1475	1447	1753
Fathers-in-law	1096	1467	1424	1720

SAMPLE AND METHODS

The overall structure of our sample is summarized in Table I. In that table we give the total number of male monozygotic (MZ), female MZ, male dizygotic (DZ), and female DZ twins and, for each twin group, the total numbers of spouses, mothers, fathers, mothers-in-law, and fathers-in-law for whom educational data are available. Table I reveals an excess of female twins, a common finding in studies which rely upon the cooperativeness of subjects; but the excess of MZ twins which is common when twins are ascertained through appeals for volunteers (Lykken *et al.*, 1978) has not occurred.

In Table II we summarize the proportion of twins and their spouses achieving a given educational level, as a function of the year of birth of the twins. Also given are the educational levels of the parents and parents-in-law of the twins. Since there were no systematic differences in educational attainment between twins of different zygosity, between the twins and their spouses, or between their parents and their parents-in-law, figures have been pooled for twins and spouses of the same sex and parents and parents-in-law of the same sex. A more detailed breakdown is available on request from the authors. It should be noted that a small proportion of the twins from the 1950–1960 cohort, born during 1960, and spouses of twins born during or after 1960 will not have had a chance to complete more than 12 years of education at the time of the questionnaire study. Educational level in this cohort is slightly underestimated in Table II.

Educational level measured as years of education completed is clearly a discontinuous variable, though it is often treated as a continuous one (e.g., Ahern *et al.*, 1983; Vogler and Fulker, 1983). For the present study educational level was classified into four ordered categories:

Table II. Educational Level by Twins' Year of Birth

Twins' year of birth	N	Years of education completed (%) 0–7	8–9	10–12	>12
1915–1939					
Male twins, spouses	3131	58	22	8	12
Female twins, spouses	3281	60	29	6	5
Mothers, mothers-in-law	4494	86	11	2	1
Fathers, fathers-in-law	4416	81	12	3	4
1940–1949					
Male twins, spouses	2787	33	36	8	23
Female twins, spouses	2888	27	52	10	11
Mothers, mothers-in-law	4130	79	17	3	1
Fathers, fathers-in-law	4065	73	17	5	5
1950–1960					
Male twins, spouses	3474	13	49	17	21
Female twins, spouses	3529	10	53	22	15
Mothers, mothers-in-law	4713	68	24	6	2
Fathers, fathers-in-law	4620	62	22	8	8

(I) 0–7 years of education completed,
(II) 8–9 years of education completed,
(III) 10–12 years of education completed, and
(IV) more than 12 years of education.

Using discontinuous data to estimate a product–moment correlation will lead to a serious bias in the estimate obtained (e.g., Olsson, 1979). The resemblance of any set of pairs of relatives can, however, be summarized by a two-way 4 × 4 contingency table. In Table III, for example, we give the overall contingency tables for twins and their spouses, pooling MZ and DZ, male and female, and first-born and second-born twins, and for their mothers and fathers, pooling the parents of the twins and the parents of their spouses. When calculating the latter table, for the parents of twins we have used only the educational levels reported by the first twin from each pair (i.e., the first twin to be included in the twin panel) or the second if the first twin did not provide this information.

The derivation of "polychoric" correlations from discontinuous data has been discussed by Olsson (1979; see also Pearson, 1900; Tallis, 1962; Eaves *et al.*, 1978). We hypothesize that underlying our discontinuous scale of level of education there is a continuous normal distribution of true educational achievement, our class boundaries corresponding to thresholds superimposed upon the latent distribution (cf. Pearson, 1900). We further hypothesize that underlying a two-way contingency table are two continuous latent variables (the true educational achievements of the

Table III. Assortative Mating for Educational Level in Two Generations

Twins and their spouses				
	Husband			
Wife	I	II	III	IV
I	1269	423	92	71
II	725	1474	299	396
III	77	258	182	263
IV	26	73	62	458
Parents of twins and parents of their spouses				
	Father			
Mother	I	II	III	IV
I	8785	1005	215	122
II	583	1081	296	318
III	58	117	153	167
IV	12	23	20	152

pairs of relatives) whose joint distribution is bivariate normal with a correlation ρ and that superimposed upon the continuous distributions of these latent variables are thresholds $t_0, t_1, \ldots, t_4$ and $t'_0, t'_1, \ldots, t'_4$, where $t_0 = t'_0 = -\infty$ and $t_4 = t'_4 = +\infty$, and the values of $t_1 \ldots t_3$ and $t'_1 \ldots t'_3$ and ρ are to be determined by model fitting. Let p_{ij} denote the probability that an observation falls into the i, jth cell of the two-way contingency table, under our hypothesis of an underlying bivariate normal distribution; and let the observed frequency of individuals in that cell be $f_{i,j}$. Immediately we see that the log-likelihood of a set of observations under our hypothesis will be

$$L = \ln(c) + \sum\sum f_{ij} \ln(p_{ij}), \tag{1}$$

where c is a constant,

$$p_{ij} = \phi(t_i, t'_j) - \phi(t_{i-1}, t'_j) - \phi(t_i, t'_{j-1}) + \phi(t_{i-1}, t'_{j-1}),$$

and ϕ is the bivariate normal distribution function with correlation ρ. Hence to obtain maximum-likelihood estimates of the correlation between relatives for educational attainment and the threshold values, we may simply maximize function (1) with respect to these parameters or, equivalently, minimize minus the log-likelihood. In practice we have minimized minus the log-likelihood, using a commercially available routine for minimization, E04JBF (Numerial Algorithms Group, 1978).

To test the goodness of fit of the bivariate normal threshold model, we may calculate a chi-square value based on $(kk' - k' - k)$ degrees of freedom,

$$\chi^2 = \sum\sum (O_{ij} - E_{ij})^2/E_{ij},$$

where O_{ij} and E_{ij} are the observed and expected frequencies of the i, jth cell. Alternatively, we may calculate the log-likelihood, L_0, obtained under a perfect-fit solution which estimates a separate probability p_{ij} for each cell in the two-way contingency table. The likelihood-ratio statistic

$$C = 2(L_0 - L)$$

is approximately distributed as chi-square with $(kk' - k - k')$ degrees of freedom. The latter statistic was preferred whenever low expected cell frequencies, which would invalidate the use of the overall chi-square test of goodness of fit, were obtained. The sampling covariance matrix of our estimates of ρ and the threshold values may be obtained as the inverse of the Fisher information matrix, whose elements are given (see Tallis, 1962; Olsson, 1979) by

$$I_{m,n} = N \sum\sum \frac{1}{p_{ij}} \left(\frac{dp_{ij}}{d\theta_m}\right) \left(\frac{dp_{ij}}{d\theta_n}\right),$$

where N is the total number of observations and $\boldsymbol{\theta}$ is the vector of parameter estimates (i.e., of the estimates of ρ and the threshold values). In the analyses reported here, all first partial derivatives were evaluated numerically by forward differences (Conte and deBoor, 1965). The standard error of our estimate of ρ is simply the square root of the corresponding diagonal element of the sampling covariance matrix.

RESULTS

Fitting the bivariate threshold model to the two contingency tables given in Table III yielded estimates of the marital correlation for educational achievement of 0.67 ± 0.009 in the twins and their spouses and 0.80 ± 0.006 in the parents of the twins and the parents of the spouses. This confirms the apparent generation difference in assortative mating reported by Ahern *et al.* (1983). Our estimates of the marital correlation are substantially greater than theirs, but estimates in the two studies are not directly comparable. The polychoric correlations which we report are correlations for the hypothesized underlying variable "educational attainment," not for the observed variable "educational level." Product–

moment correlations were also estimated, scoring educational level on a scale from 1 to 4, to use as starting values in the estimation of the polychoric correlations. (Since they were calculated using an ordinal measure of educational level, these correlations will, of course, be equivalent to Spearman's rank correlation computed ignoring ties.) These correlations were consistently found to underestimate the polychoric correlations by a substantial amount, values of 0.58 and 0.64 being obtained in the present case. The likelihood-ratio tests against a perfect-fit model (χ^2 = 163.65, df = 8, $P < 0.001$; χ^2 = 188.94, df = 8, $P < 0.001$), unfortunately, indicate that the threshold model used in computing the polychoric correlations fails to fit the data. Our hypothesis of an underlying continuous bivariate normal distribution of true educational attainment is inadequate.

One possible explanation of the failure of the bivariate normal threshold model is that we are attempting to pool heterogeneous data. The proportion of individuals achieving a given level of education varies as a function of the year of birth (Table II). We would therefore also expect the threshold values to vary according to year of birth. By pooling together data on twins and their spouses born between 1915 and 1960, we are in effect superimposing upon one another normal distributions with different means and variances. In general, a distribution which is a sum of normal distributions will not itself be normal.

Two approaches are available which would overcome this problem. We could extend the pedigreee analysis method of Lange *et al.* (1976; see also Eaves, 1978; 1980) to handle discontinuous variables and allow the threshold values to vary as a function of the year of birth. However, for the parents of the twins and the parents of their spouses, the birth date is in many cases unavailable. We have therefore followed the alternative course of subdividing our sample by the period of birth of the twins. Within any subsample, very little variation in educational attainment as a function of the year of birth would be expected.

In Table IV we summarize the polychoric correlations between the twins and their spouses, and between the parents of these twins and spouses, for twins born during the periods 1915–1924, 1925–1934, 1935–1939, 1940–1944, 1945–1949, 1950–1954, and 1955–1960. For comparison, we have also given rank correlations computed as before. For twins born throughout the period 1925–1960 our estimates of the marital correlation are remarkably consistent, falling within the range 0.59–0.64. The marital correlation for twins born in the period 1915–1924 is slightly higher (0.72 ± 0.03). In this age group, however, the marital correlation may be inflated through selective mortality. Data on the educational levels of the twins and their spouses were collected only on individuals still alive at the time of the survey. Any tendency for couples discordant for education

Table IV. Temporal Changes in the Marital Correlation for Educational Level

Twins' year of birth	Correlation coefficient and SE[a]					
	Husband–wife			Father–mother		
	N	$\rho \pm$ SE	r	N	$\rho \pm$ SE	r
1915–1924	538	0.72 ± 0.03	0.59	1259	0.85 ± 0.02	0.66
1925–1934	918	0.64 ± 0.03	0.54	1879	0.82 ± 0.02*	0.62
1935–1939	630	0.62 ± 0.03*	0.53	1267	0.81 ± 0.02	0.64
1940–1944	876	0.64 ± 0.03**	0.56	1683	0.79 ± 0.02**	0.62
1945–1949	1243	0.62 ± 0.02**	0.55	2386	0.78 ± 0.02**	0.61
1950–1954	1260	0.62 ± 0.02**	0.54	2492	0.79 ± 0.01**	0.64
1955–1960	683	0.59 ± 0.03**	0.51	2141	0.76 ± 0.01**	0.62

[a] ρ, polychoric correlation coefficient; r, Spearman's rank correlation coefficient.
* Rejected at the 1% significance level.
** Rejected at the 0.1% significance level.

to be discordant for age at death too would therefore lead to an overestimation of the marital correlation for educational attainment. Thus we can be certain that the marital correlation has not changed over a 35-year period, and it is possible that no change occurred during the previous 10 years either.

Marital correlations for the parents of the twins and for their parents-in-law, once again broken down by the year of birth of the twins themselves, again give no evidence of a substantial change in assortative mating for educational achievement. Throughout the period covered by the study, the marital correlations fall within the range 0.76–0.85, and for the parents and parents-in-law of twins born during 1925–1954, all correlations fall within the range 0.78–0.82. Many parents and parents-in-law of twins born in the later part of this period will themselves have been born in the years 1915–1924. The absence of any marked decline in the marital correlations found for the parental generation supports our interpretation that the slightly elevated correlation between twins and their spouses for twins born during 1915–1924 is a consequence of selective mortality. Reports of the educational levels of the parents were obtained regardless of whether they were alive at the time of the survey.

Though the marital correlations for the twins and their spouses and the marital correlations for their parents and parents-in-law reveal little or no decline in the intensity of assortative mating, there is a consistent and substantial difference between the two sets of correlations, similar to the difference we observed for the entire sample and to the difference between generations reported by Ahern *et al.* (1983). Since the parents of the younger twins in the sample will have been born during the same period as many of the older twins, we would expect the two sets of correlations in Table IV to converge at some point, but no such convergence occurs. This same generational difference is apparent, though less marked, if we compare the nonparametric correlations.

There are several possible explanations of this generational difference. Some of the twins and spouses in the sample have not had children, whereas their parents, by definition, have reproduced successfully. If couples with less similar educational levels are less likely to have children, we would expect the data from twins and their spouses to underestimate the intensity of assortative mating, insofar as this implies successful reproduction. It is possible, but unlikely, that twins practice less intense assortative mating than the rest of the population. Finally, the father–mother correlations may be inflated because of some systematic bias in the retrospective report of their parents' educational levels by the twins and their spouses.

Breaking down our sample into groups which are quite homogeneous with respect to year of birth considerably reduced the values of the likelihood-ratio statistic for testing against a perfect-fit model the bivariate normal threshold model. Nevertheless, as can be seen in Table IV, many of these values are still significant. It is possible that we are dealing with a genuine discontinuity, individuals completing their education at the same level being much more likely to marry than would be predicted under any continuous model. Alternatively, the failure of the multiple threshold model may merely reflect its sensitivity to outliers, our distributional assumptions being essentially correct.

Cross-Cultural Comparisons

Several data sets allow us to compare the marital correlations obtained for twins and their spouses in Norway with marital correlations for nontwin populations from other countries. From a reanalysis of U.S. census data, Warren (1966) has found marital correlations for educational level, measured on a nine-point scale from "no school" to "4+ years of college," in the range 0.55–0.63. Garrison *et al.* (1968) give two-way contingency tables, using a six-point scale of educational level, for all couples giving birth to babies in Minnesota in the years 1965 and 1966. Assuming an underlying bivariate normal distribution of educational attainment, we obtained estimates of the marital correlation of 0.63 for the 1965 data set and 0.64 for the 1966 data set. The chi-square tests of goodness of fit revealed a significant deviation from bivariate normality ($\chi^2 = 4086.39$, df = 24, $P < 0.001$; $\chi^2 = 3307.44$, df = 24, $P < 0.001$). In Table V we present data extracted from the 1971 census of Great Britain (Office of Population Censuses and Surveys, 1979; General Register Office, 1979), in which educational level was measured on a coarse three-point scale. By fitting our bivariate normal threshold model, we obtained estimates of the correlation between the true educational attainments of spouses of 0.67 for England and Wales and 0.70 for Scotland. Highly significant chi-square values were obtained for both of these data sets ($\chi^2 = 17055.18$, df = 3, $P = 0$; $\chi^2 = 4275.05$, df = 3, $P = 0$). The marital correlations which we obtained for twins and their spouses in the Norwegian twin register fit well with these values obtained in population surveys of spouses in other countries. The correlations observed for the parents and parents-in-law of the twins, in contrast, are very high. We must therefore conclude that a bias in the retrospective report of the educational attainments of the latter group is the most likely explanation of the apparent generational difference.

Table V. Assortative Mating for Educational Level in England and Wales and in Scotland[a]

England and Wales			
	Husband		
Wife	I	II	III
I	662,615	28,550	52,725
II	12,083	11,086	7,124
III	20,856	4,679	27,423
Scotland			
	Husband		
Wife	I	II	III
I	64,825	3,992	3,843
II	2,961	4,212	1,242
III	2,261	872	3,138

[a] I, above "A" level or equivalent; II, A level or equivalent; III, below A level or equivalent.

Model-Fitting Analyses

Data on the educational attainments of twins and the educational attainments which each twin reports for each of his or her parents will allow us to detect some biases in these retrospective reports. If MZ twins show greater agreement in their reports than DZ twins, we may conclude that some heritable aspect of the individual's phenotype is biasing reporting. If a twin's educational level is more highly correlated with the educational levels which he/she reports for his/her parents than with those which his/her cotwin reports, we may infer that the individual's own educational level is biasing his/her recall of his/her parents' education.

In Table VI we reproduce all possible polychoric correlations among the educational levels of the first twin (i.e., the first twin to be included in the population twin registry), second twin, mother as reported by first twin, mother as reported by second twin, father as reported by first twin, and father as reported by second twin. Separate matrices of correlations are given for male and female MZ twins and male and female DZ twins. All coefficients are polychoric correlations, obtained by fitting the bivariate normal threshold model. The two-way contingency tables from which these were calculated included only families in which educational data were available on both twins, and reports on both parents were

Table VI. Polychoric Correlations Between Educational Levels of Twins and Educational Levels Reported by Them for Their Parents, and Their Standard Errors[a]

	1	2	3	4	5	6		1	2	3	4	5	6
	Male MZ twins (N = 684)							Male DZ twins (N = 821)					
1	1.00	**0.01**	0.04	0.04	0.03	0.04	1	1.00	0.03	0.03	0.04	0.03	0.04
2	0.87	1.00	0.04	0.04	0.04	0.03	2	0.62	1.00	0.04	0.03	**0.03**	**0.03**
3	0.65	0.59	1.00	**0.01**	**0.02**	**0.03**	3	0.62	0.57	1.00	**0.01**	**0.02**	**0.03**
4	0.60	0.62	0.94	1.00	**0.03**	**0.03**	4	0.54	0.63	0.91	1.00	0.03	**0.02**
5	0.67	0.61	0.82	0.77	1.00	**0.01**	5	0.65	0.56	0.83	0.73	1.00	**0.01**
6	0.61	0.64	0.76	0.79	0.96	1.00	6	0.52	0.62	0.72	0.78	0.92	1.00
	Female MZ twins (N = 1003)							Female DZ twins (N = 1077)					
1	1.00	**0.01**	**0.03**	**0.03**	0.03	**0.03**	1	1.00	**0.02**	0.03	0.03	0.03	0.03
2	0.90	1.00	0.03	**0.03**	0.03	0.03	2	0.79	1.00	0.03	0.03	0.03	0.03
3	0.68	0.64	1.00	**0.01**	**0.02**	0.02	3	0.63	0.62	1.00	**0.01**	**0.02**	0.02
4	0.64	0.68	0.96	1.00	**0.02**	0.02	4	0.55	0.65	0.95	1.00	0.03	**0.02**
5	0.68	0.65	0.84	0.78	1.00	**0.01**	5	0.67	0.65	0.83	0.73	1.00	**0.01**
6	0.65	0.68	0.77	0.82	0.96	1.00	6	0.63	0.68	0.78	0.80	0.94	1.00

[a] Polychoric correlations are given in the lower triangle; their standard errors, in the upper triangle of each matrix. Standard errors are printed in boldface type in cases where the polychoric threshold model did not give an acceptable fit to the observed data. 1, first twin's education; 2, second twin's education; 3, mother's education (report of first twin); 4, mother's education (report of second twin); 5, father's education (report of first twin); 6, father's education (report of second twin).

provided by both twins. Polychoric correlation coefficients are given in the lower triangle in Table VI; their standard errors, in the upper triangle. Estimates of a polychoric correlation are given even when the bivariate normal threshold model failed to fit the data, but in these cases the estimated standard error is printed in boldface type.

Table VI reveals good agreement between twins in their recall of their parents' educational level: the lowest correlation between twins is for the report of maternal education by male DZ twins, for which a correlation of 0.91 was found. There is a very slight tendency in both sexes for MZ twins to show greater agreement than DZ twins. However, the correlation between the educational level of one twin and that which he/she reports for either parent is strikingly and consistently higher than the correlation between his/her own educational level and the educational level reported by his/her cotwin for the same parent. The cross-correlations between the educational level reported for one parent by one twin and that reported for the other parent by the cotwin are consistently smaller than the correlation between the educational levels reported for the two parents by the same twin. Precisely the same pattern is observed

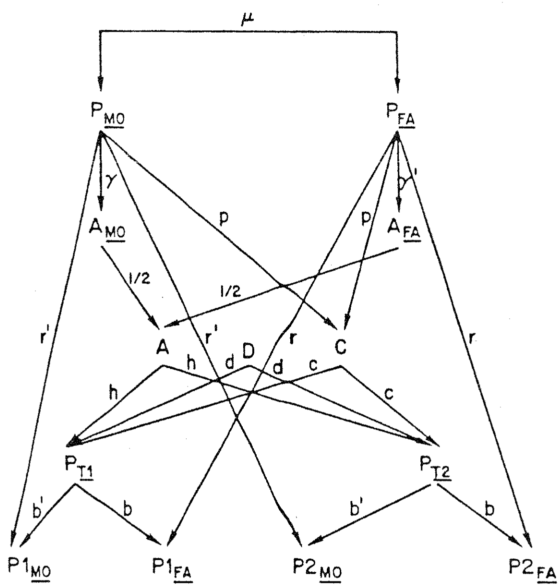

Fig. 1. Model for resemblance of MZ twins and their parents.

when Spearman's rank correlations are used, though differences between correlations are less pronounced, so these are not reproduced here. There is clear evidence of biased reporting, the more highly educated member of a twin pair tending to give a higher estimate of his/her parents' educational levels than the less-educated member.

The importance of this biased reporting of educational level can best be ascertained by model fitting. When fitting models we are compelled to use the polychoric correlations, despite the fact that in some cases the bivariate normal threshold model was rejected. Estimates of the polychoric correlation coefficient are very similar whether or not the threshold model fits, in cases where the expected correlation does not differ (see Tables IV and VI). This suggests that failure of the bivariate normal threshold model reflects its extreme sensitivity to even slight departures from bivariate normality, so that our estimates of the correlations for educational attainment will be subject to little statistical bias.

In Fig. 1 we present a simple model developed using path analysis (Wright, 1968) which represents the causes of the resemblance of MZ twins and their parents, allowing for additive gene action, dominance, familial environmental effects, a direct environmental effect of parental educational attainment on offspring environment, and primary phenotypic assortative mating for true educational attainment. These elements of the model represent a simplification of the model considered by Loehlin (1978). Alternative assumptions could be made about cultural transmission (Loehlin, 1978; Cloninger *et al.*, 1979; Heath, 1983) and about assortative mating (Rao *et al.*, 1979; Heath, 1983) but cannot be resolved using only data on twins and their parents. In addition, we distinguish between the true educational level of a parent, P, and the educational levels reported for the parent by each twin, P1 and P2. We include a direct path b (for fathers) or b' (for mothers) from the twin's phenotype P to the educational level reported by him/her for his/her parent, as well as path r (for fathers) or r' (for mothers) from the true educational level of the parent to the educational levels reported by each twin. Other symbols used in Fig. 1 are as follows: A—additive genetic or "breeding" value (Falconer, 1982; Mather and Jinks, 1982); D—dominance deviation; C—familial environmental value; h, d, and c and h', d', and c'—standardized path regressions of phenotype on additive genetic value, dominance deviation, and familial environmental value in males and in females; p and p'—standardized path regressions of offspring familial environmental value on paternal and maternal true phenotypic value; μ—primary correlation between true educational attainments of spouses, identical to the standardized path regression of the phenotype of one spouse on the phenotype of the other (Cloninger *et al.*, 1979); and γ and γ'—standardized path regressions of additive genetic value on phenotype in males and in females. $\gamma = h + ac$ and $\gamma' = h' + ac'$, where a is the genotype–environmental correlation which, at equilibrium, may be derived as

$$a = [\tfrac{1}{2}h(p + \mu p') + \tfrac{1}{2}h'(p' + \mu p)]/[1 - \tfrac{1}{2}c(p + \mu p') - \tfrac{1}{2}c'(p' + \mu p)].$$

All residual factors which contribute to the variance of a variable but not to its covariance with other variables have been omitted from Fig. 1. Some expected correlations under this general model are given in Table VII. Other expected correlations are easily derived by making appropriate substitutions in those given in Table VII.

Models were fitted to the correlations in Table VI by nonlinear generalized least squares (Lee and Jennrich, 1979), under the simplifying assumption that all correlations are independent. The standard deviations from Table VI were used as weights for model fitting. Thus we minimized, with respect to the parameters of a model, the function

$$C = \sum ([r_i - \bar{r}_i]/s_i)^2, \tag{2}$$

where r_i is the ith observed polychoric correlation, $\bar{r}_i$ is the corresponding expected correlation, and s_i is the standard deviation of that polychoric correlation. In using the empirical variances of our polychoric correlations as weights, we are assuming that sample sizes are sufficiently large for our observed correlations to be very close in value to the true population correlations and, hence, for the variances of our observed values to be

Table VII. Expected Correlations

Male MZ twins
$\rho_{MZM} = h^2 + d^2 + c^2 + 2hac$
Male DZ twins
$\rho_{DZM} = \frac{1}{2}h^2[1 + \mu(h + ac)(h' + ac')] + \frac{1}{4}d^2 + c^2 + 2hac$
Son–father (own report)
$b + \frac{1}{2}h[h + ac + \mu(h' + ac')]r + c(p + \mu p')r$
Son–mother (own report)
$b' + \frac{1}{2}h[h' + ac' + \mu(h + ac)]r' + c(p' + \mu p)r'$
Daughter–father (own report)
$b + \frac{1}{2}h'[h + ac + \mu(h' + ac')]r + c'(p + \mu p')r$
Male MZ twin–father (cotwin's report)
$b\rho_{MZM} + \frac{1}{2}h[h + ac + \mu(h' + ac')]r + c(p + \mu p')r$
Male DZ twin–father (cotwin's report)
$b\rho_{DZM} + \frac{1}{2}h[h + ac + \mu(h' + ac')]r + c(p + \mu p')r$
Mother and father (reported by same son)
$bb' + \mu rr' + \frac{1}{2}h[h + ac + \mu(h' + ac')]rb' + c(p + \mu p')rb' + \frac{1}{2}h[h' + ac' + \mu(h + ac)]r'b + c(p' + \mu p)r'b$
Mother and father (reported by male MZ twins)
$bb'\rho_{MZM} + \mu rr' + \frac{1}{2}h[h + ac + \mu(h' + ac')]rb' + c(p + \mu p')rb' + \frac{1}{2}h[h' + ac' + \mu(h + ac)]r'b + c(p' + \mu p)r'b$
Mother and father (reported by male DZ twins)
$bb'\rho_{DZM} + \mu rr' + \frac{1}{2}h[h + ac + \mu(h' + ac')]rb' + c(p + \mu p')rb' + \frac{1}{2}h[h' + ac' + \mu(h + ac)]r'b + c(p' + \mu p)r'b$
Father (reported by male MZ twins)
$b^2\rho_{MZM} + r^2 + h[h + ac + \mu(h' + ac')]rb + 2c(p + \mu p')rb$
Father (reported by male DZ twins)
$b^2\rho_{DZM}\, r^2 + h[h + ac + \mu(h' + ac')]rb + 2c(p + \mu p')rb$

very close to the variances of the population values. Provided that this assumption is justified and that the sampling distribution of the polychoric correlations is multivariate normal, as seems plausible for the very large sample sizes used here, and provided also that our correlations are indeed independent, the minimization of function (2) yields estimates of the model parameters which are asymptotically equivalent to maximum-likelihood estimates; and the minimum value of C obtained is distributed as chi-square, with the number of degrees of freedom equal to the number of observed correlations minus the number of model parameters estimated. Strictly, our correlations are not independent. Where the effects of ignoring the correlation between correlations have been examined, however, they have been found to be slight (Rao *et al.*, 1977).

The critical results of model fitting are summarized in Table VIII. With data on only twins and their parents we cannot fit a model which

Table VIII. Results of Model Fitting

	Parameter estimate (× 100)													Test of goodness of fit		
Model	h	h'	d	d'	c	c'	μ	p	p'	r	r'	b	b'	df	χ^2	P
1. $hh'dd'cc'\mu rr'bb'$	67	70	36	0	54	65	74			80	79	33	35	49	48.56	0.49
2. $hh'dd'cc'\mu$	82	86	37	0	26	41	79			100*				53	440.08	0
3. $hh'dd'cc'\mu rb$	67	70	36	0	54	66	74			79		34		51	51.43	0.46
4. $hdc\mu rr'bb'$	66		0		68		73			80	77	36	39	52	97.22	<0.001
5. $hh'cc'\mu pp'rr'bb'$	64	67			65	68	73	−17	19	79	79	38	34	49	60.91	0.12
6. $hh'cc'\mu rr'bb'$	65	68			65	67	73			79	78	35	37	51	61.79	0.14
7. $hh'dd'\mu rr'bb'$	81	89	47	35			84			98	96	−02	−02	51	177.19	0

* Parameter value fixed to unity *ex hypothesi*.

allows for additive gene action, dominance, and cultural transmission. A model which allows for sex-limited additive gene action, dominance, and familial environmental effects, with primary phenotypic assortative mating and with biases in the reporting of parental educational level which depend upon the sex of the parent (Model 1), gives an excellent fit to the data (χ^2 = 48.56, df = 49, P = 0.49). A significant positive estimate of the dominance parameter d is obtained in males. Fitting a model which allows for additive gene action, familial environmental effects, and cultural transmission but no dominance (Model 5) gives a significantly worse fit than Model 1 and yields a negative estimate for the paternal cultural transmission parameter. Any effects of cultural transmission will tend to be masked by dominance in these data. Restricting Model 1 by assuming that there is no sex limitation (Model 4), no dominance (Model 6), or no familial environmental effects (Model 7) in each case leads to expected correlations which do not fit the observed correlations. Model 3, which assumes that the regressions of reported parental educational level on true parental educational level and on twin's educational level do not depend upon either the sex of the parent or the sex of the twin, gives an excellent fit, does not give a significantly worse fit than Model 1 (χ^2 = 51.43, df = 51, P = 0.46), and is therefore preferred over Model 1.

The most critical result in Table VIII is that of fitting Model 2, which simplifies Model 1 by assuming that the educational levels reported for the parents are identical to their true educational levels. This model gives a very significantly worse fit than Model 1 (χ^2 = 388.65, df = 4, P = 0) and is rejected at a very high level of significance. Allowing for the bias in the report of parental educational levels reduces our estimate of the true marital correlation for educational level in the parents of the twins to 0.74. Allowing for this bias also has a major effect on our conclusions about the importance of genetic and environmental influences on educational attainment. Under the rejected model, Model 2, we obtained estimates of broad and narrow heritabilities (Falconer, 1982) of 81 and 67% in males and 74 and 74% in females. Under the best-fitting model, Model 3, we obtained estimates of 58 and 45% in males and 49 and 49% in females. Our estimate of the parameter b under Model 3 is 0.34, indicating that bias in the report of parental educational levels is making a sizable contribution to the parent–offspring correlation for educational attainment. Making allowance for this bias not only decreases our estimate of the narrow heritability of educational attainment, but also increases slightly our estimate of the contribution of dominance to the total genetic variance in males (17 vs. 21%).

Bootstrap Analyses

The validity of our model-fitting analyses depends critically upon two assumptions. The first of these, that our estimates of the true correlation between relatives for educational attainment are almost unbiased (in the statistical sense), we have justified briefly by noting that in cases where the expected correlation is the same, very similar polychoric correlations were estimated regardless of whether or not the threshold model fitted the data in any particular case. The second assumption, that the sampling distribution of a polychoric correlation can be assumed to be normal and that an adequate estimate of its standard deviation can be obtained from the estimated sampling covariance matrix, even when the threshold model gave a poor fit to the data, might appear to be more questionnable.

To explore this second issue, we have used the method of bootstrapping (Effron, 1982; Heath *et al.*, 1984) to approximate the sampling distribution of each of the 60 polychoric correlations used in model fitting. For each twin group, the original set of data points was sampled from at random, and with replacement, to generate 50 new data sets of identical sample size. If the sampling distributions of the original polychoric correlations are indeed normal, we would also expect the empirical sampling distributions, i.e., the distributions of the bootstrapped polychoric correlations, to be normal (Effron, 1982). For 56 of the 60 polychoric correlations, the bootstrapped distribution did not deviate significantly from normal, by Shapiro–Wilk test (Shapiro and Wilk, 1965), and for the remaining 4 polychoric correlations, the deviation was not great. It appears that the sampling distribution of the original polychoric correlations can safely be assumed to be normal.

Model 2 in Table VIII was refitted to the original set of polychoric correlations, using the standard deviations of the bootstrapped distributions to replace the theoretical standard deviations in function (2). Using the bootstrapped standard deviations, the following parameter estimates were obtained: $h = 0.68$, $h' = 0.72$, $d = 0.36$, $d' = 0.0$, $c = 0.53$, $c' = 0.64$, $\mu = 0.75$, $b = 0.32$, and $r = 0.80$. All these parameter estimates fall within ± 0.03 of the estimates obtained using the theoretical standard deviations. An approximate chi-square value of 50.22 was obtained using the empirical standard errors, which is very close to the original value of 51.43. It appears that using the theoretical standard deviation, even when the bivariate threshold model gives a poor fit, has not led to serious errors in our estimates.

DISCUSSION

From Table IV, we have seen that assortative mating for educational level has remained remarkably constant over a period of 45 years but that some factors are biasing upward our estimate of the marital correlation in the parents and parents-in-law of our twins. From our cross-cultural comparisons it does not appear that the correlations found for the twins and their spouses have been biased downward in any way. Model fitting has shown that there is a tendency for the better-educated member of a twin pair to give a higher estimate of the educational level of his parents than the less-educated twin. Allowing for this effect reduced our estimate of the marital correlation in the parental generation from 0.82 to 0.74, much closer to the value of 0.67 obtained for the total sample of twins and their spouses. The remaining difference may reflect a tendency to overestimate the similarity of the educational attainments of one's parents which is unrelated to one's own educational attainment or any other partly heritable trait. Alternatively, there may have been a genuine decrease in the marital correlation for educational attainment during the first decades of this century. Only the collection of data from the parents themselves will help resolve this issue.

Our conclusions about the familial transmission of educational attainment should not be regarded as definitive. Considering all possible relationships between the twins and their parents and their spouses and their spouses' parents in a single analysis will allow us to resolve a much wider range of alternative hypotheses about assortative mating. Obtaining data on the offspring of the MZ and DZ twins will similarly allow the resolution of dominance and a variety of mechanisms of cultural transmission (Heath, 1983). This should allow a more precise estimate of the importance of cultural and biological influences on educational attainment. Nevertheless, the magnitude of the change in parameter estimates which occurred when bias in the reporting of parental educational level was allowed for has important implications for genetic analyses and all other studies which rely upon unvalidated retrospective report data. With suitable data, as we have shown here, such reporting biases can be corrected for.

For several relationships we have found that a bivariate normal threshold model gives a very poor fit to the data. Even when only a narrow age range was considered, this problem persisted. It is probable that this reflects the great sensitivity of the multiple threshold model to even slight deviations from bivariate normality. When the empirical distributions of our polychoric correlations were determined by bootstrapping, little deviation from the theoretial distribution expected if the threshold model gave a good fit to the data was found. It seems that for data of this type, ignoring the problem of nonnormality, as we have done in most of the analyses in this paper, does not lead to a serious bias in parameter estimates.

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