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Robert M. Hauser
and
Raymond Sin-Kwok Wong

CDE Working Paper 87-4

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Robert M. Hauser
Raymond Sin-Kwok Wong

Center for Demography and Ecology
Department of Sociology
The University of Wisconsin-Madison
1180 Observatory Drive
Madison, Wisconsin 53706

CDE Working Paper 87-4

Revised August 22, 1987



This research was carried out with support from the Spencer Foundation, the National Science Foundation (SES-80-10640), the CASMIN Project (Comparative Analysis of Social Mobility in Industrial Nations) at The University of Mannheim, Federal Republic of Germany, and the Graduate School of the University of Wisconsin-Madison. Computations were performed using facilities of the Center for Demography and Ecology of the University of Wisconsin-Madison, which are supported by the Center for Population Research of the National Institute of Child Health and Human Development (HD-5876) and by the William and Flora Hewlett Foundation. Statistical tabulation and estimation were performed using SPSS-X and LISREL VI on a VAX 11/780. All of the data used in this paper are available in cited publications, are published herein, or are available from the authors. We thank Michael R. Olneck for his generosity and advice in helping us use primary data from his Kalamazoo study, and for other helpful comments. We thank Taissa Hauser, Mary Schlaefer, and Junko Sato for assistance in research. The opinions expressed herein are those of the authors.

*SIBLING RESEMBLANCE AND INTER-SIBLING EFFECTS*
*IN EDUCATIONAL ATTAINMENT*

ABSTRACT

Motivated by Benin and Johnson's (1984) study of two samples of adult Nebraska siblings, we model reciprocal influence between male siblings to determine whether there is reciprocal influence between siblings or a predominant flow of influence from older to younger siblings in the determination of educational attainment. Contrary to Benin and Johnson, our reanalysis of their data shows that differences in common family variance among types of sibling pairs are not derived from an unusually high level of fraternal resemblance, but only from the unusually low similarity between older sisters and younger brothers. While the Nebraska data do not identify a model with reciprocal effects between brothers' educational attainments, suitable data are available in Olneck's (1976) study of brothers from Kalamazoo, Michigan. These permit us to estimate reciprocal effects of brothers' educational attainments. In modeling the Kalamazoo data, we use a mixture of multiple proxy reports and self-reports to take account of random and correlated response errors in the measured variables. In brief, we find evidence of reciprocal influence of brothers' levels of educational attainment, net of the common effects of family background and the effect of each brother's mental ability on his schooling. The unconstrained effect of older brother's schooling on younger brother's schooling is larger than the reciprocal effect, but there is no significant difference between the two effects. This leaves open the questions whether reciprocal influence occurs in other types of sibling pairs and whether there is a predominant effect of older on younger siblings.

This paper reports an investigation of sibling resemblance in educational attainment. Specifically, we attempt to model reciprocal influence between male siblings in an effort to determine whether there is a predominant flow of influence from older to younger siblings. The study is motivated by Benin and Johnson's (1984) study of two samples of adult Nebraska siblings, from which they concluded that pairs of brothers resemble one another more, net of socioeconomic background, than do pairs of sisters or brother-sister pairs. Benin and Johnson (1984:18-19) say that the greater similarity of brother pairs is better explained by cross-sibling influence, predominantly from older to younger siblings, than by the possibility that some unmeasured common family factors affect brothers, but not sisters. They suggest that sibling role modeling and facilitation (social contacts and support) are sources of the greater resemblance between brothers and that "role modeling has a stronger effect on sibling resemblance than does facilitation."

After a brief review of previous research, we reanalyze the sibling data reported by Benin and Johnson (1984), which is the first published report based on all four possible combinations of sibling pairs ordered by sex and age. Contrary to Benin and Johnson, we find that differences in common family variance among types of sibling pairs are not derived from an unusually high level of fraternal resemblance, but only from the unusually low similarity between older sisters and younger brothers. Since Benin and Johnson's interpretation rests partly on the finding of a high degree of fraternal resemblance, this leads us to look for more direct evidence of mutual influence between brothers.

While the Nebraska data do not permit us to identify a model with reciprocal effects between brothers' educational attainments, suitable data are available in Michael Olneck's (1976) well-known study of brothers from Kalamazoo, Michigan.[1] In addition to data on educational attainments and parental status characteristics, the Kalamazoo data include mental test scores for each brother. These permit us to estimate reciprocal effects of brothers' educational attainments on the assumption that one man's mental ability does not affect his brother's schooling. We believe that this analysis is responsive to Benin and Johnson's (1984:19) call for the development of models for male siblings that include cross-sibling effects. In modeling the Kalamazoo data, we use a mixture of multiple proxy reports and self-reports to take account of random and correlated response errors in the measured variables. In brief, we find evidence of reciprocal influence of brothers' levels of educational attainment, net of the common effects of family background and the effect of each brother's mental ability on his schooling. The unconstrained effect of older brother's schooling on younger brother's schooling is larger than the reciprocal effect, but there is no significant difference between the two effects. This leaves open the questions whether reciprocal influence occurs in other types of sibling pairs and whether there is a predominant effect of older on younger siblings.

Families in the Stratification Process

It is a sociological truism that the family is one of the important stratifying agents in the transmission of social and economic inequality. This was well stated many decades ago by Charles Horton Cooley: "There is a certain opposition between the ideal of equal opportunity and that of family responsibility. Responsibility involves autonomy, which will produce divergence among families, which, in turn, will mean divergent conditions for the children; that is, unequal opportunities ..." (quoted from *Social Process* in

Blau and Duncan 1967:vii). The well-known models of the stratification process represent family characteristics through a handful of variables - race, education, occupation, earnings, female headship, size of sibship, and the like - that cannot very well represent the full variety or influence of family environments. Although Blau and Duncan (1967:316-28) initiated modern studies of sibling resemblance in their discussions of the effects of "family climate" on achievement and the possibility of cross-sibling influence, the emphasis in research has been on between-family differences in easily measured social and economic characteristics, rather than on differences within the - family in resources or social support, on typical patterns of activity or interaction, or on mutual influences among family members. For example, the classic status attainment model, introduced by Blau and Duncan (1967), and then elaborated as the Wisconsin model (Sewell, Haller, and Portes 1969, Sewell, Haller, and Ohlendorf 1970, Sewell and Hauser 1972, Sewell, Hauser, and Wolf 1980, Hauser, Tsai, and Sewell 1983), focuses on the ways that socioeconomic background affects individual status attainment in adulthood through social and psychological processes during adolescence.

Beyond the neglect of family factors, there have been two other distinctive weaknesses of the status attainment paradigm. One line of criticism typically centers on the contention that status attainment models underestimate the extent or persistence of inequality, the importance of allocative mechanisms controlled by elites relative to socialization and self-selection, or the importance of labor markets relative to schooling and socializing experiences of youth as sources of social stratification. To a degree, these criticisms have coalesced in the development of the so-called "new structuralism." The other line of criticism argues that status attainment models underestimate the importance of family background because of response errors, omitted variable bias, and other types of misspecification (Bowles 1972). The latter concerns sound methodological, but in principle their substantive implications could be large.

As statistical methods have become more sophisticated, some of the major concerns of Bowles and other critics have been tested. For example, one can specify measurement errors or response errors (random or correlated) in both the dependent and independent variables and then purge such effects from estimates of structural parameters. Several studies have found that response variability in proxy reports of social and economic variables do not lead to large biases in estimates of the effects of social background characteristics or of schooling on adult social and economic success (Jencks et al 1972, Bielby, Hauser, and Featherman 1977, Bielby and Hauser 1977, Olneck 1977, Hauser, Tsai, and Sewell 1983, Hauser and Mossel 1985, Hauser 1984, Hauser and Sewell 1986). At the same time, there is sufficient evidence of response variability in socioeconomic variables to imply that correction for it ought to become standard procedure in future stratification research.

If there is agreement that families are agents of stratification, questions remain about the size and mechanisms of family influence. Some believe that the family, together with other social institutions, mainly reflects and embodies class reproduction (Bourdieu and Passeron 1977; Bowles and Gintis 1976; Althusser 1971). Others, like Griliches (1979) believe that the family also has a depolarizing effect, especially as a potential income equalizer. That is, because there is heterogeneity among family members, parents may attempt to compensate for differences among offspring in order to equalize outcomes. In smaller families, parents may be more effective in equalizing outcomes, thus accounting for the greater heterogeneity in socioeconomic outcomes in larger sibships. As family size declines, between-family differences may then become relatively larger as a component of inequality; there is presumably a trade-off between these effects of lesser inequality within (smaller) families and declines in the main effects of size of sibship (Blake 1981, Heer 1985).

Even though research has found declining effects of socioeconomic background on achievement in recent cohorts (Featherman and Hauser 1978, Hout 1984, Hout 1986), it is likely that the family unit will remain an important stratifying agent. There is ample evidence of heterogeneity among families in socioeconomic life chances, that is, of sibling resemblance. For example, after correcting for measurement error, Hauser and Featherman (1976:117) find that about two-thirds of the variance in the educational attainments of cohorts of American men born between 1907 and 1951 can be explained by a common family factor, and about half the variance in the common family factor can be explained by measured socioeconomic variables. After correction for measurement error (Olneck 1976:186), Olneck's (1977:132) findings for Kalamazoo brothers imply that a common family factor explains 52 percent of the variance in mental test scores, 59 percent of the variance in educational attainment, 49 percent of the variance in status of first occupation, 37 percent of the variance in status of current occupation, and 27 percent of the variance in annual earnings. In their sample of Wisconsin brothers, Hauser and Sewell (1986) find that a common family factor explains 49 percent of the variance in mental ability, 46 percent of the variance in educational attainment, 41 percent of the variance in status of first occupation, 38 percent of the variance in status of current occupation, and 27 percent of the variance in annual earnings. The difference in findings about educational attainment between the national sample (the 1973 Occupational Changes in a Generation Survey) and the Kalamzoo and Wisconsin samples may well be an effect of restriction in range in the latter; that suggests we would find even larger family effects on ability, occupational status, and earnings, if national data on them were available.

Some researchers, like Behrman and Taubman (1976), Taubman (1976) and Behrman, Taubman, and Wales (1977) have studied sibling resemblance in an effort to resolve the old debate about 'nature' versus 'nurture' as sources of social inequality. The major emphasis in these studies is on comparing genetic and environmental sources of variation in education, occupation, and earnings by using data on the resemblance between monozygotic (MZ) and dizygotic (DZ) twins. Yet, as Jencks and Brown (1977) have argued, this heredity-environment distinction is based on a false dichotomy. A high correlation between genotype and phenotype need not imply a low effect of environmental influences on phenotype. Moreover, data on twins do not provide any more identifying restrictions than data on sibling pairs (Chamberlain 1977, Goldberger 1977) without additional assumptions that are both implausible and restrictive. Our concern in studying siblings is not to estimate how much variance genes can explain in a given outcome, but to elucidate the effects of measured and unmeasured variables in the process of stratification (see Hauser and Mossel 1985, Hauser and Sewell 1986).

In the present study, our interest is on the micro-foundation of familial equality or inequality, and in particular, sibling resemblance. There is a long-standing tradition of research on social and psychological differences among siblings, which centers on the measurement and explanation of supposed effects of ordinal position. In recent years, birth order research has developed in response to the Zajonc-Markus confluence model, which explains differences in mental ability within families by differences in the social environments faced by offspring of different ages across time (Zajonc and Markus 1975, Zajonc 1976, Zajonc, Markus, and Markus 1979, Zajonc 1983). Despite the efforts that have been devoted to this line of research, the bulk of the evidence appears to show that such effects are of minor importance, once controls have been introduced for size of sibship.[2]

A related line of research, initiated by Blau and Duncan and revived by Benin and Johnson (1984), has postulated that older siblings educational attainments affect those of younger siblings by means of role modeling or social facilitation; here, cross-sibling influence is interpreted as a source of similarity in outcomes, rather than of differentiation. The main purpose of this paper is to provide new evidence about this source of sibling resemblance.

Another Look at the Nebraska Data

As Benin and Johnson (1984) have noted, most sibling-based studies of the stratification process have included only brother-brother pairs, and thus have not addressed potential differences in the size or sources of resemblance in female or mixed-sex sibling pairs. In their analysis, they were able to obtain sample data for all four different combinations of sibling pairs: older brother-younger brother (OB-YB), older brother-younger sister (OB-YS), older sister-younger brother (OS-YB), and older sister-younger sister (OS-YS) pairs. Their data were obtained from two independent samples: the 1981 Nebraska Annual Social Indicators Survey (NASIS) and a 1976 survey in Lincoln, Nebraska (Lincoln). In each survey, adult respondents were asked to report the sex and educational attainments of their offspring aged 21 and older, along with the educational attainment of the mother and father and the father's occupation. The NASIS data are drawn from a statewide probability sample and represent the adult resident population. The Lincoln sample, on the other hand, is restricted to children from maritally stable families who had lived in the Lincoln area for 28 to 29 years. The NASIS sample covers a much broader age range among offspring (21 to 70) than the Lincoln sample (21 to 50). The final weighted sample sizes for NASIS and Lincoln data are 442 and 462 respectively. The Benin-Johnson analyses used all possible pairs formed from these data, weighted to compensate for clustering and for multiple appearances of individuals in pairs (pp. 13-15); our reanalysis uses the correlations and standard deviations reported by Benin and Johnson (1984:20-1).

Figure 1 shows a path diagram of the model of sibling's educational attainment that was used in most of Benin and Johnson's analysis. There are three exogenous variables, father's occupational status ($\xi_1$ = POPDUNC), father's educational attainment ($\xi_2$ = POPED), and mother's educational attainment ($\xi_3$ = MOMED). Occupational status is coded on the Duncan SEI (rescaled downward by a factor of 10 for convenience in reporting parameter estimates), and educational attainment is expressed in years. Each of the exogenous variables is permitted to covary with the others, as shown by the curved, two-headed arrows on the left side of the diagram, and each affects the educational attainments of each member of a sibling pair. In the diagram, the dependent variables are older sister's education ($\eta_1$ = OSED) and younger sister's education ($\eta_2$ = YSED), but there is a similar model for each type of sibling pair. Note that the same person may appear more than once in a paired subsample, not only in the same ordinal position, but also - within same-sex pairs - in the other ordinal position. To the extent that siblings' educational attainments are not explained by the measured variables, they are determined by disturbances, $\zeta_1$ and $\zeta_2$, which are assumed to be independent of the exogenous variables, but are not necessarily uncorrelated with one another. The latter possibility is indicated by the curved arrow connecting $\zeta_1$ and $\zeta_2$ at the right hand side of the path diagram.

The effects of substantive interest in the Benin-Johnson analysis are the slopes of educational attainments on the background variables, where $\gamma_{ij}$ is the effect of $\xi_j$ on $\eta_i$, and the covariances between the disturbances, denoted by $\psi_{12}$. If the $\psi_{12}$ are positive, then there must be at least one source of variance in educational attainment that is common to siblings, but not reflected in measured social background. Although this model consists of two multiple regression equations in measured variables, sharing the same set of regressors, Benin and Johnson used the LISREL program to estimate the parameters because the maximum likelihood estimation method used by LISREL permits the imposition of equality constraints on parameters across equations and across multiple samples (Jøreskog and Sørbom 1984). Thus, Benin and Johnson were able to test equalities of the form $\gamma_{11} = \gamma_{21}$, which says that the effect of father's occupational status on each siblings' educational attainments is the same. Also, they were able to test equalities of the form, $\psi_{12} = \psi^*_{12}$, where $\psi_{12}$ and $\psi^*_{12}$ pertain to disturbance covariances in different types of sibling pairs.

We have deliberately used the LISREL notation for the Benin-Johnson model because we use an elaboration of this model in subsequent analyses.[3] In its general form, the LISREL model has three sets of equations. The structural model, part of which is used in figure 1, is

$$\eta = \mathbf{B}\eta + \Gamma\xi + \zeta, \qquad (1)$$

where $\eta$ is a vector of endogenous latent variables, $\xi$ is a vector of exogenous latent variables, $\zeta$ is a vector of disturbances that are independent of the $\eta$ and $\xi$ and $\mathbf{B}$ and $\Gamma$ are parameter matrices; $\mathbf{B}$ is a square matrix, defined with zero entries on its main diagonal to indicate which is the left-hand variable in each equation. The variance-covariance matrices of the $\xi$, denoted by $\Phi$, and that of the $\zeta$, denoted by $\Psi$, are also parameter matrices of the structural model. There are two measurement models, which specify the dependence of observed variables on the latent variables. These are

$$X = \Lambda_x\xi + \delta \qquad (2)$$

and

$$Y = \Lambda_y\eta + \epsilon, \qquad (3)$$

where X and Y are vectors of observed indicators of exogenous and endogenous variables, respectively, $\Lambda_x$ and $\Lambda_y$ are parameter matrices, $\delta$ and $\epsilon$ are vectors of independently distributed errors in variables, and the variance-covariance matrices of $\delta$ and of $\epsilon$ are denoted by $\Theta^\delta$ and $\Theta^\epsilon$, respectively. In the model of figure 1, equations 2 and 3 are trivial because observables are equivalent to latent variables.

Benin and Johnson's findings are provocative and consistent across samples. First, they find that the residual covariances, $\psi_{12}$, between sibling's educational attainments differ significantly across pair types; the OB-YB pairs have the largest residual covariances and the OS-YB pairs have the smallest; their findings are shown in the lower left-hand panel of table 1. According to Benin and Johnson, this implies that "social mobility studies which use the residual correlation between brothers' educational attainments as an indicator of unmeasured family-background effects have overestimated the size of that effect for the other sex combinations" (1984:17-18).

Second, Benin and Johnson use common heredity, common environment, and most important, inter-sibling effects to explain the cross-group differences in the residual covariance. In their interpretation, the smallest residual correlation, that in OS-YB pairs, is the upper bound of a common heredity effect. Role modeling and facilitation are invoked to support the argument for inter-sibling effects. Role modeling should be greater between like-sex than mixed-sex sibling pairs. Older brothers are more likely than older sisters to become role models for younger siblings; older brothers are more likely than older sisters to contribute social contacts or resources that will facilitate the schooling of their younger siblings; older brothers will also be more likely to aid a younger brother than a younger sister (p. 12). These arguments lead to the prediction, apparently supported by the data (p. 16), that brother-brother pairs will have the largest residual covariance, that older sister-younger brother pairs will have the smallest residual covariance, and that the other two pair types will display intermediate values.

Benin and Johnson (1984:17) also re-estimate the same model for each pair type after classifying the sibling pairs by the age difference between the siblings. The overall pattern of differences in residual covariances by pair type is replicated within the classification by age differences, and there is some evidence that, within pair type, the residual covariance is larger for siblings who are separated in age by 3 or fewer years than among siblings who are more widely separated in age. Benin and Johnson say that the latter finding is more consistent with role modeling than with social facilitation as a mechanism of cross-sibling influence.

Although Benin and Johnson's work makes a novel and signficant contribution to our understanding of sibling resemblance, there are weak points in their argument and the supporting data analysis. First, Benin and Johnson have not actually estimated any cross-sibling effects on educational attainment. No such effects are identified in their model because they have no explanatory variables that can be said to affect the schooling of one sibling in a pair, but not of the other. The only evidence of simultaneity offered by Benin and Johnson is the difference in residual covariances among the groups; these covariance terms are not used in the identification or estimation of reciprocal effects.[4]

Second, although Benin and Johnson's exposition appears to invite the specification of a common factor in educational attainment, composed of measured and unmeasured family influences, they do not specify a model of this form. Figure 2 shows the path diagram of this more parsimonious model, in which the effects of the social background variables affect educational attainments by way of a common family factor. Here, the educational outcomes, $Y_1$ = OSED and $Y_2$ = YSED, appear as indicators of a common family factor, $\eta_1$, which is directly affected by the exogenous variables, $\xi_1$, $\xi_2$, and $\xi_3$. The effects of the common factor, $\eta_1$, on the siblings' educational attainments, $Y_1$ and $Y_2$, are expressed in the LISREL measurement model as loadings, $\lambda^y_{11}$ and $\lambda^y_{21}$. This is a MIMIC (multiple indicator-multiple cause) model (Hauser and Goldberger 1971, Jøreskog and Goldberger 1975). The specification of the common factor in figure 2 imposes a proportionality constraint on coefficients of the reduced form regressions of the $Y_i$ on the $\xi_j$, that is, equations of the form of those in figure 1, and the variance, $\psi_{11}$, of the disturbance of the family factor, $\zeta_1$, is equivalent to the residual covariance under the model of figure 1.

One might think this makes little difference in the analysis. Benin and Johnson are unable to reject a model in which the effects of socioeconomic variables on schooling ($\gamma_{ij}$) are the same (model 4 in table 1, p. 16), regardless of sex or ordinal position; their tests of equality in residual covariances are conditioned on this model, which is formally equivalent to a MIMIC model in which the loadings of $Y_1$ and $Y_2$ are both set equal to unity. However, this is just the problem. Although the MIMIC model does fit the Lincoln and NASIS data, it also provides greater statistical power than their model in tests of cross-group differences, and these yield different substantive findings than those of Benin and Johnson. Specifically, we find a consistent difference in effects of social background, such that younger children, regardless of sex, are less affected by parental status characteristics than their older siblings.

Furthermore, in Benin and Johnson's model, as well as in the revised model, we find that there are no significant differences between common variance components of the OB-YB, OB-YS, or OS-YS pair types; the only significant contrast is that between the OS-YB type and all others. This is inconsistent with Benin and Johnson's argument that cross-sibling effects are especially large in the OB-YB pairs. This finding occurs under Benin and Johnson's model as well as the MIMIC model, so one might wonder why we have proposed the latter. The MIMIC model is useful because it yields the finding of a signficant association of ordinal position with parental influence, and it also lays the basis for our presentation of an extended model that permits true cross-sibling effects.

Before reporting the findings of our reanalysis, we should note that we have been unable to reproduce one finding of Benin and Johnson. Our estimate of the likelihood ratio test statistic ($L^2$) for model 1 (36.83 with 18 *df*) is much larger than they have reported (11.9 in Table 1, p. 16). That is, we have to reject the hypothesis that $\Phi$, the variance-covariance matrix of parental status characteristics, is invariant across groups in the Lincoln sample. We do not know the reason for this anomaly, but it does not substantially affect other findings. The lower right-hand side of table 1 shows our estimates of the residual covariances in the Lincoln data, obtained from a model in which $\Phi$ was not constrained, and these are essentially the same as the estimates reported by Benin and Johnson from the constrained model. We have been able to reproduce their other results within the limits of rounding error in both the Lincoln and NASIS data; for example, the upper half of table 1 shows both sets of estimates of residual covariances in the NASIS data. In the remainder of our analysis, we have relied upon our findings, and we have not imposed the invariance restriction on $\Phi$ in the Lincoln data.

Table 2 reports tests of model fit in our reanalysis of the NASIS data. In the baseline MIMIC model, we specify that $\lambda^y_{11} = 1$, so within each group, $\Gamma$ gives the constrained effects of the $\xi$ on the schooling of older siblings, and $\lambda^y_{21}$ is a proportional adjustment of the elements of $\Gamma$ to express the constrained effects of the $\xi$ on the schooling of younger siblings in that pair type. The fit of the MIMIC model is exceptionally good ($L^2 = 8.33$ with 8 degrees of freedom).[5] The second model imposes the constraint that all the variances and covariances of parental background variables are the same across the four types of siblings, that is, $\Phi$ is invariant across pair types. The difference between the likelihood ratio test statistics of these two models is 6.1 with 18 degrees of freedom (*df*), so there are no signficant differences in the variances or covariances of parental characteristics across groups. The model of line C imposes the constraint that $\Theta^{\epsilon}$ is the same for all groups, regardless of sex or age. This is a sociologically interesting hypothesis; it says that factors unique to each sibling, and unrelated to common family background, contribute equally to the variance of schooling, regardless of sex or ordinal position. The contrast yields $L^2 = 11.43$ with 7 *df*, which is not statistically significant; this hypothesis cannot be rejected.

In the model of line D we pool effects of the parental background variables across the four groups, that is, $\Gamma$ is invariant. Given our normalization of $\Lambda_y$, this constraint implies that effects of social background variables are equal for older brothers and older sisters, but not necessarily for younger brother and younger sisters, for whom $\Lambda_y$ is permitted to differ across groups. The change in fit under this constraint is not statistically significant ($L^2 = 11.56$ with 9 *df*). The model in line E introduces the constraint that the effect of father's education on the common family background factor is the same as that of mother's education. The change in fit under this constraint is negligible.

In the model of line F we specify that the common family factor affects schooling of all younger siblings equally, regardless of sex. Again, the change in fit is small ($L^2 = 0.98$ with 3 *df*) we are unable to reject this model. In the model of line G we add the restriction that the effect of the common family factor on educational attainment is the same for all individuals, regardless of sex or birth order. Here, the deterioration of fit is large and statistically significant. Comparing line F and G, $L^2$ is 15.77 with only 1 *df*. The contrast is still significant, even if we compare line G with line E. That is, based on this model, we have reason to believe that common family effects differ between younger and older siblings. In our final model, for which estimates are shown in table 4, we find that the effect of the common factor on the schooling of younger siblings is only .73 times as large as its effect on older siblings, regardless of sex.[6]

In line H, we add the restriction to the model of line F that the variance of the disturbance in the common family factor, $\psi_{11}$ is equal across pair types. Violation of the corresponding restriction in the model of figure 1 is the basis for Benin and Johnson's major findings. Similarly, although the overall fit of the model is still good ($L^2$ = 49.06 with 49 *df*), the deterioration in fit relative to model F indicates that we should reject this model. However, when we compared the deviations of estimated from observed values of the residual covariances, the only group that stood out was the OS-YB pair. For this reason, in the model of line I we equate the other variances in the other 3 groups but not among OS-YB pairs. This model fits very well ($L^2$ = 40.18 with 48 *df*), and the contrasts with lines F and H clearly demonstrate that much of the discrepancy lies in the OS-YB pair. That is, model I is nested between models F and G: The closeness of fit to that of model F shows that the restriction on the common variances in the three other groups is not substantially violated by the data, and the deterioration in fit from model I to model H shows how different the common variance is among OS-YB pairs. Clearly, the lack of homogeneity between OS-YB pairs and the others is consistent with Benin and Johnson's hypotheses about the weak basis for role modeling and facilitation in such pairs. At the same time, the homogeneity of common variance in the other three groups is inconsistent with Benin and Johnson's hypothesis that cross-sibling effects are strongest between older and younger brothers.[7]

Our reanalysis of the Lincoln data follows exactly the same logic and procedures as that of the NASIS data. Thus, we focus primarily on the differences between the two. The first difference, as noted earlier, is that we reject the specification of cross group equality in the variances and covariances among exogenous variables (see line B of table 3). Second, the within-family variances of the siblings' educational attainments are not equal for all pair types; they are only equal for siblings of the same sex (line D). In Lincoln the within-family variance among men is much larger than that among women (5.340 and 2.300 respectively from table 4), while the within-family variance for among women in Lincoln is similar to that among men and women in the statewide sample (2.593). This large residual variance implies the importance of some other non-familial variables uniquely affecting the schooling of urban men; one possibility, suggested by an examination of the tables of mean years of schooling, is that men in the Lincoln sample may have been far more likely than all women or men from elsewhere in the state to have complete college or obtained some post-graduate schooling (Benin and Johnson 1984:20-21).

With these two exceptions, our findings from the NASIS data are very similar to those from the Lincoln data. First, the MIMIC model fits the data very well; background can be said to affect siblings' schooling through a common factor. Second, the effects of parental socioeconomic background can be pooled across sibling pair types. Third, the effects of father's education and mother's education are essentially the same. Fourth, the effect of the common family background factor on the younger sibling is only about 70 percent as large as its effect on his or her older sibling. Finally, the variance of the disturbance of the common family factor is much lower in OS-YB pairs than in the other 3 pair types. Common, unmeasured family factors affect OB-YB, OB-YS, and OS-YS pairs more than they affect OS-YB pairs.[8]

To summarize, the present reanalysis suggests modifications of Benin and Johnson's conclusions. While our results are not radically different from theirs, we have found, contrary to their contention, that previous empirical research has not led to significant overestimation of the effect of unmeasured common family background on schooling, except in OS-YB pairs. On the other hand, the low level of resemblance in OS-YB pairs raises an interesting sociological question. One plausible explanation, of course, is the theory of role modeling and facilitation proposed by Benin and Johnson, but their theory also predicts higher resemblance among brother pairs, for which the evidence is not persuasive. Additional empirical research will be needed to strengthen the evidence in support of Benin and Johnson's hypothesis or to provide an alternative explanation for the low level of resemblance in OS-YB pairs.

Cross-Sibling Effects on Educational Attainment

In the second part of our analysis, we have estimated models of resemblance in the educational attainment of brothers that permit direct estimates of cross-sibling effects. The data from the Kalamazoo brothers survey were originally collected and analyzed by Olneck (1976, 1977, 1979). The sample consists of men identifiable as brothers for whom there were sixth grade school records in the Kalamazoo, Michigan, public school system for the years 1928 to 1950. After the interviewing process, in which Olneck attempted to interview each member of each sibling pair, the data contained 692 individual respondents or 346 weighted pairs aged between 35 and 59 in 1973 for whom complete, self-reported data had been obtained. In the interviewing process, Olneck first chose one (primary) brother at random from each sibship to interview; the interview with this individual ascertained data on social background, current social and economic status, and the social and economic status of each of the respondent's brothers. If the selected person could not be interviewed, then another primary brother was selected. When the second or other brothers from the same sibship were interviewed, the social background data and self-reports of status variables were ascertained again, but proxy reports about siblings were not obtained from secondary brothers. Thus, for each sibship, there are reports from a primary respondent that include proxy data for his brother(s) and reports from each secondary brother that repeat measurements of social background, but do not include proxy data for other brothers.

The data for the present analysis differ slightly from those used by Olneck. First, we have verified all of the data against responses from the original questionnaires; in a very small number of cases, this has altered matches between records. We have recoded detailed occupational codes for fathers and brothers into the Duncan SEI (rescaled downward by a factor of 10 for convenience in reporting parameter estimates). Second, except in Olneck's very useful estimates of response error (1976:Ch. 4, 1977:149-50), his published analyses use only the self-reported responses, but we also use proxy reports from the primary brother. By so doing, we obtain two reports of the secondary brother's education attainment, and this permits us to estimate the structural model with corrections for response variability in self reports and proxy reports. Similarly, we use the reports of parental status characteristics by primary and secondary respondents to estimate their response variability. Third, in order to test hypotheses about directionality in cross-sibling effects, we order the members of each pair of brothers by age. This leads to a two-sample design, where the older brother is the primary respondent in one sample, and the younger brother is the primary respondent in the other sample. Finally, our effective sample size also differs from that reported by Olneck because we have followed the same weighting scheme used by Benin and Johnson. That is, we have chosen weights so the sum over pairs from each family is equal to one.[9] The weighted sample sizes are 163 where older brothers are primary respondents and 164 where younger brothers are primary respondents.

Our model, shown in figure 3, is similar in structure to that of figure 2 in hypothesizing a latent, common family background variable, $\eta_6$, that is affected by father's occupational status, $\eta_2$, father's educational attainment, $\eta_3$, mother's educational attainment, $\eta_4$, and an disturbance, $\zeta_6$. In turn, the common family factor affects the educational attainment of the primary brother, $\eta_7$, and of that of the secondary brother, $\eta_8$. Because the variables in figure 3 refer to primary or secondary respondents, while our substantive interest lies in comparisons between older and younger brothers, we sometimes refer to parameters with a sample designation. Where this is necessary, we use the superscript "(1)" to refer to parameters in the sample where older brothers are primary respondents and the superscript "(2)" where younger brothers are primary respondents. For example, in an initial specification of the model in figure 3 we normalize the effect of the common factor by fixing $\beta^{(1)}_{7,6} = \beta^{(2)}_{8,6} = 1$, where $\beta^{(1)}_{7,6}$ is the effect of the common family factor on the educational attainment of older primary brothers, and $\beta^{(2)}_{8,6}$ is the effect of the common family factor on the educational attainment of older secondary brothers.

There are three main differences between the models of figure 2 and figure 3. First, in figure 3 all of the variables in the structural model are latent constructs. Excepting the educational attainment of the primary brother (ED), there are two indicators of each construct; for that variable, we estimate the error variance in the self report by that estimated for self reports by secondary brothers (ZED). Second, in the model of figure 3, all of the constructs are specified to be endogenous variables ($\eta$) in the LISREL model, even the parental status characteristics. This is of no substantive importance, but it gives us greater freedom in specifying and testing correlated response errors (Hauser, Tsai, and Sewell 1983).

Third, and most important, we add constructs for the academic ability of each brother, $\eta_1$ and $\eta_5$, for which we use scores on the Otis or Terman group tests (IQ and ZIQ) and scores on the Metropolitan Achievement Test (EDQ and ZEDQ) as indicators.[10] Thus, the educational attainment of the primary brother, $\eta_7$, is affected directly by his academic ability, $\eta_1$, as well as by the family factor, $\eta_6$. Similarly, the educational attainment of the secondary brother, $\eta_8$, is affected directly by his academic ability, $\eta_5$, and by the family factor, $\eta_6$. Not only is the specification of each sibling's schooling improved by introducing an academic achievement construct, but that construct serves as an instrument for the estimation of simultaneous, cross sibling effects of educational attainment. That is, the primary brother's academic ability, $\eta_1$, affects the secondary brother's educational attainment, $\eta_8$, indirectly by way of his own educational attainment, $\eta_7$, and the secondary brother's academic ability, $\eta_5$, affects the primary brother's educational attainment, $\eta_7$, indirectly by way of his own educational attainment, $\eta_8$. There is no direct effect of $\eta_1$ on $\eta_8$, nor of $\eta_5$ on $\eta_7$.

Our specification of reciprocal effects has at least two serious weaknesses. First, the model is mis-specified to the degree that the academic ability of a man affects that of his brother directly or that the schooling of both brothers is affected by omitted common factors, correlated with academic ability, that explain the covariation of ability with schooling across brothers. Second, educational attainment measures the cumulative school experience of each brother; attainment occurs across time and school completion need not be ordered in time to match the ordinal position of members of the fraternal pair. Thus, if we take educational attainment to represent schooling at the time of final school leaving, there is no temporal basis for specifying either an effect of the schooling of the older brother on the younger or the reverse. For present purposes, and presumably consistent with past efforts to establish reciprocal effects of schooling among siblings, we take educational attainment as a proxy for decisions about the length of completed schooling which are not necessarily ordered in time as dates of school leaving. It will be useful for future research to test these critical assumptions by measuring additional background and intervening variables and, perhaps, by focusing on intervening variables, like aspirations, as the key endogenous variables that are subject to reciprocal influence.

Neglecting the distinction between older and younger brothers, within either subsample the structural model can be represented by the following equations:

$$\eta_6 = \beta_{62}\eta_2 + \beta_{63}\eta_3 + \beta_{64}\eta_4 + \zeta_6, \qquad (4)$$

$$\eta_7 = \beta_{71}\eta_1 + \beta_{76}\eta_6 + \beta_{78}\eta_8 + \zeta_7, \qquad (5)$$

and

$$\eta_8 = \beta_{85}\eta_5 + \beta_{86}\eta_6 + \beta_{87}\eta_7 + \zeta_8, \qquad (6)$$

where $E[\zeta_i \eta_j] = 0$ for $i = 6, 7, 8$ and $j = 1, 2, 3, 4, 5$; $E[\zeta_6 \zeta_7] = E[\zeta_6 \zeta_8] = E[\zeta_7 \zeta_8] = 0$; and $E[\zeta_7 \eta_6] = E[\zeta_8 \eta_6] = 0$. That is, $\zeta_6$, $\zeta_7$, and $\zeta_8$ are all uncorrelated with prior latent variables ($\eta_j$) on the right hand side of equations 4, 5, and 6, but $E[\zeta_7 \eta_8]$ and $E[\zeta_8 \eta_7]$ are not equal to zero because of simultaneity.[11] The critical theoretical assumption of our analysis of cross-sibling effects is that each brother's academic achievement only affects the other brother's schooling indirectly.

In our analyses of the Kalamazoo data, as in Benin and Johnson's study, each of the educational attainment variables is expressed in years of schooling, and each of the occupational status variables is expressed in the Duncan SEI. Table 5 gives the means, standard deviations, and the descriptions of variables in the two subsamples of brother pairs, and table 6 gives the correlations within each subsample. The entries in the lower left triangle of table 6 are for the subsample where the primary brother is older, and the entries in the upper right triangle are for the subsample where the secondary brother is older.

We began our analysis of the model in figure 3 by developing a measurement model that permits random measurement errors and selected correlations of measurement errors and that pools selected parameters between subsamples and/or between older and younger brothers. The loadings of both indicators of each parental variable, of the Otis or Terman test scores (but not those of the Metropolitan Achievement Tests), and of self and proxy reports of son's educational attainment are fixed at unity *a priori*. That is,

$$\lambda^y_{1,1} = \lambda^y_{3,2} = \lambda^y_{4,2} = \lambda^y_{5,3} = \lambda^y_{6,3} = \lambda^y_{7,4} = \lambda^y_{8,4} = \lambda^y_{9,5} = \lambda^y_{11,7} = \lambda^y_{12,8} = \lambda^y_{13,8} = 1 \qquad (7)$$

Since there is only one indicator of the primary brother's educational attainment, we use the estimate of error variance in the secondary brother's self report of educational attainment, $\theta^{\epsilon}_{11,11} = \theta^{\epsilon}_{12,12}$. Although the loadings of academic achievement (EDQ or ZEDQ) on the academic ability factor ($\eta_1$ or $\eta_5$) are free parameters of the model, preliminary analyses indicated that their values were quite similar, and we constrained all four of these parameters to be equal, that is,

$$\begin{aligned} &\lambda^{(1)}_{2,1} = \lambda^{(2)}_{2,1} \\ &= \lambda^{(1)}_{10,5} = \lambda^{(2)}_{10,5}. \end{aligned} \tag{8}$$

In order to obtain a satisfactory fit, it was necessary to introduce two sets of correlations between errors in variables; these are not shown in figure 3. First, as in previous research, we found clear evidence of correlation between errors in reports of mother's education and of father's education by the same respondent (Corcoran 1980, Hauser, Tsai, and Sewell 1983). Thus, we introduced free parameters for $\theta^{\epsilon}_{7,5}$ and $\theta^{\epsilon}_{8,6}$ in each subsample. Second, we found strong evidence of correlation between errors in academic achievement (EDQ or ZEDQ) and errors in reports of parental status variables. For this reason, we introduced free parameters in each subsample for $\theta^{\epsilon}_{2,3}$, $\theta^{\epsilon}_{2,4}$, $\theta^{\epsilon}_{2,5}$, $\theta^{\epsilon}_{2,6}$, $\theta^{\epsilon}_{2,7}$, $\theta^{\epsilon}_{2,8}$, $\theta^{\epsilon}_{10,3}$, $\theta^{\epsilon}_{10,4}$, $\theta^{\epsilon}_{10,5}$, $\theta^{\epsilon}_{10,6}$, $\theta^{\epsilon}_{10,7}$, $\theta^{\epsilon}_{10,8}$. These error covariances are not clearly patterned; for example, some are significant and positive, while others are significant and negative. Thus, for example, they do not appear to support either the argument that academic achievement is more highly correlated with parental status or that it is less highly correlated with parental status than one might expect from the factor model specified in figure 3. Finally, we constrained all of the error variances and covariances to be equal in the two subsamples.

The measurement model does not fit the data satisfactorily by conventional standards of statistical significance ($L^2$ = 144.47 with 99 *df*), but, given the number of observed variables in the model, their high degree of intercorrelation, the sample size, and the lack of systematic departures from fit, we find this a satisfactory baseline model. To evaluate fit, we have also used the *bic* statistic, proposed by Raftery (1986) as a rule of thumb based upon Bayesian theory for *a posteriori* tests:

$$bic = L_m^2 - df_m \times \log N, \qquad (9)$$

where $L_m^2$ is the likelihood ratio test statistic under Model *m*, $df_m$ is the degrees of freedom under model *m*, and N is the sample size. Satisfactory fit is indicated by a negative *bic* statistic, and in comparisons of fit between models, those with lower (more negative) *bic* statistics are preferred. For the measurement model, we find *bic* = -429.04, which suggests a better fit than is indicated by the nominal significance level.

Table 7 shows selected parameters of the measurement model, estimated in the final version of the structural model (line F of table 8). Our estimates of reliabilities are similar to those of Olneck (1976:186), except the validity of academic ability is less than the reliability assumed by Olneck (p. 172) on the basis of published estimates. One might expect to find a lower value in our model because we have treated measured mental ability as an indicator of a construct that also comprises academic achievement *per se*. The reliability of brothers' educational attainments is slightly less than that estimated by Olneck, and the reliability of the proxy report of education (ZEDPR) is as high as that of the self report (ZED). Excepting the correlation between a man's reports of his mother's and father's schooling, there is no evidence of correlation between errors in reports of the same variable by the same respondent in the Kalamazoo survey itself. However, there remains the puzzlingly inconsistent set of correlations between errors in academic achievement (EDQ and ZEDQ) and errors in the reports of parental status.[12]

Table 8 displays the fit of selected versions of the model of figure 3 in the two Kalamazoo subsamples. For each model, we report a brief description, the fit statistic ($L^2$), the degrees of freedom (*df*), and the *bic* statistic. As reported in line A, the initial version of the structural model yields $L^2$ = 145.96 with 103 *df*. There are 4 more degrees of freedom for error in the structural model because we estimated the measurement model by deleting $\eta_6$ from the setup of figure 3 and freeing all of the elements in the $\Psi$ matrix; that is, we ignored the restrictions in the structural model and permitted all of the latent constructs to be freely correlated. The additional four degrees of freedom come from the two proportionality constraints of the MIMIC model within each subsample.

In line B we report the fit of a model in which 26 constraints are added for consistency across subsamples within ages; that is, no age differences in parameters of the structural model are eliminated. In the **B** matrix the constraints are

$$\beta^{(1)}_{6,2} = \beta^{(2)}_{6,2},$$
$$\beta^{(1)}_{6,3} = \beta^{(2)}_{6,3}, \qquad (10)$$
and
$$\beta^{(1)}_{6,4} = \beta^{(2)}_{6,4},$$

which pertain to effects of parental statuses on the common family factor, and

$$\beta^{(1)}_{7,1} = \beta^{(2)}_{8,5},$$
$$\beta^{(1)}_{8,5} = \beta^{(2)}_{7,1},$$
$$\beta^{(1)}_{8,6} = \beta^{(2)}_{7,6}, \qquad (11)$$
$$\beta^{(1)}_{8,7} = \beta^{(2)}_{7,8},$$

and $$\beta^{(1)}_{7,8} = \beta^{(2)}_{8,7},$$

which pertain to age-specific effects of academic ability, of the common family factor, or of educational attainment on educational attainment. Within $\Psi$ we introduce the following constraints:

$$\begin{aligned}
\psi^{(1)}_{1,1} &= \psi^{(2)}_{5,5},\\
\psi^{(1)}_{5,5} &= \psi^{(2)}_{1,1},\\
\psi^{(1)}_{2,2} &= \psi^{(2)}_{2,2},\\
\psi^{(1)}_{3,3} &= \psi^{(2)}_{3,3},\\
\text{and}\quad \psi^{(1)}_{4,4} &= \psi^{(2)}_{4,4},
\end{aligned} \tag{12}$$

which pertain to variances of the exogenous variables, and

$$\begin{aligned}
\psi^{(1)}_{2,1} &= \psi^{(2)}_{5,2},\\
\psi^{(1)}_{5,2} &= \psi^{(2)}_{2,1},\\
\psi^{(1)}_{3,1} &= \psi^{(2)}_{5,3},\\
\psi^{(1)}_{5,3} &= \psi^{(2)}_{3,1},\\
\psi^{(1)}_{4,1} &= \psi^{(2)}_{5,4},\\
\psi^{(1)}_{5,4} &= \psi^{(2)}_{4,1},\\
\psi^{(1)}_{5,1} &= \psi^{(2)}_{5,1},\\
\psi^{(1)}_{3,2} &= \psi^{(2)}_{3,2},\\
\psi^{(1)}_{4,2} &= \psi^{(2)}_{4,2},\\
\text{and}\quad \psi^{(1)}_{4,3} &= \psi^{(2)}_{4,3},
\end{aligned} \tag{13}$$

which pertain to covariances among the exogenous variables, and

$$\begin{aligned}
\psi^{(1)}_{6,6} &= \psi^{(2)}_{6,6},\\
\psi^{(1)}_{7,7} &= \psi^{(2)}_{8,8},\\
\text{and}\quad \psi^{(1)}_{8,8} &= \psi^{(2)}_{7,7},
\end{aligned} \tag{14}$$

which pertain to variances of the endogenous variables. In this model, the test statistic increases only by $L^2 = 31.88$ with 26 *df*, which is not statistically significant, and the *bic* statistic decreases to -569.46; thus, in the remainder of the analysis we disregard the distinction between primary and secondary brothers and look only at age differences in parameters of the model.

In the model of line B, we find $\hat{\beta}^{(1)}_{8,7} = \hat{\beta}^{(2)}_{7,8} = .310$ with a standard error of .106; thus, the effect of older brother's educational attainment on younger brother's attainment is highly significant statistically. On the other hand, we find $\hat{\beta}^{(1)}_{7,8} = \hat{\beta}^{(2)}_{8,7} = .013$ with a standard error of .130, which is obviously not statistically significant. This would appear to support Benin and Johnson's contention that the predominant flow of influence is from older to younger brothers. However, the data cannot support this finding. In line C of table 8, we report the fit of a model that differs from line B only by introducing the constraint $\beta_{87} = \beta_{78}$ in both groups. This increases the test statistic only by 3.25 with 1 *df*, which is not statistically significant at the .05 level. The pooled estimate of cross-sibling influence is $\hat{\beta}_{87} = \hat{\beta}_{78} = .166$ with a standard error of .081. Thus, the data appear to support a finding of reciprocal influence in the educational attainments of brothers, sampling variability is too great to sustain a finding about the predominant direction of cross-sibling influence.

In the preceding analysis, we have not conditioned our test of equality in the reciprocal effects of brothers' educational attainments on any other equalities between older and younger brothers in parameters of the structural model; thus, we have tried selecting backward, from an unconstrained structural model. Now, as an additional test for consistency in our finding, we estimate a model of complete equality between older and younger brothers in the parameters of the structural model, and then we relax the assumption of equal reciprocal effects in educational attainment. The fit of these two models is reported in lines 4 and 5 of table 8. Relative to the model of line C, seven additional constraints are needed to impose complete equality between older and younger brothers in parameters of the structural model. These constraints can be written without reference to the group structure of the data:

$$\begin{aligned} \beta_{76} = \beta_{86} &= 1, \\ \beta_{71} &= \beta_{85}, \\ \psi_{11} &= \psi_{55}, \\ \psi_{21} &= \psi_{52}, \\ \psi_{31} &= \psi_{53}, \\ \psi_{41} &= \psi_{53}, \\ \text{and} \quad \psi_{77} &= \psi_{88}. \end{aligned} \tag{15}$$

These additional constraints lead to a non-significant deterioration in fit; comparing the models of lines 3 and 4, we find $L^2 = 12.01$ with 7 *df*. In the model of line D, the *bic* statistic declines to -600.54, which is preferable to that of any earlier model. One interesting constraint in this set is that of equality between older and younger brothers in the effect of the common family factor; recall that in the Nebraska data, this slope was about 70 percent as large for younger as for older siblings. In the Kalamzoo data, the point estimate was similar, $\hat{\beta}^{(1)}_{8,6} = \hat{\beta}^{(2)}_{7,6} = .686$; however, the standard error of this estimate was .255, which implies that the point estimate of the slope is not significantly different from unity.

Relative to the model of line D, we can eliminate the restriction $\beta_{87} = \beta_{78}$ to test the hypothesis of equality in reciprocal effects of educational attainment. The fit of this model is shown in line E, for which $L^2$ declines only by 1.71 with 1 *df*; thus, our finding of a non-signficant difference between the effect of the older brother's education on the younger brother's education and that of the younger on the older is confirmed by forward selection as well as backward selection.

Finally, drawing on our finding in the Nebraska data that the effects of mother's and father's educational attainments were not significantly different from one another, we add a similar constraint in the model of line F. The fit changes by $L^2 = 2.00$ with 1 *df*, so we confirm this finding in the Kalamazoo data.

The parameter estimates of the final structural model are reported in table 9. In the Kalamazoo data, parental status and academic ability account for 90 percent of the variance in the common family education factor. This may be compared with about 54 percent of the common variance explained by a vector of social background variables among male cohorts in the 1973 Occupational Changes in a Generation survey (Featherman and Hauser 1976:117), with 36 percent in the NASIS data, and with 53 percent in the Lincoln data. Evidently, the addition of academic ability to the model substantially increases the share of common family influence that can be explained. The common factor, in turn, accounts for 60 percent of the variance in individual educational attainment; the latter estimate is similar to those of Hauser and Featherman in OCG cohorts. However, we note that in our estimates for the Nebraska data, the share of common variance in educational attainment of brothers may be lower: 67 and 49 percent among older and younger brothers in the NASIS data and 44 and 31 percent among older and younger brothers in the Lincoln data.

In the final model, we estimate $\hat{\beta}_{87} = \hat{\beta}_{78} = .174$ with a standard error of .081. Thus, each increase of a year in a man's schooling leads directly to an increase of about one sixth of a year of schooling for his brother. When we take account of the causal loop specified in the model (see Duncan 1975), we find that the total effect of one brother's schooling on the other's is thus $.174/(1 - .174^2) = .179$, which appears to be substantially larger than the effect of either parent's educational attainment.[13] Of course, parental status is also represented in this model by father's occupational status. In any event, although the estimated reciprocal effect is subject to substantial sampling variability, the finding is surely strong and important enough to justify replication and extension of the present model to include additional antecedents of educational attainment.

Discussion

With regard to both major issues in the present analysis - the existence of cross sibling effects and their predominant direction - there are obvious needs and opportunities for additional research. The indirect evidence of mutual influence marshalled by Benin and Johnson appears also in our extension of their analysis, yet we find less evidence in the Nebraska data that cross sibling effects are larger among male siblings. Like the Benin-Johnson data for Nebraska, the Kalamazoo data appear to confirm a process of mutual influence between brothers in the attainment of schooling, but they provide evidence that is no more than suggestive with regard to a predominant flow of influence from older to younger siblings.

Another, less central issue in the analysis also deserves further attention. There is remarkable agreement across the three samples (NASIS, Lincoln, and Kalamazoo) that the effect of parental statuses on younger offspring is about 30 percent less than that among older offspring. This differential is statistically significant in the NASIS and Lincoln samples, but not in the Kalamazoo sample. We think there should be further tests of this finding; it may indicate a birth order effect quite different from the additive effects that are usually investigated. Why should parental status matter less for younger children? Do parents use more status-linked resources in the upbringing of older children? Are parents less concerned with the maintenance of status among younger children? Or is this evidence that cohort trends in the effects of social background on educational attainment, like those in mean levels of schooling (Hauser and Sewell 1985), occur within the family as well as between families?

Our findings about measurement error in the Kalamazoo data reinforce the evidence from earlier studies that proxy reports of status variables by adult offspring about one another are just about as accurate as self reports. However, both types of reports are subject to response variability, and there appears to be more response variability in proxy reports of parents' educational attainments than in those of adult offspring.[14] Thus, it would appear wise to continue efforts to collect sibling data (and other survey-based measurements of social stratification) using research designs that generate independent, multiple measurements of status variables.

Given the level of sampling variability in both the Nebraska and Kalamazoo data, it will be useful to replicate the present analysis among Wisconsin brothers and other available samples of sibling data.[15] For example, in data from the Wisconsin Longitudinal Study, it will be possible to estimate models like that of figure 3 for mixed sex pairs ordered by age, as well as among brothers. It will also be useful to add more social background variables to the models and to incorporate explicit measurements of adolescent aspiration. As noted above, these elaborations should make it possible to test assumptions of the simultaneous equation model used in the present analysis.

Finally, in future research, we may want to ask whether mutual influence among siblings may also be a source of resemblance in post-schooling success. For example, do men or women find ways to improve the employment or occupational chances of their siblings or to provide opportunities for better earnings? As this outline for future research becomes longer, we ought not to forget there may be more direct ways of learning about cross sibling influence: Better models and methods may help us to understand the sources of familial resemblance, but we ought also to ask people about their relationships with siblings.

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# FOOTNOTES

1. We are grateful to Michael Olneck for making his primary data available to us for this reanalysis. We take sole responsibility for any errors in their use herein.

2. For example, see Hauser and Sewell (1985) and sources cited therein.

3. All of our estimates have been obtained by maximum likelihood using procedures for multiple group analyses in LISREL VI (Jøreskog and Sørbom 1984).

4. See Duncan (1975:Chapters 5 and 6) for a useful, elementary exposition of identification in simultaneous equation models.

5. The likelihood ratio test statistic, $L^2$, is distributed as $\chi^2$ in large samples under the assumption of multivariate normality, and the difference between likelihood ratio test statistics of two nested models will be distributed as $\chi^2$ with degrees of freedom equal to the difference of those under the two models.

6. We have also estimated effects of sex on the loadings of educational attainments on the common family factor, and these are not significant in either the NASIS data or the Lincoln data.

7. We believe that our analysis is generous in its treatment of the hypothesis of cross group differences in the variance of the unmeasured family factor. We have found that our findings are very sensitive to the order in which tests of homogeneity are carried out, and in some of our analyses, there were no significant differences among groups in the variance of the unmeasured family factor.

8. Inspection of Table 4 suggests the possibility of pooling the estimation of many of the parameters across all eight groups. This would give yet more power to the statistical tests that we have carried out.

9. The weighting scheme depends on the number of siblings in each family, or to be more exact, the number of male siblings included in the data. The downward correction permits us to treat the data as if they were obtained by simple random sampling.

10. We have rescaled all of the test scores downward by a factor of 10 for convenience in reporting parameter estimates.

11. Again, note that one of $\beta_{76}$ and $\beta_{86}$ is initially fixed at unity within each subsample.

12. In retrospect, we believe these errors are patterned, namely, that the correlations occur only between errors in measured academic achievement and reports of parental status by the same individual. For example, if we revise the final model to delete all correlations between errors in one brother's academic achievement and errors in reports of parental status by the other brother, the change in fit is not statistically significant ($L^2 = 5.01$ with 6 *df*). A puzzle remains, which is that the remaining within-brother error correlations are opposite in sign for primary and secondary brothers. In the present analysis we have permited correlations with errors in reports of parental status by both brothers, and we do not believe this affects other findings in any significant way.

13. Because the variances of schooling among older and younger brothers are the same by construction, the estimated reciprocal effects are the same when the variables are expressed in raw or standardized form. That is, in the case of the reciprocal effects, the structural coefficients are also standardized

coefficients.

14. Compare the diagonal elements of table 7 pertaining to statuses of parents and offspring.

15. These might include the samples of siblings from the original Parnes NLS samples, those from the 1979 Ohio State-National Opinion Research Center sample of American youth, and sibling samples drawn from the High School and Beyond Study.

Table 1

Residual Family Covariances and Correlations:
Benin-Johnson Estimates and Re-Estimates, Nebraska Sibling Data

| Sample | Sibling Pair Combination | Benin-Johnson Estimates | | Re-Estimates | |
|---|---|---|---|---|---|
| | | Covariance | Correlation | Covariance | Correlation |
| NASIS (weighted data) | OB-YB | 2.485 | .464 | 2.445 | .459 |
| | OB-YS | 1.630 | .305 | 1.632 | .306 |
| | OS-YB | .487* | .091* | .486* | .091* |
| | OS-YS | 1.666 | .311 | 1.666 | .313 |
| Lincoln (weighted data) | OB-YB | 2.649 | .394 | 2.640 | .392 |
| | OB-YS | 1.803 | .268 | 1.791 | .266 |
| | OS-YB | .633* | .094* | .636* | .094* |
| | OS-YS | .922 | .137 | .925 | .137 |

Note: Coefficients marked with "*" are not statistically signficant at the .05 level. See text for explanation.

Table 2

Summary of Model Selection for
Comparisons between Sibling Pair Types: NASIS Sample

| | Model | $L^2$ | *df* | p | Contrasts | $L^2$ | *df* | p |
|---|---|---|---|---|---|---|---|---|
| A. | MIMIC model | 8.33 | 8 | <.40 | | | | |
| B. | A + $\Phi$ invariant | 14.33 | 26 | <.97 | (B)-(A) | 6.10 | 18 | <.99 |
| C. | B + $\Theta^{\epsilon}$ invariant and equal | 25.76 | 33 | <.81 | (C)-(B) | 11.43 | 7 | <.10 |
| D. | C + $\Gamma$ invariant | 37.32 | 42 | <.68 | (D)-(C) | 11.56 | 9 | <.20 |
| E. | D + $\gamma_{12} = \gamma_{13}$ | 37.58 | 43 | <.71 | (E)-(D) | .26 | 1 | <.90 |
| F. | E + $\Lambda_y$ invariant | 38.56 | 46 | <.77 | (F)-(E) | .98 | 3 | <.80 |
| G. | F + $\lambda^y_{21} = 1$ | 54.33 | 47 | <.22 | (G)-(F) | 15.77 | 1 | <.01 |
| | | | | | (G)-(E) | 16.75 | 4 | <.01 |
| H. | F + $\Psi$ invariant | 49.06 | 49 | <.47 | (H)-(F) | 10.50 | 3 | <.01 |
| I. | F + $\Psi$ invariant for OB-YB, OB-YS, and OS-YS pairs | 40.18 | 48 | <.78 | (I)-(F) | 1.62 | 2 | <.30 |
| | | | | | (H)-(I) | 8.88 | 1 | <.01 |

Note: See text for explanation.

Table 3

Summary of Model Selection for
Comparisons between Sibling Pair Types: Lincoln Sample

| | Model | $L^2$ | *df* | p | Contrasts | $L^2$ | *df* | p |
|---|---|---|---|---|---|---|---|---|
| A. | MIMIC model | 1.59 | 8 | <.99 | | | | |
| B. | A + $\Phi$ invariant | 38.42 | 26 | <.06 | (B)-(A) | 36.83 | 18 | <.01 |
| C. | A + $\Theta^{\epsilon}$ invariant and equal | 27.17 | 15 | <.03 | (C)-(A) | 25.58 | 7 | <.01 |
| D. | A + $\Theta^{\epsilon}$ equal for same sex | 6.84 | 14 | <.95 | (D)-(A) | 5.25 | 6 | <.50 |
| E. | D + $\Gamma$ invariant | 12.90 | 23 | <.95 | (E)-(D) | 6.06 | 9 | <.70 |
| F. | E + $\gamma_{12} = \gamma_{13}$ | 15.90 | 24 | <.89 | (F)-(E) | 3.00 | 1 | <.10 |
| G. | F + $\Lambda_y$ invariant | 20.36 | 27 | <.82 | (G)-(F) | 4.46 | 3 | <.30 |
| H. | G + $\lambda^y_{21} = 1$ | 30.67 | 28 | <.33 | (H)-(G) | 10.31 | 1 | <.01 |
| | | | | | (H)-(F) | 14.77 | 4 | <.01 |
| I. | G + $\Psi$ invariant | 29.73 | 30 | <.48 | (I)-(G) | 9.37 | 3 | <.05 |
| J. | G + $\Psi$ invariant for OB-YB, OB-YS, and OS-YS pairs | 23.14 | 29 | <.77 | (J)-(G) | 2.78 | 2 | <.30 |
| | | | | | (I)-(J) | 6.59 | 1 | <.01 |

Note: See text for explanation.

Table 4

Estimates of Structural Parameters in Models of Sibling Resemblance in Educational Attainment: Nebraska Data

| Sample and Pair Type | $\psi_{11}$ | $\theta^{\epsilon}_{11}$ | $\theta^{\epsilon}_{22}$ | $\gamma_{12} = \gamma_{13}$ | $\gamma_{11}$ | $\lambda^{y}_{21}$ |
|---|---|---|---|---|---|---|
| NASIS: | | | | | | |
| OB-YB, OB-YS, and OS-YS | 2.623 (.386) | 2.593 (.175) | 2.593 (.175) | .228 (.023) | .143 (.046) | .733 (.057) |
| OS-YB | .977 (.373) | 2.593 (.175) | 2.593 (.175) | .228 (.023) | .143 (.046) | .733 (.057) |
| Lincoln: | | | | | | |
| OB-YB | 2.018 (.378) | 5.340 (.426) | 5.340 (.426) | .192 (.029) | .242 (.055) | .768 (.066) |
| OB-YS | 2.018 (.378) | 5.340 (.426) | 2.300 (.233) | .192 (.029) | .242 (.055) | .768 (.066) |
| OS-YS | 2.018 (.378) | 2.300 (.233) | 2.300 (.233) | .192 (.029) | .242 (.055) | .768 (.066) |
| OS-YB | .645* (.375) | 2.300 (.233) | 5.340 (.426) | .192 (.029) | .242 (.055) | .768 (.066) |

Note: All estimates are significant at the .05 level except the one marked with "*." Parenthetic entries are estimated standard errors. See text for explanation.

Table 5

Means, Standard Deviations and Descriptions of Variables:
Kalamazoo Brother Pairs

| | | Older Primary | | Older Secondary | | |
|---|---|---|---|---|---|---|
| $Y_i$ | Name | Mean | Std Dev | Mean | Std Dev | Description |
| $Y_1$ | IQ | 10.110 | 1.430 | 9.811 | 1.457 | Otis Test Score |
| $Y_2$ | EDQ | 9.839 | 1.226 | 9.409 | 1.184 | Metropolitan Score |
| $Y_3$ | POPDUNC | 3.856 | 2.243 | 3.716 | 2.200 | Father's Occupation |
| $Y_4$ | ZPOPDUNC | 3.852 | 2.190 | 3.742 | 2.147 | Father's Occupation |
| $Y_5$ | POPED | 9.909 | 3.288 | 9.432 | 3.369 | Father's Education |
| $Y_6$ | ZPOPED | 9.852 | 3.436 | 9.355 | 3.135 | Father's Education |
| $Y_7$ | MOMED | 10.851 | 2.908 | 10.059 | 3.304 | Mother's Education |
| $Y_8$ | ZMOMED | 10.536 | 2.971 | 10.116 | 3.023 | Mother's Education |
| $Y_9$ | ZIQ | 10.254 | 1.640 | 10.023 | 1.617 | Otis Test Score |
| $Y_{10}$ | ZEDQ | 9.761 | 1.317 | 9.759 | 1.322 | Metropolitan Score |
| $Y_{11}$ | ED | 13.020 | 2.490 | 12.570 | 2.515 | Education |
| $Y_{12}$ | ZED | 13.646 | 2.916 | 13.087 | 2.779 | Education |
| $Y_{13}$ | ZEDPR | 13.703 | 2.794 | 12.869 | 2.689 | Education by Proxy |

NOTE: Father's occupational status is in Duncan's SEI (scaled downward by a factor of 10). Educational attainments are in years of schooling. Otis and Metropolitan scores have also been scaled downward by a factor of 10. Names of variables pertaining to the secondary brother begin with "Z."

Table 6

Correlation Matrices: Kalamazoo Brothers

| | $Y_1$ | $Y_2$ | $Y_3$ | $Y_4$ | $Y_5$ | $Y_6$ | $Y_7$ | $Y_8$ | $Y_9$ | $Y_{10}$ | $Y_{11}$ | $Y_{12}$ | $Y_{13}$ |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| $Y_1$ | 1.000 | .814 | .275 | .256 | .166 | .201 | .228 | .210 | .432 | .451 | .464 | .357 | .344 |
| $Y_2$ | .814 | 1.000 | .248 | .240 | .294 | .293 | .322 | .281 | .423 | .454 | .509 | .371 | .367 |
| $Y_3$ | .278 | .200 | 1.000 | .780 | .525 | .481 | .440 | .405 | .274 | .377 | .401 | .417 | .391 |
| $Y_4$ | .222 | .251 | .738 | 1.000 | .485 | .481 | .414 | .385 | .229 | .350 | .429 | .397 | .435 |
| $Y_5$ | .276 | .319 | .494 | .479 | 1.000 | .776 | .681 | .518 | .385 | .465 | .403 | .468 | .435 |
| $Y_6$ | .223 | .180 | .483 | .428 | .754 | 1.000 | .507 | .509 | .302 | .330 | .405 | .442 | .413 |
| $Y_7$ | .208 | .288 | .422 | .380 | .573 | .444 | 1.000 | .741 | .336 | .376 | .326 | .398 | .388 |
| $Y_8$ | .276 | .310 | .419 | .339 | .421 | .526 | .769 | 1.000 | .272 | .221 | .297 | .353 | .307 |
| $Y_9$ | .523 | .532 | .313 | .271 | .266 | .190 | .320 | .315 | 1.000 | .867 | .412 | .575 | .547 |
| $Y_{10}$ | .556 | .612 | .200 | .264 | .221 | .109 | .285 | .351 | .844 | 1.000 | .514 | .631 | .644 |
| $Y_{11}$ | .630 | .629 | .409 | .367 | .407 | .420 | .365 | .450 | .394 | .449 | 1.000 | .578 | .597 |
| $Y_{12}$ | .515 | .477 | .431 | .344 | .395 | .341 | .454 | .436 | .600 | .590 | .570 | 1.000 | .877 |
| $Y_{13}$ | .510 | .471 | .402 | .332 | .362 | .277 | .424 | .418 | .612 | .586 | .570 | .916 | 1.000 |

Note: Entries in upper right triangle are for the subsample with older secondary brothers, while entries in lower left triangle are for the subsample with younger secondary brothers. See Table 5 for detailed descriptions of variables.

Table 7

Selected Parameters of Final Measurement Model:
Educational Attainments of Kalamazoo Brothers

| | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 | 9 | 10 | 11 | 12 | 13 |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Theta Epsilon ($\Theta^{\epsilon}$): | | | | | | | | | | | | | |
| 1 | 0.44 | -- | -- | -- | -- | -- | -- | -- | -- | -- | -- | -- | -- |
| 2 | -- | 0.22 | -2.05 | -0.14 | 2.39 | 0.37 | 2.53 | 1.67 | -- | -- | -- | -- | -- |
| 3 | -- | -0.30 | 1.13 | -- | -- | -- | -- | -- | -- | -0.92 | -- | -- | -- |
| 4 | -- | -0.02 | -- | 1.19 | -- | -- | -- | -- | -- | 1.23 | -- | -- | -- |
| 5 | -- | 0.36 | -- | -- | 2.45 | -- | 5.45 | -- | -- | 0.43 | -- | -- | -- |
| 6 | -- | 0.05 | -- | -- | -- | 2.73 | -- | 2.13 | -- | -1.89 | -- | -- | -- |
| 7 | -- | 0.36 | -- | -- | 0.62 | -- | 2.48 | -- | -- | -0.18 | -- | -- | -- |
| 8 | -- | 0.25 | -- | -- | -- | 0.24 | -- | 2.14 | -- | -0.87 | -- | -- | -- |
| 9 | -- | -- | -- | -- | -- | -- | -- | -- | 0.52 | -- | -- | -- | -- |
| 10 | -- | -- | -0.15 | 0.20 | 0.07 | -0.31 | -0.03 | -0.15 | -- | 0.17 | -- | -- | -- |
| 11 | -- | -- | -- | -- | -- | -- | -- | -- | -- | -- | 0.82 | -- | -- |
| 12 | -- | -- | -- | -- | -- | -- | -- | -- | -- | -- | -- | 0.82 | -- |
| 13 | -- | -- | -- | -- | -- | -- | -- | -- | -- | -- | -- | -- | 0.78 |
| Reliability ($\theta^{\epsilon}_{ii}/\sigma_{ii}$): | | | | | | | | | | | | | |
| | 0.79 | 0.85 | 0.78 | 0.75 | 0.77 | 0.77 | 0.71 | 0.76 | 0.81 | 0.91 | 0.87 | 0.90 | 0.90 |

Note: Diagonal entries are elements of $\Theta^{\epsilon}$, that is, error variances. Below-diagonal entries are correlations between errors in variables; above-diagonal entries are t-statistics of corresponding error correlations. Variable identifications are 1 = IQ, 2 = EDQ, 3 = POPDUNC, 4 = ZPOPDUNC, 5 = POPED, 6 = ZPOPED, 7 = MOMED, 8 = ZMOMED, 9 = ZIQ, 10 = ZEDQ, 11 = ED, 12 = ZED, 13 = ZEDPR. See table 5 for description of variables. Estimates and standard errors are based on the final structural model (line F of table 8). There is one additional parameter of the measurement model, $\lambda_{2,1} = \lambda_{10,5} = .865$ with a standard error of .029.

Table 8

Summary of Model Selection:
Kalamazoo Brother Pairs

| Model | *df* | $L^2$ | $L^2/df$ | *bic* |
|---|---|---|---|---|
| A. Baseline MIMIC model | 103 | 145.96 | 1.42 | -450.72 |
| B. Model A plus within-age constraints | 129 | 177.87 | 1.38 | -569.46 |
| C. Model B plus $\beta_{76} = \beta_{67}$ | 130 | 181.09 | 1.39 | -572.00 |
| D. Model C plus equality across ages | 137 | 193.10 | 1.41 | -600.54 |
| E. Model D with $\beta_{76} \neq \beta_{67}$ | 136 | 191.39 | 1.41 | -596.46 |
| F. Model D plus $\beta_{63} = \beta_{64}$ | 138 | 193.39 | 1.40 | -606.05 |

Note: See text for explanation.

Table 9

Estimated Parameters of the Structural Model of Sibling Resemblance and Cross-Sibling Effects on Educational Attainment: Kalamazoo Brothers

| | Dependent Variable | | |
|---|---|---|---|
| Independent Variable | $\eta_6$ | $\eta_7$ | $\eta_8$ |
| $\eta_1$ = older brother's academic ability | | .866 (.082) | |
| $\eta_2$ = father's occupational status | .156 (.062) | | |
| $\eta_3$ = father's educational attainment | .091 (.025) | | |
| $\eta_4$ = mother's educational attainment | .091 (.025) | | |
| $\eta_5$ = younger brother's academic ability | | | .866 (.082) |
| $\eta_6$ = common family education factor | | 1.000* | 1.000* |
| $\eta_7$ = older brother's educational attainment | | | .174 (.081) |
| $\eta_8$ = younger brother's educational attainment | | .174 (.081) | |
| $\psi_{ii}$ = disturbance variance | .051 (.479) | 2.482 (.466) | 2.482 (.466) |
| $R^2$ | .902 | .606 | .606 |

Note: Values in parentheses are standard errors. Fixed coefficients, for which there are not standard errors, are indicated by "*."

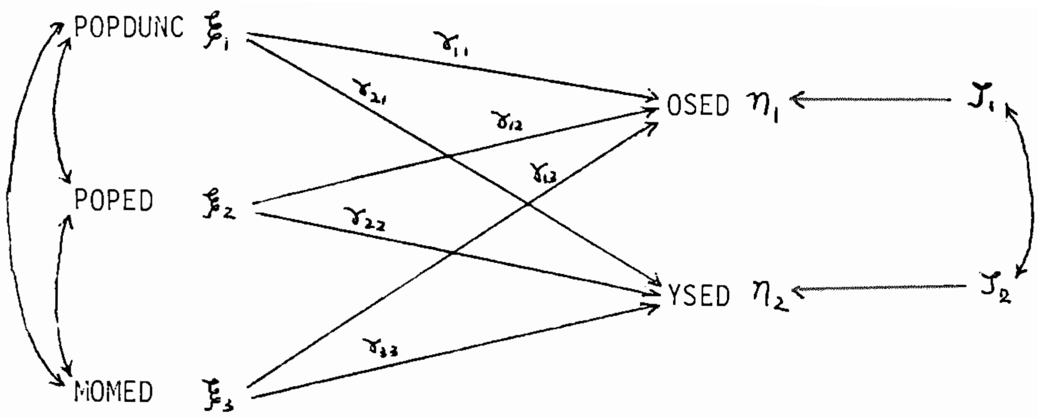


Fig. 1 Structural Model of Sibling Resemblance in Educational Attainment as Hypothesized by Benin and Johnson

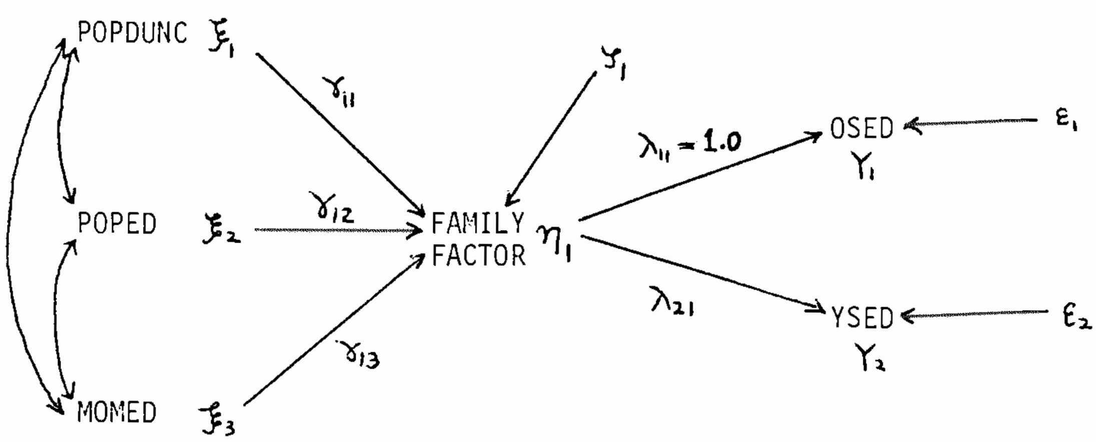


Fig. 2 Alternative Structural Model of Sibling Resemblance in Educational Attainment

where OSED = Older sibling's educational attainment
YSED = Younger sibling's educational attainment
POPDUNC = Father's occpuational status
POPED = Father's educational attainment
MOMED = Mother's educational attainment

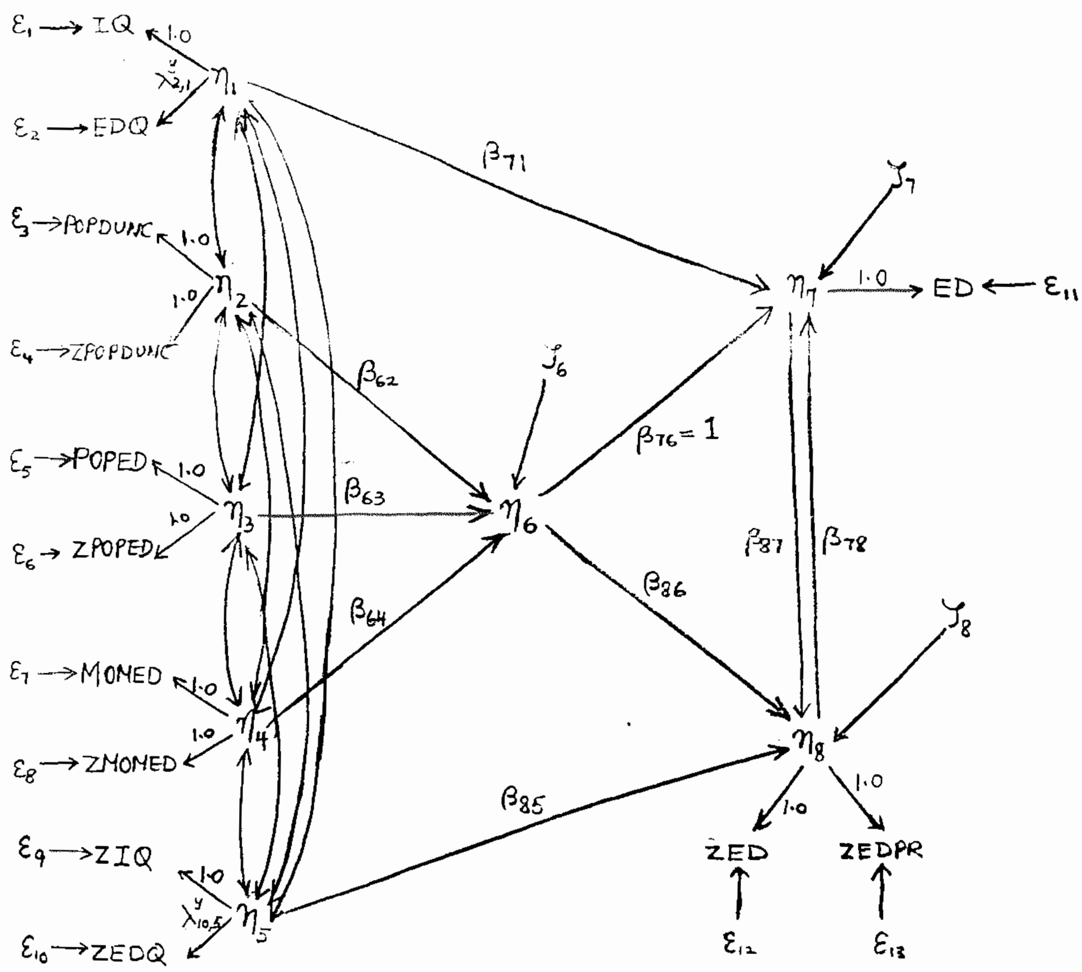


Figure 3. Structural Model of Sibling Resemblance and Cross Sibling Effects on Educational Attainment

Mailing Address:

Center for Demography and Ecology
University of Wisconsin
1180 Observatory Drive
Madison, Wisconsin 53706-1393
U.S.A.