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ABSTRACT

This study uses measurements of social background variables, mental ability, educational attainment, occupational status, and earnings among male, Wisconsin high school graduates and a random sample of their brothers to develop and interpret simple models of socioeconomic achievement. The study was designed to contribute to the data on the influence of education on social mobility. Maximum likelihood estimates for pooled samples of fraternal pairs from the Wisconsin Longitudinal Study where both brothers were interviewed (532 pairs) and where only one brother was interviewed (N=928) were obtained. No evidence was found that the effects of family background lead to bias in the effect of mental ability on schooling or in the effects of schooling on occupational status or earnings. These findings are confirmed by a reanalysis of another set of data on 346 fraternal pairs from Kalamazoo, Michigan. Family background was found to have a large independent effect on ability, schooling, and, to a lesser degree, socioeconomic attainment. Although the models used are simple, it is felt that they provide a useful framework for the development of more complete models and comparative analyses of family effects on offspring who differ in sex, age, or ability. Several tables display the data. (Author/CG)

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ABSTRACT

Data from fraternal pairs are used to develop simple structural equation models of the effects of measured and unmeasured family background factors, mental ability, and schooling on occupational status and earnings. These models incorporate corrections for response variability, and they permit direct comparisons of within- and between-family regressions. We obtain maximum likelihood estimates of these models for pooled samples of fraternal pairs from the Wisconsin Longitudinal Study where both brothers were interviewed (N = 532) and where only one brother was interviewed (N = 928). We find no evidence that the effects of family background lead to bias in the effect of mental ability on schooling or in the effects of schooling on occupational status or earnings. These findings are confirmed in a reanalysis of Olneck's data for fraternal pairs (N = 346) from Kalamazoo, Michigan. At the same time, family background does have large independent effects on ability, schooling, and - to a lesser degree - socioeconomic attainment. Although the present models are very simple, we believe that they provide a useful framework for the specification of more complete models and comparative analyses of family effects on offspring who differ in sex, age or ability.

1.0 INTRODUCTION

Sociologists and economists have long recognized the importance of measuring the effects of schooling. Its influence on such measures of success as occupational status and earnings serves on the one hand as an indicator of the role of educational institutions in fostering (or hampering) social mobility and on the other hand as an indicator of the productivity of personal and public investments in schooling. At the same time, it is well known that social and economic success may depend directly upon personal characteristics and conditions of upbringing that also affect the length and quality of schooling.[1] For these reasons, it is by no means obvious that an association of schooling with social or economic success can be interpreted in causal terms, and many studies have attempted to determine the degree to which such causal inferences are warranted.

The effects of background, broadly conceived, on achievement can be taken into account by modeling the similarity of siblings. That is, a research design based upon sibling pairs (or n-tuples) permits a decomposition of the cross-sibling variance-covariance matrix into "between-family" and "within-family" components. If fraternal differences in schooling lead to differences in adult success, we can be confident that the association of schooling with success is not merely an artifact of the tendency of school success to run in families that are also economically successful. Statistical controls for common family influences are by no means sufficient to eliminate problems of omitted-variable bias in the measured effects of schooling; other, sibling-specific factors that jointly determine schooling and economic success must also be controlled. Still, the prospect of controlling common family influences has helped to motivate a number of studies of the stratification process that are based upon samples of siblings, rather than of the general population, beginning with Blau and Duncan (1967:316-328) and most notably in the two major studies by Jencks and his associates (Jencks et al. 1972, 1979). Moreover, the use of sibling data has called attention to the larger issue of the role of families in the stratification process.

In his important review of sibling models and data in economics, Griliches (1979) has noted a potentially significant methodological twist in the use of sibling-based research designs (also, see Griliches 1977). In a regression, say, of earnings on schooling, random response variability in schooling leads to more (downward) bias in the within-family estimator than in a naive regression that ignores family effects.[2] This occurs in a random effects model because measurement error will increase the variability of individual responses, but not of their family components. Consequently, a fixed component of error variance in schooling is larger relative to within-family variance than to total variance, resulting in a proportionately larger downward bias in the within-family regression than in the total regression. Since we expect the omission of family background to lead to upward bias in the schooling coefficient, the biases attributable to omitted background variables and to measurement error are probably opposite in effect, and it is necessary to correct both at the same time.

In the late 1960s, little was known about the sensitivity of estimated parameters of models of the stratification process to response variability. Since then, there have been a number of efforts to measure the reliability or validity of survey reports of socioeconomic variables,[3] and -- contrary to some expectations (Bowles 1972) -- they have not led to massive downward revisions in estimates of the effects of schooling on occupational or economic success. At the same time, Griliches' argument shows it is important to correct for response variability in within-family regressions of socioeconomic success on schooling.[4]

The present analysis uses measurements of social background variables, mental ability, educational attainment, occupational status, and earnings among male, Wisconsin high school graduates and a random sample of their brothers to develop and interpret simple models of socioeconomic achievement that incorporate a family variance component structure and that also correct for response variability. The models include parameters for within-family and between-family regressions, and thus make it straightforward to test for heterogeneity in those regressions.[5] While Griliches' work has made it difficult for other researchers to ignore the problem of response error, most of his model building efforts have been based upon trade-offs between the specification of parameters for response variability and for lagged effects of background variables in the structural equation model. We have avoided this problem by obtaining multiple measurements of most variables in our models for one or both members of each sibling pair and embedding the structural model of achievement within a model of response variability.

Another methodological problem that has plagued studies of sibling resemblance is lack of statistical power in estimating the schooling coefficient and possible biases in it. Sibling samples have often been small, perhaps as a corollary to the fact that they have so often been drawn incidentally to some other research purpose. For example, consider Gorseline's (1932) sample of Indiana brothers (N = 156), Jencks et al.'s (1979:Appendix A2.1) samples of brothers from Project Talent (N = 99) and from an National Opinion Research Center survey (N = 150), Brittain's (1977) set of Cleveland brothers (N = 60), the brother pairs that Chamberlain and Griliches (1977) drew from the National Longitudinal Studies (N = 161), and those that Corcoran and Datcher (1981) have drawn from the Panel Study of Income Dynamics (N = 206). We doubt that any other active area of social scientific research is so dependent on such meagre scraps of survey data. Even if there were a consensus on model specification, and even if selectivity, coverage, and data quality problems were nonexistent, we would still not be surprised to find that "different samples appear to be telling different stories" (Griliches 1979:S39). This problem is revealed forcefully in Hauser's (1984a) analyses of the regression of occupational status on schooling, where large (20 to 40 percent) family biases prove not to be statistically significant until they are pooled across more than 4000 sibling pairs.

Our contribution to the solution of this problem - aside from having collected data for a relatively large number of sibling pairs - is to suggest ways of using available data more efficiently. In the present analysis, we pool maximum likelihood estimates of models of fraternal resemblance in ability, schooling, occupational status, and earnings across two subsamples of brother pairs from the Wisconsin Longitudinal Study (Sewell and Hauser 1980). In one set of pairs (N = 928), we have complete data for a primary respondent, but only his proxy reports about schooling and occupational status of the other brother. In the other set of pairs (N = 532), we have complete, self-reported data for both members of each pair, plus the proxy reports about the other brother. Using the sample with complete data to estimate response variability in the proxy reports and - in some cases - assuming error variances are the same for primary respondents and their brothers, we have been able to use all of the information in the sample with incomplete data and thus obtain more efficient estimates. This method will make it possible to increase the efficiency of existing samples.[6] It may also lead to more efficient and less expensive ways of collecting sibling data.

At the same time, we have not attempted here to solve all of the methodological problems in modeling sibling resemblance. Readers will see that our models are obviously incomplete or misspecified in several ways.[7] To provide a baseline for further work, we have made our analyses strictly comparable to those of Olneck (1977). Thus, we have used only father's education, father's occupational status, and number of siblings to measure family background, and we have not included a measure of experience in our earnings equations.[8] Moreover, we have not seriously attempted - beyond our corrections for response variability in schooling - to solve the problem of correlation between schooling and the disturbance in earnings (or occupational status); that is, we have done nothing about the possibility of simultaneity between schooling and earnings (or occupational status).[9] We do not think these problems are intractable, given our data and modeling framework, but we also think that the present, modest analyses will be useful in their own right.

Following a brief description of the Wisconsin data, we describe our measurement model and report initial estimates of bias due to social background and mental ability in the effects of schooling on occupational status and earnings. These estimates incorporate corrections for response variability in measured social background characteristics, but specify no other common family factors.[10] Next, we present our structural model of family background and achievement and report some results based upon it, including comparisons of estimates from our complete data sample and from our pooled sample of complete and incomplete data. Last, we cross-validate the Wisconsin findings using 346 pairs of Kalamazoo brothers with complete data.

2.0 THE WISCONSIN SIBLING DATA

The Wisconsin Longitudinal Study has followed a random sample of more than 10,000 men and women who were seniors in the state's public, private, and parochial high schools in 1957 (Sewell and Hauser 1980). Late in the senior year, detailed information was collected on the social origins and the educational and occupational aspirations of the students. These data were supplemented by reports of father's occupation and parents' incomes from state income tax records, by mental ability scores from the State Testing Service, and by ranks in high school class supplied by the individual schools. There were successful follow-up surveys of the total sample (with approximately 90 percent response rates) in 1964 and in 1975. The first follow-up, a mail survey of the parents of the primary respondents, yielded educational histories and reports of marital status, occupation, and military service.

The 1975 telephone survey, conducted when the respondents were about 36 years old, yielded additional first-hand reports of social background characteristics, educational and occupational experiences, marital and fertility histories, and social participation. The questionnaire included a roster of siblings of the respondent, including date of birth, sex, and educational attainment. For a randomly selected sibling, full name, current address and occupation were ascertained, along with the name and location of the last high school that person attended in Wisconsin. In 1977, telephone interviews were conducted with a highly stratified subsample of these selected siblings.[11] Of 879 brothers of male primary respondents in this subsample, telephone interviews were completed with 749 (85.2 percent).[12] In addition, using identifying information from the 1975 and 1977 interviews, we were able to locate mental ability scores for almost 80 percent of the subsample of siblings in the records of the State Testing Service.

For the present analysis, we have selected two samples of brother pairs. In the subsample where brothers were interviewed in 1977, 532 pairs meet our criteria for inclusion in the analysis; hereafter, we refer to this as the complete sample. There is a second sample of pairs in which supplementary information about a brother was collected in the 1975 survey, but no mental ability scores were collected, nor any self-reported data from the second brother in the pair. In this incomplete sample, there are 928 pairs of brothers. In the complete sample, the brother of the primary respondent must have been 20 to 55 years old in 1975, and each brother must have met the following criteria: not enrolled in school at his survey date, employed within the past 5 years, positive earnings in the past year, worked 10 or more weeks in the past year, usually worked 10 or more hours per week, and an imputed wage rate greater than $2.00 per hour. In the incomplete sample, the same criteria were applied to the primary respondent, but the brother need only have been 20 to 55 years old with an occupation reported in 1975. Within each sample we have treated item nonresponse by pairwise-present estimation of the moments.[13] Although the sample with complete data is evidently more highly selected, there are few differences in measured characteristics between the two samples.[14]

3.0 THE MEASUREMENT MODEL

Table 1 lists the variables used in the present analysis and gives the marginal sample sizes, means, standard deviations and estimated reliabilities.[15] The reliabilities are maximum likelihood estimates that were obtained by pooling across the complete and incomplete samples.[16] They are based upon an unrestricted reduced form model of schooling, occupational status, and earnings, which is shown schematically by a path diagram in Figure 1. For the moment, we need only consider the measurement equations of the model, which are represented in the figure by arrows linking ξs with Xs in the case of exogenous variables and by arrows linking ηs with Ys in the case of endogenous variables. For variables that were not measured in the incomplete sample, the paths from unbservables (ξ or η) to observables (X or Y) are shown as dotted lines. To simplify this analysis, we have specified uncorrelated response errors throughout. In earlier work we had postulated correlations between errors in the reports of some of the indicators (Hauser, Tsai and Sewell 1983, Hauser and Mossel 1984), but there was little evidence of correlated errors among variables used in the present analysis.

Throughout the measurement model, we have specified equalities between selected parameters pertaining to the primary respondent and his brother. First, we have always chosen reference indicators - that is, variables with unit loadings on the true variables - which have been measured on the same scale and using the same questions and coding procedures for each brother. This guarantees that the unobservable true variables are in the same metric for each member of each brother pair. Second, in each of these cases we have specified that the error variances are the same for the primary respondent and his brother. This is not necessary to insure comparability of slopes between brothers, but it is logically consistent with the "borrowing" of error variance estimates, when they are identified for only one member of the pair.

In some of the models estimated for this paper, but not in the model for which reliability estimates are reported in Table 1, we have also specified equality between brothers in corresponding variances and covariances of unobservable variables. We refer to such equalities as symmetries and to their absence as asymmetries. The specification of symmetry in the structural model ignores some known differences between the populations of primary respondents and brothers. Primary respondents were almost all born in 1939, and all graduated from high school; their brothers varied widely in age and were selected only incidentally for high school completion. In earlier analyses (Hauser and Mossel 1984, Hauser 1984a) we looked closely for asymmetries in structure between primary respondents and siblings, and we found little evidence of them. Thus, although our models provide a useful analytic framework for investigating differential treatment of offsprin (Hauser 1984b), we have chosen to simplify the analysis at some points by ignoring asymmetries between brothers.

There are three indicators of father's education in years, X_1 and X_2, which were reported by the primary respondent in 1957 and 1975, respectively, and X_3, which was reported by the selected brother in 1977. The measurement equations are

$$X_1 = \lambda_{11}\xi_1 + \delta_1, \quad (1)$$

$$X_2 = \xi_1 + \delta_2, \quad (2)$$

and

$$X_3 = \xi_1 + \delta_3, \quad (3)$$

where the errors, δ_i, are uncorrelated with one another and uncorrelated with the true variable, ξ_1, and the error variances are subject to the restriction, $\theta_2^{(\delta)} = \theta_3^{(\delta)}$. Note also that ξ_1 has a unit slope both in equations 2 and 3; we specified these restrictions because the schooling items were identical in the 1975 and 1977 surveys.

In the case of father's occupational status, X_4 was obtained from state tax records around 1957, and coded using materials from the 1950 Census. The other two reports, X_5 and X_6, were ascertained from the primary respondent and the selected brother in the 1975 and 1977 surveys, respectively; they were classified using materials from the 1970 Census. All occupation codes were translated into the Duncan SEI (Duncan 1961, Hauser and Featherman 1977:Appendix B).[17] Our specification of father's occupational status parallels that of father's education:

$$X_4 = \lambda_{42}\xi_2 + \delta_4, \quad (4)$$

$$X_5 = \xi_2 + \delta_5, \quad (5)$$

and

$$X_6 = \xi_2 + \delta_6, \quad (6)$$

where $\theta_5^{(\delta)} = \theta_6^{(\delta)}$, and we again impose two normalizing restrictions on the slopes of ξ_2.

We assumed that the total number of siblings, X_7, was measured without error:

$$X_7 = \xi_3. \quad (7)$$

Olneck (1976) reports a reliability of 0.943 for this variable in his Kalamazoo sample, and our variable is based upon a heavily edited combination of responses to questions about the number of older and younger siblings of each sex, plus our sibling roster.

Our initial measurement of ability among primary respondents, Y_1, was the Henmon-Nelson (1954) Test of Mental Ability; this test was for many years administered to all Wisconsin high school students in their junior year. There was indirect evidence that the Henmon-Nelson test was highly reliable in a sample as heterogeneous as ours (Hauser, Tsai and Sewell 1983:36). To obtain direct evidence of longer-term stability, we searched archival records of the State Testing Service and were able to find freshman year scores on the same test, Y_2, for almost two-thirds of the primary respondents whose siblings were interviewed in 1977. We also located test scores, Y_3, for about 80 percent of the brothers in this same sample of pairs. Our specification of errors in measured mental ability is based on these three measurements:

$$Y_1 = \eta_1 + \epsilon_1, \quad (8)$$

$$Y_2 = \eta_1 + \epsilon_2, \quad (9)$$

and

$$Y_3 = \eta_2 + \epsilon_3, \quad (10)$$

where we assume that all three tests have the same (unit) loadings. The error variances are the same for selected brothers as for primary respondents in the junior year, $\theta_1^{(\epsilon)} = \theta_3^{(\epsilon)}$. This equality constraint is not an overidentifying restriction; the error variance of the brother's test score is not identified except in highly restrictive specifications of our structural model, for example, where we assume there is no family bias in the effects of ability on schooling or some socioeconomic outcome.

As shown in Table 1, there are two indicators of the educational attainment of the primary respondent (Y_4, Y_5) and of his brother (Y_6, Y_7). The first member of each pair is a proxy report and the second is a self-report. In the case of the primary respondent, the proxy report (Y_4) was coded from the educational history provided by a parent in the 1964 follow-up, and in that of the brother, the proxy report (Y_6) was given by the primary respondent in the 1975 survey. The self-reports are from the 1975 and 1977 surveys for the primary respondent and selected brother, respectively. In both cases there is some slippage in time between the self and proxy reports, and consequently some true educational mobility may appear as response variability in later models. To minimize this problem, as well as that of classifying post-graduate education in years, we have followed the U.S. Census practice of truncating years of schooling at or beyond 17 years. The measurement model for educational attainment is

$$Y_4 = \lambda_{43}\eta_3 + \epsilon_4, \tag{11}$$

$$Y_5 = \eta_3 + \epsilon_5, \tag{12}$$

$$Y_6 = \eta_4 + \epsilon_6, \tag{13}$$

and

$$Y_7 = \eta_4 + \epsilon_7, \tag{14}$$

where we have fixed unit loadings for the self-reports by each brother to normalize the unobservables and, in addition, restricted the proxy report of the selected brother's education to have a unit slope. Also, we have equated the error variances of the two self-reports of schooling, $\theta_5^{(\epsilon)} = \theta_7^{(\epsilon)}$.

In the case of occupational status, we have self-reports of the first, full-time occupation after leaving school for the last time from the 1975 and 1977 surveys for the primary respondent, Y_8, and for the selected brother, Y_{11}:

$$Y_8 = \eta_5 + \epsilon_8 \tag{15}$$

and

$$Y_{11} = \eta_{10} + \epsilon_{11}. \tag{16}$$

Because there is only one indicator of each of these variables, the error variances, $\theta_8^{(\epsilon)}$ and $\theta_{11}^{(\epsilon)}$ are not identified, and we have "borrowed" estimates of them from the measurement model for current occupation. In the latter case, there is one self-report for the primary respondent, Y_9, but there is a proxy report from the 1975 survey, Y_{12}, as well as a self-report, Y_{13}, for the selected brother from the 1977 survey. The measurement equations are

$$Y_9 = \eta_7 + \epsilon_9, \tag{17}$$

$$Y_{12} = \eta_8 + \epsilon_{12}, \tag{18}$$

and

$$Y_{13} = \eta_8 + \epsilon_{13}. \tag{19}$$

The error variances $\theta_{12}^{(\epsilon)}$ and $\theta_{13}^{(\epsilon)}$ are both identified, and we "borrowed" the latter to estimate the error variances of Y_8, Y_{11}, and Y_9, that is, $\theta_{13}^{(\epsilon)} = \theta_8^{(\epsilon)} = \theta_{11}^{(\epsilon)} = \theta_9^{(\epsilon)}$.[18] Note that the me...res of the selected brother's occupational status were obtained two years apart; thus, our concept of response variability in occupational status is inclusive of true, short-run changes in occupational status.

In the case of earnings, we have only self-reports for primary respondents and their brothers from the 1975 and 1977 surveys, Y_{10} and Y_{14}, respectively:

$$Y_{10} = \eta_9 + \epsilon_{10} \qquad (20)$$

and

$$Y_{14} = \eta_6 + \epsilon_{14}. \qquad (21)$$

The error variance in earnings cannot be identified from these data. Based on econometric evidence (Heckman and Polachek 1974), we estimated semi-logarithmic earnings functions; thus, we were able to borrow an estimate of the error variance in earnings as reported in the Current Population Survey, from which we had also borrowed the questions used to measure earnings in our 1975 and 1977 surveys.[19]

We fitted an unrestricted reduced form model to the pooled, complete and incomplete samples; in this model, there are no constraints on the variance-covariance matrix of the unobservable variables. As shown in Figure 1, there are distinct recursive models for each sibling, but all subsequent variables depend in common on the three socioeconomic background constructs, ξ_1, ξ_2, and ξ_3. Mental ability, η_1 or η_2, depends only on the background variables;[20] schooling, η_3 or η_4, depends also on ability; and each of the three socioeconomic outcomes, $\eta_5, \ldots, \eta_{10}$, depends on all of the preceding variables that pertain to the same sibling.

Exogenous variables are uncorrelated with the disturbances (ζs) of endogenous variables, but disturbances among endogenous variables are freely correlated, except where there is a corresponding slope. Thus, all of the disturbances in the socioeconomic outcomes are freely intercorrelated, as are the all of the cross-sibling covariances in endogenous variables.

Despite the lack of identification of some parameters and our consequent borrowing of estimates from various sources, the measurement model is highly overidentified. It yields a likelihood ratio test statistic of $L^2 = 366.84$ with 217 degrees of freedom (df).[21] This is nominally a highly significant statistic, but it is also quite typical of those that are considered acceptable in models with large numbers of variables and of observations. From our examination of residuals from this model, we think that we are justified in basing our analysis on it.

Table 2 displays correlations among the constructs in the measurement model. The variables are arranged in three blocks: characteristics of the family (ξ_1, ξ_2, ξ_3), characteristics of the primary respondent (η_1, η_3, η_5, η_7, η_9), and characteristics of the selected brother (η_2, η_4, η_{10}, η_8, η_6). This arrangement makes it possible to see the substantial symmetries between correlations for primary respondents and brothers. For example, the two blocks of within-sibling correlations (in the center of the table near the main diagonal) are similar, as are the blocks of correlations between family and sibling characteristics (the first three columns below the main diagonal at the left) and the cross-sibling, cross-variable correlations (symmetric about the underlined entries).[22] The underlined entries are cross-sibling correlations of the same construct, and they approximate the percentages of variance that occur between families. Clearly, there is substantial homogeneity of families in mental ability, schooling, and socioeconomic outcomes, but homogeneity is greater in the cases of ability and schooling than of the later variables.

4.0 REDUCED FORM COEFFICIENTS

To provide a baseline for our estimates of bias in family models, we varied the specification of lagged effects of social background and ability in the reduced form equations of the socioeconomic outcomes. Table 3 summarizes these analyses for each of the outcome variables under specifications of asymmetry and symmetry between the primary respondent and his brother. In each case, we estimated the effect of education without any controls, then controlling social background, and, finally, controlling social background and ability. There is no sign of background or ability bias in the slopes of status of first occupation on educational attainment. In the case of current occupation, there is a very small reduction in the slope when social background is controlled, and a substantial reduction when ability is also controlled. For example, in the symmetric model, the schooling coefficient falls from .602 to .519 when ability is controlled.[23] In the case of earnings, the schooling coefficients are all rather low; recall that we have not controlled work experience. Moreover, there are relatively large reductions in the schooling coefficient when either social background or mental ability is controlled. For example, the symmetric slope estimate of earnings on schooling is 0.609 without controls, 0.520 with background controlled, and 0.426 with background and ability controlled.

These estimates of bias due to measured social background are relative low for three reasons: First, for all of the primary respondents and most of their brothers, most variability in schooling occurs beyond high school graduation, and there is evidence that family bias is greater for primary and secondary schooling than fc post-secondary schooling (Olneck 1979:159-190). Second, the present specification of socioeconomic background is very limited; for example, we have not included mother's education or parents' income. Third, family bias appears to be less among these fraternal pairs than in other subsets of the Wisconsin sample (Hauser 1984a); for example, among 3,411 male respondents to the 1975 Wisconsin survey, Sewell, Hauser, and Wolf (1980:571,581) found biases of 13.7 percent in the case of first, full-time civilian occupation and of 32.9 percent in the case of current occupation.

5.0 FAMILY FACTORS IN SIBLING RESEMBLANCE

Figure 2 shows our model of family factors in fraternal resemblance in occupational status, which we have used as a prototype in our analysis. The measurement model is the same as in the reduced form, so we shall say no more about it. The specification of the structural model is as follows. For each variable on which siblings have distinct values, we postulate a decomposition into a family factor and individual or within-family factors. Thus, in the case of mental ability, the decompositions are

$$\eta_1 = \eta_2 + \zeta_4 \qquad (22)$$

and $$\eta_3 = \eta_2 + \xi_5 \qquad (23)$$

for primary respondent and sibling, respectively. Note that there are no disturbances (ζ) in these equations; rather, we specify that the within-family factors (ξs) have the properties usually associated with disturbances in a classical factor model, that is, $E[\eta_2\xi_4] = E[\eta_2\xi_5] = E[\xi_4\xi_5] = 0$. This permits us, using the LISREL model, to write auxiliary regressions in the family and within-family constructs. Thers are similar decompositions of educational attainment,

$$\eta_5 = \eta_6 + \eta_4 \qquad (24)$$

and $$\eta_7 = \eta_6 + \eta_8, \qquad (25)$$

where $E[\eta_6\eta_4] = E[\eta_6\eta_8] = E[\eta_4\eta_8] = 0$, and of occupational status,

$$\eta_{10} = \eta_{11} + \eta_9 \qquad (26)$$

and $$\eta_{12} = \eta_{11} + \eta_{13}, \qquad (27)$$

where $E[\eta_{11}\eta_9] = E[\eta_{11}\eta_{13}] = E[\eta_9\eta_{13}] = 0$. Although the specification of unit slopes on the within-family factors (ξ_4, ξ_5, η_4, η_8, η_9, and η_{13}) is trivial, there are overidentifying restrictions in our specification of unit slopes for η_2 in both of equations 22 and 23, for η_6 in both of equations 24 and 25, and for η_{11} in both of equations 26 and 27. For example, in the model of Figure 2, the ratios of the slopes of η_1 and η_3 on η_2 are identified by their common dependence on ξ_1, ξ_2 and ξ_3.

In the upper and lower portions of the path diagram in Figure 2, we show within-family regressions of schooling on mental ability, and of occupational status on schooling and mental ability for the primary respondent and the selected brother, respectively:

$$\eta_4 = \gamma_{44}\xi_4 + \zeta_4 \qquad (28)$$

and $$\eta_9 = \gamma_{94}\xi_4 + \beta_{94}\eta_4 + \zeta_9, \qquad (29)$$

where $E[\xi_4\zeta_4] = E[\xi_4\zeta_9] = E[\zeta_4\zeta_9] = 0$, for the primary respondent;

and $$\eta_8 = \gamma_{85}\xi_5 + \zeta_8 \qquad (30)$$

and $$\eta_{13} = \gamma_{13,5}\xi_5 + \beta_{13,8}\eta_8 + \zeta_{13}, \qquad (31)$$

where $E[\xi_5\zeta_8] = E[\xi_5\zeta_{13}] = E[\zeta_8\zeta_{13}] = 0$, for the selected brother.[24] In addition, there are constraints on the cross-sibling covariances of the disturbances:

$$E[\zeta_4\zeta_8] = E[\zeta_4\zeta_{13}] = E[\zeta_8\zeta_9] = E[\zeta_9\zeta_{13}] = 0. \qquad (32)$$

In addition to the three slopes in the within-family equations for each brother, there are three variances to be estimated, $Var[\xi_4] = \phi_4$, $Var[\zeta_4] = \Psi_4$, and $Var[\eta_9] = \Psi_9$ in the case of the primary respondent and $Var[\xi_5] = \phi_5$, $Var[\zeta_8] = \Psi_8$, and $Var[\eta_{13}] = \Psi_{13}$ in the case of the selected brother. Each of these parameters is identified but we are particularly interested in the case where symmetry holds, that is,

$$\begin{aligned} \gamma_{44} &= \gamma_{85} \\ \gamma_{94} &= \gamma_{13,5} \\ \beta_{94} &= \beta_{13,8} \qquad (33) \\ \phi_4 &= \phi_5 \\ \Psi_4 &= \Psi_8 \end{aligned}$$

and $$\Psi_9 = \Psi_{13}.$$[25]

Finally, the social background characteristics affect the characteristics of siblings only through the common, unmeasured family factors of the endogenous variables. There is an equation for each of the common family factors:

$$\eta_2 = \gamma_{21}\xi_1 + \gamma_{22}\xi_2 + \gamma_{23}\xi_3 + \zeta_2, \qquad (34)$$

$$\eta_6 = \gamma_{61}\xi_1 + \gamma_{62}\xi_2 + \gamma_{63}\xi_3 + \beta_{62} + \zeta_6, \qquad (35)$$

$$\eta_{11} = \gamma_{11,1}\xi_1 + \gamma_{11,2}\xi_2 + \gamma_{11,3}\xi_3$$

and

$$+ \beta_{11,2} + \beta_{11,6} + \zeta_{11}, \qquad (36)$$

where disturbances are uncorrelated with one another or with the exogenous variables. Here, we are particularly interested in the specification of homogeneity in the corresponding between- and within-family regressions:

$$\beta_{62} = \gamma_{44} = \gamma_{85}, \qquad (37)$$

$$\beta_{11,2} = \gamma_{9,4} = \gamma_{13,5}, \qquad (38)$$

and

$$\beta_{11,6} = \beta_{94} = \beta_{13,8}. \qquad (39)$$

If the restrictions in equations 37 to 39 cannot be rejected, then family bias in the ability and schooling coefficients is attributable only to the social background constructs.[26]

The model leaves open the number of distinct family factors; there may be as many as the number of variables ascertained for each pair of siblings. However, that hypothesis may be rejected in favor of a smaller number of common family factors. For example, if we cannot reject the restriction $\mathrm{Var}[\zeta_{11}] = \Psi_{11} = 0$, then we need not posit a distinct family factor for occupational status. In fact, we find that the variance of the disturbance in the family factor for the status of current occupation is not statistically significant among Kalamazoo brothers; the estimate is $\hat{\Psi}_{11} = .247$ with a standard error of .182. Also, among Wisconsin brothers, the family factor for the status of first occupation is not statistically significant; the estimate is $\hat{\Psi}_{11} = .166$ with a standard error of .122.[27] In all cases, we find distinct family factors for mental ability, educational attainment, and earnings.

6.0 NET FAMILY BIAS IN THE WISCONSIN SAMPLES

Table 4 displays estimates of the within- and between-family slopes of schooling on mental ability and of each socioeconomic outcome on mental ability and on schooling in the pooled Wisconsin samples. All of these findings are based upon the assumption of symmetry as specified in equations 33. In the first column of the table, which pertains to the regression of mental ability on schooling, the estimates are from a model in which a canonical variate in all three socioeconomic outcomes has been substituted for occupational status in Figure 2.[28] The remaining columns of the table are from models in which each of the socioeconomic outcomes in turn was the ultimate endogenous variable in Figure 2.

In the upper portion of each column, we report estimates of the within- and between-family slopes, and in the lower portion, we report estimates of the slope in a model differing only in the specification of the restrictions in equations 37 to 39. The likelihood ratio test statistics pertain to the fit of each model as a whole, and the difference between the two test statistics reported in each column tests the hypothesis that the within- and between-family slopes in a given model are homogeneous, against the alternative that there are distinct within- and between family slopes. The reported test statistics do not test only the hypothesis that the specified slope, say that of mental ability (MA) on schooling (Ed) is homogeneous; rather, they pertain to the global hypothesis that the slopes of schooling on mental ability, of the socioeconomic outcome on mental ability, and of the socioeconomic outcome on schooling are homogeneous in the model including that socioeconomic outcome. It would be inappropriate to report these test statistics as pertaining to the hypothesis of homogeneity in just one of the slopes, except the differences in the test statistics are in each case so small that the null hypothesis would not be rejected even in a test with one degree of freedom.

In short, there is absolutely no evidence of net family bias in the slopes of schooling on mental ability, of any of the socioeconomic outcomes on mental ability, or of any of the socioeconomic outcomes on schooling.[29] The pattern of within- and between-family slopes reported in the upper panel of Table 4 adds further to the impression that nothing is going on here: the between-family slopes are in some cases larger, in some cases smaller, and in some cases about the same as the within-family slopes. At the same time, we do find evidence of ability bias, at least in the homogeneous slope estimates. The effect of mental ability on schooling is highly significant, as are the lagged effects of mental ability on each of the socioeonomic outcomes.[30]

In the complete data sample, we have tested for asymmetries in the slopes for primary respondents and their brothers and for distinct family biases in each population of brothers. Table 5 reports the slopes of schooling on ability and of current occupation and earnings on mental ability and schooling. The lower two panels of Table 5 correspond exactly to Table 4, and the results are exactly the same; there is no evidence of family bias, net of social background.

It is also interesting to compare the standard errors of slope estimates in Table 5 with those of the corresponding, pooled estimates in Table 4. If we had complete data for every pair of brothers and added the 928 observations in the incomplete data sample to the 532 observations in the complete sample, we would expect the standard errors to decline by approximately the ratio $[532/(532 + 928)]^{1/2} = .604$. In no case is this gain achieved, but there are advantages to pooling. The gains vary across coefficients and models, but they appear to be largest in the symmetric, homogeneous model, where the ratios of standard errors of pooled to unpooled slope estimates range from 0.69 to 0.75.

In the upper panel of Table 5, we report estimates of coefficients of the model of Figure 2 without the symmetry constraints of equations 33. There are apparent differences in the within-family slopes for primary respondents and their brothers. In each case the slope is larger for the selected brother than for the primary respondent; moreover, in the model for earnings but not that for occupational status, a global comparison of the fit of this asymmetric model with that of complete symmetry yields a significant test statistic.[31] However, neither in the case of earnings nor of occupational status are the asymmetries due to significant differences between the slopes for primary respondents and their brothers. When we add constraints on each pair of within-family slopes in turn to the asymmetric models, we find no significant decrements in fit.[32] Thus, it is appropriate to condition our tests of homogeneity on symmetry (or at least on symmetry in the within-family slopes).

7.0 HOMOGENEITY OF FAMILIES

The finding of negligible net family bias should not be construed as a suggestion either that families do not have substantial effects on socioeconomic achievement or that their effects are explained by the background variables in our model.

Table 6 reports decompositions of the variances in mental ability, schooling, and the three socioeconomic outcomes in our preferred models of symmetric, homogeneous effects. Our model permits a four-fold decomposition of variances into explained and unexplained within- and between-family components.[33]

We have already noted the substantial equality of within- and between-family components of variance in ability and schooling and the gradient in between-family components of variance as we consider first occupation, current occupation, and earnings. The models explain most of the between-family variance in first and in current occupations,[34] but not in ability, schooling, or earnings. Thus, families do have large effects on achievement. They have relatively larger direct effects on the two most important antecedents of adult socioeconomic success - ability and schooling - than on occupational status or earnings, and their effects on those antecedents of achievement are by no means explained by socioeconomic background.

Even more striking is the failure of the models to account for much of the within-family variation in schooling or in earnings. Obviously, these findings suggest there is room for improvement in the specification of the within-family equations. Our earlier work suggests some fruitful ways in which these parts of the model can be elaborated (Hauser, Tsai and Sewell 1983, Hauser 1979), for example, by introducing other social psychological antecedents of completed schooling and by adding work experience to the earnings equations. Again, we have not elaborated the present analysis in these ways because we have sought to maintain comparability with Olneck's (1977) analysis.

We turn now to a cross validation of our findings, which 's based upon a reanalysis of Olneck's data with corrections for attenuation.

8.0 CROSS-VALIDATION USING THE KALAMAZOO DATA

Although his published analyses (Olneck 1976, Olneck 1977, Olneck 1979) are based upon a small sample of brother pairs with complete data (weighted N = 346), we believe Olneck's pioneering study compares favorably with other efforts to model resemblance among American men in the size and heterogeneity of the sample and the completeness and quality of the data. For this reason, we think that our findings about fraternal resemblance in schooling and socioeconomic attainment ought to be compared with those of Olneck.[35]

Olneck (1977:145-150) reports an extensive analysis of bias in the effects of ability and schooling attributable to measured and unmeasured family factors, but these analyses do not introduce corrections for errors in variables. He concludes:

> Controlling measured socioeconomic variables does not fully eliminate biases due to background in the effects of test scores on educational attainment, and in the effects of education on current occupational status and earnings. The effect of measured ability on earnings among brothers is, however, the same as it is among unrelated individuals. This result is anomolous [sic], and may well be due to sampling error.

In reporting the analyses without corrections for attenuation, Olneck (1977) did not report direct tests of the significance of family bias; rather, he followed the common practice of offering prima facie interpretations of changes in within-family slope estimates as various changes were made in the specification of the model. However, Olneck (1976:107-139) does report appropriate tests of the differences of between- and within-family slopes.

Although Olneck (1976:Ch. 4) was very much aware of the problem of errors in variables and made a serious effort to assemble estimates of error variances from various sources and to correct his analyses for them, he did not apply statistical tests of family bias to his corrected models. Because Olneck's reported findings are in apparent conflict with ours on the significance of family bias, we think it is useful to reassess those findings in the present modeling framework, using his estimates of error variance. Olneck reported correlations between true scores and indicators, and he used these to estimate standardized path coefficients from corrected correlations. We have translated the estimates of validity into error variance components and estimated our measurement and structural models from the variance-covariance matrix.[36]

Table 7 reports our estimates of the models of Table 4 in the Kalamazoo data. Because he had initially sampled families, rather than persons, Olneck treated the members of each brother pair symmetrically; he entered each brother as both the first and second member of each pair when the moments were estimated. Consequently, because the measurement model is also trivial, the present models are just-identified when the socioeconomic outcomes are treated one at a time, and distinct within- and between-family slopes are specified.

Our reading of the coefficients in Tables 4 and 7 is that they are remarkably similar, including such details as the negligible within-family slope of status of first occupation on mental ability and the anomalous, but non-significant negative between-family slope of earnings on mental ability. Indeed, if we construct approximate standard errors of differences between slopes in the two samples as

$$SE(\hat{\beta}_{iw} - \hat{\beta}_{ik}) = \{[SE(\hat{\beta}_{iw})]^2 + [SE(\hat{\beta}_{ik})]^2\}^{1/2}, \quad (40)$$

where $\hat{\beta}_{ij}$ is the estimate of the i^{th} parameter in the j^{th} sample, there appear to be only two significant differences in slope between the samples. Both the within-family and homogeneous coefficients of status of first occupation on schooling are significantly larger in the Wisconsin samples than in the Kalamazoo sample.[37] Otherwise, none of the differences in estimated slopes is as large as two standard errors of the difference.

If we consider the Kalamazoo estimates by themselves, there is little evidence of net family bias. As in Table 4, we have reported global fit statistics for each model and supplemented these with tests of homogeneity in specific coefficients. The first column of Table 7 reports estimates of the effect of mental ability on schooling in a model where the socioeconomic outcomes are treated canonically. Here, it is interesting that the global test of homogeneity yields a smaller test statistic, $L^2 = 47.52 - 44.69 = 2.83$ with 3 degrees of freedom, han do any of the subsequent global tests. This test statistic would not be significant even on one degree of freedom, yet it provides a global test of net family bias in the effects of mental ability on schooling and in average effects of schooling and of mental ability on occupational status and earnings.

In the other six columns of Table 7, each of the test statistics for homogeneity is large enough to contain a significant contrast, so we have separately tested the homogeneity of each slope by imposing one constraint at a time on the model of heterogeneous within- and between-family slopes. These test statistics are reported in the bottom row. In only one case, that of the improbable, negative between-family slope of earnings on mental ability, is there statistically significant evidence of heterogeneity. In the absence of any rationale for that finding, we are inclined to disregard it. Our conclusion about the Kalamazoo data, after correction for attenuation, is that they provide no more evidence of net family bias in the effects of mental ability or of schooling than do the Wisconsin data.

REFERENCES

Allison, Paul D. 1982. "Maximum likelihood estimation of linear models when data are missing." Revised version of a paper presented at the 1981 Annual Meeting of the American Sociological Association.

Allison, Paul D., and Robert M. Hauser. 1984. "Reducing bias in estimates of linear models by remeasurement of a random subsample." Manuscript (March).

Behrman, J., P. Taubman, and T. Wales. 1977. "Controlling for and measuring the effects of genetics and family environment in equations for schooling and labor market success." Pp. 35-96 in Paul Taubman (ed.) Kinometrics: Determinants of socioeconomic success within and between families. Amsterdam: North-Holland.

Bielby, William T., and Robert M. Hauser. 1977. "Response error in earnings functions for nonblack males." Sociological Methods and Research 6 (November):241-80.

Bielby, William T., Robert M. Hauser and David L. Featherman. 1977. "Response errors of black and nonblack males in models of the intergenerational transmission of socioeconomic status." American Journal of Sociology 82 (May):1242-88.

Blau, Peter M., and Otis Dudley Duncan. 1967. The American occupational structure New York: Wiley and Sons.

Bound, John, Zvi Griliches, and Bronwyn H. Hall. 1984. "Brothers and sisters in the family and the labor market." Manuscript.

Bowles, Samuel. 1972. "Schooling and inequality from generation to generation." Journal of Political Economy 80 (May/June):S219-51.

Brittain, John A. 1977. The inheritance of economic status. Washington, D.C.: Brookings Institution.

Chamberlain, Gary, and Zvi Griliches. 1977. "More on brothers." Pp. 97-107 in Paul Taubman (ed.), Kinometrics: Determinants of socioeconomic success within and between families. Amsterdam: North-Holland.

Corcoran, Mary, and Linda P. Datcher. 1981. "Intergenerational status transmission and the process of individual attainment." Pp. 169-206 in Martha S. Hill, Daniel H. Hill, and James N. Morgan (eds.), Five thousand American families - Patterns of economic progress, Volume IX. Ann Arbor: Institute for Social Research, The University of Michigan.

Duncan, Otis Dudley. 1961. "A socioeconomic index for all occupations." Pp. 109-38 in Albert J. Reiss, Jr. (ed.), Occupations and social status. New York: Free Press.

Featherman, David L., and Robert M. Hauser. 1978. Opportunity and Change New York: Academic Press.

Gorseline, D.W. 1932. The effect of schooling upon income. Bloomington: Indiana University Press.

Griliches, Zvi. 1977. "Estimating the returns to schooling: Some econometric problems." Econometrica 45 (January):1-22.

Griliches, Zvi. 1979. "Sibling models and data in economics: Beginnings of a survey." Journal of Political Economy 87 (October):S37-64.

Griliches, Zvi. 1983. "Data problems in econometrics." Draft,Chapter 25, Handbook of Econometrics.

Hauser, Robert M. 1984a. "Some cross-population comparisons of family bias in the effects of schooling on occupational status." Social Science Research 13 (1984):159-87.

Hauser, Robert M. 1984b. "A note on two models of sibling resemblance." Working Paper 84-3, Center for Demography and Ecology, The University of Wisconsin-Madison.

Hauser, Robert M., and David L. Featherman. 1977. The process of stratification. New York: Academic Press.

Hauser, Robert M., and Arthur S. Goldberger. 1971. "The treatment of unobservable variables in path analysis." Pp. 81-117 in Herbert L. Costner (ed.) Sociological Methodology, 1971. San Francisco: Jossey Bass.

Hauser, Robert M., and Peter A. Mossel. 1984. "Fraternal resemblance in educational attainment and occupational status." American Journal of Sociology (forthcoming).

Hauser, Robert M., William H. Sewell and Brian R. Clarridge. 1982. "The influence of family structure on socioeconomic achievement: A progress report." Working Paper 82-59, Center for Demography and Ecology, The University of Wisconsin-Madison.

Hauser, Robert M., Shu-Ling Tsai and William H. Sewell. 1983. "A model of stratification with response error in social and psychological variables." Sociology of Education 56 (January):20-46.

Heckman, James, and Solomon Polachek. 1974. "Empirical evidence on the functional form of the earnings-schooling relationship." Journal of the American Statistical Association 69 (350-54).

Henmon, V.A.C., and M.J. Nelson. 1954. The Henmon-Nelson Tests of Mental Ability. Manual for Administration. Revised 1954. Boston: Houghton-Mifflin Company.

Jencks, Christopher, Marshall Smith, Henry Acland, Mary Jo Bane, David Cohen, Herbert Gintis, Barbara Heyns and Stephan Michelson. 1972. Inequality. New York: Basic Books.

Jencks, Christopher, Susan Bartlett, Mary Corcoran, James Crouse, David Eaglesfield, Gregory Jackson, Kent McClelland, Peter Mueser, Michael Olneck, Joseph Schwartz, Sherry Ward and Jill Williams. 1979. Who gets ahead? The determinants of economic success in America. New York: Basic Books.

Joreskog, Karl G., and Dag Sorbom. 1978. LISREL: Analysis of linear structural relationships by the method of maximum likelihood. User's Guide, Version IV. University of Uppsala.

Joreskog, Karl G., and Dag Sorbom. 1983. LISREL VI: Supplement to the LISREL V manual Manuscript (March).

Mossel, Peter A. 1984. Family effects on the returns to schooling: Theory and estimation. Unpublished doctoral thesis. Department of Sociology, The University of Wisconsin-Madison.

Olneck, Michael R. 1976. The determinants of educational attainment and adult status among brothers: The Kalamazoo study. Unpublished doctoral thesis, Harvard Graduate School of Education.

Olneck, Michael R. 1977. "On the use of sibling data to estimate the effects of family background, cognitive skills, and schooling: Results from the Kalamazoo brothers study." Pp. 125-62 in Paul Taubman (ed.) Kinometrics: Determinants of socioeconomic success within and between families. Amsterdam: North-Holland.

Olneck, Michael R. 1979. "The effects of education." Pp. 159-190 in Christopher Jencks, et al., Who gets ahead? The determinants of economic success in America. New York: Basic Books.

Sewell, William H., and Robert M. Hauser. 1975. Education, occupation, and earnings. New York: Academic Press.

Sewell, William H., and Robert M. Hauser. 1980. "The Wisconsin longitudinal study of social and psychological factors in aspirations and achievements." Research in Sociology of Education and Socialization 1:59-99.

Sewell, William H., Robert M. Hauser and Wendy C. Wolf. 1980. "Sex, schooling, and occupational status." American Journal of Sociology 86 (November):551-83.

FOOTNOTES

1. For example, Sewell and Hauser (1975:72,81,93,98) and Sewell, Hauser and Wolf (1980:571,581) have found substantial biases in effects of post-secondary schooling on the occupational success of Wisconsin youth. However, Olneck (1979:159-90) has found much smaller biases in effects of post-secondary schooling than of primary or secondary schooling.

2. Griliches (1979:S61-S62) does not apply this same argument to observed differences in the within- and between-family slopes of schooling on ability, which he attributes instead to the family's compensatory treatment of offspring with greater and lesser ability.

3. These have been reviewed by Hauser, Tsai and Sewell (1983) and by Griliches (1983).

4. For pioneering efforts to do this, see Jencks et al. (1972:Appendix B), Olneck (1976:166-198), Olneck (1977:149-150), and Behrman, Taubman and Wales (1977:80-81). Bielby, Hauser and Featherman (1977), Bielby and Hauser (1977), and Hauser, Tsai and Sewell (1983) have estimated models of socioeconomic achievement with extensive corrections for errors in variables, but these have not incorporated global family effects.

5. The findings reported here pertain only to brother pairs in the Wisconsin sample, but Hauser (1984a) has also analyzed the regression of occupational status on schooling among Wisconsin sister pairs and sister-brother pairs.

6. Bound, Griliches, and Hall (1984) have independently used this method to exploit incomplete data for mixed-sex sibling pairs in the National Longitudinal Study. It is intriguing that Olneck

(1976) collected a full set of proxy reports of education, occupation, and earnings for the Kalamazoo brothers from the first member of each sibship that he interviewed, yet his published analyses are based only upon pairs with self-reports of all variables.

7. Mossel (1984) has used data from brother pairs in the Wisconsin Longitudinal Study to replicate well-known economic specifications of the earnings function.

8. Olneck (1977) did not enter work experience in his published earnings equations because he found no effects of that variable in the Kalamazoo sample (personal communication).

9. See Griliches (1977, 1979) for an exposition of the simultaneity problem.

10. Note that corrected measures of parental social and economic characteristics are unmeasured common family factors.

11. We obtained complete, self-reported data for approximately 750 brothers of male respondents, 750 sisters of female respondents, 250 sisters of male respondents, and 250 brothers of female respondents.

12. There is reason to believe that the achieved sample of brother pairs adequately reflects the composition of the sample of primary respondents (and their brothers) from which it was drawn (Hauser, Sewell and Clarridge 1982:7-13).

13. The sample sizes are shown in Table 1. We have also repeated and confirmed much of the analysis using listwise-present estimates of moments in the complete sample.

14. A test of homogeneity between the variance-covariance matrices of the 13 variables that were measured in both samples

and used in the present analysis yielded a likelihood-ratio test statistic of 84.24 with 91 degrees of freedom, which is not statistically significant.

15. All standard deviations have been rescaled so there is one significant digit to the left of the decimal point. This simplifies the presentation of numeric findings in a fixed format. It also improves the performance of the LISREL program.

16. All of the estimates reported in this paper were obtained using Joreskog and Sorbom's (1983) LISREL VI program (also, see Joreskog and Sorbom 1978). Allison (1982) shows how LISREL can be used to obtain ML estimates when data are missing; also, see Allison and Hauser (1984). Briefly, the models are estimated in multiple samples with restrictions on parameters across samples. In the sample(s) with missing moments, one specifies arbitrary variances and zero covariances as data for the missing indicators, and one specifies zero loadings, free error variances and no correlated errors involving the missing indicators. It is necessary to correct the degrees of freedom reported by LISREL to take account of the non-existent covariances.

17. Detailed data on industry and class of worker were used to refine the scale values reported by Hauser and Featherman for certain occupation lines.

18. All four of these variables were ascertained using the same, detailed Census-type questions. One may question the assumption that error variances in the status of current occupation and of first occupation are equal because of the variable periods of recall. However, among nonblack U.S. men aged 25 to 64, Bielby, Hauser and Featherman (1977:1258) estimated

error variances in parallel measures of first occupations that were, if anything, slightly smaller than that in a self-report of current occupation. On the basis of this and related evidence, we believe that the major problem in reporting occupations is not recall, but the inherent difficulty of describing the same job in the same way on two different occasions.

19. Bielby and Hauser (1977:262) estimated this variance as 3.1684; note that we have shifted the decimal point two places to the right. In future extensions of this work, we hope to pool the Bielby-Hauser data with those from the Wisconsin sample, but for the present analysis, we have treated their estimate as a constant.

20. The causal ordering of social background and abili should not be taken too literally; it is more of a convenience in writing the models than a substa ve assumption. Nothing would be altered in our statistical analysis if ability were regarded as merely correlated with measured or unmeasured family or social background.

21. This test statistic contrasts the measurement model with a completely unrestricted model of the observed variance-covariance matrix. It is distributed as χ^2 in large samples under the assumption of multivariate normality.

22. There are 40 possible symmetry restrictions between primary respondents and their brothers in variances and covariances of the unobservables, and the simultaneous imposition of all of these raises the likelihood ratio test statistic by 103.89; thus, the imposition of symmetry leads to a greater decrement in fit, relative to degrees of freedom, than does the

measurement model.

23. In our scaling of the variables, this says that an additional year of schooling is worth 5.19 points of occupational status on the Duncan scale.

24. It may seem odd to write these equations in the disturbances of the family factor model, rather than in the total constructs for each sibling, for example, in ξ_4, η_4, and η_9, rather than η_1, η_5, and η_{10}. Hauser (1984b) has shown there are logical and statistical advantages to the present specification.

25. Given the selection of brothers into the sample through primary respondents, we should probably have conditioned all of our models on $\phi_4 \neq \phi_5$, $\gamma_{44} \neq \gamma_{85}$, and $\Psi_4 \neq \Psi_8$, and imposed only the other three conditions of equation 32 as specifications of symmetry.

26. There is a stronger version of the hypothesis of no family bias, which says that the between- and within-family regressions are homogeneous and, in addition, there are no lagged effects of the social background variables on family factors following ability in the model. We have not tested the latter hypothesis in this analysis. However, the present test of family bias is more stringent than a test based upon a larger set of family background variables.

27. These estimates pertain to the models shown in the bottom row of Table 4.

28. That is, the effects of social background, mental ability, and schooling on the three outcomes are subject to a proportionality constraint in the reduced form (Hauser and Goldberger 1971, Hauser 1973). The full model has many parameters

and takes a great deal of computer time to estimate. For this reason we did not enter the socioeconomic outcomes simultaneously except in this model and the reduced forms. Our canonical treatment of the three outcomes is unreasonably restrictive in the final equations of the model, but it has the advantage of producing a pooled estimate of the slope of schooling on ability. The estimates of that coefficient vary slightly among the other models.

29. Recall that there are lagged effects of the socioeconomic outcomes in each of these equations, so we have tested a hypothesis of no net family bias, not of no bias at all. The failure to find any net bias is an invitation to test the stronger hypothesis that there is no family bias whatever, but we have not done so here. However, Hauser (1984a; also, see Hauser 1984b) was not able to reject the hypothesis of no family bias in the regressions of occupational status on schooling in the Wisconsin and Kalamazoo samples.

30. There is a substantial gain in the precision of the homogeneous slope estimates, relative either to the within- or between-family estimates in the heterogeneous model.

31. In the case of earnings the global test of symmetry yields a test statistic of $L^2 = 185.63 - 144.35 = 41.28$ with 6 degrees of freedom, which is highly significant. In the case of occupational status, the corresponding test statistic is only $L^2 = 209.04 - 197.92 = 11.12$.

32. The same finding holds for the slope of schooling on ability, which we estimated in the model where earnings was the ultimate endogeneous variable.

33. Of course, there is no decomposition of the within-family variance in mental ability, which is exogenous.

34. Recall that the unexplained between-family variance component in status of first occupation is not statistically significant.

35. The Kalamazoo brothers were selected from the rolls of sixth graders in public schools from 1928 to 1950 and were followed up in 1973. Because of sample attrition and item nonresponse, the Kalamazoo data pertain to roughly one quarter of the men originally selected by Olneck. We have analyzed moments that have been reported in several of the cited sources, and we have not reproduced them here.

36. The estimated error variance components, rescaled as the variables in Table 1, are father's education, 2.482; father's occupational status, 1.189; number of siblings, 0.3659; mental ability, 0.2347, years of schooling, 0.5269; status of first occupation, 0.3028; status of current occupation, 0.8409; earnings, 2.509. Olneck's model estimates the within-family coefficients from total variates, rather than from within-family deviations, and this changes the interpretation of the family factors in important ways (Hauser 1984b). Thus, it is difficult to compare Olneck's corrected estimates directly with ours.

37. The larger share of Kalamazoo men with less than 12 years of schooling probably accounts for this difference; regressions of occupational status on schooling are less steep at grades 0 to 12 than beyond grade 12 (Featherman and Hauser 1978:268, Olneck 1979:159-190).

Table 1. Description of variables: Complete and incomplete samples of Wisconsin brother pairs

		Complete sample	Incomplete sample
X_1	EDFA57Q (Father's education, Primary respondent, 1957 School Survey)		
	Mean:	10.245	10.283
	Dev.:	3.100	3.147
	N:	482	847
	Rel.:	.733	.733
X_2	EDHHYR (Father's education, Primary respondent, 1975 Survey)		
	Mean:	9.656	9.520
	Dev.:	3.361	3.387
	N:	512	882
	Rel.:	.777	.777
X_3	XEDHHYR (Father's education, Selected sibling, 1977 Survey)		
	Mean:	9.869	—
	Dev.:	3.210	—
	N:	490	—
	Rel.:	.777	—
X_4	OCSF57 (Father's occupation, Parent, State tax records)		
	Mean:	29.472	29.059
	Dev.:	2.1473	2.1885
	N:	460	825
	Rel.:	.609	.609
X_5	OCSH57 (Father's occupation, Primary respondent, 1975 Survey)		
	Mean:	32.491	32.976
	Dev.:	2.1934	2.2511
	N:	518	909
	Rel.:	.749	.749
X_6	XOCSH57 (Father's occupation, Selected sibling, 1977 Survey)		
	Mean:	32.069	—
	Dev.:	2.1747	—
	N:	525	—
	Rel.:	.749	—

(continued next page)

Table 1, continued.

		Complete sample	Incomplete sample
X_7	SIBSTT (Number of sibings, Primary respondent, 1975 Survey)		
	Mean:	3.477	3.517
	Dev.:	2.382	2.558
	N:	532	928
	Rel.:	1.000	1.000
Y_1	IQHNSCRQ (Respondent's ability, Primary respondent, State testing service)		
	Mean:	100.575	100.665
	Dev.:	1.4812	1.4625
	N:	532	928
	Rel.:	.778	.778
Y_2	IQFSCOR (Respondent's ability, Primary respondent, State testing service)		
	Mean:	100.926	—
	Dev.:	1.4123	—
	N:	336	—
	Rel.:	.851	—
Y_3	XIQSCOR (Sibling's mental ability, Selected sibling, State testing service)		
	Mean:	102.000	—
	Dev.:	1.4493	—
	N:	411	—
	Rel.:	.773	—
Y_4	EDAT64 (Respondent's education, Parent, 1964 Survey)		
	Mean:	13.454	13.494
	Dev.:	1.890	1.918
	N:	474	826
	Rel.:	.846	.846
Y_5	EDEQYR (Respondent's education, Primary respondent, 1975 Survey)		
	Mean:	13.600	13.665
	Dev.:	2.356	2.375
	N:	532	928
	Rel.:	.940	.940
Y_6	SSBED (Sibling's education, Primary respondent, 1975 Survey)		
	Mean:	13.402	13.262
	Dev.:	2.552	2.617
	N:	530	923
	Rel.:	.896	.896

(continued next page)

Table 1, continued.

		Complete sample	Incomplete sample
Y_7	XEDEQYR (Sibling's education, Selected sibling, 1977 Survey)		
	Mean:	13.560	—
	Dev.:	2.590	—
	N:	532	—
	Rel.:	.949	—
Y_8	OCSX1 (Respondent's first job, Primary respondent, 1975 Survey)		
	Mean:	39.683	39.386
	Dev.:	2.6483	2.6438
	N:	532	919
	Rel.:	.813	.813
Y_9	OCSXCR (Respondent's current job, Primary respondent, 1975 Survey)		
	Mean:	49.525	49.638
	Dev.:	2.3863	2.4230
	N:	532	928
	Rel.:	.775	.775
Y_{10}	YRER74L (Respondent's earnings, Primary respondent, 1975 Survey)		
	Mean:	5.023	5.028
	Dev.:	4.05	4.08
	N:	532	928
	Rel.:	.809	.809
Y_{11}	XOCSX1 (Sibling's first job, Selected sibling, 1977 Survey)		
	Mean:	39.876	—
	Dev.:	2.6663	—
	N:	530	—
	Rel.:	.817	—
Y_{12}	OCSSIB (Sibling's occupation, Primary respondent, 1975 Survey)		
	Mean:	48.048	44.241
	Dev.:	2.4687	2.4987
	N:	532	928
	Rel.:	.815	.815
Y_{13}	XOCSXCR (Sibling's current job, Selected sibling, 1977 Survey)		
	Mean:	49.069	—
	Dev.:	2.5366	—
	N:	532	—
	Rel.:	.795	—

(continued next page)

Table 1, continued.

		Complete sample	Incomplete sample
Y_{14}	XYRER76L (Sibling's earnings, Selected sibling, 1977 Survey)		
	Mean:	5.075	—
	Dev.:	5.09	—
	N:	532	—
	Rel.:	.879	—

Note: See text for definition of complete and incomplete samples. Standard deviations have been rescaled to lie between 1 and 10; see text for explanation. Parenthetic entries give the name of the construct, the person reporting the variable, and the survey or other source. Reliabilities of earnings are based upon an error variance estimated by Bielby and Hauser (1977) from the March 1973 Current Population Survey and Income Reinterview Program. Other reliabilities are maximum likelihood estimates from unrestricted reduced form equations, pooled across complete and incomplete samples.

Table 2. Disattenuated correlations among variables: Wisconsin brothers

	ξ_1	ξ_2	ξ_3	η_1	η_3	η_5	η_7	η_9	η_2	η_4	η_{10}	η_8	η_6
ξ_1	1.000	.671	-.274	.366	.424	.414	.356	.243	.355	.405	.370	.328	.214
ξ_2	.671	1.000	-.274	.326	.341	.380	.389	.334	.348	.380	.393	.361	.213
ξ_3	-.274	-.274	1.000	-.176	-.193	-.191	-.177	-.101	-.160	-.260	-.202	-.217	-.061
η_1	.366	.326	-.176	1.000	.540	.530	.512	.297	.491	.325	.293	.278	.164
η_3	.424	.341	-.193	.540	1.000	.859	.656	.357	.304	.459	.375	.320	.178
η_5	.414	.380	-.191	.530	.859	1.000	.772	.367	.353	.418	.382	.349	.155
η_7	.356	.389	-.177	.512	.656	.772	1.000	.446	.345	.340	.336	.355	.199
η_9	.243	.334	-.101	.297	.357	.367	.446	1.000	.163	.256	.253	.233	.300
η_2	.355	.348	-.160	.491	.304	.353	.345	.163	1.000	.533	.497	.547	.313
η_4	.405	.380	-.260	.325	.459	.418	.340	.256	.533	1.000	.895	.738	.362
η_{10}	.370	.393	-.202	.293	.375	.382	.336	.253	.497	.895	1.000	.801	.351
η_8	.328	.361	-.217	.278	.320	.349	.355	.233	.547	.738	.801	1.000	.473
η_6	.214	.213	-.061	.164	.178	.155	.199	.300	.313	.362	.351	.473	1.000

Note: Variables are: ξ_1 = father's education, ξ_2 = father's occupational status, ξ_3 = number of siblings. η_1 = mental ability of primary respondent, η_2 = mental ability of brother, η_3 = education of primary respondent, η_4 = education of brother, η_5 = first job of primary respondent, η_6 = earnings of brother, η_7 = current job of primary respondent, η_8 = current job of brother, η_9 = earnings of primary respondent, η_{10} = first job of brother. Underlined entries are correlation ratios. See text for explanation.

Table 3. Disattenuated regressions of socioeconomic achievement on schooling: Wisconsin brother pairs

	Occupational status		
	First	Current	Earnings
Asymmetric model ($L^2(217) = 366.84$)			
Primary respondent:			
1. No controls	.891 (.021)	.605 (.024)	.567 (.045)
2. Social background controlled	.858 (.023)	.554 (.026)	.464 (.050)
3. Social background and ability controlled	.821 (.027)	.471 (.030)	.384 (.058)
Brother of primary respondent:			
4. No controls	.866 (.028)	.669 (.021)	.696 (.083)
5. Social background controlled	.863 (.032)	.641 (.024)	.638 (.094)
6. Social background and ability controlled	.854 (.037)	.556 (.031)	.505 (.109)
Symmetric model ($L^2(257) = 470.73$)			
All brothers:			
7. No controls	.887 (.017)	.641 (.016)	.609 (.040)
8. Social background controlled	.862 (.019)	.602 (.018)	.520 (.045)
9. Social background and ability controlled	.833 (.022)	.519 (.021)	.426 (.051)

Note: Occupational status is expressed in 10-point intervals on the Duncan scale. Earnings are expressed in natural logs, and coefficients have been multipled by 10. Entries are maximum likelihood estimates of slopes (with standard errors in parentheses) from reduced form equations, pooled across complete and incomplete samples.

Table 4. Selected slope estimates and measures of fit: Wisconsin brothers

Antecedent variable:	MA	MA	Ed	MA	Ed	MA	Ed
Dependent variable:	Ed	1st Occ	1st Occ	Cur Occ	Cur Occ	Earn	Earn
Symmetric, heteroganeous slopes:							
Within-family	0.816	0.036	0.860	0.306	0.559	0.575	0.313
	(0.084)	(0.095)	(0.039)	(0.089)	(0.036)	(0.222)	(0.092)
Between-family	0.800	0.256	0.753	0.354	0.442	-0.031	0.718
	(0.120)	(0.145)	(0.076)	(0.122)	(0.069)	(0.348)	(0.186)
Fit: LR statistic	538.63	272.27	272.27	300.64	300.64	282.01	282.01
Degrees of freedom	267	152	152	178	178	152	152
Symmetric, homogeneous slopes:							
Within- and between-family	0.811	0.118	0.826	0.318	0.525	0.360	0.437
	(0.046)	(0.042)	(0.022)	(0.046)	(0.022)	(0.101)	(0.053)
Fit: LR statistic	540.33	273.51	273.51	302.75	302.75	284.95	284.95
Degrees of freedom	270	155	155	181	181	155	155

Note: Variables are MA = mental ability, Ed = years of schooling, 1st Occ = status of first, full-time occupation, Cur Occ = status of current occupation, Earn = natural log of annual earnings. Findings are based upon maximum likelihood estimates, corrected for attenuation and pooled across brother pairs with complete (N = 532) and incomplete (N = 928) data. Parenthetic entries are approximate standard errors.

Table 5. Selected slope estimates and measures of fit: Wisconsin brothers in complete data sample (N = 532)

Antecedent variable:	MA	MA	Ed	MA	Ed
Dependent variable:	Ed	Cur Occ	Cur Occ	Earn	Earn
Asymmetric, heterogeneous slopes:					
Primary respondent	0.689	0.345	0.467	0.359	0.207
	(0.122)	(0.138)	(0.078)	(0.283)	(0.164)
Brother	0.931	0.406	0.618	1.001	0.382
	(0.162)	(0.151)	(0.060)	(0.409)	(0.162)
Between-family	0.734	0.341	0.462	-0.122	0.767
	(0.155)	(0.158)	(0.099)	(0.373)	(0.232)
Fit: LR statistic	144.35	197.92	197.92	144.35	144.35
Degrees of freedom	91	106	106	91	91
Symmetric, heterogeneous slopes:					
Within-family	0.802	0.361	0.568	0.599	0.340
	(0.098)	(0.104)	(0.050)	(0.248)	(0.119)
Between-family	0.742	0.344	0.458	-0.044	0.730
	(0.155)	(0.158)	(0.098)	(0.376)	(0.235)
Fit: LR statistic	185.63	209.04	209.04	185.63	185.63
Degrees of freedom	97	112	112	97	97
Symmetric, homogeneous slopes:					
Within- and between-family	0.781	0.350	0.537	0.368	0.459
	(0.061)	(0.061)	(0.032)	(0.143)	(0.075)
Fit: LR statistic	187.68	210.14	210.14	187.68	187.68
Degrees of freedom	100	115	115	100	100

Note: Variables are MA = mental ability, Ed = years of schooling, Cur Occ = status of current occupation, Earn = natural log of annual earnings. Parenthetic entries are approximate standard errors.

Table 6. Components of variance in selected variables: Wisconsin brothers

Relative component of variance	MA	Ed	1^{st} Occ	Cur Occ	Earn
Within-family:	50.7	54.4	58.6	62.4	72.8
Explained	—	9.9	39.1	24.4	5.8
Unexplained	—	44.6	19.5	38.0	67.1
Between-family:	49.3	45.6	41.4	37.6	27.2
Explained	15.2	26.1	38.5	29.0	11.8
Unexplained	34.0	19.5	2.9	8.7	15.4
Total	100.0	100.0	100.0	100.0	100.0
Variance component	(1.686)	(5.707)	(5.812)	(4.795)	(15.992)

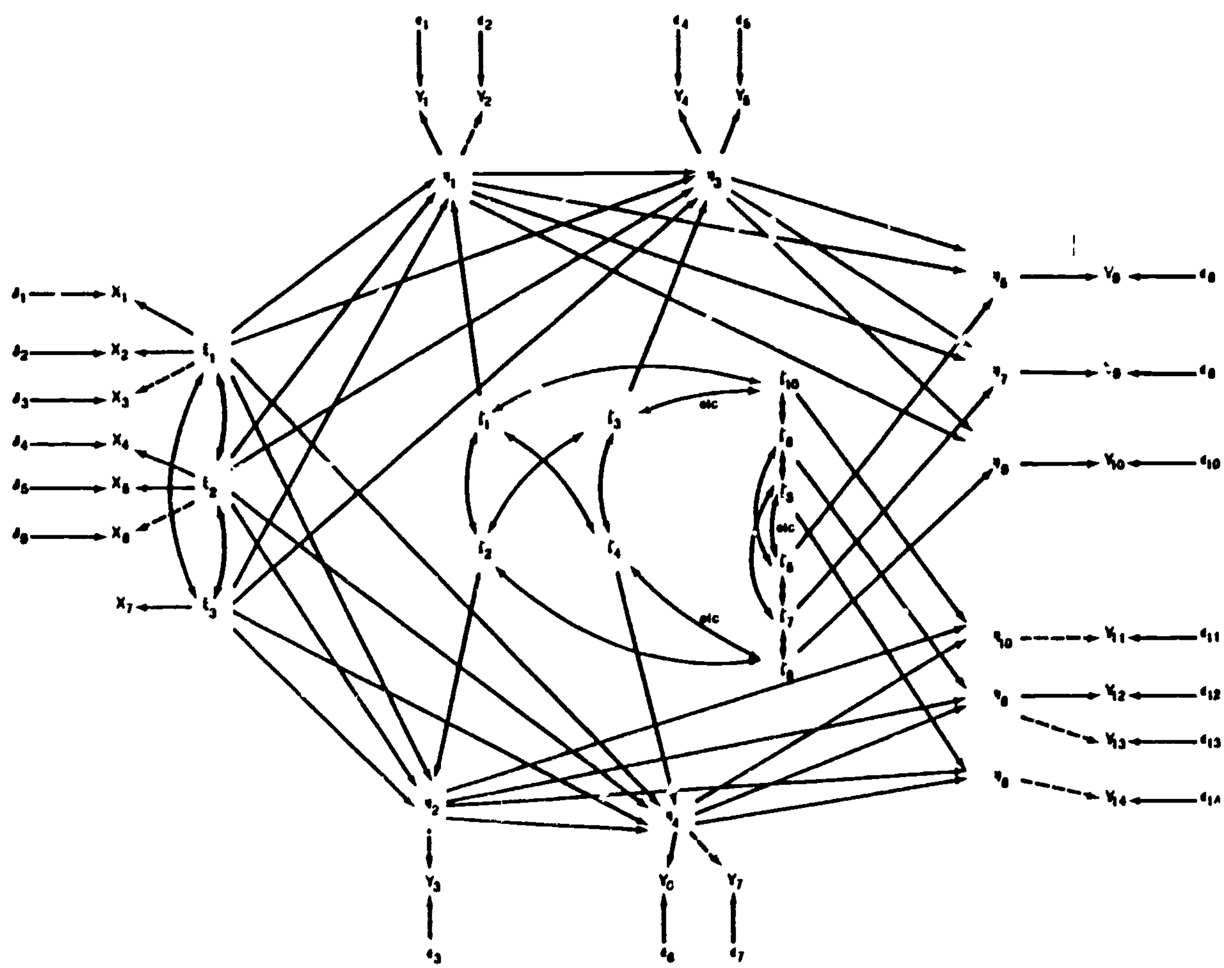
Note. Variables are MA = mental ability, Ed = years of schooling, 1^{st} Occ = status of first, full-time occupation, Cur Occ = status of current occupation, Earn = natural log of annual earnings.

Table 7. Selected slope estimates and measures of fit: Kalamazoo brothers

Antecedent variable:	MA	MA	Ed	MA	Ed	MA	Ed
Dependent variable:	Ed	1st Occ	1st Occ	Cur Occ	Cur Occ	Earn	Earn
Symmetric, heterogeneous slopes:							
Within-family	0.728	-0.066	0.680	0.235	0.414	1.068	0.321
	(0.102)	(0.107)	(0.063)	(0.121)	(0.071)	(0.258)	(0.151)
Between-family	1.[illegible]6	0.166	0.490	0.180	0.607	-0.484	1.070
	(0.1[illegible]1)	(0.199)	(0.126)	(0.227)	(0.145)	(0.545)	(0.344)
Fit: LR statistic	44.69	0.00	0.00	0.00	0.00	0.00	0.00
Degrees of freedom	14	0	0	0	0	0	0
Symmetric, homogeneous slopes:							
Within- and between-family	0.860	-0.006	0.624	0.242	0.478	0.596	0.511
	(0.065)	(0.063)	(0.038)	(0.070)	(0.043)	(0.155)	(0.095)
Fit: LR statistic	47.52	4.15	4.15	4.74	4.74	8.25	8.25
Degrees of freedom	17	3	3	3	3	3	3
LR statistic (with 1 df) for homogeneity of given slope:	na	0.72	1.25	0.03	1.02	5.41	3.28

Note: Variables are MA = mental ability, Ed = years of schooling, 1st Occ = status of first, full-time occupation, Cur Occ = status of current occupation, Earn = natural log of annual earnings. Findings are based upon maximum likelihood estimates for Kalamazoo brother pairs (N = 346), corrected for attenuation using Olneck's (1976:165) estimates of reliability. Parenthetic entries are approximate standard errors.

Figure 1. Reduced form model of socioeconomic achievement: Wisconsin brothers



Note. Paths from $\xi_1, \ldots, \xi_3$ to $\eta_5, \ldots, \eta_{10}$ are not shown, nor are certain disturbance covariances. See text for explanation.

Figure 2. Structural model of fraternal resemblance in occupational status: Wisconsin brothers