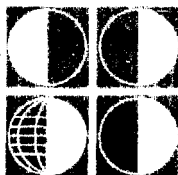


Gender, Family Configuration, and the Effect of Family Background on Educational Attainment



Hsiang-Hui Daphne Kuo and Robert M. Hauser

*Center for Demography and Ecology
Department of Sociology
University of Wisconsin
Madison, Wisconsin 53706*

ABSTRACT: A comprehensive model of family influences on educational resemblance of siblings expands the traditional sibling pair model to a full sibship model in order to investigate how gender, gender composition of sibships, and a measure of ordinal position moderate the effect of social origins on educational attainments of siblings. One common family factor is sufficient to explain the variation of educational attainment among brothers and sisters. Although effects of social origin variables on brothers are larger than on sisters, the relative effects of measured social origins are virtually the same among sisters and brothers. The disparity between educational attainments of brothers and sisters persists across sex composition and family size. Ordinal position does not alter the effects of social origins on educational attainment nor does it directly affect educational attainment. Father's and mother's education are equally important for all siblings regardless of birth order, gender composition, and family size.

We propose a social-structural model of family influence on educational resemblance among siblings. It expands the traditional sibling pair model to include full sibships. In this way, we can show how gender, gender composition of sibship, and birth order moderate the effect of social origins on educational attainment. Thus, the model builds on recent studies of sibling resemblance to analyze effects of family configuration.

Sibling resemblance and the effects of family configuration have long fascinated social scientists. Studies of sibling resemblance and differentiation can answer two kinds of questions (Sewell and Hauser, 1977). First, researchers are interested in distinguishing variation within a family from variation between families; they study how much more siblings are similar to each other than to unrelated persons (Benin

and Johnson, 1984; Hauser and Wong, 1989; DeGraaf and Huinink, 1992). Second, researchers are interested in differences between siblings; they look at the influence of variables on which siblings do not have common values, for example, birth order, sex, and birth spacing (Adams, 1972; Hauser and Sewell, 1985; Retherford and Sewell, 1991). Models of sibling resemblance address the first question while models of family configuration address the second.

Most sibling resemblance models focus on modeling and identifying different components of social origins, e.g., within-family variation in ability or between-family variation in social and economic standing, rather than investigating effects of other elements of family structure, such as birth order and sex of a sibling. About a decade ago, researchers first looked for differ-

ent effects of common family background on status attainment of members of the same sibship, for example, differences between brothers and sisters in the effect of social origins on educational attainment.

Benin and Johnson (1984) reported that social origins had larger effects on the educational attainments of sisters than on those of brothers in two small Nebraska samples, but Hauser and Wong's (1989) reanalysis of the Nebraska data showed that the gender differences were explained by the lower variability in women's schooling. That is, invariant effects of background on the schooling of brothers and sisters explained a larger share of the variance in women's schooling because there was less variance in sisters' than in brothers' schooling within families.

Previous studies have also examined variation in the effects of social background by relative ordinal position within sibling pairs.¹ In some populations, the effect of family background on educational attainment has been less among younger than among older siblings (Hauser and Wong, 1989; Dronkers, 1988), but in others there has been no birth order difference in the effect of family background (De Graaf and Huinink, 1992; Hsueh, 1992).

These findings are incomplete. *Family environment includes all ele-*

ments of family configuration, but the limitation of analyses to sibling pairs ignores some possible effects, e.g., that of the gender composition of the sibship or those of specific positions in the birth order. Here, we employ a data set with information on educational attainment in full sibships to model the resemblance among siblings and look for some effects of family configuration.

SIBLING RESEMBLANCE MODELS: METHODOLOGICAL ISSUES

The distinct advantage of the sibling resemblance model is methodological. The unit of analysis in the classical status attainment model is an individual in the general population (Blau and Duncan, 1967), so the model cannot properly specify either within-family or between-family effects. For example, effects of birth order are typically estimated in samples of persons from different families, rather than from the same family, so birth order may be confounded with other, between-family effects. In the Wisconsin Longitudinal Study (WLS), variations in educational attainment by birth order were far different in full sibships than among the original respondents who graduated from high school in 1957 (Hauser and Sewell, 1985, pp. 9-11).

At the same time, as noted by Bowles (1972), because individual data do not allow the complete specification of the relevant social background of individuals, existing estimates of the role of schooling in the intergenerational transfer of economic status may be biased upward. That is, no matter how

¹These studies have contrasted effects of family background on older and younger siblings within each possible pair of siblings; thus, they do not pertain to effects of birth order, strictly defined. Because of the very large number of combinations of birth order and gender within large sibships, it is very difficult to use a strict definition of birth order in these analyses. Throughout this analysis, we have followed a similar convention. Our references to "birth order" actually pertain to relative ordinal position among siblings of the same gender.

many social background variables—paternal and maternal schooling, occupation, income, race, region, etc.—one puts into a model, some relevant common family factors are probably left out.

By specifying one or more common, unmeasured family factors, a model of sibling resemblance can meet this criticism. However, problems of unreliable measurement loom larger in such models (Hauser and Mossel, 1985; Hauser and Mossel, 1987), and other omitted variable problems remain. For example, a within-family regression of occupational status on educational attainment may be biased upward if ability is not controlled.

Olneck (1976, 1977, 1979) has applied sibling resemblance models with a latent common family factor to data from the 1962 Occupational Changes in a Generation Survey (OCG) and from his survey of Kalamazoo brothers, and he finds relatively small biases in the effects of educational attainment on occupational status and earnings. Similarly, using Wisconsin sibling data, Hauser and Mossel (1985, 1987) have found little family bias in the effect of educational attainment on occupational status, and Hauser and Sewell (1986) reconfirmed these findings, both in the Wisconsin and Kalamazoo data, while extending the Wisconsin findings to include earnings as well as occupational status. However, these analyses have been limited to similarities between pairs of brothers.

SEX DIFFERENCES BETWEEN SIBLINGS

Although brother pairs are sufficient to identify models of sibling resemblance, it is necessary to estimate models among pairs of sisters or of

mixed sex in order to increase the generality of previous findings and to look for effects of gender and birth order. By comparing the residual covariances between siblings' educations across groups of sibling pairs that differ in gender and birth order composition, Benin and Johnson (1984) argued that brother pairs resemble one another more than do sister pairs or brother-sister pairs. Through role-modeling and facilitation, they argued, like-sex siblings would influence one another more than opposite-sex siblings, and older brothers would have more influence than older sisters. Thus, pairs of older and younger brothers should show the greatest resemblance, net of social background, while pairs of older sisters and younger brothers should show the least resemblance.

Benin and Johnson's analysis of two Nebraska sibling samples suffers from methodological and substantive problems (Hauser and Wong, 1989, pp. 152–156). First, a common family factor is not specified in their model, but only in their verbal proposition. Their analysis was actually based on unrestricted regressions of educational attainment on social background in each sibling group. Second, Benin and Johnson's evidence was both inappropriate and weak. Their cross-group comparisons among residual covariances could not support their argument about "cross-sibling effects," because the covariances are irrelevant to the identification of cross-sibling effects.

Hauser and Wong (1989, pp. 156–160) reanalyzed the Nebraska data using a MIMIC (multiple-indicator, multiple-cause) model and found that the differences in residual covariances between sibling pairs of different gen-

der composition were insignificant, excepting a low level of resemblance among pairs of older sisters and younger brothers. They also analyzed Dutch and German sibling samples and found no evidence to support the Nebraska findings. Finally, they analyzed data for Kalamazoo brother pairs, using academic ability and achievement as instrumental variables, and they directly estimated reciprocal influences of brothers' educational attainments. While Benin and Johnson had assumed a predominant flow of influence from older to younger siblings, and Hauser and Wong found this pattern in the Kalamazoo data, their estimate was not significantly larger than the reverse effect from younger to older brothers.²

Because the data were limited to pairs of brothers, Hauser and Wong's analysis of the Kalamazoo data could not address differences in cross-sibling effects between like-sex and opposite-sex pairs. To address this limitation, Lee (1989) analyzed groups of sibling pairs drawn from the Wisconsin Longitudinal Study, where the groups of pairs were constructed as by Benin and Johnson. She estimated a model similar to that of Hauser and Wong, but she used measured ability alone as an instrumental variable to estimate cross-sibling influence. Her analysis was limited to a subsample of about 2,000 pairs in the WLS data in which test score data had been collected, and the sibling had been interviewed directly. The subsampling design, when combined with survey and item nonresponse of

both the original respondent and the sibling, led to a substantial loss of statistical power. Lee found no reciprocal effects between older siblings and the younger brother, but positive reciprocal effects between older siblings and the younger sister. A common family factor had the same effect on all sibling pairs, except the all-sister pair. Finally, the effect of measured ability on a brother's educational attainment was significantly larger than its effect on a sister's attainment.

FAMILY CONFIGURATION

When we use data for sibling pairs from sibships with different gender compositions, the effect of gender is confounded with effects of other elements of family configuration. First, among randomly selected pairs of siblings, we cannot distinguish the effect of birth order from that of gender in a mixed-sex pair. Second, the likelihood of choosing a pair of brothers in a random sample is, of course, higher for families with sibships with more brothers, while the opposite holds true for sister pairs. Thus, differences among brother pairs, sister pairs, and brother-sister pairs may result from the differences in the size and gender composition of sibships or from differences in ordinal position.

FAMILY SIZE AND BIRTH ORDER

Many studies focus on the effect of birth order and family size on intelligence, personality traits, or educational outcomes. Most early work had two serious flaws: use of small samples not selected from the general population and a failure to control other variables of family configuration which are

²In this analysis, we specify only reduced-form models of the effects of social origins on schooling. In principle, these effects could be extended to include cross-sibling effects in the manner of Hauser and Wong (1989).

confounded with sibling position (Adams, 1972). Since the 1970's, there have been new theories of birth-order and family-size effects, along with better data and research designs.

The confluence theory, proposed by Zajonc and Markus (1975; Zajonc, 1976), argues that the quality of the intellectual environment of a given child is a complex function of the intelligence of other family members and consequent opportunities to learn from and teach other siblings. Short birth intervals and a large family have negative effects on the average intellectual environment of a child. Even though the intellectual environment of the only child or last child is relatively high, because other family members have higher intelligence than s/he has, the absence of a chance to teach younger siblings depresses intellectual development.³

Lindert (1977) argues that when birth spacing is controlled, the investment of parental resources, e.g., time and money, in a child varies by birth order and thus influences the child's achievement. Because of the absence of other competing children, first-born, last-born, and only children do better than other children in the family, and this difference decreases with closer birth spacing. The argument is known as resource dilution theory. The theory explains the findings of Blau and Duncan's (1967) study using sib-

ling data from the 1962 Occupational Changes in a Generation (OCG) survey. First-born and last-born men in large families (with 3 or more siblings) have greater educational and occupational attainments than their brothers in the middle of the sibship. Lindert's own study, based on a sample of 1,087 siblings from a nonrandom sample of New Jersey executives (Hermalin, 1969), confirmed this linkage between sibling position and education.

After Lindert, there has been no strong and consistent empirical support of birth order effects on educational and occupational attainment. Wright (1977) used 1962 OCG data to test the confluence and resource dilution theories, and her regression analysis supports neither of them. She found a small effect of birth order only on educational attainment, and this tendency increases with sibship size for last-born children. Similarly, Olneck and Bills (1979) found no support for either the confluence theory or the resource dilution theory in analyses of the Kalamazoo brother data. They conclude that the family-size effect persists net of the effect of socioeconomic background, but the birth-order effect disappears when brothers are compared with one another.

Examining 1975 Wisconsin high school graduates and their siblings, Hauser and Sewell (1985) found a substantial negative effect of sibship size, but no significant or systematic effects of birth order on schooling when social origins were controlled. When family size is controlled, years of schooling increase with birth order. However, the gain of schooling coincides with inter-cohort gains of educational attainment in the general population between 1930

³Several studies evaluating the confluence theory have failed to confirm Zajonc et al.'s (1975, 1976) findings. Zajonc and his colleagues have attributed their failure to wrong methods and data, but we believe that evidence strongly favors the critics of the confluence theory. See Steelman (1985, 1986) versus Zajonc (1986) and Retherford and Sewell (1991, 1992) versus Zajonc et al. (1991).

and 1950. They conclude that there are virtually no birth-order effects.

Powell and Steelman (1990) analyzed data from the High School and Beyond survey and found that closer spacing strengthens the negative effect of sibship size. However, their failure to control for birth order, which is one of the concerns of the confluence model and resource dilution theory, confounds the effects of other family configuration variables with those of birth order, including the disproportionate occurrence of lower birth orders in a sample of the general population.

GENDER COMPOSITION

Gender composition is another important characteristic of family configuration. Brim (1958) found that the gender composition of siblings in a two-child family influenced the personality traits of both children. Because the sample is limited to two-child families, Brim cautions against the application of his findings to siblings from families of other sizes.

There are two theories to illustrate the process of influence: role models and role expectations. The former argues that parents and siblings of the same gender serve as role models for a child; that is, a child assimilates his or her behaviors to the parent or sibling of the same gender. On the other hand, role expectation theory suggests that a child interacts with others according to others' (social) expectations of his/her behaviors, and siblings on both sides of the interaction are aware of the expectations overtly or covertly. Thus, a setting with all other siblings of the same sex is different from the setting with all other siblings of the opposite sex.

Lee's (1989) findings are consistent with the role expectation theory: The traditional female role is submissive and passive to authority, the older siblings. Thus, the younger sister rather than the younger brother is influenced by older siblings, and the expectation is mutual. In Powell and Steelman (1990), two variables are constructed to represent gender composition: one is number of brothers, and the other is number of sisters. The number of brothers has a negative effect on children's educational achievement, and that of sisters has an inconsistent effect.⁴

Taken as a whole, most of the available evidence shows that the effects of social origins on brothers and sisters are very similar. One exception is Lee's finding that sisters are less influenced by family background than their brothers. The question of reciprocal influence between siblings remains unresolved, and the present analysis will not attempt to provide further evidence about it. There is no consistent evidence of effects of birth order on educational attainment, and there has been no definitive test of the effect of gender composition of sibships on educational attainment.

In this paper, our goal is to look more closely at effects of social origins on educational attainment in reduced-form models, using the identifying information in data from full sibships to provide new evidence about the factorial structure of measured social background and schooling and stronger evidence about differences in family

⁴However, Powell and Steelman did not test whether the effects of number of brothers and of number of sisters were significantly different from one another.

background effects by gender, gender composition, birth order, and size of sibship. We begin by asking whether there is more than one common family factor in the educational attainment of siblings, and whether those factors may differ between brothers and sisters.

ONE-FACTOR AND TWO-FACTOR MODELS OF SIBLING RESEMBLANCE

If there were different effects of social origins on brothers and sisters, then there might, but need not be, two distinct family factors in the educational attainments of sisters and brothers.⁵ A one-factor model is identified with only one measured background variable in a set of sibling pairs whose educational attainments are known, and it is thus possible to distinguish between-family variation from within-family variation. With a second measured background variable, a single factor model can be rejected (Hauser and Goldberger, 1971), but in an analysis of sibling pairs it may not be clear whether the rejection could be explained by sex differences in the effect of background. However, with two or more background variables and measures of a single outcome for each member of sibships larger than two, it is possible both to reject a one-factor model and to determine whether sex

differences account for the emergence of a second factor.

Without loss of generality, consider the case of two social background variables, ξ_1 and ξ_2 , and sibling outcomes, $\eta_1, \eta_2, \dots, \eta_m$, where $\eta_1, \eta_2, \dots, \eta_m$ pertain to sisters, and $\eta_{m+1}, \eta_{m+2}, \dots, \eta_m$ pertain to brothers, arrayed by birth order within sex. In the one-factor model, the reduced form equations,

$$\begin{aligned} \eta_1 &= \pi_{1,1} \xi_1 + \pi_{1,2} \xi_2 + \varepsilon_1 \\ \eta_2 &= \pi_{2,1} \xi_1 + \pi_{2,2} \xi_2 + \varepsilon_2 \\ &\vdots \\ \eta_m &= \pi_{m,1} \xi_1 + \pi_{m,2} \xi_2 + \varepsilon_m \end{aligned} \quad (1)$$

will satisfy the constraints,

$$\frac{\pi_{1,1}}{\pi_{1,2}} = \frac{\pi_{2,1}}{\pi_{2,2}} = \dots = \frac{\pi_{m,1}}{\pi_{m,2}} = \kappa \quad (2)$$

This model can be rejected with as few as two background variables and an outcome for two siblings. However, rejection of the model does not tell us, for example, whether there is a consistent difference in $\pi_{i,1}/\pi_{i,2}$ between brothers and sisters, nor can it tell us whether the gender composition of the sibship affects the ratios of the reduced form coefficients.

Rejection of the constraint on the ratios of reduced-form coefficients in Equation 2 need not imply that similar constraints do not hold for subgroups of siblings. For example, let κ_s and κ_b be the ratios of reduced-form coefficients among sisters and brothers. It could be that

$$\begin{aligned} \frac{\pi_{1,1}}{\pi_{1,2}} &= \dots = \frac{\pi_{m^*,1}}{\pi_{m^*,2}} = \kappa_s \\ \text{and} \quad \frac{\pi_{m^*+1,1}}{\pi_{m^*+1,2}} &= \dots = \frac{\pi_{m,1}}{\pi_{m,2}} = \kappa_b \end{aligned} \quad (3)$$

⁵Chamberlain and Griliches (1977) suggested a two-factor model of sibling resemblance in earnings among the young men in the National Longitudinal Survey of Labor Market Experience, who were 14 to 24 years old in 1966. Because they used sibling pair data, they had to impose arbitrary identifying constraints to identify their model. Our model is less restrictive because we observe full sibships. Also, we are interested in differences in the effect of family background between brothers and sisters, while Chamberlain and Griliches (1977) were attempting to identify multiple family factors in the achievement of brothers.

where $\kappa_b \neq \kappa_s$, which says that a proportionality constraint holds among sisters and among brothers, but the two constraints are not the same. For example, the effect of education of the same-sex parent may be larger than that of the opposite-sex parent, both among brothers and among sisters. In this situation, a one-factor model of family background would be rejected in full sibships, even though it would hold for sibling pairs that were homogeneous in gender composition.⁶

Note that Equation 2 does not imply that the effect of the background variables is the same for brothers and sisters. For example, suppose that $\pi_{i,1} = \gamma_{i,1}$ and $\pi_{i,2} = \gamma_{i,2}$ for $i \leq m^*$, while $\pi_{i,1} = \gamma_{i,1}\lambda_b$ and $\pi_{i,2} = \gamma_{i,2}\lambda_b$ for $i > m^*$. In this case, the model of Equation 2 should not be rejected, but the effect of social background is either uniformly more or less for sisters or brothers, depending on the value of λ_b . That is, for any sister, s , or brother, b , in any birth order,

$$\begin{aligned} \eta_s &= \gamma_{1,1} \xi_1 + \gamma_{1,2} \xi_2 + \varepsilon_s \\ \eta_b &= \gamma_{1,1} \lambda_b \xi_1 + \gamma_{1,2} \lambda_b \xi_2 + \varepsilon_b \end{aligned} \quad (4)$$

and the effects of background differ by λ_b , even though $\kappa_s = \kappa_b$.

Up to this point, our reference to one- and two-factor models pertains only to the number of factors required to describe the effects of social background variables on educational attainment. That is, we have ignored the covariance structure of the disturbances

(ε) in individual educational attainment. If there are two or more family background variables, it would be possible, in sibships of three or more, to reject the hypothesis that there is a single common factor in educational attainment, even when there is only one common family background factor.

For example, consider the following model of attainment, where ξ_1 and ξ_2 are social background variables; y_1 , y_2 , and y_3 are the educational attainments of three siblings; ε_1 , ε_2 , and ε_3 are disturbances in attainment; and η_1 is a common family factor:

$$\begin{aligned} \eta_1 &= \gamma_{11} \xi_1 + \gamma_{12} \xi_2 \\ y_1 &= \eta_1 + \varepsilon_1 \\ y_2 &= \lambda_{21} \eta_1 + \varepsilon_2 \\ y_3 &= \lambda_{31} \eta_1 + \varepsilon_3 \end{aligned} \quad (5)$$

where we normalize the coefficients by setting $\lambda_{11} = 1$. This model is over-identified with two restrictions, even if we place no constraints on $\theta^{e_{ij}} = \text{cov}(\varepsilon_i, \varepsilon_j)$. However, if we respecify the model to introduce an unmeasured common family factor, ζ_1 , so

$$\eta_1 = \gamma_{11} \xi_1 + \gamma_{12} \xi_2 + \zeta_1, \quad (6)$$

where $\theta^{e_{12}} = \theta^{e_{13}} = \theta^{e_{23}} = 0$, then the model forces η_1 both to mediate all effects of social background and to account for the covariances among the siblings' educational attainments, which adds two more over-identifying restrictions, relative to the model of Equation 5. Thus, one could reject the hypothesis that the family background factor accounts entirely for the covariances among siblings' levels of completed schooling, even when there is a single family background factor. We will also test this stronger version of the single-factor model.

⁶If $\kappa_s \neq \kappa_b$ in Equation 3, then the model could be rejected in an analysis of opposite sex pairs, but that could not in itself tell us whether gender differences were responsible because gender and birth order could be confounded.

OVERVIEW OF THE ANALYSIS

In the next section, we introduce the full sibship data from the Wisconsin Longitudinal Study and describe the variables used in this analysis (See figure 1). Then, we introduce a multiple-group factor model of schooling in order to carry out one test of the hypothesis that there is more than one common factor in siblings' educational attainments. Next, we add social background variables and specify a multiple-group Multiple-Indicator-and-Multiple-Cause (MIMIC) model of schooling. This model permits a second test of the one-factor hypothesis, against the alternative that more than one factor is required to mediate the effects of social background on schooling. We have, also, refined this model to specify differences in the effect of social origins by gender and birth order. Then, we present estimates of several parameters for sibships of size 3, 4, and 5: effects of exogenous variables, variances of shared unmeasured family background, and nonshared variances of siblings' educational attainments by gender.

In the next section, we expand this model to contrast the one-factor hypothesis with a specific alternative, namely, that the family background factors differ between brothers and sisters. Finally, we estimate effects of birth order and gender composition on mean levels of educational attainment. In the final section, we summarize our findings.

MATERIALS AND METHODS

The Wisconsin Longitudinal Study (WLS) began with a survey of all high-school seniors in Wisconsin's class of

1957. Later, one-third of the original respondents were selected at random for further study. In 1964, a brief follow-up was conducted by mailing parents a postcard questionnaire in order to update the social and economic situation of 1957 respondents. In 1975, the original respondents were interviewed by telephone in a second follow-up, which collected information about social background, occupation, education, marriage, children, and social activities. Data on age, sex, and educational attainment were collected for all living siblings of the respondent. At the same time, one sibling was randomly selected from each full sibship roster. In 1977, a subsample of about 2,000 of the selected siblings were interviewed, using essentially the same questionnaire as in 1975. Most of the previous studies of sibling resemblance in the WLS have used only data for this subsample of sibling pairs (Hauser and Mossel, 1985, 1987; Hauser, 1984, 1988; Hauser and Wong, 1989; Lee, 1989). Hauser and Sewell (1985) used the WLS sibling rosters to study the effects of birth order on levels of educational attainment, but they did not attempt to model sibling resemblance.

We use the WLS data here because they contain the educational attainment of all siblings of the primary respondents who were living in 1975. To test the two-factor model, data for more than two siblings are needed for model identification. To investigate effects of family configuration, data from full sibship rosters permit us to map the effects of birth order and gender composition. It is well established that the educational data from the WLS are highly reliable, both for the original respondents and for their siblings

FIGURE 1
VARIABLE DESCRIPTIONS

Variable	Source	Description
Offspring's Education	CFOED1, CFOED2, CFOED3, CFOED4. (and CFOED5, if sibship size = 5) 1975 survey	Years of school completed
Family Income	Constructed Variable. (BMPIN2) Taking information from Wisconsin tax data (PI5760) first; if not available, taking from 1975 report (YFML57)	Per \$1000. Truncated at \$15,000. Best Measure of Parental Income 1957
Father's Occupation	Constructed variable. (BMFOC1) Taking from 1975 report (OCSH57) first; if not available, taking from 1957 report (OCSF57)	Duncan Scores Best Measure of Household Head's Occupation
Father's Education	Constructed variable. (BMFAED) Taking from 1975 report (EDHHYR) first; if not available, taking from 1957 report (EDFA57)	Years of School Completed Best Measure of Father's Education
Mother's Education	Constructed variable. (BMMAED) Taking from 1975 report (EDMOYR), first; if not available, taking from 1957 report (EDMO57). (BMMAED)	Years of School Completed Best Measure of Mother's Education
Catholic	Recoded variable from RELFML (1975 survey)	1: Catholic (1, 2, and 3 in RELFML); 0: otherwise
Farm Background	Recoded variable from OCMH57 (1957 survey)	1: Farmer/farm managers/farm laborers and foremen (16 and 17 in OCMH57); 0: otherwise

Note: Variable mnemonics refer to public use file for the Wisconsin Longitudinal Study.

(Hauser et al., 1983; Hauser and Mossel, 1985, 1987). Also, because the WLS data have been used extensively in previous studies of sibling resemblance, we can easily compare our results with previous findings.

Table 1 shows the distribution of sibship size and the gender composition of 9,081 WLS respondents in the 1975 survey who reported their numbers of siblings.⁷ Sixty-two per cent of respondents have 2 or more siblings. Almost half of respondents who answer the sibling roster have 2 to 4 siblings; that is, 49 per cent of the sample

is from sibships of size 3, 4, or 5. Note that, even in sibships of 5 or 6 persons, the number of families (of original respondents) with all boys or all girls is quite small. We start the analyses with sibship size 3, which is the minimum sibship size needed to identify the two-factor model. Owing to the wide range of siblings' ages, we limit our analysis to siblings aged 21 to 55 in 1975. This limitation may eliminate some effects of wide birth spacing, but it also eliminates persons who probably had not completed their schooling or who were unlikely to be biological siblings of the original respondents. Finally, we have a sample of 1,790 cases for sibship size 3; 1,178 for sibship size 4; and 785 for

⁷In this analysis, sibship size includes the respondent. It is one more than the respondent's number of brothers and sisters.

TABLE 1
DISTRIBUTION OF SIBSHIP SIZE AND GENDER COMPOSITION: WISCONSIN LONGITUDINAL STUDY^a

NUMBER OF SISTERS	Sibship Size							Total
	1	2	3	4	5	6	7+	
0.....	320 (46.9)	425 (21.3)	241 (12.1)	105 (7.2)	29 (2.8)	19 (2.8)	15 (1.2)	1,154 (12.7)
1.....	362 (53.1)	1,051 (52.6)	742 (37.3)	367 (25.2)	169 (16.4)	67 (9.8)	41 (3.3)	2,799 (30.8)
2.....	522 (26.1)	772 (38.8)	542 (37.1)	321 (31.1)	154 (22.4)	122 (9.9)	2,433 (26.8)
3.....	233 (11.7)	368 (25.2)	313 (30.3)	191 (27.8)	243 (19.7)	1,348 (14.8)
4.....	77 (5.3)	169 (16.4)	170 (24.8)	305 (24.7)	721 (7.9)
5.....	31 (3.0)	73 (10.6)	256 (20.7)	360 (4.0)
6+.....	12 (1.7)	254 (20.6)	266 (2.9)
Total.....	682 (7.5)	1,998 (22.0)	1,988 (21.9)	1,459 (16.1)	1,032 (11.4)	686 (7.6)	1,236 (13.6)	9,081 (100.0)

^aPer cents in parentheses.

sibship size 5. These numbers represent 90 per cent of the original respondents in sibship size 3; 81 per cent in sibship size 4; and 76 per cent in sibship size 5. We have not included sibships of 6 or more in this analysis because of the small number of observations in many of the gender combinations.

We have defined subgroups and endogenous variables for the analysis by the gender composition of the sibship and by the arrangement of relative birth order within sex. We first group the sample of each sibship size by gender composition. Thus, there are four subgroups for sibship size 3, five for sibship size 4, and six for sibship size 5. Next, we divide siblings in each group by sex, and within each sex, place them by the ascending order of birth, i.e., from the oldest to the youngest. For example, in a two-sister sibship of size four, the order is the oldest sister, the youngest sister, the oldest brother, and

the youngest brother. Thus, our design does not identify effects of birth order, per se, but only of relative ordinal position within same-sex siblings, except in the case of all-female or all-male sibships. We use this simplified specification of birth order because of the very large number of possible combinations of birth order with gender composition. Figure 2 gives the example of our model for sibship size 4.

Table 2 gives means and standard deviations of the measured endogenous variables, that is, the years of schooling of siblings, by size of sibship (vertical panel), gender (horizontal panels), and relative birth order (rows within horizontal panels). Thus, in all-male sibships of size 3, shown in the first column of the table, the mean years of schooling completed are 13.87 for first-born sons, 13.68 for second-born sons, and 13.62 for last-born sons. In sibships of the same size with one

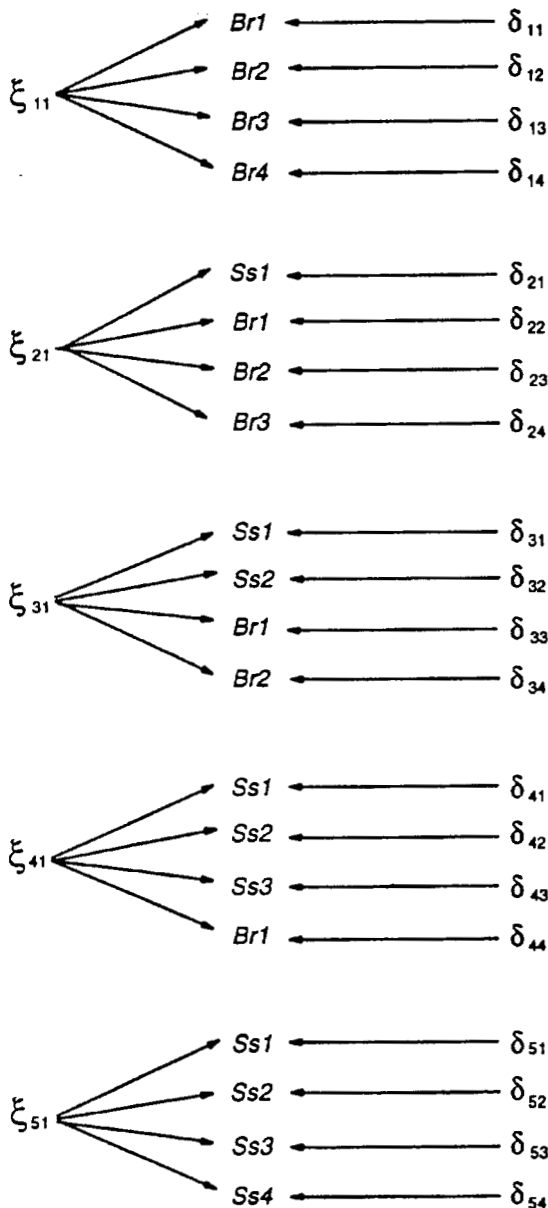


Fig. 2.—Multiple Group One-Factor Model (sibship size = 4)

TABLE 2
DESCRIPTIVE STATISTICS OF VARIABLES*

VARIABLES	SIBSHIP SIZE = 3				SIBSHIP SIZE = 4					SIBSHIP SIZE = 5					
	1	2	3	4	1	2	3	4	5	1	2	3	4	5	6
Endogenous (Yrs of School)															
Sister 1	13.30	13.11	13.12	...	13.08	12.93	12.90	12.76	...	12.94	12.62	12.82	12.51	12.37	
	(2.02)	(2.05)	(2.05)	...	(2.04)	(2.11)	(2.07)	(2.01)	...	(1.87)	(1.93)	(2.01)	(1.87)	(1.92)	
Sister 2	13.21	13.06	13.00	12.81	12.86	12.90	12.90	12.58	12.56	
	...	(2.09)	(2.08)	(1.88)	(2.12)	(1.91)	(1.83)	(2.00)	(1.94)	(1.60)	
Sister 3	13.17	12.95	12.70	12.87	12.65	12.74	
	(2.12)	(1.91)	(1.63)	(1.87)	(1.93)	(1.61)	
Sister 4	12.90	12.72	13.04	
	(1.65)	(1.75)	(1.93)	
Sister 5	12.63	
	(1.76)	
Brother 1	13.87	13.96	13.61	...	12.94	13.45	13.36	13.39	...	12.27	12.55	13.01	13.18	13.13	...
	(2.94)	(2.86)	(2.71)	...	(2.17)	(2.70)	(2.79)	(2.57)	...	(2.90)	(2.54)	(2.82)	(2.64)	(2.37)	...
Brother 2	13.68	13.95	12.97	13.65	13.46	12.73	12.95	12.90	13.53
	(2.76)	(2.63)	(2.33)	(2.84)	(2.59)	(2.25)	(2.29)	(2.32)	(2.46)
Brother 3	13.62	13.56	13.41	12.59	13.23	13.22
	(2.37)	(2.65)	(2.30)	(2.46)	(2.44)	(2.25)
Brother 4	13.10	12.91	13.22
	(2.66)	(2.39)	(2.44)
Brother 5	13.55
	(1.99)
Exogenous															
Income	5.84	6.52	6.18	6.14	5.62	6.03	5.88	5.91	5.76	4.95	5.05	5.31	5.39	5.84	4.86
	(2.88)	(3.36)	(3.31)	(3.17)	(2.92)	(3.33)	(3.33)	(3.15)	(2.71)	(2.16)	(3.17)	(3.04)	(2.95)	(3.35)	(2.38)
Father's Occup.	3.49	3.96	3.65	3.59	3.08	3.54	3.23	3.46	3.58	2.32	3.21	3.16	3.14	3.14	2.58
	(2.28)	(2.49)	(2.38)	(2.17)	(1.78)	(2.35)	(2.21)	(2.28)	(2.24)	(1.17)	(2.16)	(2.27)	(2.25)	(2.22)	(1.80)
Father's Educ.	9.87	10.36	10.16	9.86	9.17	9.75	9.50	9.75	9.35	8.32	9.41	9.02	9.47	9.50	9.67
	(3.42)	(3.55)	(3.40)	(3.15)	(2.86)	(3.65)	(3.37)	(3.24)	(3.21)	(2.19)	(3.51)	(3.27)	(3.36)	(3.50)	(3.20)
Mother's Educ.	10.72	10.93	10.66	10.28	10.00	10.60	10.39	10.29	10.54	9.50	10.13	10.09	10.45	10.03	9.78
	(2.63)	(2.75)	(2.90)	(2.70)	(2.77)	(2.81)	(2.86)	(2.69)	(2.42)	(3.02)	(3.14)	(2.80)	(2.64)	(2.79)	(2.22)
Catholic	0.39	0.38	0.38	0.41	0.52	0.43	0.45	0.41	0.48	0.50	0.37	0.47	0.45	0.49	0.56
	(0.49)	(0.49)	(0.49)	(0.49)	(0.50)	(0.50)	(0.50)	(0.49)	(0.50)	(0.51)	(0.48)	(0.50)	(0.50)	(0.50)	(0.51)
Farmer	0.19	0.15	0.18	0.16	0.23	0.20	0.20	0.21	0.19	0.23	0.29	0.26	0.23	0.23	0.26
	(0.39)	(0.36)	(0.38)	(0.37)	(0.42)	(0.40)	(0.40)	(0.41)	(0.40)	(0.43)	(0.46)	(0.44)	(0.42)	(0.42)	(0.45)

*Standard deviations in parentheses.

sister, the mean sister's education is 13.30 years, while the mean schooling levels are just below 14 years both for older and younger brothers. Our design does not distinguish among the three possible ordinal positions of the sister in this configuration.

Inspection of Table 2 suggests the consistency of the data with previous findings. Siblings from smaller families finish more schooling, and within family sizes and configurations, men usually finish more schooling than women. Also, there is consistently less variability in the schooling of sisters than of brothers. There is no clear pattern to educational attainment by birth order or relative birth order, although there is a hint of a positive relationship between birth order and schooling among gender-homogeneous sibships of size 5. With one exception (in sibship size 5), the education of sisters in families with only one daughter is larger than the education of sisters in all other sibling configurations of the same size, but this effect, if any, is quite small.

The exogenous variables include family income (in thousands of dollars), father's and mother's education (in years), father's occupational status (Duncan SEI score), Catholic upbringing, and farm background.

RESULTS

FACTOR MODELS

Our strategy for the comparison between two-factor and one-factor models, first, is to test the multiple-group factor model with one common family factor. Second, we report analyses of the effects of family configuration in

the one-factor MIMIC model.⁸ Third, we evaluate the findings in the first two sections by developing and testing a two-factor MIMIC model of educational attainment. The equation of the single factor model of educational attainment is

$$X = \Lambda^x \xi + \delta \quad (7)$$

where X is a vector of siblings' educational attainments; ξ is the common family factor; Λ^x is the matrix of loadings of X 's on ξ ; and δ is a vector of unmeasured unique factors. In matrix form, the model is⁹

$$\begin{bmatrix} x_{k1} \\ x_{k2} \\ x_{k3} \\ x_{k4} \end{bmatrix} = \begin{bmatrix} \lambda_{k,1,1} \\ \lambda_{k,2,1} \\ \lambda_{k,3,1} \\ \lambda_{k,4,1} \end{bmatrix} \begin{bmatrix} \xi_1 \end{bmatrix} + \begin{bmatrix} \delta_{k1} \\ \delta_{k2} \\ \delta_{k3} \\ \delta_{k4} \end{bmatrix} \quad (8)$$

where k indexes groups defined by gender composition, e.g., the five distinct groups in sibships of size 4. We report the test statistics in Table 3.

The baseline specification, Model A in Table 3, is a one-factor model in which the effect of one common family factor on educational attainment of brothers differs from that of sisters, and the within-family variance in schooling also differs between brothers and sisters. In all other respects, the parameters of the model are invariant. They do not differ by gender composition of sibship or by birth order

⁸In addition to L^2 , we use the *bic* statistic to evaluate goodness of fit. The *bic* statistic is based on Bayesian theory for a posteriori tests: $bic = L^2 - df \times \ln(N)$, where df are the degrees of freedom under the tested model or contrast, and N is the sample size. Satisfactory fit is indicated by a negative value of *bic*, and models with lower *bic* statistics are preferred (Raftery, 1986, 1993, 1995).

⁹Without loss of generality, we have ignored the structure of variable means. We have used the case where sibship size is 4 to illustrate model specification.

TABLE 3
GOODNESS OF FIT STATISTICS, ONE-FACTOR MODEL

Model	L ²	df	bic	Contrast	p
Sibship Size = 3 (n = 1790)					
A. Baseline Model.....	51.09	20	-98.71	...	0.000
B. A + λ ^s 's INVARIANT BY SEX.....	56.88	21	-100.41	5.79 1	0.016
C. A + Θ ^δ INVARIANT BY SEX.....	208.96	21	51.67	157.87 1	0.000
D. A - Θ ^δ ACROSS-GROUP CONSTRAINT.....	46.64	16	-73.20	-4.45 4	0.349
E. D - Θ ^δ WITHIN-GROUP CONSTRAINT.....	27.03	10	-47.87	-19.61 6	0.003
F. A - λ ^s 's ACROSS-GROUP CONSTRAINT.....	49.61	19	-92.70	-1.48 2	0.224
G. F - λ ^s 's WITHIN-GROUP CONSTRAINT.....	35.85	13	-61.52	-13.76 5	0.017
H. A - Φ ACROSS-GROUP CONSTRAINT.....	50.83	17	-76.50	-0.26 3	0.967
Sibship Size = 4 (n = 1178)					
A. Baseline Model.....	89.07	46	-236.22	...	0.000
B. A + λ ^s 's INVARIANT BY SEX.....	113.19	47	-219.17	24.12 1	0.000
C. A + Θ ^δ INVARIANT BY SEX.....	188.64	47	-143.72	99.57 1	0.000
D. A - Θ ^δ ACROSS-GROUP CONSTRAINT.....	79.16	40	-203.70	-9.91 6	0.128
E. D - Θ ^δ WITHIN-GROUP CONSTRAINT.....	43.90	28	-154.10	-35.26 12	0.001
F. A - λ ^s 's ACROSS-GROUP CONSTRAINT.....	87.81	44	-223.34	-1.26 2	0.533
G. F - λ ^s 's WITHIN-GROUP CONSTRAINT.....	53.62	32	-172.67	-34.19 11	0.000
H. A - Φ ACROSS-GROUP CONSTRAINT.....	82.33	42	-214.68	-6.74 4	0.150
Sibship Size = 5 (n = 785)					
A. Baseline Model.....	183.93	86	-389.82	...	0.000
B. A + λ ^s 's INVARIANT BY SEX.....	193.74	87	-386.17	9.81 1	0.002
C. A + Θ ^δ INVARIANT BY SEX.....	285.64	87	-294.27	101.71 1	0.000
D. A - Θ ^δ ACROSS-GROUP CONSTRAINT.....	162.06	78	-357.86	-21.87 8	0.005
E. D - Θ ^δ WITHIN-GROUP CONSTRAINT.....	121.39	58	-265.22	-40.67 20	0.004
F. A - λ ^s 's ACROSS-GROUP CONSTRAINT.....	180.35	83	-372.90	-3.58 3	0.309
G. F - λ ^s 's WITHIN-GROUP CONSTRAINT.....	149.69	63	-270.25	-30.66 19	0.044
H. A - Φ ACROSS-GROUP CONSTRAINT.....	183.23	81	-357.60	-0.70 5	0.983

within gender. For example, in sibships of size 4, we impose the following constraints: $\lambda_{111}^x = \lambda_{121}^x = \lambda_{131}^x = \lambda_{141}^x = \lambda_{221}^x = \lambda_{231}^x = \lambda_{241}^x = \lambda_{331}^x = \lambda_{341}^x = \lambda_{441}^x = 1$ and $\lambda_{211}^x = \lambda_{311}^x = \lambda_{321}^x = \lambda_{411}^x = \lambda_{421}^x = \lambda_{431}^x = \lambda_{511}^x = \lambda_{521}^x = \lambda_{531}^x = \lambda_{541}^x$; $\theta_{111}^{\delta} = \theta_{121}^{\delta} = \theta_{131}^{\delta} = \theta_{141}^{\delta} = \theta_{221}^{\delta} = \theta_{231}^{\delta} = \theta_{241}^{\delta} = \theta_{331}^{\delta} = \theta_{341}^{\delta} = \theta_{441}^{\delta}$ and $\theta_{211}^{\delta} = \theta_{311}^{\delta} = \theta_{321}^{\delta} = \theta_{411}^{\delta} = \theta_{421}^{\delta} = \theta_{431}^{\delta} = \theta_{511}^{\delta} = \theta_{521}^{\delta} = \theta_{531}^{\delta} = \theta_{541}^{\delta}$; $\phi_{111} = \phi_{211} = \phi_{311} = \phi_{411} = \phi_{511}$, where the θ_{kij}^{δ} are covariances of the δ_{ki} , and the ϕ_{k11} are variances of ξ_{k1} in the several groups. By adding or releasing constraints, we test specific

hypotheses about the effects of family background on educational attainments of brothers and sisters, all conditioning on the one-factor specification.¹⁰

In the baseline model, the goodness of fit test yields $L^2 = 51.09$ with 20 *df* for size 3, $L^2 = 89.07$ with 46 *df* for size 4, and $L^2 = 183.93$ with 86 *df*. Each of these test statistics is nominally statis-

¹⁰That is, with the exception of the additional restrictions in Models B and C, we are reporting a forward process of model selection. However, we have also selected backward from a completely unrestricted one-factor model, and this process yields the same preferred model, namely, the baseline model of Table 3.

tically significant. However, in each case, the corresponding *bic* statistic is negative, $bic = -98.71$, $bic = -236.22$ and $bic = -389.82$ for sizes 3, 4, and 5 respectively. That is, the baseline model can nominally be rejected for each size of sibship, but the fit is not bad enough to justify the loss of parsimony that rejection would entail.

Because it incorporates several restrictions, the fit of the baseline model does not provide a global test of the hypothesis that there is only one family factor. For this reason, we also estimated a completely unrestricted one-factor model in sibships of sizes 4 and 5; in sibships of size 3, the completely unrestricted model is just-identified, so there is no test of its goodness-of-fit. In sibships of size 4, the completely unrestricted model yields $L^2 = 12.75$ with 10 *df*, and in sibships of size 5, $L^2 = 65.92$ with 31 *df*. The first of these statistics does not approach statistical significance, and the second is nominally significant, but *bic* is large and negative for both models. Although there is room for a 2-factor model to improve fit, at least in sibships of size 5, there does not appear to be strong evidence that a second family factor is needed to fit the data.

Model B and model C are more restrictive than model A. Although the constraints in model A are analogous to the findings from most studies of sibling resemblance, there are two more hypotheses worth considering before we test models with fewer constraints. First, we specify that gender does not alter the effect of family background on educational attainment. Second, we specify that gender does not affect the within-family variance of schooling. We test the first hypothesis in Model B;

that is, we specify $\lambda^x_{kij} = 1$ for all i , j , and k . The contrasts of fit statistics between model A and model B are $L^2 = 5.79$ for size 3, $L^2 = 24.12$ for size 4, and $L^2 = 9.81$ for size 5 with 1 *df* for each. They are all nominally significant at the 0.02 level or beyond. We have chosen to reject this hypothesis, partly because of the consistency of the finding in each size of sibship.¹¹ We find $\lambda^x = 0.883$, 0.762, and 0.829 in sibships of size 3, 4, and 5, respectively. This implies that the effect of the common family factor on educational attainment among brothers, arbitrarily normed at $\lambda^x = 1.0$, is larger than that among sisters, regardless of sibship size or gender composition.

The second restrictive hypothesis is that the within-family variances in schooling are equal between brothers and sisters. We test this hypothesis by imposing the constraint that all elements in Θ^b are equal and contrasting the fit of model C with that of model A. This hypothesis is easily rejected: for size 3, the contrast between models C and A yields $L^2 = 157.87$; for size 4, the contrast yields $L^2 = 99.57$; and for size 5, the contrast yields $L^2 = 101.71$, each with 1 *df*. Also, there are corresponding increases in *bic*. These contrasts confirm the well-established finding that, in the cohorts of the 1950's, educational attainment is far more variable among men than among women (Hauser and Wong, 1989, p. 158-159; Sewell et al., 1980, p. 557).

¹¹Except in sibships of size 3, a positive *bic* statistic for the contrast also suggests rejection of the alternative hypothesis. Raftery (1993) suggests that changes in *bic* of less than 10 should not be taken very seriously. Thus, the fit of this particular contrast remains ambiguous. We have in this case chosen to take the gender difference seriously, despite the weak evidence for it.

For example, under model A, in sibships of size 3, the estimated within-family standard deviation of schooling is 2.204 among brothers and 1.505 among sisters; in sibships of size 4, the estimated within-family standard deviation of schooling is 2.098 among brothers and 1.593 among sisters; in sibships of size 5, the estimated within-family standard deviation of schooling is 2.018 among brothers and 1.491 among sisters.

Having rejected two more restrictive hypotheses, we selectively release constraints in model A in order to test less restrictive hypotheses about the effect of family background on educational attainment that are still consistent with the one-factor specification. In model D, we release the cross-group constraints on within-family variance of schooling for brothers and for sisters. The hypothesis suggests that gender composition affects the within-family variance of schooling of brothers and of sisters. The contrasts between model D and model A are $L^2 = 4.45$ with 4 *df* for size 3, $L^2 = 9.91$ with 6 *df* for size 4, and $L^2 = 21.87$ with 8 *df* for size 5. The fit of size 5 is on the borderline, but the *bic* statistic increases by 31.96. Thus, model D is rejected for sibships of all sizes. All the same, in model E we use model D as a point of comparison in order to test the hypothesis that within-family variances differ by birth order within gender. That is, we release all of the remaining equality constraints on Θ^b . The contrasts of fit statistics are $L^2 = 19.61$ with 6 *df* for size 3, $L^2 = 35.26$ with 12 *df* for size 4, and $L^2 = 40.67$ with 20 *df* for size 5; they are each significant at the 0.003 level. However, the *bic* statistics increase by 25.33 for size 3, by

49.60 for size 4, and by 92.64 for size 5. Since we already rejected model D, we also reject model E.

The hypothesis in model F tests cross-group equality in the loadings of the common family factor on educational attainment for brothers and for sisters. That is, we release the cross-group constraints on λ^x . The change of fit is negligible for sibships of all three sizes: $L^2 = 1.48$ with 1 *df* for size 3; $L^2 = 1.26$ with 2 *df* for size 4; and $L^2 = 3.58$ with 3 *df* for size 5. We reject this hypothesis. Analogous to the comparison between model D and model E, in model G we use model F as a point of comparison in order to test the hypothesis that within-family loadings differ by birth order within gender. That is, we release all of the remaining equality constraints on Λ^x . The contrasts of fit statistics are $L^2 = 13.76$ with 5 *df* for size 3; $L^2 = 34.19$ with 11 *df* for size 4; and $L^2 = 30.66$ with 19 *df* for size 5; they are significant at the 0.044 level. Although the contrasts between models G and F are stronger than the contrasts between model D and E, the *bic* statistics increase substantially, as in the comparison between models D and E; we reject model G as well.

In model H, we release the cross-group constraints on Φ . The model suggests that the variance of the common family factor varies by gender composition of sibship, even though gender composition does not affect the within-family variance of schooling or the effect of family background on siblings' schooling. The contrasts between model H and model A yield $L^2 = 0.26$ with 3 *df* for size 3; $L^2 = 6.74$ with 4 *df* for size 4; and $L^2 = 0.70$ with 5 *df* for size 5. The less constrained hypothesis,

model H, is obviously rejected. The estimated variances of the common family factor are 2.554, 2.490, and 2.030 for size 3, 4, and 5 respectively, which suggests that smaller families are more heterogeneous than larger families.¹²

Model A is our preferred model. The preceding findings and fit statistics indirectly support the conclusion that, among sibships of 3, 4, and 5, a constrained one-factor model satisfactorily explains the covariation of sisters' and brothers' educational attainments. We also find that gender composition has essentially no effects on the factorial structure of siblings' educational attainments. It does not affect the loadings of educational attainments on the common family factors, the within-family variances in education, or the variance in the common family factor. At the same time, we do find substantial effects of gender: Within-family variances in schooling are much larger among brothers than among sisters, and the effects of the common family factor are larger on brothers than on sisters.

In the next section, we analyze differences in the effects of family background, gender, family configuration, and relative birth order on educational attainment. Then, we use our preferred model of effects of background and gender to test the hypothesis that there are distinct family background factors for brothers and for sisters. Finally, we incorporate means into the

structural model in order to examine to what extent the common family factor, educational attainment of siblings, and social origins are different between brothers and sisters, and across groups defined by gender composition.

EFFECTS OF FAMILY BACKGROUND

We apply a modified MIMIC model to test whether the family configurations moderate the effect of family background characteristics. Figure 3 displays the path diagrams of a multiple-group model of sibling resemblance. It modifies slightly the conventional multiple-group MIMIC model, which has been used in previous analyses of birth order effects on sibling resemblance, e.g., Hauser and Wong (1989, p. 156).¹³ In the conventional MIMIC model, the structure is

$$\eta = \Gamma \xi + \zeta \quad (9)$$

where η is the endogenous latent variable, ξ is a vector of exogenous latent variables, ζ is a vector of disturbances (with variance-covariance matrix Ψ) that are independent of η and ξ , and Γ is a parameter matrix. In each group, there is only one η , a latent factor which carries the effect of social origins

¹³The dashed, curved paths among the error terms (ϵ) in Figure 3 show an alternative MIMIC model, which is not preferred, but which we use to distinguish between the fit of the factor model per se and the constraints on loadings imposed by effects of the exogenous variables (Hauser and Goldberger, 1971). If we eliminate the disturbance in the common family factor, we can free all of the within-family error covariances, and the contrast between the fit of this model and the MIMIC model without error covariances provides a test of the consistency between the loadings implied by social background effects and those implied by a single factor model of siblings' educational attainments, as well as a test of the overidentifying restrictions implied by the single-factor model of educational attainment.

¹²We had thought, from other analyses of family effects, that larger families were more heterogeneous than smaller families. Our finding raises the possibility that this heterogeneity may be an artifact of composition, e.g., of the greater likelihood of heterogeneity in gender in larger sibships, rather than of an intrinsic increase in heterogeneity within larger families.

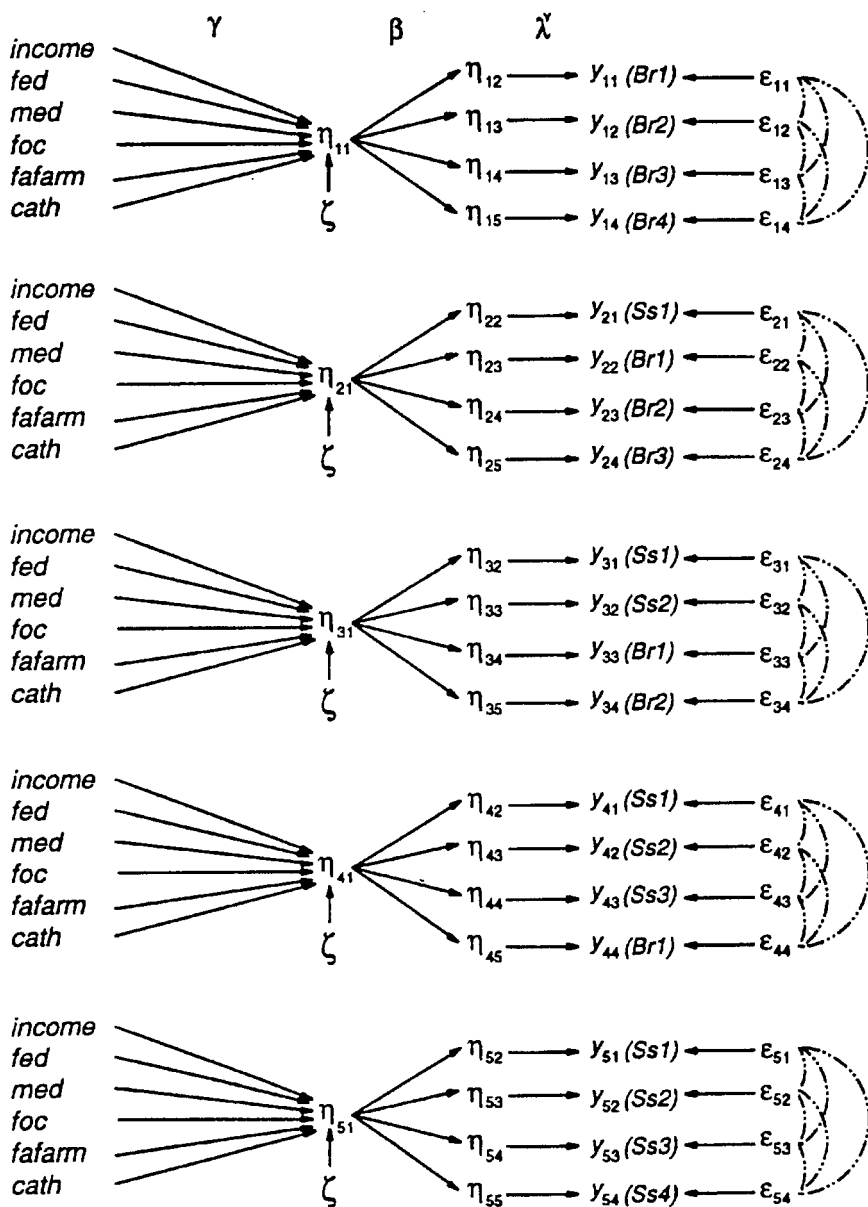


FIG. 3.—Multiple Group Revised MIMIC Model (sibship size = 4)

on educational attainments of offspring, and it may or may not contain a random disturbance, ζ , which is independent of measured background.¹⁴ The measurement models for latent variables are

$$x = \Lambda^x \xi + \delta \tag{10}$$

$$y = \Lambda^y \eta + \varepsilon \tag{11}$$

where x and y are vectors of observable variables, i.e., indicators of ξ and η ; Λ^x , and Λ^y are parameter matrices; and δ and ε are vectors of disturbances with variance-covariance matrices Θ^δ and Θ^ε . Because each exogenous variable only has one indicator, all elements in Λ^x are equal to one and all of the elements in Θ^δ , the variance-covariance matrix of the errors in x 's, are equal to zero; that is, all ξ 's are perfectly measured by corresponding x 's. Equations 11 and 9 in matrix form are

$$\begin{bmatrix} y_1 \\ y_2 \\ y_3 \\ y_4 \end{bmatrix} = \begin{bmatrix} \lambda_{11} \\ \lambda_{21} \\ \lambda_{31} \\ \lambda_{41} \end{bmatrix} [\eta_1] + \begin{bmatrix} \varepsilon_1 \\ \varepsilon_2 \\ \varepsilon_3 \\ \varepsilon_4 \end{bmatrix} \tag{12}$$

$$[\eta_1] = [\gamma_{11} \ \gamma_{12} \ \gamma_{13} \ \gamma_{14} \ \gamma_{15} \ \gamma_{16} \ \gamma_{17}] \begin{bmatrix} x_1 \\ x_2 \\ x_3 \\ x_4 \\ x_5 \\ x_6 \\ x_7 \end{bmatrix} + [\zeta_1] \tag{13}$$

where λ_{11} in equation 12 is normalized as 1 and the other λ 's are proportional to it. However, in our model, we have

elaborated the structure of the y 's (educational attainments) as follows:

$$\begin{bmatrix} \eta_2 \\ \eta_3 \\ \eta_4 \\ \eta_5 \end{bmatrix} = \begin{bmatrix} \beta_{21} \\ \beta_{31} \\ \beta_{41} \\ \beta_{51} \end{bmatrix} [\eta_1] \tag{14}$$

and

$$\begin{bmatrix} y_1 \\ y_2 \\ y_3 \\ y_4 \end{bmatrix} = \begin{bmatrix} \lambda^y_{12} & 0 & 0 & 0 \\ 0 & \lambda^y_{23} & 0 & 0 \\ 0 & 0 & \lambda^y_{34} & 0 \\ 0 & 0 & 0 & \lambda^y_{45} \end{bmatrix} \begin{bmatrix} \eta_2 \\ \eta_3 \\ \eta_4 \\ \eta_5 \end{bmatrix} + \begin{bmatrix} \varepsilon_1 \\ \varepsilon_2 \\ \varepsilon_3 \\ \varepsilon_4 \end{bmatrix} \tag{15}$$

where the β 's are effects of the common family factor, η_1 , on latent individual factors, η_2, \dots, η_5 ; λ^y 's are now effects of the latent individual factors, η_2, \dots, η_5 , on educational attainments of siblings, y_1, \dots, y_4 , and ε 's are disturbances in the y 's, i.e., measurement errors and/or nonshared factors in educational attainment. When we specify $\beta_{2,1} = \beta_{3,1} = \beta_{4,1} = \beta_{5,1} = 1.0$, this model is formally equivalent to the conventional MIMIC model. However, in the modified model, the effect of family background on siblings' educational attainments can be partitioned into two different elements by specifying β 's and λ^y 's separately (Sörbom and Jöreskog, 1981); later, we use this feature of the modified model to specify effects of relative birth order and of gender.¹⁵

In the first panel of Tables 4, 5, and 6, we report the test statistics for several versions of this structural model. The baseline model is highly

¹⁴This setup changes the notation for the common family factor. In the model of Equation 7, ξ was the common family factor. In Equations 9 to 11, η is a common family factor, and ξ represents a vector of social background variables. For convenience, we have suppressed notation for the groups defined by gender composition.

¹⁵Another advantage to this specification is that the extra latent variables (η s) permit us to introduce direct effects of the exogenous variables on the endogenous variables (Kuo and Hauser, 1990). Only the first advantage is important for the purposes of this paper.

TABLE 4
GOODNESS OF FIT STATISTICS (SIBSHIP SIZE = 3, $n = 1,790$)

Model	L^2	df	bic	Contrast	p
Structural Model					
A. Baseline Model	153.48	87	-498.15	...	0.000
B. A - θ^e cross-group constraints	150.68	83	-470.99	2.80 4	0.592
C. B - θ^e within-group constraints	128.69	77	-448.04	21.99 6	0.001
D. A - λ^y cross-group constraints	151.14	86	-493.00	2.34 1	0.126
E. D - λ^y within-group constraints	140.87	80	-458.33	10.27 5	0.068
F. A - $\gamma_{k,2} = \gamma_{k,3}$ constraint	153.23	86	-490.91	0.25 1	0.617
G. F - Γ cross-group constraints	137.80	68	-371.52	15.43 18	0.632
H. A - Ψ cross-group constraints	150.65	84	-478.51	2.83 3	0.419
I. A + birth order constraints	151.84	86	-492.30	1.64 1	0.200
Two-factor Model					
J. Baseline Model, Model A	153.48	87	-498.15	...	0.000
K. A - Γ within-groups constraints	134.51	82	-479.67	18.97 5	0.002
L. A - Ψ within-group constraints	140.80	85	-495.85	12.68 2	0.002
Mean-Structure Model					
M. A + τ^y within-group constraints	157.40	93	-539.17	3.92 6	0.688
N. M + τ^y cross-group constraints	164.96	97	-561.57	7.56 4	0.109
O. N + $\tau_{sister} = \tau_{brother}$	265.96	98	-468.06	101.00 1	0.000
P. N + τ^z cross-group constraints	181.79	115	-679.56	16.83 18	0.535
Q. P + $\tau_{k,2} = \tau_{k,3}$	238.91	116	-629.93	57.12 1	0.000
R. P - $\alpha_{k,1}$ cross-group constraints	180.08	112	-658.80	1.71 3	0.635

constrained: λ^y 's of brothers are normalized to one and λ^y 's of sisters are equated within and across groups defined by gender composition; θ^e 's of brothers are equal across groups, as are those of sisters; Γ 's and Ψ 's are invariant across groups; and γ 's of father's and mother's educational levels are equal. For example, in sibship size 4, the constraints we impose are $\lambda^y_{112} = \lambda^y_{123} = \lambda^y_{134} = \lambda^y_{145} = \lambda^y_{223} = \lambda^y_{234} = \lambda^y_{245} = \lambda^y_{334} = \lambda^y_{345} = \lambda^y_{445} = 1$ and $\lambda^y_{212} = \lambda^y_{312} = \lambda^y_{323} = \lambda^y_{412} = \lambda^y_{423} = \lambda^y_{434} = \lambda^y_{512} = \lambda^y_{523} = \lambda^y_{534} = \lambda^y_{545}$; $\theta^e_{111} = \theta^e_{122} = \theta^e_{133} = \theta^e_{144} = \theta^e_{222} = \theta^e_{233} = \theta^e_{244} = \theta^e_{333} = \theta^e_{344} = \theta^e_{444}$ and $\theta^e_{211} = \theta^e_{311} = \theta^e_{322} = \theta^e_{411} = \theta^e_{422} = \theta^e_{433} = \theta^e_{511} = \theta^e_{522} = \theta^e_{533} = \theta^e_{544}$; $\psi_{111} = \psi_{211} = \psi_{311} = \psi_{411} = \psi_{511}$; $\gamma_{111} = \gamma_{211} = \gamma_{311} = \gamma_{411} = \gamma_{511}$; $\gamma_{112} = \gamma_{212} = \gamma_{312} = \gamma_{412} = \gamma_{512} = \gamma_{113} = \gamma_{213} = \gamma_{313} = \gamma_{413} = \gamma_{513}$; $\gamma_{114} = \gamma_{214} = \gamma_{314} = \gamma_{414} = \gamma_{514}$,

$\gamma_{115} = \gamma_{215} = \gamma_{315} = \gamma_{415} = \gamma_{515}$, and $\gamma_{116} = \gamma_{216} = \gamma_{316} = \gamma_{416} = \gamma_{516}$.¹⁶ The model suggests that, first, gender composition does not affect the influence of the common family factor on the

¹⁶As in the case of the single factor model, we have also estimated a completely unconstrained version of the multiple-group MIMIC model. The goodness-of-fit statistics are $L^2 = 85.65$ with 48 df for size 3, $bic = -379.34$; $L^2 = 149.46$ with 100 df for size 4, $bic = -787.96$; and $L^2 = 248.56$ with 174 df for size 5, $bic = -1328.50$. The fit statistics are each highly significant statistically, but the bic statistics are satisfactory. While the contrasts between these models and the corresponding baseline models are each statistically significant, the bic statistics are much smaller in the more constrained baseline model. Further, when we eliminate the disturbance in the common family factor and free the covariances among within-family errors, the test statistics for the contrasts between these models and the unconstrained MIMIC model are $L^2 = 17.17$ with 8 df for size 3, $bic = -60.33$; $L^2 = 49.11$ with 25 df for size 4, $bic = -185.24$; and $L^2 = 126.48$ with 54 df for size 5, $bic = -362.95$. Thus, the MIMIC model without error covariances is preferable to the model with covariances, and the (more constrained) baseline model is yet more preferable.

TABLE 5
GOODNESS OF FIT STATISTICS (SIBSHIP SIZE = 4, $n = 1178$)

Model	L^2	df	bic	Contrast	p
Structural Model					
A. Baseline Model	248.04	161	-890.48	...	0.000
B. A - θ^c cross-group constraints	237.33	155	-858.76	10.71 6	0.098
C. B - θ^c within-group constraints	207.46	143	-803.77	29.87 12	0.003
D. A - Λ^v cross-group constraints	247.12	159	-877.26	0.92 2	0.631
E. D - Λ^v within-group constraints	216.29	147	-823.23	30.83 11	0.001
F. A - $\gamma_{k,2} = \gamma_{k,3}$ constraint	247.59	160	-883.86	0.45 1	0.502
G. F - Γ cross-group constraints	222.43	136	-739.30	25.16 24	0.397
H. A - Ψ cross-group constraints	240.88	157	-869.36	7.16 4	0.128
I. A + birth order constraints	242.08	160	-889.37	5.96 1	0.015
Two-Factor Model					
J. Baseline Model, Model A	248.04	161	-890.48	...	0.000
K. A - Γ within-groups constraints	228.80	156	-874.37	19.24 5	0.002
L. A - Ψ within-group constraints	247.47	159	-876.91	0.57 2	0.752
Mean Structure Model					
M. A + τ^v within-group constraints	258.24	173	-965.14	10.20 12	0.599
N. M + τ^v cross-group constraints	262.99	179	-1002.82	4.75 6	0.576
O. N + $\tau_{sister} = \tau_{brother}$	318.68	180	-954.20	55.69 1	0.000
P. N + τ^c cross-group constraints	282.28	203	-1153.25	19.29 24	0.736
Q. P + $\tau_{k,2} = \tau_{k,3}$	364.72	204	-1077.88	82.44 1	0.000
R. P - $\alpha_{k,1}$ cross-group constraints	279.46	199	-1127.78	2.82 4	0.588

TABLE 6
GOODNESS OF FIT STATISTICS (SIBSHIP SIZE = 5, $n = 785$)

Model	L^2	df	bic	Contrast	p
Structural Model					
A. Baseline Model	392.72	261	-1347.02	...	0.000
B. A - θ^c cross-group constraints	372.74	253	-1313.68	19.98 8	0.010
C. B - θ^c within-group constraints	333.08	233	-1220.02	39.66 20	0.006
D. A - Λ^v cross-group constraints	389.80	258	-1329.95	2.92 3	0.403
E. D - Λ^v within-group constraints	362.14	238	-1224.29	27.66 19	0.088
F. A - $\gamma_{k,2} = \gamma_{k,3}$ constraint	392.30	260	-1340.78	0.42 1	0.507
G. F - Γ cross-group constraints	360.17	230	-1172.94	32.13 30	0.300
H. A - Ψ cross-group constraints	387.58	256	-1318.84	5.14 5	0.392
I. A + birth order constraints	389.72	260	-1343.36	3.00 1	0.077
Two-Factor Model					
J. Baseline Model	392.72	261	-1347.02	...	0.000
K. A - Γ within-group constraints	385.26	256	-1321.16	7.46 5	0.180
L. A - Ψ within-group constraints	381.60	259	-1344.81	11.12 2	0.004
Mean Structure Model					
M. A + τ^v within-group constraints	423.78	281	-1449.28	31.06 20	0.054
N. M + τ^v cross-group constraints	433.87	289	-1492.51	10.09 8	0.266
O. N + $\tau_{sister} = \tau_{brother}$	461.71	290	-1471.34	27.84 1	0.000
P. N + τ^c cross-group constraints	484.43	319	-1641.92	50.56 30	0.011
Q. P + $\tau_{k,2} = \tau_{k,3}$	564.18	320	-1568.84	79.75 1	0.000
R. P - $\alpha_{k,1}$ cross-group constraints	479.22	314	-1613.80	5.21 5	0.399

educational attainments of sisters and brothers, the within-family variances of schooling for sisters and brothers, the effect of social origins on the common family factor, or the variance of the common family factor. Second, gender may affect the influence of the common family factor on educational attainment and the within-family variances of schooling among sisters and brothers. From the findings of the previous section, we consider these constraints appropriate in our baseline model. Third, we equate the effect of father's educational level on the common factor with that of mother's educational level; that is, the effect of father's education on the educational attainment of offspring is equal to that of mothers (Hauser and Wong, 1989, pp. 159, 167). Sewell et al. (1980) found that effects of father's education and mother's education were very close in the process of educational attainment for Wisconsin women, but only father's education affected the educational attainment for Wisconsin men. Tsai (1983) found that, controlling for measurement errors, effects of father's education and mother's education on educational attainment were the same for Wisconsin women and men. Lee (1989) used sibling pair data and also found that the effects on a common family education factor were equal.

In the baseline model, the goodness-of-fit statistics are $L^2 = 153.48$ with 87 *df* for size 3, $L^2 = 248.04$ with 161 *df* for size 4, and $L^2 = 392.72$ with 261 *df* for size 5. They are highly significant statistically, but the *bic* statistics are quite satisfactory: *bic* = -498.15, *bic* = -890.48 and *bic* = -1347.02 for sizes 3, 4, and

5 respectively.¹⁷ As in the last section, we selectively release constraints in model A in order to test less restrictive hypotheses about the effect of family background on educational attainment that are still consistent with the specification of a single family factor.¹⁸ In model B, we test the hypothesis that the within-family variances of schooling for brothers and sisters vary with gender composition; that is, we release across-group constraints on Λ^y for brothers and for sisters. The contrasts of fit statistics are $L^2 = 2.80$ with 4 *df* for size 3, $L^2 = 10.71$ with 6 *df* for size 4, and $L^2 = 19.98$ with 8 *df* for size 5. This is nominally significant only in sibships of size 5, and *bic* increases in each test. Thus, we reject this hypothesis. We find that gender composition does not affect the within-family variance of schooling of brothers and sisters. In model C, we use model B as a point of comparison in order to test the hypothesis that within-family variances differ by birth order within gender. That is, we release all of the remaining equality constraints on Θ^e . The contrasts of fit statistics yield $L^2 = 21.99$ with 6 *df* for size 3, $L^2 = 29.87$ with 12 *df* for size 4,

¹⁷Note that, when we use *bic* as a criterion, these highly restricted models each fit better than the corresponding unrestricted models.

¹⁸As in our analysis of the one-factor model, we have tried both forward and backward selection, and each procedure led to the same preferred models, namely, Model A in Tables 4, 5, and 6. We have also tested models, within each size of sibship, which constrain the variance-covariance matrices of the exogenous (social background) variables to be invariant with respect to gender composition. In sibships of sizes 3 and 4, the fit of this highly constrained model is not significantly different from that of the baseline model. While the nominal improvement in fit is significant in sibships of size 5, *bic* is smaller under this constrained model than in our baseline model. Thus, we think that gender composition is not selective with respect to the joint distributions of social background variables.

and $L^2 = 39.77$ with 20 *df* for size 5. They are significant at 0.006 level and beyond. However, the *bic* statistics increase by 22.95 for size 3, 54.99 for size 4, and 93.66 for size 5. We have already rejected model B, and thus we reject model C as well.

The hypothesis in model D tests whether gender composition of the sibship affects the influence of the common family factors on the educational attainments of siblings. We release the cross-group constraints on Λ 's. The contrasts between model D and model A yield $L^2 = 2.34$ with 2 *df* for size 3, $L^2 = 0.92$ with 3 *df* for size 4, and $L^2 = 2.92$ with 4 *df* for size 5. None of these contrasts approaches statistical significance. The hypothesis is rejected; that is, gender composition of the sibship does not alter the effect of the common family factor on the educational attainment of offspring. Analogous to the comparison between models B and C, we use the specification in model D as a point of comparison in order to test the hypothesis that the effects of the common family factor on the educational attainments of siblings differ by birth order within gender. When we release all of the remaining equality constraints on Λ , the contrasts of fit between model D and model E are $L^2 = 10.27$ with 5 *df* for size 3, $L^2 = 30.83$ with 11 *df* for size 4, and $L^2 = 27.66$ with 19 *df* for size 5. However, the *bic* statistics increase by 44.67, 54.03, and 105.66 for size 3, 4, and 5 respectively. We also reject model E.

In model F and model G, we test hypotheses relevant to specific family background variables. First, model F tests that the effects of father's education and of mother's education on ed-

ucational attainments of offspring differ. We reject this hypothesis because the contrasts of fit between model A and model F are negligible: $L^2 = 0.25$, $L^2 = 0.45$, and $L^2 = 0.42$ with 1 *df* for sibships of sizes 3, 4, and 5. The failure of this hypothesis supports findings from previous studies: The effect of father's education on educational attainments of offspring is as important as that of mother's education (Hauser and Wong, 1989; Lee, 1990). By releasing the other constraints on coefficients of family background variables, with model F as the point of comparison, we test the hypothesis that effects of measured family variables on educational attainments of offspring vary across groups. The contrasts of fit statistics yield $L^2 = 15.43$ with 18 *df* for size 3, $L^2 = 25.16$ with 24 *df* for size 4, and $L^2 = 32.13$ with 30 *df* for size 5. None of these contrasts approaches statistical significance, and we reject this hypothesis.

In model H, we release the final constraint on model A, cross-group equality in Ψ . Model H suggests that the variances of the unmeasured family factor differ by gender composition. The change of fit is insignificant for sibship of all three sizes: $L^2 = 2.83$ with 3 *df* for size 3, $L^2 = 7.16$ with 4 *df* for size 4, and $L^2 = 5.34$ with 5 *df* for size 5. We fail to accept model H.

In a final model, we partition the effect of family background into effects of gender (Λ) and ordinal position within gender (B); then both matrices are estimated with the slopes for brothers and oldest siblings normalized as one. That is, we distinguish the first sibling from others in matrix B: $\beta_{121} = \beta_{221} = \beta_{231} = \beta_{321} = \beta_{341} = \beta_{421} = \beta_{451} = \beta_{521} = 1.0$ and $\beta_{131} = \beta_{141} =$

$\beta_{151} = \beta_{241} = \beta_{251} = \beta_{331} = \beta_{351} = \beta_{431} = \beta_{441} = \beta_{531} = \beta_{541} = \beta_{551}$ for sibships of size 4. That is, the effect of the family factor on educational attainment is equal to one for eldest brothers, β_{131} for other brothers, λ_{212} for eldest sisters and $(\lambda_{212} \times \beta_{131})$ for other sisters. Compared to model A, the current model has one *df* less but L^2 does not decrease significantly in sizes 3 and 5 (1.64 and 3.00); in size 4, the contrast of L^2 is 5.96, yet the *bic* statistic increases only by 1.11. Thus, we reject the hypothesis that the oldest brother differs from other brothers and the oldest sister differs from other sisters in the effect of family background. Again, this finding should be read as pertaining only to our definition of relative birth order within gender.

PARAMETERS OF THE MIMIC MODEL

Table 7 gives parameter estimates of our preferred models of sibling resemblance in educational attainment. The within-family variances of brothers' schooling (4.71, 4.39, and 4.05) vary inversely with size of sibship while the variances of sisters' attainments (2.35, 2.54, and 2.24) barely differ. Thus, the differences of within-family or non-shared variances (θ^e) of schooling between brothers and sisters decline along with increases in size of sibship. The variances (Ψ) of the unmeasured, shared family factor decrease with increases in size of sibship. They are 1.57, 1.42, and 1.26 for sizes 3, 4, and 5, respectively.¹⁹ The effects of the

family factors on sisters, 0.84, 0.76, and 0.81, are smaller than on brothers (for whom the effects are normalized as one); that is, the common family factor—including both measured and unmeasured family characteristics—influences sisters one-fifth to one-quarter less than brothers. Possibly excepting Catholic upbringing, which lowers schooling in sibships of size 3 and has insignificant positive coefficients in sibships of size 4 or 5, there do not appear to be any substantial variations in the effect of measured social background by size of sibship.

TWO-FACTOR MODEL OF SIBLING RESEMBLANCE

Figure 4 shows the path diagram of our multiple-group two-factor model. As stated in the introduction, despite the different loadings of education on the family background variable, the one factor model may not capture all of the differences in effects of social origins between brothers and sisters.²⁰ In the model of Figure 4, one additional latent factor is added to Equation 9, that is, in matrix form,

$$\begin{bmatrix} \eta_1 \\ \eta_2 \end{bmatrix} = \begin{bmatrix} \gamma_{11} & \gamma_{12} & \gamma_{13} & \gamma_{14} & \gamma_{15} & \gamma_{16} \\ \gamma_{21} & \gamma_{22} & \gamma_{23} & \gamma_{24} & \gamma_{25} & \gamma_{26} \end{bmatrix} \begin{bmatrix} \xi_1 \\ \xi_2 \\ \xi_3 \\ \xi_4 \\ \xi_5 \\ \xi_6 \end{bmatrix} + \begin{bmatrix} \epsilon_1 \\ \epsilon_2 \end{bmatrix} \quad (16)$$

$$\begin{bmatrix} \eta_3 \\ \eta_4 \\ \eta_5 \\ \eta_6 \end{bmatrix} = \begin{bmatrix} \beta_{31} & \beta_{32} \\ \beta_{41} & \beta_{42} \\ \beta_{51} & \beta_{52} \\ \beta_{61} & \beta_{62} \end{bmatrix} \begin{bmatrix} \eta_1 \\ \eta_2 \end{bmatrix} \quad (17)$$

¹⁹These variances are, of course, smaller than those in the single factor model with no exogenous variables because they do not include the variance associated with measured background variables. All the same, we find the same pattern of declining heterogeneity with increased family size.

²⁰Although it is a two-factor model, because the factors are gender-specific, only one factor appears in the all-sister and all-brother groups.

TABLE 7
PARAMETER ESTIMATES OF PREFERRED MODELS

	Sibship Size 3		Sibship Size 4		Sibship Size 5	
Structural Coefficients						
Effect of the Common Family Factor on (Λ^y)						
Sister	0.84	(0.04)	0.76	(0.80)	0.81	(0.04)
Brother	1.00	...	1.00	...	1.00	...
Loadings of Exogenous Variables on the Common Family Factor (Γ)						
Income	0.09	(0.02)	0.10	(0.02)	0.11	(0.02)
Father's education	0.13	(0.01)	0.13	(0.01)	0.10	(0.01)
Mother's education	0.13	(0.01)	0.13	(0.01)	0.10	(0.01)
Father's SEI	0.15	(0.02)	0.15	(0.03)	0.14	(0.03)
Farmer	0.26	(0.12)	0.35	(0.13)	0.31	(0.13)
Catholic	-0.22	(0.08)	0.16	(0.10)	0.11	(0.10)
Variances						
Unmeasured Family Variance (ψ).....	1.57	(0.13)	1.42	(0.13)	1.26	(0.13)
Within-Family Variances (θ^e)						
Sister	2.35	(0.09)	2.54	(0.09)	2.24	(0.08)
Brother	4.71	(0.16)	4.39	(0.15)	4.05	(0.15)
Means						
Means of Schooling (τ^y)						
Sister	13.20	(0.05)	12.93	(0.05)	12.73	(0.05)
Brother	13.77	(0.06)	13.39	(0.06)	13.06	(0.07)
Means of Exogenous Variables (τ^e)						
Income	6.23	(0.08)	5.89	(0.09)	5.21	(0.10)
Father's Education	10.14	(0.08)	9.56	(0.10)	9.18	(0.12)
Mother's Education	10.72	(0.07)	10.40	(0.08)	10.16	(0.10)
Father's SEI	3.72	(0.06)	3.35	(0.06)	2.98	(0.07)
Farmer	0.17	(0.01)	0.20	(0.01)	0.26	(0.02)
Catholic	0.38	(0.01)	0.44	(0.01)	0.44	(0.02)

where $\psi_{12} \neq 0$, η_1 and η_2 are the family factors for sisters and brothers, and η_3 , η_4 , η_5 , and η_6 function as η_2 , η_3 , η_4 , and η_5 in Equation 12. In a multiple-group model, $\beta_{k,b(i),1} = \beta_{k,s(i),2} = 0$, where k indicates group (gender composition), and $s(i)$ and $b(i)$ indicate the positions of sisters and brothers who do not appear in this sibship. This model is equivalent to the one-factor MIMIC model when we specify that the two common factor disturbances have equal variances and are perfectly correlated and that there are identical loadings of observable exogenous variables on each factor. By releasing re-

strictions on Γ and Ψ , we can test alternatives to the one-factor MIMIC model of sibling resemblance. However, the two-factor model is under-identified in sibships with one or fewer sisters and with one or fewer brothers. For example, in sibship size 4, both factors are identified only in the third group, which contains two brothers and two sisters. We use cross-group constraints to identify properties of the latent factor indicated by only one sibling. That is, we use information from the groups with more than two siblings of each gender to identify the parameters of the factor model for groups in

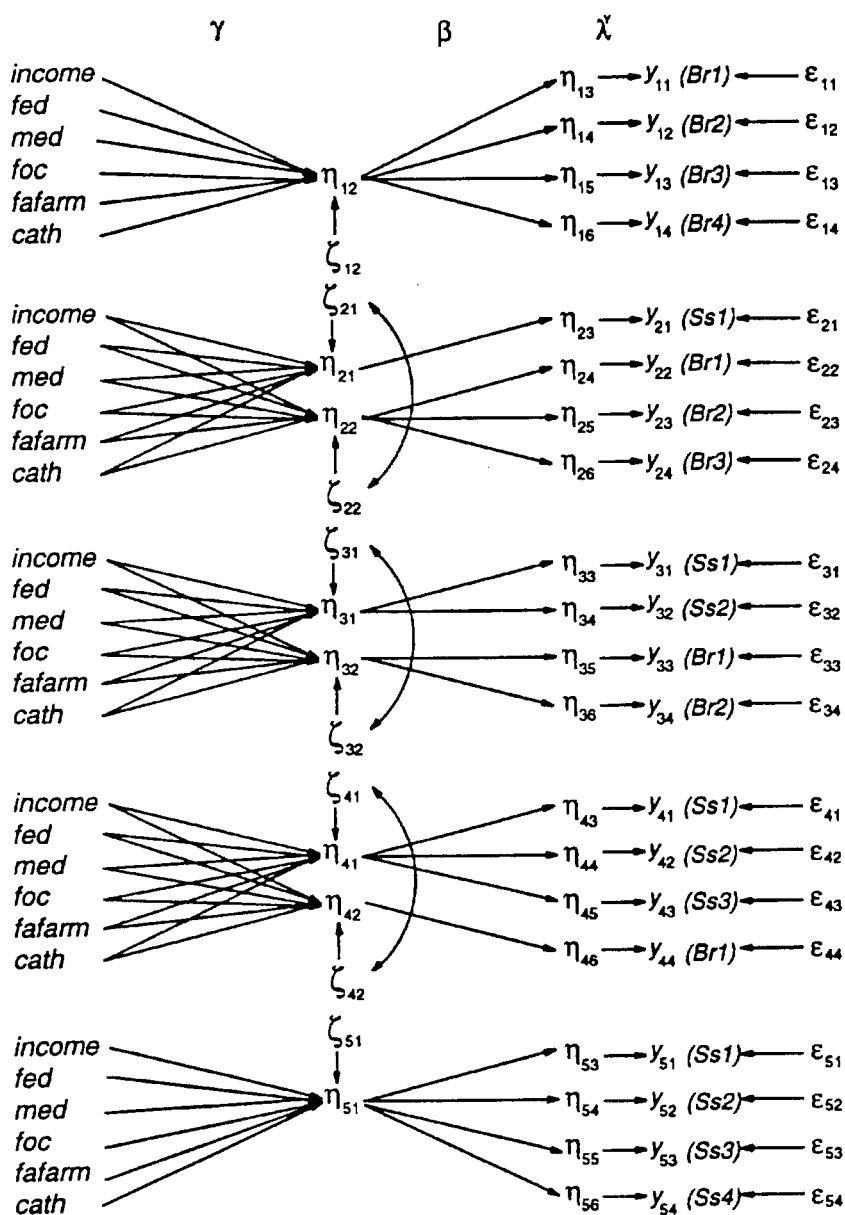


FIG. 4. — Multiple Group Two-Factor MIMIC Model (sibship size = 4)

which there is only one brother or only one sister. We also equate the parameters for men and women in the three gender configurations for each sex in which the gender-specific factor model is identified. That is, the single factor model for men is identified in the first three groups, while the single factor model for women is identified in the last three groups.

Model J, a nominal two-factor model, is equivalent to model A of Tables 4, 5, and 6. Thus, the fit statistics are identical: $L^2 = 153.48$ with 87 *df* for size 3, $L^2 = 248.04$ with 161 *df* for size 4, and $L^2 = 392.72$ with 261 *df* for size 5. In model K, we release the within-group constraints on Γ and, in model L, we release the within-group constraints on Ψ . The contrasts between model K and model J are $L^2 = 18.97$ with 5 *df* for size 3, $L^2 = 19.24$ with 5 *df* for size 4, and $L^2 = 7.46$ with 5 *df* for size 5. Although the contrast for sizes 3 and 4 are significant at the 0.002 level, the *bic* statistic increases by 18.48 and 16.11. Thus, we reject the hypothesis that there are distinctive effects of the measured background variables on brothers and sisters; that is, there are no gender differences in the relative effects of family income, mother's and father's education, father's occupational status, farm origin, and Catholic religious origin. The only significant differences in these effects between brothers and sisters are their systematically lower values among women than among men.

In model L, we release the constraints on Ψ to test the hypothesis that there are two distinct unmeasured factors, one for brothers and the other for sisters. That is, we release the constraints of equality among ψ_{k21} , ψ_{k11} ,

and ψ_{k22} for each gender composition group, while retaining the constraints of equality in corresponding parameters across groups. The contrasts between this model and model J are $L^2 = 12.68$, $L^2 = 0.57$, and $L^2 = 11.12$ with 2 *df* for sizes 3, 4, and 5, respectively. The largest contrast, that for size 3, is significant at 0.002 level, but the *bic* statistics only decrease by 2.30. Thus, we reject this hypothesis as well. Our failure to accept model K and model L support the assumption of the one-factor model in previous sibling resemblance studies, e.g., Hauser and Wong (1989).

STRUCTURAL MODEL WITH MEANS

The preceding structural models all ignore the means of the variables; that is, we have ignored differences among groups in mean levels of social background and in mean levels of attainment. We have also ignored differences in mean levels of attainment between men and women, as well as other possible differences, e.g., by relative birth order. We now drop this simplification and estimate means of educational attainments of siblings, family background variables, and latent common family factors. We are interested not only in estimating the means, but also in testing cross- and within-group constraints on the means of measured and latent variables, that is, testing hypotheses that parallel our previous tests of differences in slopes and variance components.

The mean of a latent variable is under-identified. To estimate the effects of latent variables on observable variables, we have to normalize one of the effects and estimate other effects in proportion to the normalized effect;

likewise, to estimate the means of latent variables, we have to normalize one of the means as zero and estimate its difference from others.²¹ The MIMIC model is now defined by the following equations:

$$\eta = \alpha + \Gamma \xi + \zeta \quad (18)$$

$$x = \tau^x + \Lambda^x \xi + \delta \quad (19)$$

$$y = \tau^y + \Lambda^y \eta + \varepsilon \quad (20)$$

where α , τ^x , τ^y are vectors of constant intercept terms. Owing to the specification of structure of endogenous variables, Equations 13 and 14 can be rewritten in matrix form as follows:

$$[\eta_1] = [\alpha_1] + [\gamma_{11} \gamma_{12} \gamma_{13} \gamma_{14} \gamma_{15} \gamma_{16} \gamma_{17}] \begin{bmatrix} x_1 \\ x_2 \\ x_3 \\ x_4 \\ x_5 \\ x_6 \\ x_7 \end{bmatrix} + [\zeta_1] \quad (21)$$

$$\begin{bmatrix} \eta_2 \\ \eta_3 \\ \eta_4 \\ \eta_5 \end{bmatrix} = \begin{bmatrix} \alpha_2 \\ \alpha_3 \\ \alpha_4 \\ \alpha_5 \end{bmatrix} + \begin{bmatrix} \beta_{21} \\ \beta_{31} \\ \beta_{41} \\ \beta_{51} \end{bmatrix} [\eta_1] \quad (22)$$

where $\beta_{2,1} = \beta_{3,1} = \beta_{4,1} = \beta_{5,1} = 1.0$.

The factorial structure of this model is identical to model A, except that constraints of $\tau^y = 0$ and $\tau^x = 0$ are removed. In model M, we test the hypothesis that the educational attainments of sisters are equal to each other and that those of brothers are equal to each other within each group defined by gender composition. That is, we equate the mean of schooling (τ^y) for

brothers and for sisters within each group, for example, in sibships of size 4, $\tau^y_{111} = \tau^y_{122} = \tau^y_{133} = \tau^y_{144}$, $\tau^y_{222} = \tau^y_{233} = \tau^y_{244}$, $\tau^y_{311} = \tau^y_{322}$, $\tau^y_{333} = \tau^y_{344}$, $\tau^y_{411} = \tau^y_{422} = \tau^y_{433}$, and $\tau^y_{511} = \tau^y_{522} = \tau^y_{533} = \tau^y_{544}$. The contrasts of L^2 between model M and model A are 3.92 with 6 *df* for size 3, 10.20 with 12 *df* for size 4, and 31.06 with 20 *df* for size 5. We fail to reject this hypothesis; the differences of years of schooling among sisters and those among brothers in a family are not statistically significant at even the 0.05 level. Next, we impose constraints of cross-group equality (within each gender) in model N, for example, in size 4, $\tau^y_{111} = \tau^y_{122} = \tau^y_{133} = \tau^y_{144} = \tau^y_{222} = \tau^y_{233} = \tau^y_{244} = \tau^y_{333} = \tau^y_{344} = \tau^y_{444}$ and $\tau^y_{211} = \tau^y_{311} = \tau^y_{322} = \tau^y_{411} = \tau^y_{422} = \tau^y_{433} = \tau^y_{511} = \tau^y_{522} = \tau^y_{533} = \tau^y_{544}$. This constraint says that gender composition does not affect mean educational attainments of sisters and of brothers. The contrasts of fit between model M and this model are $L^2 = 7.56$ with 4 *df* for size 3, $L^2 = 4.75$ with 6 *df* for size 4, and $L^2 = 10.09$ with 8 *df* for size 5; these are all insignificant. We fail to reject this hypothesis as well. In model O, all means of years of schooling are specified to be equal, regardless of gender and gender composition. This model is rejected. With an increase of one *df*, the contrasts of L^2 between model O and model N are 101.00, 55.69, and 27.84 for sizes 3, 4, and 5 respectively. In sum, differences of educational attainment between sisters and brothers persist, but there are no effects of family configuration or of relative birth order.

We equate the means of each family background variable across gender composition in model P; that is, τ^x (the

²¹This strictly parallels the use of omitted categories in dummy-variable regression analysis.

matrix of means of family background variables) is invariant across groups. The contrasts of L^2 (with model N) are insignificant: 16.83 with 18 *df*, 19.29 with 24 *df*, and 50.56 with 30 *df*. In model Q, we test the hypothesis that the mean of father's education is equal to that of mothers. The contrasts of fit yield $L^2 = 57.12$, $L^2 = 82.44$, and $L^2 = 79.75$ with one *df* for each size; the model is rejected.

In model R, we condition on model P to test hypotheses about means of the latent factors. The constraints on α 's are lifted to estimate the difference in means of the family factor across gender composition. In model R, we specify that the first α in the first group is zero and estimate the first α 's in the other groups. This hypothesis tests whether the means of the common family factors differ from each other. The contrasts of fit between model R and model P yield $L^2 = 1.71$ with 3 *df* for size 3, $L^2 = 2.82$ with 4 *df* for size 4, and $L^2 = 5.21$ with 5 *df* for size 5. The nonsignificance of these test statistics implies that the means of the common family factors do not vary with gender composition. In sum, in our preferred model (P), within each size of sibship, there are no differences in mean levels of family background or educational attainment by gender composition, nor are there differences in educational attainment by relative birth order within gender; the only significant differences in means are those between brothers' and sisters' educational attainments.

The more substantial differences in means occur across sibship sizes. The mean of years of schooling for both sisters and brothers consistently decreases with increasing sibship size,

that is, 13.20 in size 3, 12.93 in size 4, and 12.73 in size 5 for sisters, and 13.77 in size 3, 13.39 in size 4, and 13.06 in size 5 for brothers. Within each size of sibship, brothers obtain more schooling than their sisters. It also appears that the gaps of educational attainment among different sibship sizes are larger for brothers than for sisters, that is, brothers benefit more from small sibship size than sisters do, and mean educational differences between brothers and sisters decrease with increasing sibship sizes.

DISCUSSION

As the study shows, sisters' educational attainments differ from those of their brothers, with respect to the level of schooling completed, the dependence of schooling on social background, and the variability in school completion. However, these differences follow a relatively simple pattern. First, sisters have less education than their brothers. Second, the absence of competition for resources with brothers, that is, the all-sister family, does not improve educational attainments of girls. Third, the negative effect of sibship size on education may be moderated by gender; size matters less for girls. Fourth, gender composition does not affect inequality of education; that is, the variance in the common family effect on schooling does not vary by gender composition. Fifth, there is less inequality in educational attainment among women than among their brothers. This result has two sources: Family background has less influence on the educational attainments of sisters than on brothers. Also, there is less variation in education

within families among women than among their brothers.

How can we interpret persistent inequality between women and their brothers, along with substantial equality within gender? According to the maximization assumption (Becker, 1980), along with the increase of family size, parents are more likely to invest in a certain child, the most gifted, a boy, or the oldest, to maximize their return. On the other hand, the compensation hypothesis says that parents try to equalize outcomes, so they tend to "allocate resources equally between their children and to compensate, to some extent, for the handicaps of the children with lower natural endowments" (Griliches, 1979). Both arguments are partially supported. Parents might invest more in boys than in girls, but within gender, parents invest equally, at least with respect to the characteristic measured here, relative birth order. There may be other characteristics of siblings, not included in our models, such as differences among children in health, ability, and motivation, that may tend to attract or discourage parental investment. However, our negative findings with respect to relative birth order and gender composition tend to rule out the influence of factors that might be highly correlated with them.

We also reject the hypothesis that mother's education has a larger effect than father's education on sons or on daughters. That is, in none of our one-factor models do we reject the hypothesis that the effect of maternal schooling is equal to that of paternal schooling, nor do we find that the relative effects of social background vari-

ables differ between brothers and sisters.²² Thus, our findings agree with those of Tsai (1983) and Lee (1989) but slightly differ from those of Sewell et al. (1980).

We do not find that birth order affects educational attainment, nor does it change the effect of family background on educational attainment. In the full Wisconsin sample, Sewell and Hauser (1986; Retherford and Sewell, 1993) also did not find any birth order effects on educational attainment, but in the 1962 Occupational Changes in a Generation survey (OCG), Blau and Duncan (1967) found an advantage for the eldest and youngest in a large family. This discrepancy may be due to differences in population definition. The OCG is sampled from a number of cohorts in the general population, but the WLS always has at least one sibling graduating from high school. Second, Sewell and Hauser studied both brothers and sisters, while Blau and Duncan only investigated brothers. Lee (1989) found that birth order affects the influence of family background only in sister pairs, while we find birth order has no influence at all.²³

From the findings of our study, we think that it may be useful to develop new analyses of the influence of size of sibship on educational and other socioeconomic outcomes. We believe that the full sibship model is an appropriate

²²Recall that the latter hypothesis is tested by the constraints on Γ in the two-factor model.

²³Again, differences in population definition may explain this discrepancy. From Table 1, we calculate that more than three-quarters of potential sister pairs occur in families of six or more siblings, which are not included in our analyses.

and powerful way to study the effect of family configuration on the resemblance and variation among siblings. At the same time, our findings are essentially negative with respect to hypotheses that depend on our use of data from full sibships. Thus, we do not think that there is a great deal to be lost in future research that may be limited to hypotheses that can only be conveniently tested using data for sibling pairs. Moreover, in using data for full sibships, we have had to limit our analyses to those data for individual members of sibships that were available for every member, in this case, only gender, educational attainment, and the position of the sibling in his or her own family structure.

Since the 1960's, American family structure has changed dramatically. In the Wisconsin sample, less than 10 per cent of our primary respondents reported having grown up in a nonintact family, and widows were as common as other female heads. In the 1980's, at least, the educational achievement of children from families of widowed parents did not differ from that in intact families (McLanahan and Sandefur, 1994). Thus, the increasing share of divorced or never-married heads of one-parent families has altered family structure in ways that might affect educational outcomes. Also, declining fertility after the 1960's may have changed the effects of gender composition. Here, we only study sibship sizes three to five, which are now unusually large.

We believe that a great deal more can be learned about family resemblance by bringing in more individual and family variables, which is not fea-

sible in most studies, including the Wisconsin study, for more than two siblings in each family. For example, following Olneck (1977, 1979), Sewell and Hauser (1986) have brought individual measures of academic ability into models of education, occupational status, and income among brothers in the Wisconsin Longitudinal Study, and the 1992-93 round of the WLS will provide new individual data for a much larger sample of brother and sister pairs (Hauser et al., 1994).

ACKNOWLEDGMENTS

This research was supported by a grant from the Spencer Foundation, by the Vilas Estate Trust, and by the National Institute on Aging (AG-9775). It was carried out using facilities of the Center for Demography and Ecology at the University of Wisconsin-Madison, for which core support comes from the National Institute of Child Health and Human Development. We thank William H. Sewell, Judith Seltzer, and Robert Kaufman for helpful advice.

Statistical tabulation and estimation were performed using SPSS-X and LISREL 7.20 on a VAXstation 3100. All of the summary data used in this article are either published herein, or are available from the authors. A public-use version of unit record data from the Wisconsin Longitudinal Study is available from the Data Program and Library Service, University of Wisconsin-Madison, 1180 Observatory Drive, Madison, Wisconsin 53706 or from the Inter-university Consortium for Political and Social Research, P.O. Box 1248, Ann Arbor, Michigan 48106.

The opinions expressed herein are those of the authors. Correspondence should be directed to Robert M. Hauser, Department of Sociology, University of Wisconsin-Madison, 1180 Observatory Drive, Madison, Wisconsin 53706 or HAUSER@SSC.WISC.EDU.

REFERENCES

- ADAMS, B. N. 1972. Birth order: A critical review. *Sociometry* 35:411-439.
- BECKER, G. S. 1980. A treatise on the family. Harvard Univ. Press, Cambridge, MA.
- BENIN, M. H., and D. R. JOHNSON. 1984. Sibling similarities in educational attainment: A comparison of like-sex and cross-sex sibling pairs. *Sociol. Educ.* 57(Jan.):11-21.
- BLAU, P. M., and O. D. DUNCAN. 1967. *The American occupational structure*. John Wiley and Sons, New York.
- BOWLES, S. 1972. Schooling and inequality from generation to generation. *J. Polit. Econ.* 80(June):S219-251.
- BRIM, O. G., JR. 1958. Family structure and sex role learning by children: A further analysis of Helen Koch's data. *Sociometry* 21:1-16.
- CHAMBERLAIN, G., and Z. GRILICHES. 1977. More on brothers, p. 97-107. In P. Taubman (ed.), *Kinometrics: Determinants of socioeconomic success within and between families*. North-Holland, Amsterdam.
- DE GRAAF, P. M., and J. J. HUIJINK. 1992. Trends in measured and unmeasured effects of family background on educational attainment and occupational status in the Federal Republic of Germany. *Soc. Sci. Res.* 21:84-112.
- DRONKERS, J. 1993. Is the importance of family decreasing? Evidence regarding Dutch sibling data and educational attainment. In W. Meeus, M. de Goede, W. Kox, and K. Hurlmann (eds.), *Adolescence, careers and cultures*. De Gruyter, Berlin.
- GRILICHES, Z. 1979. Sibling models and data in economics: Beginnings of a survey. *J. Polit. Econ.* 87(Oct.):S37-S64.
- HAUSER, R. M. 1984. Some cross-population comparisons of family bias in the effects of schooling on occupational status. *Soc. Sci. Res.* 13(June):159-187.
- . 1988. A note on two models of sibling resemblance. *Amer. J. Sociol.* 93(May):1401-1423.
- HAUSER, R. M., D. CARR, T. S. HAUSER, J. A. HAYES, M. KRECKER, D. KUO, J. PRESTI, D. SHINBERG, M. SWEENEY, T. THOMPSON-COLON, S. C. N. UHRIG, and J. R. WARREN. 1994. The class of 1957 after 35 years: Overview and preliminary findings. CDE Working Paper 93-17. Center for Demography and Ecology, University of Wisconsin-Madison.
- HAUSER, R. M., and A. S. GOLDBERGER. 1971. The treatment of unobservable variables in path analysis, p. 81-117. In H. L. Costner (ed.), *Sociological methodology*. Jossey-Bass, San Francisco.
- HAUSER, R. M., and P. A. MOSSEL. 1985. Fraternal resemblance in educational attainment and occupational status. *Amer. J. Sociol.* 91(Nov.):650-673.
- . 1987. Some structural equation models of sibling resemblance in educational attainment and occupational status, p. 108-137. In P. Cuttance and R. Ecob (eds.), *Structural modeling by example: Applications in educational, sociological, and behavioral research*. Cambridge Univ. Press, Cambridge, MA.
- HAUSER, R. M., and W. H. SEWELL. 1985. Birth order and educational attainment in full sibships. *Amer. Educ. Res. J.* 32(Spring):1-23.
- . 1986. Family effects in simple models of education, occupational status, and earnings: Findings from the Wisconsin and Kalamazoo studies. *J. Labor Econ.* 4(Spring):S83-S115.
- HAUSER, R. M., and R. S-K. WONG. 1989. Sibling resemblance and inter-sibling effects in educational attainment. *Sociol. Educ.* 62(July):149-171.
- HAUSER, R. M., S-L. TSAI, and W. H. SEWELL. 1983. A model of stratification with response error in social and psychological variables. *Sociol. Educ.* 56(Jan.):20-46.
- HERMALIN, A. I. 1969. The homogeneity of siblings on education and occupation. Doctoral thesis. Department of Sociology, Princeton University.
- HSUEH, C-T. 1992. Sibling resemblance in educational attainment. Doctoral thesis. Department of Sociology, University of Wisconsin-Madison.
- KUO, H-H. D., and R. M. HAUSER. 1990. Model selection: On black self-concept. Unpublished ms.
- LEE, M-L. 1989. Some structural models of family and inter-sibling effects on educational attainment: Findings from Wisconsin data. Master's thesis. Department of Sociology, University of Wisconsin-Madison.
- LINDERT, P. 1977. Sibling position and achievement. *J. Hum. Res.* 12:198-219.
- McLANAHAN, S., and G. D. SANDEFUR. 1994. *Growing up with a single parent: What hurts, what helps*. Harvard Univ. Press, Cambridge, MA.
- OLNECK, M. R. 1976. The determinants of educational attainment and adult status among brothers: The Kalamazoo Study. Doctoral thesis. Harvard Graduate School of Education.
- . 1977. On the use of sibling data to estimate the effects of family background, cognitive skills, and schooling: Results from the Kalamazoo Brothers Study, p. 125-162. In P. Taubman (ed.), *Kinometrics: Determinants of socioeconomic success within and between families*. North-Holland, Amsterdam.
- . 1979. The effects of education, p. 159-190. In C. S. Jencks (ed.), *Who gets ahead? The determinants of economic success in America*. Basic Books, New York.

- OLNECK, M. R., and D. B. BILLS. 1979. Family configuration and achievement: Effects of birth order and family size in a sample of brothers. *Soc. Psychol. Quart.* 42:135-148.
- POWELL, B., and L. C. STEELMAN. 1990. Beyond sibship size: Sibling density, sex composition, and educational outcomes. *Soc. Forc.* 69(1)(Sept.):181-206.
- RAFTERY, A. E. 1986. Choosing models for cross-classifications (Comment on Grusky and Hauser). *Amer. Sociol. Rev.* 51(1)(Feb.):145-146.
- . 1993. Bayesian model selection in structural equation models, p. 163-180. In K. A. Bollen and J. S. Long (eds.), *Testing structural equation models*. Sage Publications, Newbury Park, CA.
- . 1995. Bayesian model selection in social research, p. 111-163. In P. V. Marsden (ed.), *Sociological methodology 1995* (Vol. 25). Blackwell, Cambridge.
- RETFERFORD, R. D., and W. H. SEWELL. 1991. Birth order and intelligence: Further tests of the confluence model. *Amer. Sociol. Rev.* 56(2)(April):141-158.
- . 1993. The effects of birth order and gender on educational and occupational attainment. Presented at the 1993 meetings of the Population Association of America, Cincinnati, Ohio.
- SEWELL, W. H., and R. M. HAUSER. 1977. On the effects of families and family structure on achievements, p. 255-283. In P. Taubman (ed.), *Kinometrics: The determinants of educational attainment, mental ability, and occupational success within and between families*. North Holland, Amsterdam.
- SEWELL, W. H., R. M. HAUSER, and W. C. WOLF. 1980. Sex, schooling and occupational status. *Amer. J. Sociol.* 86(Nov.):551-583.
- STEELMAN, L. C. 1985. A tale of two variables: A review of the intellectual consequences of sibship size and birth order. *Rev. Educat. Res.* 55(3):353-386.
- . 1986. The tale retold: A response to Zajonc. *Rev. Educat. Res.* 56(3):373-377.
- TSAI, S-L. 1983. Sex differences in the process of stratification. Doctoral thesis. University of Wisconsin-Madison.
- WRIGHT, A. 1977. The effects of sibship size and birth order on education, occupation and earnings. Working Paper. Center for Demography and Ecology. University of Wisconsin-Madison.
- ZAJONC, R. B. 1976. Family configuration and intelligence. *Science* 192:227-236.
- ZAJONC, R. B., and G. B. MARKUS. 1975. Birth order and Intellectual development. *Psychol. Rev.* 82:74-88.
- ZAJONC, R. B., G. B. MARKUS, M. L. BERBAUM, J. A. BARGH, and R. R. MORELAND. 1991. One justified criticism plus three flawed analyses equals two unwarranted conclusions: A reply to Retherford and Sewell. *Amer. Sociol. Rev.* 56(2)(April):159-165.