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GENDER, FAMILY CONFIGURATION, AND THE EFFECT OF FAMILY BACKGROUND ON EDUCATIONAL ATTAINMENT

Hsiang-Hui Daphne Kuo

and

Robert M. Hauser

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Hsiang-Hui Daphne Kuo
Robert M. Hauser

Department of Sociology
The University of Wisconsin-Madison

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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 family background 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 family background variables on brothers are larger than on sisters, the relative effects of measured family background variables 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 family background 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 family background 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 those of family configuration address the second.

Most sibling resemblance models focus on modeling and identifying different components of family background, 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 different 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 family background on educational attainment. Benin and Johnson (1984) reported that family background 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 those 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.[2] 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 elements 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

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[2] These 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.

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 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 many social background variables -- paternal and maternal schooling, occupation, income, race, region, etc. -- one puts into a model, some relevant common family background factors are probably left out. By specifying one or more common, unmeasured family background 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; Olneck 1977; Olneck 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 background 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 gender 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.

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 2000 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 composition, 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 confounded with sibling position (Adams 1972). Since the 1970s, 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 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 sibling 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 non-random 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

[3] Several 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; Steelman 1986) versus Zajonc (1986) and Retherford and Sewell (1991; Retherford and Sewell 1992) versus Zajonc et al. (1991).

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 family background was controlled. When family size is controlled, years of schooling increase with birth order; however, the gain of schooling coincides with intercohort gains of educational attainment in the general population between 1930 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 was number of brothers, and the other was 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 family background variables on brothers and sisters are very similar. One exception is Lee's finding that sisters are less influenced by family background than their

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

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 family background variables 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 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 family background variables 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

[5] Chamberlain 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.

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, $\eta_1, \eta_2, \ldots, \eta_m$, where $\eta_1, \eta_2, \ldots, \eta_{m^*}$ pertain to sisters, and $\eta_{m^*+1}, \eta_{m^*+2}, \ldots, \eta_m$ pertain to brothers, arrayed by birth order within sex. In the one-factor model, the reduced form equations,

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

will satisfy the constraints,

$$
\frac{\pi_{1,1}}{\pi_{1,2}} = \frac{\pi_{2,1}}{\pi_{2,2}} = \ldots = \frac{\pi_{m,1}}{\pi_{m,2}} = \kappa \tag{2}
$$

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

$$\frac{\pi_{1,1}}{\pi_{1,2}} = \ldots = \frac{\pi_{m*,1}}{\pi_{m*,2}} = \kappa_s$$

and

$$\frac{\pi_{m*+1,1}}{\pi_{m*+1,2}} = \ldots = \frac{\pi_{m,1}}{\pi_{m,2}} = \kappa_b \tag{3}$$

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

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[6] If $\kappa_s \neq \kappa_b$ 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.

sisters or brothers, depending on the value of λ_b. That is, for any sister, s, or brother, b, in any birth order,

$$\eta_s = \gamma_{1,1}\xi_1 + \gamma_{1,2}\xi_2 + \varepsilon_s$$
$$\eta_b = \gamma_{1,1}\lambda_b\xi_1 + \gamma_{1,2}\lambda_b\xi_2 + \varepsilon_b \tag{4}$$

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:

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

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 $\theta^{\varepsilon}_{ij} = \mathrm{cov}(\varepsilon_i,\varepsilon_j)$. However, if we respecify the model to introduce an unmeasured common family factor, ζ_1, so

$$\eta_1 = \gamma_{11}\xi_1 + \gamma_{12}\xi_2 + \zeta_1, \tag{6}$$

where $\theta^{\varepsilon}_{12} = \theta^{\varepsilon}_{13} = \theta^{\varepsilon}_{23} = 0$, 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; this 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.

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. 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 family background variables 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 non-shared 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.

Data and Variables

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; Hauser 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 (Hauser, Tsai and Sewell 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 9081 WLS respondents in the 1975 survey who reported their numbers of siblings.[7] Sixty-two percent of respondents have 2 or more siblings. Almost half of respondents who answer the sibling roster have 2 to 4 siblings; that is, 49 percent 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 three, 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

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

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 sibship size 5, that is, 90 percent of the original respondents in sibship size 3, 81 percent in sibship size 4, and 76 percent in sibship size 5. We have not included sibships of six 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 three, five for sibship size four, and six for sibship size five. 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 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.[8]

[8] We also report correlation matrices for all groups of each sibship size in Appendix A.

Analyses and Findings

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.[9] Third, we evaluate the findings in the first two sections by developing and testing a 2-factor MIMIC model of educational attainment. The equation of the single factor model of educational attainment is

$$X = \Lambda^x \xi + \delta \tag{7}$$

where X is a vector of siblings' educational attainments; ξ is the common family factor; Λ^x is the matrix of loadings of Xs on ξ; and δ is a vector of unmeasured unique factors. In matrix form, the model is[10]

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

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[9] In addition to χ^2, we use the *bic* statistic to evaluate goodness of fit. The *bic* statistic is based on Bayesian theory for *a posteriori* tests: $bic = \chi^2 - df \times ln[N \times (p + q)]$, where *df* are the degrees of freedom under the tested model or contrast, N is the sample size, p is the number of observed exogenous variables, and q is the number of observed endogenous variables. Satisfactory fit is indicated by a negative value of *bic*, and models with lower *bic* statistics are preferred (Raftery 1986, 1993).

[10] Without 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.

where k indexes groups defined by gender composition, e.g., the 5 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 within gender. For example, in sibships of size 4, we impose the following constraints: $\lambda^x_{111} = \lambda^x_{121} = \lambda^x_{131} = \lambda^x_{141} = \lambda^x_{221} = \lambda^x_{231} = \lambda^x_{241} = \lambda^x_{331} = \lambda^x_{341} = \lambda^x_{441} = 1$ and $\lambda^x_{211} = \lambda^x_{311} = \lambda^x_{321} = \lambda^x_{411} = \lambda^x_{421} = \lambda^x_{431} = \lambda^x_{511} = \lambda^x_{521} = \lambda^x_{531} = \lambda^x_{541}$; $\theta^\delta_{111} = \theta^\delta_{121} = \theta^\delta_{131} = \theta^\delta_{141} = \theta^\delta_{221} = \theta^\delta_{231} = \theta^\delta_{241} = \theta^\delta_{331} = \theta^\delta_{341} = \theta^\delta_{441}$ and $\theta^\delta_{211} = \theta^\delta_{311} = \theta^\delta_{321} = \theta^\delta_{411} = \theta^\delta_{421} = \theta^\delta_{431} = \theta^\delta_{511} = \theta^\delta_{521} = \theta^\delta_{531} = \theta^\delta_{541}$; $\phi_{111} = \phi_{211} = \phi_{311} = \phi_{411} = \phi_{511}$. 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.[11] In the baseline model, the goodness of fit test yields $\chi^2 = 51.09$ with 20 *df* for size 3, $\chi^2 = 89.07$ with 46 *df* for size 4, and $\chi^2 = 183.93$ with 86 *df*. Each of these test statistics is nominally statistically significant; however, in each case, the corresponding *bic* statistic is negative, *bic* = -120.68, *bic* = -299.99 and *bic* = -527.73 for sizes 3, 4 and 5 respectively. That is, the baseline model can nominally be rejected for each size of

[11] That 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.

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 χ^2 = 12.75 with 10 *df*, and in sibships of size 5, χ^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 χ^2 = 5.79 for size 3, χ^2 = 24.12 for size 4, and χ^2 = 9.81 for size 5 with 1 *df* for each. They are all nominally significantly 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.[12] We find λ^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^δ 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 $\chi^2 = 157.87$; for size 4, the contrast yields $\chi^2 = 99.57$; and for size 5, the contrast yields $\chi^2 = 101.71$, each with 1 *df*. Also, there are corresponding increases in *bic*. These contrasts confirm the well-established finding that, in the cohorts of the 1950s, educational attainment is far more variable among men than among women (Hauser and Wong 1989, pp. 158-9; Sewell, Hauser, and Wolf 1980, p. 557). 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

[12] Except 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.

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 χ^2 = 4.45 with 4 *df* for size 3, χ^2 = 9.91 with 6 *df* for size 4 and χ^2 = 21.87 with 8 *df* for size 5. The fit of size 5 is on the borderline, but the *bic* statistic increases by 44.33. 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^δ. The contrasts of fit statistics are χ^2 = 19.61 with 6 *df* for size 3, χ^2 = 35.26 with 12 *df* for size 4, and χ^2 = 40.67 with 20 *df* for size 5; they are each significant at the 0.003 level. However, the *bic* statistics increase by 31.92 for size 3, by 66.23 for size 4, and by 124.83 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: χ^2 = 1.48 with 2 *df* for size 3; χ^2 = 1.26 with 3 *df* for size 4; and χ^2 = 3.58 with 4 *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 χ^2 = 13.76 with 5 *df* for size 3; χ^2 = 34.19 with 11 *df* for size 4; and χ^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 χ^2 = 0.26 with 3 *df* for size 3; χ^2 = 6.74 with 4 *df* for size 4; and χ^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.[13]

[13] We had thought, from other analyses of family effects, that larger families were more heterogeneous than smaller families. Our finding raises the possibility that this 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.

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 family background variables 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).[14] 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 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 family background variables on educational attainments of offspring, and it may or may not contain a random disturbance, ζ, which is independent of measured background.[15] The measurement models for latent variables are

$$x = \Lambda^x \xi + \delta \quad (10)$$

$$y = \Lambda^y \eta + \varepsilon \quad (11)$$

[14] The 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 sibling's educational attainments, as well as a test of the overidentifying restrictions implied by the single-factor model of educational attainment.

[15] For convenience, we have suppressed notation for the groups defined by gender composition.

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. 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, are equal to zero; that is, all ξs are perfectly measured by corresponding xs. Equations 11 and 9 in matrix form are

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

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

where λ_{11} in equation 12 is normalized as 1 and then other λs are proportional to it. However, in our model, we have elaborated the structure of the ys (educational attainments) as follows:

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

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 βs are effects of the common family factor, η_1, on latent individual factors, $\eta_2,\ldots,\eta_5$; λ^ys are now effects of the latent individual factors, $\eta_2,\ldots,\eta_5$, on educational attainments of siblings, $y_1,\ldots,y_4$; and εs are disturbances in the ys, i.e., measurement errors and/or non-shared 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 βs and λ^ys separately (Sörbom and Jöreskog 1981); later, we use this feature of the modified model to specify effects of relative birth order and of gender.[16]

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 constrained: λ^ys of brothers are normalized to one and λ^ys of sisters are equated within and across groups defined by gender composition; Θ^εs of brothers are equal across groups, as are those of sisters; Γs and Ψs are invariant across groups; and γs of father's and mother's educational levels are equal. For example, in sibship size 4, the constraints

[16] Another advantage to this specification is that the extra latent variables (ηs) 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.

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 $\lambda^y_{212} = \lambda^y_{312} = \lambda^y_{323} = \lambda^y_{412} = \lambda^y_{423} = \lambda^y_{434} = \lambda^y_{512} = \lambda^y_{523} = \lambda^y_{534} = \lambda^y_{545}$; $\theta^\varepsilon_{111} = \theta^\varepsilon_{122} = \theta^\varepsilon_{133} = \theta^\varepsilon_{144} = \theta^\varepsilon_{222} = \theta^\varepsilon_{233} = \theta^\varepsilon_{244} = \theta^\varepsilon_{333} = \theta^\varepsilon_{344} = \theta^\varepsilon_{444}$ and $\theta^\varepsilon_{211} = \theta^\varepsilon_{311} = \theta^\varepsilon_{322} = \theta^\varepsilon_{411} = \theta^\varepsilon_{422} = \theta^\varepsilon_{433} = \theta^\varepsilon_{511} = \theta^\varepsilon_{522} = \theta^\varepsilon_{533} = \theta^\varepsilon_{544}$; $\psi_{111} = \psi_{211} = \psi_{311} = \psi_{411} = \psi_{511}$; $\gamma_{111} = \gamma_{211} = \gamma_{311} = \gamma_{411} = \gamma_{511}$, $\gamma_{112} = \gamma_{212} = \gamma_{312} = \gamma_{412} = \gamma_{512} = \gamma_{113} = \gamma_{213} = \gamma_{313} = \gamma_{413} = \gamma_{513}$, $\gamma_{114} = \gamma_{214} = \gamma_{314} = \gamma_{414} = \gamma_{514}$, $\gamma_{115} = \gamma_{215} = \gamma_{315} = \gamma_{415} = \gamma_{515}$, and $\gamma_{116} = \gamma_{216} = \gamma_{316} = \gamma_{416} = \gamma_{516}$.[17] The model suggests that, first, gender composition does not affect the influence of the common family factor on the educational attainments of sisters and brothers, the within-family variances of schooling for sisters and brothers, the effect of family background variables 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

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[17] As 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 χ^2 = 85.65 with 48 *df* for size 3, *bic* = -379.34; χ^2 = 149.46 with 100 *df* for size 4, *bic* = -787.96; and χ^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 χ^2 = 17.17 with 8 *df* for size 3, *bic* = -60.33; χ^2 = 49.11 with 25 *df* for size 4, *bic* = -185.24; and χ^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.

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, Hauser, and Wolf (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 χ^2 = 153.48 with 87 *df* for size 3, χ^2 = 248.04 with 161 *df* for size 4, and χ^2 = 394.14 with 261 *df* for size 5. They are highly significant statistically, but the *bic* statistics are quite satisfactory: *bic* = -689.31, *bic* = -1261.20 and *bic* = -1971.45 for sizes 3, 4, and 5 respectively.[18] 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.[19]

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

[19] As 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

In model B, we test the hypothesis that the within-family variances of schooling for brother and for 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 χ^2 = 2.80 with 4 *df* for size 3, χ^2 = 10.71 with 6 *df* for size 4, and χ^2 = 20.18 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 of 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^ϵ. The contrasts of fit statistics yield χ^2 = 21.99 with 6 *df* for size 3, χ^2 = 29.87 with 12 *df* for size 4, and χ^2 = 39.52 with 20 *df* for size 5; they are significant at 0.006 level and beyond. However, the *bic* statistics increase by 36.14 for size 3, 82.62 for size 4, and 141.76 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^ys. The contrasts between model D and model A yield χ^2 = 2.34 with 2 *df* for size 3, χ^2 = 0.92 with 3 *df* for size 4, and χ^2 = 3.02 with 4 *df* for size 5; none of these contrast approaches statistical significance.

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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.

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^y, the contrasts of fit between model D and model E are χ^2 = 10.27 with 5 *df* for size 3, χ^2 = 30.83 with 11 *df* for size 4, and χ^2 = 27.75 with 19 *df* for size 5; however, the *bic* statistics increase by 38.16, 72.29, and 144.46 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 educational attainments of offspring differ. We reject this hypothesis because the contrasts of fit between model A and model F are negligible: χ^2 = 0.25, χ^2 = 0.45, and χ^2 = 0.44 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 χ^2 = 15.43 with 18 *df* for size 3, χ^2 = 25.16 with 24 *df* for size 4, and χ^2 = 33.53

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: $\chi^2 = 2.83$ with 3 *df* for size 3, $\chi^2 = 7.16$ with 4 *df* for size 4, and $\chi^2 = 5.20$ 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^y) 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 ($\lambda_{212} * \beta_{131}$) for other sisters. Compared to model A, the current model has one *df* less but χ^2 does not decrease significantly in sizes 3 and 5 (1.64 and 3.13); in size 4, the contrast of χ^2 is 5.96, yet the *bic* statistics increases by 3.41; 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 4.04) 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^{ε}) 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 1.27 for sizes 3, 4, and 5, respectively.[20] The effects of the family factors on sisters, 0.84, 0.76 and 0.80, 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

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[20] These 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.

background variable, the one factor model may not capture all of the differences in effects of family background variables between brothers and sisters.[21] In the model of figure 4, one additional latent factor is added to equation 9, that is, in matrix form,

$$\begin{bmatrix} \eta_1 \\ \eta_2 \end{bmatrix} = \begin{bmatrix} \gamma_{11} & \gamma_{12} & \gamma_{13} & \gamma_{14} & \gamma_{15} & \gamma_{16} \\ \gamma_{21} & \gamma_{22} & \gamma_{23} & \gamma_{24} & \gamma_{25} & \gamma_{26} \end{bmatrix} \begin{bmatrix} \xi_1 \\ \xi_2 \\ \xi_3 \\ \xi_4 \\ \xi_5 \\ \xi_6 \end{bmatrix} + \begin{bmatrix} \zeta_1 \\ \zeta_2 \end{bmatrix} \tag{16}$$

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

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

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

are identical loadings of observable exogenous variables on each factor. By releasing restrictions on Γ and Ψ, we can test alternatives to the one-factor MIMIC model of sibling resemblance. However, the two-factor model is underidentified 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 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: χ^2 = 153.48 with 87 *df* for size 3, χ^2 = 248.04 with 161 *df* for size 4, and χ^2 = 394.14 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 χ^2 = 18.97 with 5 *df* for size 3, χ^2 = 19.24 with 5 *df* for size 4, and χ^2 = 7.59 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 29.47 and 27.63. 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 χ^2 = 12.68, χ^2 = 0.57, and χ^2 = 11.08 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 still increase by 6.7. 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.[22] The MIMIC model is now defined by the following equations:

$$\eta = \alpha + \Gamma \xi + \zeta \tag{18}$$

$$x = \tau^x + \Lambda^x \xi + \delta \tag{19}$$

$$y = \tau^y + \Lambda^y \eta + \varepsilon \tag{20}$$

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:

[22] This strictly parallels the use specification of omitted categories in dummy-variable regression analysis.

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

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

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 χ^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 31.09 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 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 $\chi^2 = 7.56$ with 4 *df* for size 3, $\chi^2 = 4.75$ with 6 *df* for size 4, and $\chi^2 = 9.99$ 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 χ^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 matrix of means of family background variables) is invariant across groups. The contrasts of χ^2 (with model N) are insignificant: 16.83 with 18 *df*, 19.29 with 24 *df*, and 50.42 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 $\chi^2 = 57.12$, $\chi^2 = 82.44$ and $\chi^2 = 90.69$ 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 αs 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 αs 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 $\chi^2 = 1.71$ with 3 *df* for size 3, $\chi^2 = 2.82$ with 4 *df* for size 4, and $\chi^2 = 5.14$ with 5 *df* for size 5; the non-significance 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 12.72 in size 5 for sisters, and 13.77 in size 3, 13.39 in size 4 and 13.05 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 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 variables differ between brothers and sisters.[23] Thus, our findings agree with those of Tsai (1983) and Lee (1989) but slightly differ from those of Sewell, Hauser, and Wolf (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

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

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.[24]

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 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.

We believe that a great deal more can be learned about family resemblance by bringing in more individual and family variables, and this is not feasible in most studies, including the Wisconsin study, for more than two siblings in each family.

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[24] Again, 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.

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 1993).

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Table 1. Distribution of Sibship Size and Gender Composition: Wisconsin Longitudinal Study

Sibship Size	1	2	3	4	5	6	7+	Total
Number of Sisters								
0	320	425	241	105	29	19	15	1154
	46.9	21.3	12.1	7.2	2.8	2.8	1.2	12.7
1	362	1051	742	367	169	67	41	2799
	53.1	52.6	37.3	25.2	16.4	9.8	3.3	30.8
2		522	772	542	321	154	122	2433
		26.1	38.8	37.1	31.1	22.4	9.9	26.8
3			233	368	313	191	243	1348
			11.7	25.2	30.3	27.8	19.7	14.8
4				77	169	170	305	721
				5.3	16.4	24.8	24.7	7.9
5					31	73	256	360
					3.0	10.6	20.7	4.0
6+						12	254	266
						1.7	20.6	2.9
Total	682	1998	1988	1459	1032	686	1236	9081
	7.5	22.0	21.9	16.1	11.4	7.6	13.6	100.0

	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 Variables (Years of Schooling)															
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 Variables															
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 Occupation	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 Education	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 Education	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)	(0.14)	(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)

Table 3. Goodness of Fit Statistics, One-Factor Model

Model		χ^2	*df*	*bic*	Contrast		*p*
Sibship Size = 3 (n = 1790)							
A.	Baseline Model	51.09	20	-120.68			0.000
B.	A + λ^xs invariant by sex	56.88	21	-123.48	5.79	1	0.016
C.	A + Θ^δ invariant by sex	208.96	21	28.60	157.87	1	0.000
D.	A - Θ^δ across-group constraint	46.64	16	-90.78	-4.45	4	0.349
E.	D - Θ^δ within-group constraint	27.03	10	-58.86	-19.61	6	0.003
F.	A - λ^xs across-group constraint	49.61	18	-104.98	-1.48	2	0.477
G.	F - λ^xs within-group constraint	35.85	13	-75.80	-13.76	5	0.017
H.	A - Φ across-group constraint	50.83	17	-95.18	-0.26	3	0.967
Sibship Size = 4 (n = 1178)							
A.	Baseline Model	89.07	46	-299.99			0.000
B.	A + λ^xs invariant by sex	113.19	47	-284.33	24.12	1	0.000
C.	A + Θ^δ invariant by sex	188.64	47	-208.88	99.57	1	0.000
D.	A - Θ^δ across-group constraint	79.16	40	-259.15	-9.91	6	0.128
E.	D - Θ^δ within-group constraint	43.90	28	-192.92	-35.26	12	0.001
F.	A - λ^xs across-group constraint	87.81	43	-275.88	-1.26	3	0.739
G.	F - λ^xs within-group constraint	53.62	32	-217.03	-34.19	11	0.000
H.	A - Φ across-group constraint	82.33	42	-272.90	-6.74	4	0.150
Sibship Size = 5 (n = 785)							
A.	Baseline Model	183.93	86	-527.73			0.000
B.	A + λ^xs invariant by sex	193.74	87	-526.20	9.81	1	0.002
C.	A + Θ^δ invariant by sex	285.64	87	-434.30	101.71	1	0.000
D.	A - Θ^δ across-group constraint	162.06	78	-483.40	-21.87	8	0.005
E.	D - Θ^δ within-group constraint	121.39	58	-358.57	-40.67	20	0.004
F.	A - λ^xs across-group constraint	180.35	82	-498.21	-3.58	4	0.466
G.	F - λ^xs within-group constraint	149.69	63	-371.64	-30.66	19	0.044
H.	A - Φ across-group constraint	183.23	81	-487.05	-0.70	5	0.983

Table 4. Goodness of Fit Statistics (Sibship Size = 3, n = 1790)

Model		χ^2	*df*	*bic*	Contrast		*p*
Structural Model							
A.	Baseline Model	153.48	87	-689.31			.000
B.	A - θ^{ε} cross-group constraints	150.68	83	-653.36	2.80	4	.592
C.	B - θ^{ε} within-group constraints	128.69	77	-617.22	21.99	6	.001
D.	A - Λ^{y} cross-group constraints	151.14	85	-672.27	2.34	2	.310
E.	D - Λ^{y} within-group constraints	140.87	80	-634.11	10.27	5	.068
F.	A - $\gamma_{k,2}=\gamma_{k,3}$ constraint	153.23	86	-679.87	0.25	1	.617
G.	F - Γ cross-group constraints	137.80	68	-520.93	15.43	18	.632
H.	A - Ψ cross-group constraints	150.65	84	-663.07	2.83	3	.419
I.	A + birth order constraints	151.84	86	-681.26	1.64	1	.200
Two-factor Model							
J.	Baseline Model, Model A	153.48	87	-689.31			.000
K.	A - Γ within-groups constraints	134.51	82	-659.84	18.97	5	.002
L.	A - Ψ within-group constraints	140.80	85	-682.61	12.68	2	.002
Mean Structure Model							
M.	A + τ^{y} within-group constraints	157.40	93	-743.51	3.92	6	.688
N.	M + τ^{y} cross-group constraints	164.96	97	-774.70	7.56	4	.109
O.	N + $\tau_{sister} = \tau_{brother}$	265.96	98	-683.39	101.00	1	.000
P.	N + τ^{x} cross-group constraints	181.79	115	-932.24	16.83	18	.535
Q.	P + $\tau_{k,2} = \tau_{k,3}$	238.91	116	-884.80	57.12	1	.000
R.	P - $\alpha_{k,1}$s cross-group constraints	180.08	112	-904.89	1.71	3	.635

Table 5. Goodness of Fit Statistics (Sibship Size = 4, n = 1178)

Model		χ^2	*df*	*bic*	Contrast		*p*
Structural Model							
A.	Baseline Model	248.04	161	-1261.20			.000
B.	A - θ^{ϵ} cross-group constraints	237.33	155	-1215.66	10.71	6	.098
C.	B - θ^{ϵ} within-group constraints	207.46	143	-1133.04	29.87	12	.003
D.	A - Λ^{y} cross-group constraints	247.12	158	-1234.00	0.92	3	.821
E.	D - Λ^{y} within-group constraints	216.29	147	-1161.71	30.83	11	.001
F.	A - $\gamma_{k,2}=\gamma_{k,3}$ constraint	247.59	160	-1252.28	0.45	1	.502
G.	F - Γ cross-group constraints	222.43	136	-1052.46	25.16	24	.397
H.	A - Ψ cross-group constraints	240.88	157	-1230.86	7.16	4	.128
I.	A + birth order constraints	242.08	160	-1257.79	5.96	1	.015
Two-Factor Model							
J.	Baseline Model, Model A	248.04	161	-1261.20			.000
K.	A - Γ within-groups constraints	228.80	156	-1233.57	19.24	5	.002
L.	A - Ψ within-group constraints	247.47	159	-1243.02	0.57	2	.752
Mean Structure Model							
M.	A + τ^{y} within-group constraints	258.24	173	-1363.49	10.20	12	.599
N.	M + τ^{y} cross-group constraints	262.99	179	-1414.98	4.75	6	.576
O.	N + $\tau_{sister} = \tau_{brother}$	318.68	180	-1368.67	55.69	1	.000
P.	N + τ^{x} cross-group constraints	282.28	203	-1620.67	19.29	24	.736
Q.	P + $\tau_{k,2} = \tau_{k,3}$	364.72	204	-1547.61	82.44	1	.000
R.	P - $\alpha_{k,1}$s cross-group constraints	279.46	199	-1586.00	2.82	4	.588

Table 6. Goodness of Fit Statistics (Sibship Size = 5, n = 785)

Model		χ^2	*df*	*bic*	Contrast		*p*
Structural Model							
A.	Baseline Model	394.14	261	-1971.45			.000
B.	A - θ^{ε} cross-group constraints	373.96	253	-1919.13	20.18	8	.010
C.	B - θ^{ε} within-group constraints	334.44	233	-1777.37	39.52	20	.006
D.	A - Λ^{y} cross-group constraints	391.12	257	-1938.22	3.02	4	.554
E.	D - Λ^{y} within-group constraints	363.37	238	-1793.76	27.75	19	.088
F.	A - $\gamma_{k,2}=\gamma_{k,3}$ constraint	393.70	260	-1962.83	0.44	1	.507
G.	F - Γ cross-group constraints	360.17	230	-1724.45	33.53	30	.300
H.	A - Ψ cross-group constraints	388.94	256	-1931.34	5.20	5	.392
I.	A + birth order constraints	391.01	260	-1965.52	3.13	1	.077
Two-Factor Model							
J.	Baseline Model	394.14	261	-1971.45			.000
K.	A - Γ within-group constraints	386.55	256	-1933.73	7.59	5	.180
L.	A - Ψ within-group constraints	383.06	259	-1964.41	11.08	2	.004
Mean Structure Model							
M.	A + τ^{y} within-group constraints	425.23	281	-2121.64	31.09	20	.054
N.	M + τ^{y} cross-group constraints	435.22	289	-2184.15	9.99	8	.266
O.	N + $\tau_{sister} = \tau_{brother}$	463.06	290	-2165.38	27.84	1	.000
P.	N + τ^{x} cross-group constraints	485.64	319	-2405.64	50.42	30	.011
Q.	P + $\tau_{k,2} = \tau_{k,3}$	576.33	320	-2324.02	90.69	1	.000
R.	P - $\alpha_{k,1}$s cross-group constraints	480.50	314	-2365.46	5.14	5	.399

Table 7. Parameter Estimates of Preferred Models

	Sibship Size 3		*Sibship Size 4*		*Sibship Size 5*	
Structural Coefficents						
Effect of the Common Family Factor on (Λ^y)						
Sister	.84	(.04)	.76	(.80)	.80	(.04)
Brother	1.00		1.00		1.00	
Loadings of Exogenous Variables on the Common Family Factor (Γ)						
Income	.09	(.02)	.10	(.02)	.11	(.02)
Father's education	.13	(.01)	.13	(.01)	.11	(.01)
Mother's education	.13	(.01)	.13	(.01)	.11	(.01)
Father's SEI	.15	(.02)	.15	(.03)	.14	(.03)
Farmer	.26	(.12)	.35	(.13)	.29	(.14)
Catholic	-.22	(.08)	.16	(.10)	.09	(.10)
Variances						
Unmeasured Family Variance (ψ)	1.57	(.13)	1.42	(.13)	1.27	(.13)
Within-Family Variances (θ^ϵ)						
Sister	2.35	(.09)	2.54	(.09)	2.24	(.08)
Brother	4.71	(.16)	4.39	(.15)	4.04	(.15)
Means						
Means of Schooling (τ^y)						
Sister	13.20	(.05)	12.93	(.05)	12.72	(.05)
Brother	13.77	(.06)	13.39	(.06)	13.05	(.07)
Means of Exogenous Variables (τ^x)						
Income	6.23	(.08)	5.89	(.09)	5.19	(.10)
Father's Education	10.14	(.08)	9.56	(.10)	9.15	(.11)
Mother's Education	10.72	(.07)	10.40	(.08)	10.13	(.01)
Father's SEI	3.72	(.06)	3.35	(.06)	2.97	(.07)
Farmer	.17	(.01)	.20	(.01)	.26	(.02)
Catholic	.38	(.01)	.44	(.01)	.44	(.02)

Figure 1.
Variable Description

Variable[1]	Source	Description[2]
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 avaialbe, taking from 1975 report (YFML57).	per $100. 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.
Farmer Background	Recoded variable from OCMH57 (1957 survey).	1: Farmer/farm managers/farm laborers and foremen (16 and 17 in OCMH57); 0: otherwise.

[1]Variables are in the order of correlation matrix (Appendix A.)

[2]See Table 2 for Standard Deviation and Mean.

Figure 2. Multiple-Group One-Factor Model (Sibship Size = 4).

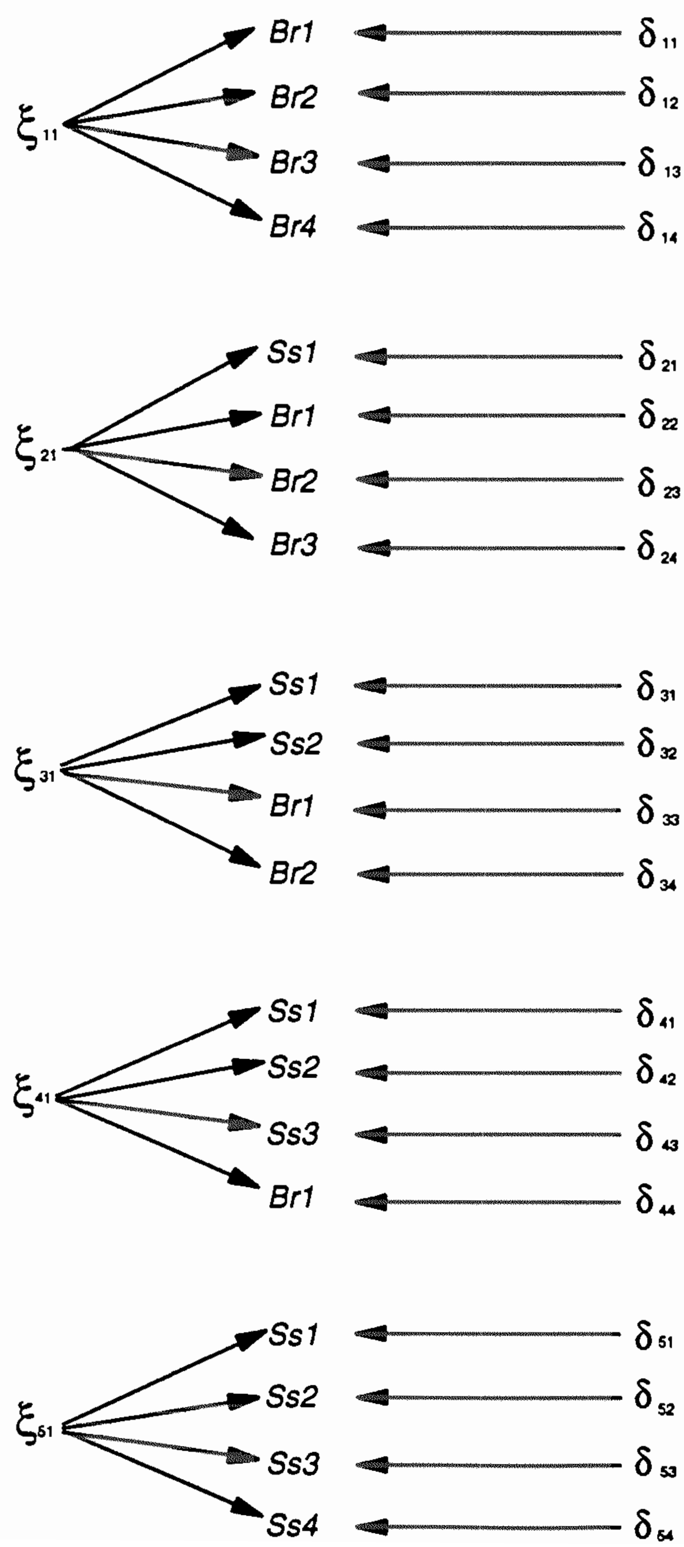


Figure 3. Multi-Group Revised MIMIC Model (Sibship Size = 4).

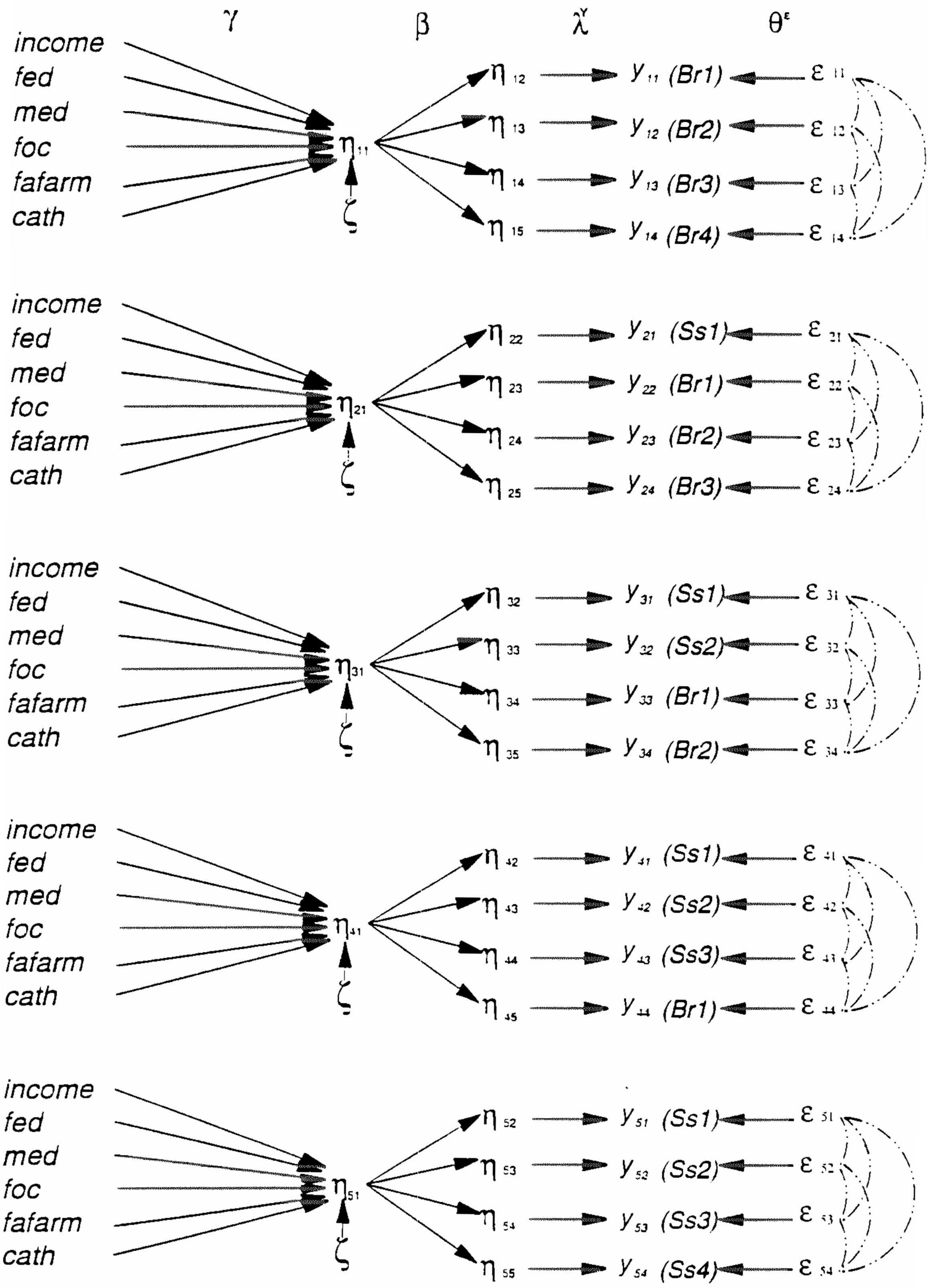
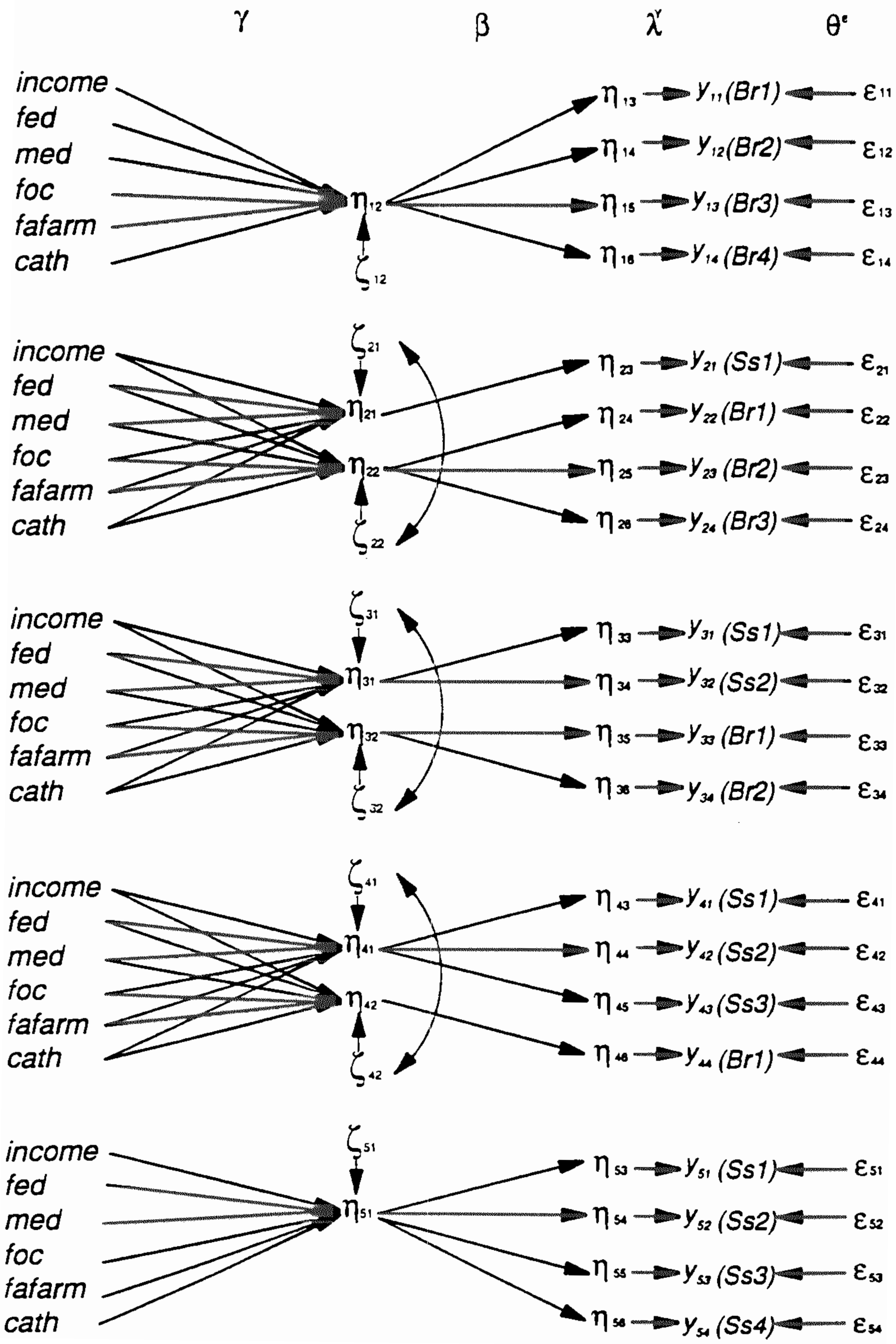


Figure 4. Multi-Group Two-Factor MIMIC Model (Sibship Size = 4).



Appendix A. Correlation Matrices

a) Sibship Size = 3

1. Number of Sisters = 0 (N = 222)

M1ED	M2ED	M3ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
1.0000									
.2999	1.0000								
.3924	.4022	1.0000							
.1515	.1964	.2707	1.0000						
.3160	.2861	.2167	.4272	1.0000					
.2766	.2150	.2286	.2358	.5039	1.0000				
.1591	.1751	.2457	.1361	.3443	.4752	1.0000			
-.1111	-.1152	-.0636	-.3351	-.4147	-.0926	-.0565	1.0000		
.0260	-.0777	.0060	.0184	.1268	-.1223	-.0690	-.1322	1.0000	
-.0923	-.1103	-.0685	-.3152	-.4514	-.0967	-.1143	.8807	-.1716	1.0000
Mean									
13.8694	13.6757	13.6216	5.8405	3.4918	9.8739	10.7162	.1802	.3874	.1892
Standard Deviation									
2.9377	2.7560	2.3747	2.8822	2.2828	3.4210	2.6304	.3852	.4883	.3925

2. Number of Sisters = 1 (N = 672)

F1ED	M1ED	M2ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
1.0000									
.4769	1.0000								
.3524	.4041	1.0000							
.3308	.3419	.2832	1.0000						
.3499	.3813	.3540	.5472	1.0000					
.3601	.4294	.3292	.4381	.5626	1.0000				
.3259	.3247	.3083	.3003	.3745	.4770	1.0000			
-.0890	-.1154	-.1966	-.3286	-.3957	-.2254	-.0922	1.0000		
-.0889	-.0788	-.0458	.0293	-.0087	-.0821	-.1039	-.1116	1.0000	
-.1043	-.1177	-.2276	-.3467	-.4400	-.2223	-.1053	.8384	-.0972	1.0000
Mean									
13.3006	13.9583	13.9539	6.5204	3.9561	10.3601	10.9256	.1488	.3795	.1548
Std. Dv.									
2.0242	2.8580	2.6267	3.3598	2.4876	3.5459	2.7472	.3562	.4856	.3619

3. Number of Sisters = 2 (N = 695)

F1ED	F2ED	M1ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
1.0000									
.5069	1.0000								
.3545	.3390	1.0000							
.2543	.2945	.2905	1.0000						
.3243	.2789	.3130	.5214	1.0000					
.3856	.3416	.3261	.4061	.5367	1.0000				
.2972	.2961	.3086	.3052	.3634	.4678	1.0000			
-.0107	-.0383	-.1293	-.3484	-.3621	-.1993	-.0449	1.0000		
-.0979	-.0531	-.0466	.0026	.0036	-.0636	-.1110	-.1027	1.0000	
-.0844	-.0640	-.1593	-.3571	-.4444	-.2293	-.0788	.8222	-.0846	1.0000
Mean									
13.1050	13.2144	13.6101	6.1758	3.6466	10.1554	10.6590	.1755	.3784	.1784
Standard Deviation									
2.0510	2.0905	2.7064	3.3082	2.3838	3.3993	2.8974	.3807	.4853	.3831

4. Number of Sisters = 3 (N = 201)

F1ED	F2ED	F3ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
1.0000									
.5118	1.0000								
.4002	.3992	1.0000							
.2144	.2049	.1153	1.0000						
.2887	.3820	.1131	.4584	1.0000					
.4044	.3715	.1455	.2924	.5179	1.0000				
.4007	.3802	.3162	.2458	.2609	.5156	1.0000			
-.0188	-.0585	.0479	-.3078	-.3547	-.1875	-.0400	1.0000		
-.0832	-.1216	-.0733	.1325	.0938	.0188	-.1685	-.0845	1.0000	
-.0188	-.1175	.0543	-.3628	-.4565	-.2092	-.0097	.8513	-.1675	1.0000
Mean									
13.1194	13.0597	13.1741	6.1388	3.5873	9.8557	10.2786	.1592	.4080	.1592
Standard Deviation									
2.0483	2.0776	2.1153	3.1732	2.1730	3.1534	2.7004	.3668	.4927	.3668

b) Sibship Size = 4

1. Number of Sisters = 0 (N = 82)

	M1ED	M2ED	M3ED	M4ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
	1.0000										
	.3641	1.0000									
	.2784	.1404	1.0000								
	.1524	.1573	.2143	1.0000							
	.1529	.2425	.3460	.0760	1.0000						
	.2088	.2349	.2180	.2113	.3032	1.0000					
	.2299	.0775	.0898	.2292	.1005	.2589	1.0000				
	.2053	.1092	.0944	.1072	.0707	.4514	.3726	1.0000			
	-.0273	-.0223	.0282	-.0075	-.3043	-.4372	-.1470	-.2407	1.0000		
	.0070	-.0974	.0267	.0443	-.1356	-.0311	.0743	-.2575	.0051	1.0000	
	-.0380	-.1430	-.0182	.0125	-.3317	-.5225	-.1650	-.2943	.8599	.0600	1.0000
Mean	12.9390	12.8659	13.5610	13.0976	5.6183	3.0805	9.1707	10.0000	.2073	.5244	.2317
Standard Deviation	2.1736	2.3294	2.6485	2.6649	2.9151	1.7771	2.8623	2.7666	.4079	.5025	.4245

2. Number of Sisters = 1 (N = 305)

	F1ED	M1ED	M2ED	M3ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
	1.0000										
	.3195	1.0000									
	.3993	.4419	1.0000								
	.3139	.3539	.4594	1.0000							
	.2461	.3335	.1941	.2074	1.0000						
	.2991	.3816	.3322	.3492	.5526	1.0000					
	.3188	.3416	.3571	.3940	.4045	.5335	1.0000				
	.3450	.2916	.3229	.2537	.2472	.3269	.4614	1.0000			
	-.0564	-.1460	-.0620	-.1251	-.3166	-.4041	-.1974	-.0165	1.0000		
	-.0547	.0974	.0670	.0648	.1272	.1198	.0688	-.0866	-.1799	1.0000	
	-.1048	-.1865	-.0820	-.1472	-.3574	-.4561	-.2471	-.0799	.8207	-.1886	1.0000
Mean	13.0820	13.4492	13.6459	13.4131	6.0344	3.5394	9.7508	10.5967	.2131	.4328	.2000
Standard Deviation	2.0366	2.7032	2.8434	2.2984	3.3265	2.3493	3.6487	2.8141	.4102	.4963	.4007

3. Number of Sisters = 2 (N = 440)

	F1ED	F2ED	M1ED	M2ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
	1.0000										
	.4016	1.0000									
	.4046	.3817	1.0000								
	.3572	.3778	.3677	1.0000							
	.2442	.2879	.2903	.2620	1.0000						
	.2305	.3479	.3616	.3094	.5478	1.0000					
	.2783	.3147	.3220	.2441	.3751	.5302	1.0000				
	.2945	.2129	.2997	.2125	.2818	.3185	.4465	1.0000			
	.0234	-.0563	-.1565	-.1622	-.3114	-.3559	-.2245	-.0755	1.0000		
	.0359	.0679	.1435	.0425	.1579	.1515	.0739	-.0715	-.1061	1.0000	
	.0187	-.0813	-.1731	-.1645	-.3146	-.4101	-.2466	-.0693	.8138	-.1485	1.0000
Mean	12.9318	12.9955	13.3636	13.4568	5.8848	3.2284	9.4955	10.3909	.1750	.4523	.1955
Standard Deviation	2.1108	1.8778	2.7927	2.5910	3.2814	2.2106	3.3654	2.8647	.3804	.4983	.3970

4. Number of Sisters = 3 (N = 288)

	F1ED	F2ED	F3ED	M1ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
	1.0000										
	.3300	1.0000									
	.3146	.4102	1.0000								
	.4053	.3309	.3046	1.0000							
	.4122	.3217	.2335	.3272	1.0000						
	.3173	.1940	.1894	.2625	.5210	1.0000					
	.3952	.3467	.2499	.4019	.3602	.4707	1.0000				
	.3374	.3400	.2741	.3456	.3325	.2838	.5050	1.0000			
	-.0670	-.0192	-.0556	-.1112	-.3776	-.4422	-.1467	-.0745	1.0000		
	.0061	.0746	.0529	.0389	.1370	.0280	-.0082	-.0188	-.0695	1.0000	
	-.0786	-.0668	-.0551	-.1619	-.3867	-.4656	-.2032	-.0909	.8624	-.0413	1.0000
Mean	12.8993	12.8090	12.9549	13.3924	5.9083	3.4633	9.7500	10.2917	.2049	.4063	.2083
Standard Deviation	2.0669	2.1221	1.9113	2.5680	3.1541	2.2772	3.2449	2.6853	.4043	.4920	.4068

5. Number of Sisters = 4 (N = 63)

	F1ED	F2ED	F3ED	F4ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
	1.0000										
	.5031	1.0000									
	.3457	.5607	1.0000								
	.1917	.2769	.3956	1.0000							
	.2401	.1581	.1739	.1430	1.0000						
	.3863	.2722	.2700	.4383	.5834	1.0000					
	.6461	.2924	.3124	.2796	.2583	.4774	1.0000				
	.5460	.3625	.4008	.2387	.1755	.3199	.5432	1.0000			
	-.3064	-.1342	-.0095	-.0458	-.3926	-.4493	-.1926	-.1594	1.0000		
	-.1250	-.0624	.0206	.0748	.1918	-.0304	-.0646	-.1480	-.1387	1.0000	
	-.2255	-.0915	-.0095	.0528	-.4437	-.4912	-.1419	-.0921	.8971	-.1387	1.0000
Mean	12.7619	12.8571	12.6984	12.9048	5.7619	3.5844	9.3492	10.5397	.1905	.4762	.1905
Std. Dev	2.0138	1.9080	1.6327	1.6531	2.7134	2.2385	3.2137	2.4218	.3958	.5034	.3958

c) Sibship Size = 5

1. Number of Sisters = 0 (N = 22)

M1ED	M2ED	M3ED	M4ED	M5ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
1.0000											
.8078	1.0000										
.5503	.4344	1.0000									
.4715	.4204	.7787	1.0000								
.2615	.4593	.4842	.4909	1.0000							
.5041	.2151	.3980	.2681	-.0892	1.0000						
.2344	.2365	.3713	.3083	.5884	-.0049	1.0000					
.4958	.2309	.4844	.5973	.1874	.3781	.2167	1.0000				
.1687	-.0561	.2081	.2971	.1978	.1075	.3665	.4427	1.0000			
.1393	.2153	.0471	-.1648	.0709	-.2302	-.3122	-.1820	-.2757	1.0000		
-.1284	-.2067	-.0567	.1558	-.4201	.2416	-.3253	.1487	.0462	-.5423	1.0000	
.1393	.2153	-.1332	-.1648	.0152	-.1633	-.4351	-.2833	-.0551	.7412	-.3254	1.0000
Mean											
12.2727	12.7273	12.5909	12.9091	13.5455	4.9546	2.3205	8.3182	9.5000	.2273	.5000	.2273
Standard Deviation											
2.8980	2.2505	2.4623	2.3886	1.9935	2.1571	1.1742	2.1906	3.0198	.4289	.5118	.4289

2. Number of Sisters = 1 (N = 130)

F1ED	M1ED	M2ED	M3ED	M4ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
1.0000											
.4226	1.0000										
.3500	.2342	1.0000									
.3297	.2867	.3353	1.0000								
.3819	.4308	.2503	.3875	1.0000							
.1755	.2900	.3072	.2699	.2436	1.0000						
.1754	.2012	.1038	.3068	.2579	.5456	1.0000					
.2734	.1891	.2834	.2782	.1871	.2914	.3807	1.0000				
.2634	.1824	.2365	.2772	.1417	.2693	.2444	.4916	1.0000			
.0031	-.1942	-.0464	-.1236	-.1006	-.3884	-.5431	-.2682	-.1946	1.0000		
-.1290	.0215	-.0895	-.0988	-.1356	-.0329	.0733	-.0345	-.1340	-.2114	1.0000	
.0031	-.2075	-.0761	-.1375	-.1214	-.3638	-.5453	-.3021	-.1946	.9256	-.2114	1.0000
Mean											
12.9385	12.5538	12.9538	13.2308	13.2231	5.0523	3.2059	9.4077	10.1308	.2923	.3692	.2923
Standard Deviation											
1.8667	2.5398	2.2853	2.4417	2.4440	3.1683	2.1605	3.5123	.1386	.4566	.4845	.4566

3. Number of Sisters = 2 (N = 252)

F1ED	F2ED	M1ED	M2ED	M3ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
1.0000											
.3707	1.0000										
.3702	.3637	1.0000									
.3246	.4071	.5417	1.0000								
.2533	.2986	.3159	.4479	1.0000							
.2551	.2640	.3015	.3257	.2704	1.0000						
.3085	.2173	.2209	.2342	.2780	.4699	1.0000					
.3604	.2365	.3423	.3590	.2989	.3292	.4721	1.0000				
.2680	.1355	.2181	.2049	.2597	.1274	.1736	.4563	1.0000			
-.1215	-.0131	-.1521	-.1069	-.1142	-.3272	-.4300	-.2346	-.1035	1.0000		
.1238	.0705	.0877	.0540	.1518	.2125	.0984	.0955	-.0833	-.1114	1.0000	
-.1331	-.0412	-.1208	-.0884	-.1596	-.3578	-.4674	-.2395	-.1028	.8426	-.1352	1.0000
Mean											
12.6190	12.8968	13.0079	12.9008	13.2222	5.3131	3.1575	9.0159	10.0873	.2460	.4683	.2579
Standard Deviation											
1.9284	1.8286	2.8228	2.3196	2.2455	3.0363	2.2663	3.2651	2.8016	.4316	.5000	.4384

4. Number of Sisters = 3 (N = 234)

F1ED	F2ED	F3ED	M1ED	M2ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
1.0000											
.3789	1.0000										
.3909	.5228	1.0000									
.2880	.3064	.3436	1.0000								
.1291	.2705	.3093	.2417	1.0000							
.2137	.2657	.2262	.2616	.1386	1.0000						
.2266	.2931	.2944	.3933	.2637	.4485	1.0000					
.2390	.2769	.2846	.2792	.2933	.3795	.5125	1.0000				
.3063	.3495	.3376	.2643	.2195	.3629	.3507	.5116	1.0000			
.0148	-.0320	.0268	-.1231	-.1010	-.2766	-.3804	-.1465	-.0163	1.0000		
.0534	.0316	.1270	.0931	.1692	.1639	.1648	.0773	-.0084	-.1113	1.0000	
-.0217	-.0477	-.0559	-.1659	-.1244	-.2795	-.4269	-.1552	-.0147	.8183	-.1027	1.0000
Mean											

12.8162 12.9872 12.8718 13.1838 13.5256 5.3906 3.1445 9.4701 10.4487 .2308 .4530 .2265
Standard Deviation
2.0119 2.0010 1.8693 2.6361 2.4550 2.9543 2.2530 3.3573 2.6351 .4222 .4989 .4195

5. Number of Sisters = 4 (N = 120)

F1ED	F2ED	F3ED	F4ED	M1ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
1.0000											
.4751	1.0000										
.3524	.3664	1.0000									
.3004	.4880	.3576	1.0000								
.4392	.4381	.3993	.3765	1.0000							
.3004	.4867	.2561	.4847	.2898	1.0000						
.2945	.4178	.3232	.4514	.3417	.6088	1.0000					
.4234	.3885	.3149	.3888	.3119	.3866	.5436	1.0000				
.4489	.4046	.2328	.4189	.3196	.3587	.3507	.4615	1.0000			
-.0629	-.1740	.0400	-.1517	-.1245	-.3779	-.3833	-.1384	-.1229	1.0000		
-.0447	.0611	.0143	.1213	-.0132	.0077	.0400	.0790	-.0569	.1112	1.0000	
-.0554	-.1847	-.0226	-.1135	-.1812	-.4380	-.4414	-.1923	-.1327	.8814	.0880	1.0000

Mean
12.5083 12.5750 12.6500 12.7167 13.1333 5.8417 3.1419 9.5000 10.0250 .2083 .4917 .2333
Standard Deviation
1.8697 1.9388 1.9301 1.7546 2.3726 3.3506 2.2210 3.4979 2.7879 .4078 .5020 .4247

6. Number of Sisters = 5 (N = 27)

F1ED	F2ED	F3ED	F4ED	F5ED	BMPIN2	BMFOC1	BMFAED	BMMAED	RURAL	CATH	FAFAM
1.0000											
.7045	1.0000										
.4301	.6259	1.0000									
.4205	.4657	.7962	1.0000								
.0649	.1169	.5093	.5142	1.0000							
.1437	.1900	.0410	.1081	-.0885	1.0000						
-.0050	.3442	.2527	.0417	.4146	.0443	1.0000					
.0833	.3229	.2892	.1142	.0046	.3979	.5161	1.0000				
.1278	.2735	.1769	.2258	.2143	.1409	.3672	.4541	1.0000			
.1525	.1135	-.0635	-.0562	-.1180	-.3247	-.1761	-.1526	-.0559	1.0000		
-.0219	-.1581	-.0053	.0175	-.1489	-.0998	-.0518	.0475	-.3301	-.1512	1.0000	
-.2055	-.3167	-.2778	-.1900	-.0200	-.4547	-.3958	-.3949	-.2108	.8071	-.1512	1.0000

Mean
12.3704 12.5556 12.7407 13.0370 12.6296 4.8556 2.5841 9.6667 9.7778 .2593 .5556 .2593
Standard Deviation
1.9245 1.6013 1.6075 1.9312 1.7574 2.3846 1.8038 3.1986 2.2246 .4466 .5064 .4466

Mailing address:
Center for Demography and Ecology
University of Wisconsin
1180 Observatory Drive #4412
Madison, WI 53706-1393
USA