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Black-White Differentials in the Effect of Family Background

on the Educational Attainment of U.S. Men Born from 1907 to 1946:

A Study of Sibling Resemblance

ABSTRACT

Among black and white men born in the first half of this century, family background accounts for half or more of the variance in educational attainment. Parental schooling, father's occupational status, size of sibship, intact family, farm background, and Southern birth account for no more than half of this common family effect. Family background has smaller effects of the attainments of younger than of older brothers. The effects of measured and unmeasured family background characteristics on the educational attainments of black and white men have also declined. Among blacks, there appears to have been a similar decline in the effects of all measured characteristics of the family of orientation while, among whites, the decline is limited to farm background, intact family, and Southern birth. The global, secular decline in inequality of educational attainment cuts across all sources of variation in schooling -- measured and unmeasured common family characteristics and unmeasured individual characteristics.

In this paper, we use structural equation models of sibling resemblance to estimate racial differentials in the effect of family background on educational attainment and to examine changes in the effect across cohorts of white and black men born from 1907 to 1946. Differences in educational attainment between blacks and whites are well-documented. Census data show that the gap in years of schooling has narrowed since 1940 (Hauser 1993a, p. 276). The high school completion rate of young blacks aged 25-29 is only around 75 percent while that of young whites aged 25-29 is higher than 85 percent (Hauser 1993a, p. 284). About half of white college entrants, but only one third of blacks, complete 16 or more years of school (Hauser 1991, p. 8). Despite African Americans' persistently lower levels of completed schooling, there is mounting evidence that, among persons of similar social background, blacks complete more years of schooling than whites (Hauser 1993b; Hauser and Phang 1993). Thus, it is important to measure black-white differentials in schooling when social background is controlled and to consider whether effects of social background differ between whites and blacks.

The family is one of the important agents in the transmission of educational inequality. Usually, we specify social background using a vector of measured social and economic variables, such as, parental schooling, occupational status, or income, along with race, family size and structure, and geographic location. However, these need not and, in fact, do not exhaust the effect of family background, which includes all other family, school, or neighborhood characteristics -- social, economic, psychological, or biological -- which may be shared by siblings. In this study, we use data on the educational attainments of brothers to estimate effects of family background in this broader sense as well as to improve estimates of the effects of measured background characteristics.

Several studies have examined black-white differentials in the effect of social and economic background on educational attainment (Wolfle 1985; Gottfredson 1981;

Kerckhoff and Campbell 1977; Portes and Wilson 1976; Hauser and Featherman 1976; Hout and Morgan 1975; Porter 1974; Hauser 1993b; Hauser and Phang 1993). However, most of these have only addressed the effects of measured variables on educational attainment and have not considered the broader concept of family background. Hauser and Featherman (1976) is one of the few exceptions, but their methods are long-since outdated . By estimating appropriate models of sibling resemblance, it is possible to estimate the total effect of family background on educational attainment and compare the contribution of measured and unmeasured family factors to black-white differentials in schooling.

Specific Research Questions

In this paper, not only do we examine black-white differentials in the effects of family background and of measured background variables, but we also address other issues such as the confounded effects of family structure and race and changes in the effects of background on schooling over time. We pose the following questions: Do effects of family background on educational attainments of blacks differ from those of whites? Do between- and within-family components of education differ between blacks and whites? Do the effects of family background decrease over time? Have the effects of measured background variables, such as parental education and occupation, declined? Are there black-white differentials in the changing effects of measured and unmeasured family background?

Race, Family Background and Educational Attainment

The global effect of family background on educational attainment has been widely discussed and analyzed. Blau and Duncan (1967) initiated research on family background using the concept of family climate, which they initially specified by the educational attainment of the brother of a 1962 OCG respondent. They later suggested that the brother's educational attainment might better be used as a second indicator of a common family factor. However, they did not incorporate a latent family factor into a model that also included measured socioeconomic origins. Not long after Blau and Duncan's initial work, Jencks (1972) and Duncan, Featherman, and Duncan (1972) each made pioneering efforts -- by combining information from several bodies of data -- to construct models that included measured and unmeasured family background variables, but they did not carry out separate analyses for blacks and whites.

There have been relative few fully elaborated efforts to model status attainment of blacks, either alone, or contrasting blacks and whites. For example, a handful of studies have examined differences in the process of educational attainment between blacks and whites, using a social psychological model of attainment (Wolfle 1985; Gottfredson 1981; Kerckhoff and Campbell 1977; Portes and Wilson 1976; and Porter 1974).

Hauser and Featherman (1976) estimated correlations between educational attainments of brothers in the 1973 Occupational Changes in a Generation (OCG) survey in order to measure the total effect of family background, that is, shared background. They found that correlations between the education attainments of men and their oldest brothers ranged from 0.53 to 0.59 across cohorts of men born from 1907 to 1951 and that measured family origins explained more than half of this correlation.[2] With plausible corrections for error in the measurement of schooling, the observed correlation between brothers was estimated to exceed 0.65. Thus, common family factors might account for almost two thirds of the variance in completed schooling among American men. While Hauser and Featherman (1976;

[2] Later, more sophisticated analyses of sibling resemblance, for example, those of Olneck (1976) and of Hauser and Sewell (1986), also found that about 60 percent of the common variance in the educational attainments of brothers is explained by measured social and economic background variables.

also, see Featherman and Hauser 1978: Ch. 3) reported separate regression analyses of measured social background and schooling among blacks and whites, they did not carry out separate analyses of sibling resemblance.

Despite the advantages of sibling-based research designs in analyses of the effects of family background on educational attainment, race has not been a focus of interest in earlier sibling studies, mainly due to the small size of most sibling samples and the small proportion of minority populations in those samples. With the development of national data resources for the general population, it has become possible to use sibling studies to examine black-white differences in schooling and economic success. Unfortunately, Hauser and Featherman's (1976) analysis failed to take advantage of this possibility within the 1973 OCG sample, and the major purpose of this paper is to make up for that omission. Hsueh (1992) uses one of the few sibling data sets with a large black subsample, the National Longitudinal Survey of Youth (NLSY) to study the effect of family structure on schooling. He treats race as one of many exogenous variables and finds that being black has a positive effect on education and a negative effect on ability. However, his study does not examine black-white differentials in the effects of family background on educational attainment.

By estimating sibling resemblance models in several cohorts, we can also investigate change over time in the effect of families on educational attainment. That is since school completion is in principle cumulative and irreversible, intercohort changes in schooling may be observed in cross-sectional age comparisons.[3] This design has been used extensively in European sibling studies. De Graaf and Huinink

[3] This is perhaps a less satisfying assumption than it used to be. We now know that survey respondents tend increasingly to exaggerate their levels of completed schooling as they age and that school now tends to continue to older ages than in the past.

(1992) studied educational attainments and occupational statuses of all siblings from 1653 West German families and found that the effects of family background declined across the cohorts born in the 1930s, 1940s, and 1950s. Within like-sex and cross-sex sibling pairs, they also find that the effect of father's occupational status grew weaker, but that of parental education was stable over cohorts. Dronkers (1993) found that the resemblance of Dutch siblings declined across fraternal birth cohorts born before 1930, from 1930 to 1940 and after 1940. Toka and Dronkers (1993) studied Hungarian sibling pairs from five birth cohorts, before 1932, 1932 to 1944, 1945 to1952, 1953 to 1959, and after 1959, and reported that the Hungarian data failed to repeat the trend of decreasing sibling resemblance that had been found in Germany and Netherlands.

In the U.S., a few studies have examined changes in the effects of measured or unmeasured family background on educational attainment. In their regression analyses, Hauser and Featherman (1976) found that both the variability in schooling attributable to differences in social background and the variability independent of social background appear to have declined. They reported similar increases in educational equality among blacks and whites. Their comparison of correlations between educational attainments of brothers also suggested that the resemblance in schooling attributable to the effect of measured family background was declining over cohorts.

After controlling for heterogeneity due to the unmeasured family factor, and also using the 1973 OCG data, Mare (1992) found that the effect of father's schooling on siblings' schooling was essentially invariant over cohorts.[4] Moreover, the total

[4] Mare's analysis is based on a latent class model of school transitions, in which the functional form is log-linear. For this reason, Mare's findings about increasing or decreasing effects are not comparable to those of Hauser and Featherman or those in our analysis, where educational attainment is treated as a quantitative variate.

effect of family background also varied little, and, the effect of father's schooling was a only modest part of the overall effect of common factors on school continuation. Those findings reinforce the need for sibling resemblance studies. However, Mare, like Hauser and Featherman, did not control for race, so the study does not measure any black-white variation in the total effect of family background. Furthermore, Mare uses only father's and son's schooling as instrumental variables and thus ignored other measured variables as sources of sibling resemblance.

In order to study black-white differentials of effects of family background on educational attainment, we have to take account of family structure as well as the social and economic characteristics of parents. Single-parent families have socioeconomic consequences, and they are confounded with race. Jaynes and Williams (1989) report that, since the 1960s, the trends in marital status, fertility marital stability and child rearing for both blacks and whites have been similar. However, black women have had higher a total fertility rate than white women, due to the higher proportion of births to unmarried mothers. Black's marital fertility rate has been down at the same level as that of whites since 1980. The number of black children living in female-headed families as well as the number of black female-headed families have increased faster than among whites (Jaynes and Williams 1989; Garfinkel and McLanahan 1986).

The effect of family structure on educational attainment has been studied extensively. Duncan and Duncan (1969) reported that the effect of schooling on occupational success were greater among men raised in two-parent families than those raised in single-parent families, and the effect came from their higher educational attainment. According to Featherman and Hauser (1976), elimination of single-parent families would have reduced the difference in educational attainments between black and white men by 27.2 percent in the birth cohort of 1937-1946 and by

39.4 percent in the birth cohort of 1947-1951. Hsueh (1992) analyzed data from the NLSY and showed that, controlling family structure, blacks had more education than whites but lower ability. Astone and McLanahan (1991) examined an array of educational outcomes, including high school dropout, in the High School and Beyond panel. They found that large effects of single-parent and step-parent families were not explained by differences in parental involvement and behavior. Hauser and Phang (1993) found large effects of non-intact family on high school dropout in the 1970s and 1980s, even when a large set of measured household and family characteristics were controlled; the effects of non-intact family were significantly larger among whites than among blacks.

Size of sibship and birth order are two other aspects of family structure that often appear in analyses of educational attainment. The negative effect of larger sibships on educational attainment is very well established, both among whites and blacks (Hauser and Featherman 1976; Blake 1989), and it enters our analysis both as an exogenous variable and as factor that influences the research design through the number of brothers that appear in connection with each OCG respondent. There is a large and current body of work on birth order, which focuses almost exclusively on differences in levels of ability or of completed schooling (Zajonc and Markus 1975; Zajonc 1976; Lindert 1977). Recent evidence shows little, if any, effect of birth order on educational attainment (Olneck and Bills 1979; Hauser and Sewell 1985; Steelman 1985; Steelman 1986; Retherford and Sewell 1991), and we have not included birth order as a measured background factor. However, because birth order affects the measurement of educational attainment in the OCG data, it does potentially appear as a factor moderating the influence of social and economic background. That is, to a limited degree, we can examine whether effects of family background are greater or less among older or younger brothers. Kuo and Hauser (1993) analyze a rare set of full sibship data from the Wisconsin Longitudinal Study and find no effects of relative birth order within gender on the educational attainment of brothers and sisters, controlling sibship size and gender composition. However, Hauser and Wong (1989) found that the effects of family background were larger among older than younger brothers in Olneck's Kalamazoo brothers sample and in a Dutch sibling sample.

Data and Methods

The data used in this study are from the 1973 Occupational Changes in a Generation Survey (*OCG*).[5] We use the multiple indicator multiple cause (MIMIC) model as an analytic framework for our analysis (Hauser and Goldberger 1971; Jöreskog and Goldberger 1975; Hauser and Wong 1989, p.156). The MIMIC model specifies that exogenous social background characteristics affect educational attainments of siblings through a single, unmeasured common family factor, which accounts for the resemblance in schooling among them. Although this model has only one type of outcome, i.e., educational attainment, it can still provide insights about the total effect of family background on educational attainment, the relationship of each observed variable with the measured family factor, and the relative importance of shared and non-shared sources of variation in attainment.

The OCG Data

OCG 1973, a supplement to the March Current Population Survey (CPS) in 1973, covered a national sample of men aged 20 to 65. The OCG supplement ascertained both the educational attainment of a man's oldest brother who lived to age 25 and that of the youngest brother who lived to age 25. The questions used to assess educational attainment of the brother(s) are:

[5] For details of the study design and a summary of findings, see Featherman and Hauser (1978).

5c. How many brothers did you have?

5d. How many of these brothers were older than you (born earlier)?

6a. Did any of your brothers live to age 25?

6b. Please indicate the highest grade of school completed by the OLDEST of your brothers who lived to age 25.

6c. Please indicate the highest grade of school completed by the YOUNGEST of your brothers who lived to age 25.

The data also include information on educational attainments of respondents, social and economic background, such as parental education and occupation, farm background, family size and family structure. Unfortunately, OCG 1973 does not include measured ability for the respondent and his brothers, and it covers only male respondents and their brothers. However, this is the only set of data for siblings in the general population that permits the study of black-white differentials across cohorts, controlling for family background.

Because the age distribution of the respondents is from 20 to 64, we only know the relative birth order of brothers through the information about the number of older brothers and the age of respondents. After selecting out the respondents younger than 25 and older than 64, there remain 9,184 respondents with two brothers reported and 5,527 respondents with one brother reported. In the current study, we divide the data into two subsamples: brother pairs and brother triads.[6] Within brother triads we can unambiguously address the issue of relative birth order by comparing outcomes between "oldest" and "youngest" brothers, but even relative birth order remains ambiguous within many of the brother pairs. Also, respondents

[6] Hereafter, we use "size of sibship" to refer to the number of brothers for whom we have observed educational attainment, including the respondent. That is, we do not use it to refer to the actual number of brothers or sisters in the respondent's family of orientation.

in brother triads typically, though not necessarily, come from larger families than respondents with only one brother.

It is not straightforward to specify a design for analysis of the OCG brother data. Both birth order and family structure, each of which is of substantive interest, also affect the array of exogenous and endogenous variables. A key problem in identifying relative birth order is that both an "oldest" and "youngest" brother might be older than the respondent. As shown in Table 1, owing to the format of the questions about siblings, among valid cases,[7] we could only determine relative birth order among 84.4 percent of white brother pairs and 79.3 percent of black pairs, that is, pairs in which only an "oldest" or only a "youngest" brother occurred. Relative birth order was identifiable among all three members of 56.2 percent of white brother triads and 47.2 percent of black triads. While the elimination of pairs or triads for which relative birth order could not be determined might not affect analyses of white brothers, because the sample size is quite large, the elimination of the indeterminate cases in the black sample is more serious because the sample is much smaller. More important, the proportion of indeterminate birth orders is very different among pairs and triads. Thus, we are concerned that an effort to focus the analysis on cases for which relative birth order is known could reduce statistical power and lead to bias in our findings.

Having rejected an effort to array the OCG brother data in terms of relative birth order, we identify educational outcomes in relation to the respondent, that is, following the references to "oldest" and "youngest" brother in the questions. The disadvantage of this arrangement of the data is that the effect of birth order can not be studied, except in certain family configurations. Where relative birth order is

[7] Here, valid cases include respondents who have siblings living up to 25, report the education of at least one sibling, and report all family background variables.

unambiguous, namely, in the distinction between "oldest" and "youngest" brothers in triads, we have attempted to determine whether the birth order moderates the effect of family background on education. However, it should be kept in mind that these cases are selected in a complex way, namely, that they represent families with at least three male offspring.

The questions about education and occupation of the household head complicate the analyses of effects of family structure on education, and, possibly, affect the definition of socioeconomic background. The questions asked are as follows:

Q8a. Where you living with both your parents most of the time up to age 16?

Q8b. if "No" above, who was the head of your family?

Q11. Now we would like to find out what kind of work your father did when you were about 16 years old. If you were not living with your father, please answer for the person marked in question 8b.

Q11a. In what kind of business or industry did he work? ...

Q12. What is the highest grade of school your father (or the person marked in 8b) completed? If you are not sure, please make a guess.

Q13. What is the highest grade of school your mother completed? If you are not sure, please make a guess.

In short, the report of "head's" education and occupation might refer to a father, mother, or someone else. Thus, for example, if the household head is the mother, we would have two mothers' education, one reported as the education of household head (Q12) and the other reported as the education of mother (Q13). In Table 2, for non-intact family groups, mothers are the largest share of household heads, fathers are a smaller share of household heads, and the remainder are other relatives. "Other male relative," may include grandfathers, step-fathers, or yet other relatives. We have found that most respondents report the same household head's education and mother's education, regardless of family structure and the identity of the household head; however, some respondents from mother-headed, non-intact families report the household head's education and mother's education differently. It is difficult to figure out whose education and occupation were reported by respondents from non-intact families.[8] Thus, we decide to use the term of "household head" throughout our analysis, rather than"father;" and the reports of household head's education and occupation consist of mother's, father's, male relatives' and female relatives' education and occupation. This is comparable to use of the OCG data in previous studies. However, the use of data for different individuals could influence the effect of socioeconomic background variables on education of offspring in intact and non-intact families.

If we were to analyze the data separately for each combination of family structure, number of brothers, race, and cohort, the design would become excessively complicated and unwieldy, and the number of cases in each of the cells of the design could become quite small. As shown in table 2, the sample of non-intact families is small, and we have sought to simplify the design by collapsing across family types. Thus, before turning to our analysis of intercohort change among blacks and whites, we first examined the effect of family structure on the process of educational attainment. That is, we compared the effect of father's education and occupation in intact families with those of household head's education and occupation in non-intact families. In this analysis, we used all of the cases of sibling pairs (from respondents with one or with two or more brothers) in which relative birth order could be

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[8] We also suspect some respondents probably report mother's education by borrowing father's education, since in the question we encourage them "make a guess".

identified. Thus, we were able to estimate the effect of relative birth order, though we did not pursue it later in the analysis.

Basic Sibling Resemblance Model

Figure 1 displays the path diagram of the MIMIC model used in earlier sibling resemblance models (Hauser and Wong 1989). The general model is

$$\eta = \Gamma \xi + B \eta + \zeta \tag{1}$$

where η is the endogenous latent variable; ξ is a vector of exogenous latent variables with variance-covariance matrix Φ; ζ is a vector of disturbances, independent of η and ξ, with variance covariance matrix Ψ; Γ is a parameter matrix of effects of ξ on η; and B is a parameter matrix of effects of η on η. In the model of Figure 1, there is only one η, which carries the effect of family background variables to the educational attainment of offspring; thus, B is non-existent. However, in a later generalization of the model, we use B to express effects of family background that differ across cohorts. In general, the measurement models for the latent variables are

$$X = \Lambda^x \xi + \delta \tag{2}$$

$$Y = \Lambda^y \eta + \varepsilon \tag{3}$$

where X and Y are vectors of observable variables, i.e., indicators of ξ and η, respectively; Λ^x and Λ^y are parameter matrices giving the effects of ξ and η on X and Y, respectively; and Θ^δ and Θ^ϵ are the variances of δ and $\in$, respectively. In our specification, equation 2 is not shown in the diagram because we assume that each exogenous variable only has one indicator, all elements in Λ^x are equal to one and Θ^δ is equal to zero; that is, all ξs are perfectly measured by corresponding *X*s. Equations 3 and 1 in matrix form for sibling pairs are

$$\begin{bmatrix} y_{k1} \\ y_{k2} \end{bmatrix} = \begin{bmatrix} \lambda_{k11} \\ \lambda_{k21} \end{bmatrix} \begin{bmatrix} \eta_{k1} \end{bmatrix} + \begin{bmatrix} \varepsilon_{k1} \\ \varepsilon_{k2} \end{bmatrix} \tag{4}$$

$$\begin{bmatrix} \eta_{k1} \end{bmatrix} = \begin{bmatrix} \gamma_{k11} & \gamma_{k12} & \gamma_{k13} & \gamma_{k14} & \gamma_{k15} & \gamma_{k16} \end{bmatrix} \begin{bmatrix} x_{k1} \\ x_{k2} \\ x_{k3} \\ x_{k4} \\ x_{k5} \\ x_{k6} \end{bmatrix} + \begin{bmatrix} \zeta_{k1} \end{bmatrix} \tag{5}$$

where x_{k1}, ..., x_{k6} are the exogenous variables, y_{k1} and y_{k2} are educational attainments of older and younger brothers, and the index *k* indicates groups, e.g., family type, race, or cohorts. In equation 4, $\lambda_{k11} = 1$ is a normalizing constraint, so λ_{k21} is a proportional rescaling of the γ_{k1j} to effects in the metric of y_2.

The effects of family background variables on the education attainment of each brother in the constrained reduced form are the products of the corresponding γs and λ^ys. By testing hypotheses of equality between λ^ys within each cohort, family structure or sibship, we are comparing the regressions of sibling's education on the family background variables. By testing hypotheses on the equality of λ_k or γ_k between (or among) groups, we are studying the interaction between the effects of family background on education and criteria for specifying groups. As stated before, the advantage of the sibling resemblance model is that we can study the variation of unmeasured family component of educational attainment. In Figure 1, ζ stands for the unmeasured but common family components or factors in the education of brothers, i.e., the co-variation of education between brothers that cannot be explained by the measured variables of family background. It can also be described as the unmeasured between-family (shared) variation. The εs stand for the variation of a brother's education which cannot be explained by the common family factor (η); it is also known as the within-family (non-shared) variation.

Family Structure and Birth Order

We divide all sibling pairs in which relative birth order was identified into groups defined by family structure within each race, that is, white intact-family, white non-intact family, black intact-family and black non-intact-family. The endogenous variables are educational attainments of siblings, arrayed by their relative birth order; the exogenous variables are household head's education, household head's occupation, mother's education, farm background, number of siblings and southern origin.

In the baseline model, we constrain all parameters except the variance-covariance matrix of the exogenous variables to be equal between groups (family types) for whites and for blacks.[9] That is, the process of educational attainment, including the effects of family background on educational attainment of siblings, within-family variances, and between-family variances, do not vary by family structure within race. As Table 3 shows, the model with cross-group constraints fits fairly well. For blacks, the χ^2 statistics is 33.64 with 20 degrees of freedom (*df*); for whites, the χ^2 statistics is 112.49 with 10 *df*. The *bic* statistics are -149.56 for blacks, and -117.11 for whites.[10] Relative to the saturated model, in which the *bic* statistic is

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[9] Also, none of the models estimated here or later in this paper constrain the mean structure of the variables. That is, mean levels of social background and of educational attainment vary freely across groups.

[10] 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: $bic = \chi^2 - df \times \sum ln[N_k \times (p_k + q_k)]$, where *df* are the degrees of freedom under the tested model or contrast, N_k is the sample size in the kth group, p_k is the number of observed exogenous variables in the kth group, and q_k is the number of observed endogenous variables in the kth group. Satisfactory fit is indicated by a negative value of *bic*, and models with lower *bic* statistics are preferred (Raftery 1986, 1993). We use *bic* as a secondary criterion of model fit when χ^2 is nominally statistically significant in order to avoid rejecting null

zero, the baseline model is acceptable. Next, we test the overall hypothesis that the process of educational attainment in the intact family is different from that in the non-intact family. That is, we lift all across-group constraints on the parameters of Γ, Λ^y, Θ^ϵ, and Ψ. The χ^2 statistics of this family structure hypothesis is 28.38 with 10 degree of freedom for blacks, and 57.52 with the same degrees of freedom for whites; the *bic* statistics are -63.22 and -57.28, correspondingly. That is, while these models are acceptable (using *bic* as the criterion), they are less attractive than the constrained models because *bic* is less negative. Thus, we do not choose to reject the null hypothesis that family structure has no effect on the process of educational attainment.[11] Consistent with these test statistics, the coefficients in the baseline model are very similar to those of intact families and of non-intact families in the second model. These findings lead us to combine the two family types in our analyses of inter-cohort change.

In this preliminary analysis, we are interested not only in examining the effect of family structure, but also in the effect of relative birth order. As stated in the previous section, the unidentifiability of birth order in almost one third of the sample unevenly reduces the sample size between racial groups and between sibsize groups. Given that the observations of brothers' schooling will not be arranged by relative birth order in the analysis of inter-cohort change, we have used this subsample to test the influence of relative birth order on family effects on schooling. The null hypothesis is that the effects of family background variables on older siblings do not differ from those on younger siblings. Literally, we impose the equality constraint,

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in favor of trivial alternative hypotheses.

[11] Again, recall that our tests pertain to the effects of measured and unmeasured family background characteristics, not to mean differences in schooling between the children of intact and non-intact families.

$\lambda_{k11} = \lambda_{k21} = 1$. Under the constrained model, the χ^2 statistics are 83.06 with 21 *df* for blacks, and 405.19 with 21 *df* for whites; the *bic* statistics are -108.98 and 164.11, correspondingly. When we contrast the constrained and unconstrained models, the χ^2 statistics are 49.42 with 1 *df* for blacks, and 292.70 with 1 *df* for whites, for which the *bic* statistics are 40.58 and 281.22. Thus, we reject the hypothesis that family background affects older and younger brothers equally. For blacks, the effect on the younger brother is only 69 percent of that on the older brother; for whites, the effect on the younger brother is 77 percent of that on the older brother.

Family Background and Inter-Cohort Changes

In the remainder of the analysis, we arrange the samples of blacks and whites into eight groups by birth cohorts of respondents and number of brothers for whom educational attainment has been reported: sibship size 3 age 25-34, size 3 age 35-44, size 3 age 45-54, size 3 age 55-64, size 2 age 25-34, size 2 age 35-44, size 2 age 45-54 and size 2 age 55-64. Table 4 shows descriptive statistics of the independent and dependent variables. Figure 2 displays the data arrangement and the schematic framework of the model. The order of endogenous variables is respondent (y_1), the oldest brother (y_2), and the youngest brother (y_3). Note that the "oldest" might be either the the oldest brother in the sibship or the brother born next after the respondent; the "youngest" might be either the youngest brother in the sibship or the brother born just before the respondent. Thus, the relationship between the oldest and youngest brothers is in the right relative birth order; but the relationship between the respondent and either the oldest or the youngest brother is unclear.

Some Notes on Methods

We used methods for the estimation of structural equation models with missing data that were introduced in Allison (1987) and Allison and Hauser (1991) to combine data for brother pairs and brother triads. By combining equality constraints across subsamples with the introduction of null parameters for missing variables, multiple group models can be estimated simultaneously for groups defined by the presence or absence of certain observations.[12] That is, we introduce pseudo-values for the moments of missing variables and specify innocuous fixed values for their factor loadings and variances.

To elaborate, equation 4 for brother pairs is re-parameterized as

$$\begin{bmatrix} y_{k1} \\ y_{k2} \\ y^{*}_{k3} \end{bmatrix} = \begin{bmatrix} \lambda_{k11} \\ \lambda_{k21} \\ \lambda^{*}_{k31} \end{bmatrix} \begin{bmatrix} \eta_{k2} \end{bmatrix} + \begin{bmatrix} \epsilon_{k1} \\ \epsilon_{k2} \\ \epsilon^{*}_{k3} \end{bmatrix} \tag{6}$$

where k indicates the groups (sibship size and cohort) and for k = 5, ..., 8 (the groups of brother pairs), $\lambda^{y^*}_{k31} = 0$ and $\theta^{\epsilon^*}_{k31} = 1$. The values of $\theta^{\epsilon^*}_{k3}$ and λ^{y}_{k3} are arbitrary because the endogenous variable (y_{k3}) has not been observed and the fixed values do not affect model fit or other parameter estimates. We also add a second latent variable (η) to the model of Figure 1 in order to specify inter-cohort change in the global effect of family background on education. The structural model becomes

$$\eta_{k1} = \gamma_{k11}\,\xi_{k1} + \dots + \gamma_{k17}\,\xi_{k7} \tag{7}$$

$$\eta_{k2} = \beta_{k21}\,\eta_{k1} + \zeta_{k22} \tag{8}$$

where, unlike equation 5, there is no disturbance for η_{k1} in equation 7. In equation 8, when Γ and Λ^{y} are invariant across cohorts, the specification that $\beta_{121} = 1$ and $\beta_{k21} \neq 1$ for k = 2, ... ,8 permits global intercohort changes in the effect of family background; otherwise, $\beta_{121} = \beta_{k21} = 1$ for all k. It is actually a matter of indifference whether we

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[12] This treatment of missing moments is required within LISREL; other computer programs, for example, EQS, do not require this kind of model specification to estimate a model with missing variables.

permit the disturbance to appear in equation 7 or in equation 8, but only one such common variance term is identified.

Because the sample is selected from the general population, the age of respondent's siblings ranges widely. That is, the cohorts are based on OCG respondents, but their brothers may belong to other cohorts. The age spread could produce larger variations in education among brothers in triads than in pairs. Table 4 also shows that the social and economic characteristics of pairs differ from those of triads, both among blacks and whites. In general, brothers appearing in pairs had more favorable social background and completed more schooling than brothers appearing in triads. That is, brothers in pairs had parents with more schooling, household heads with higher status occupations, were less likely to be of farm origin, and -- of course -- came from smaller sibships. The one interesting exception to this pattern is that there was almost no difference between the pairs and triads in the chance of having been raised in a non-intact family (1 = intact, 2 = non-intact), presumably because the effect of family dissolution on births offsets that of social deprivation on family dissolution.

To comprehend the variations among brothers, and to ensure the consistency of analyses of inter-cohort change between the pairs and triads, we first examine hypotheses related to the number of brothers that were observed. That is, in Figure 2, we compare parameters in the groups shown at the left-hand side of the page (pairs) with those in the groups shown at the right-hand side of the page (triads), assuming for the moment that there is no variation in parameters across cohorts. Our purpose is to select a model with satisfactory specifications within cohorts as a baseline model for the analysis of inter-cohort changes. Table 5 and Table 6 show the goodness of fit and model selection for whites and blacks. Hypotheses about restrictions within cohorts are reported in the first panel. Hypotheses about inter-cohort change in parameters are reported in the second panel. The selection of our preferred final model is reported in the third panel. Table 7 and Table 8 report estimates of parameters of the preferred model. Finally, we have decomposed the components of variance for each brother into measured between-family variance, unmeasured between-family variance and within-family variance, and these estimates are reported in Table 9 and Table 10.

Within-Cohort Restrictions

The baseline model permits no intercohort variation in parameters, but places only two sets of equality constraints on parameters of pairs and triads within the same cohort. Specifically, within pairs and within triads, all parameters are the same in each cohort, but the effects of exogenous variables on the common family factor (Γ) and the global effect of the common family factor on education ($\beta_{k21} = 1$) are constrained to be equal regardless of the number of brothers and of cohort. Further, the loadings (λ^y_{k11}) of the respondents' education on the common family factor in triads (but not in pairs) are normalized to unity. Thus, in this specification, a significant departure of λ^y_{k12} from unity among the brother pairs would suggest that there is an overall difference in the effect of measured social background on the educational attainment of respondents who appeared in pairs versus those who appeared in triads. The likelihood-ratio test statistics are $\chi^2 = 1163.65$ for whites and 430.68 for blacks, each with 158 *df*; the *bic* statistics are -763.95 and -1130.36 correspondingly. Thus, while we would nominally reject this baseline model, the *bic* statistics suggest that the fit is not bad, given the size of the samples. The question remains whether significantly better fit can be achieved without compromising the objective of parsimonius intercohort comparison.

We next test hypotheses about the relationship between the family factor and respondents' education. While the baseline model states that number of brothers observed affects the dependence of school on the family factor, assuming that the effects of specific variables on the common family factor do not vary, the next hypothesis, Line 1 in Tables 5 and 6, says that the effects of family background on the schooling of respondents do not vary within cohorts. This model fits negligibly worse than the baseline model; with an increase of one *df*, the contrast in χ^2 is only equal to 1.53 for whites and 0.02 for blacks.

Next, conditioning on model 1, we further add constraints of equality in the with-family variances of respondents' schooling, $\theta^{\varepsilon}_{p11} = \theta^{\varepsilon}_{t11}$, where *p* and *t* pertain to pairs and triads. The second model suggests that, with the equal effects of the family factor on the education of respondents between brother pairs and triads, the within-family variances are also invariant between brother pairs and triads. Both models 1 and 2 are aimed at testing the homogeneity of respondents' education between brother pairs and triads. Model 2 does not fit better for whites than model 1: χ^2 = 1220.81 with 160 *df* and the *bic* statistic is -731.19. However, model 2 does fit better for blacks: χ^2 = 431.27 with 160 *df* and the *bic* statistic is -1149.53. That is, black respondents in the pairs and triads are more similar to each other than are white respondents.

Model 3 specifies that the unmeasured between-family variances do not vary between brother pairs and triads. The constraint fits the data well; with an increase of one *df*, χ^2 only increases by 3.01 for whites and by 1.25 for blacks. The hypothesis cannot be rejected. Up to now, we can claim that the data of respondents are fairly homogeneous between brother pairs and triads with the exception of the within-family variances of white respondents. Next, models 4 and 5 test hypotheses about the education of oldest brothers. Conditioning on model 3, model 4 constrains the loadings of educational attainment of oldest brothers to be the same in pairs and triads, that is, $\lambda^{y}_{p22} = \lambda^{y}_{t22}$. Similarly, model 5 conditions on model 3, but adds the constraint that within-family variances are equal in pairs and triads, that is, $\theta^{\varepsilon}_{p22} = \theta^{\varepsilon}_{t22}$. Briefly, we reject model 4 but accept model 5. That is, the loadings of the family factor for oldest brothers are signficantly different between brother pairs and triads for both blacks and whites, while the within-family variances are very similar. It is understandable that the effects on education of the oldest brother differ between pairs and triads. The oldest in the triads is more likely to be older than the respondent than the oldest in a pair, given that the former come from larger families, and the spacing between the respondents and the oldest is probably larger in triads than in pairs.

Thus, the final model in this stage of the analysis is model 5. We conclude that, in general, the data between brother pairs and triads are homogeneous but the within-family variances of white respondents are significantly different between brother pairs and triads, and the effects on the oldest brothers are significantly different among both blacks and whites. In the next stage of the analysis, model 5 is the baseline model that we use to analyze inter-cohort change.

Inter-Cohort Changes

We first test hypotheses about the loadings of the education of each brother on the common family factor across cohorts, and then test hypotheses about the within-family variances of each brother, about the between-family variances, about the global effect of family background on education, and, finally, about the effects of specific family background variables. Throughout these tests, we maintain the specifications of model 5 within cohorts. For example, when the inter-cohort constraints on the unmeasured between-family variances (ψ_{22}) are released, the equality constraint on the variances between the brother pairs and triads of the same cohort are maintained because, in the previous stage of the analysis, we accepted the model with that equality constraint. Unlike the previous stage of the analysis, we did not nest the tests of these hypotheses. In each test, we relax one of the restrictions on intcohort change that was imposed by model 5. We want to select a better model than model 5 as the new baseline to answer questions related to intercohort change, and these tests help us choose hypotheses of substantive interest.

Model A through C in the second panel of Tables 5 and 6 test the effects of intercohort change in effects of the common family factor on the education of each brother (Λ^y). Models without those inter-cohort equality constraints fit negligibly better than model 5: with three *df* less, χ^2 only decreases by 7.14, 11.60, and 13.12 for blacks and by 20.53, 26.74 and 20.56 for whites. Thus, *bic* increases to -1146.44, -1149.90 and -1151.42 among blacks and to -777.04, -783.25 and -777.07 among whites. That is, although the relative effects of family background on education of each sibling relative to those among respondents differ between brother pairs and triads, they do not change across cohorts.

Model D through G test hypotheses about change in within-family variances across cohorts for respondents and their brothers. Models without inter-cohort equality constraints on the within-family variances fit well: the likelihood-ratio test statistics with 159 *df* for each model are 354.12, 389.87 and 392.34 among blacks; with 158 *df*, the test statistics are 1112.83, 1110.84 1138.50 and 1108.40 for whites. The *bic* statistics are -1216.80, -1181.05 and -1178.58 for blacks and -814.77, -816.76, -789.10 and -819.20 for whites.[13] The fit of model F for whites is marginal (compared to model 5), but since we rejected the hypothesis of constant within-family variances in all other cases, we decided to do so here as well.

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[13] We estimated one more model for whites than for blacks because, in the previous section we found that the within-family variances for white respondents are significantly different between brother pairs and triads, while those for black respondents are not significantly different between those groups.

Substantively, models H, I and J are more interesting. They test hypotheses directly related to "family background" or the "common family factor". The hypothesis in model H is that the variance in the unmeasured common family factor ($\psi_{p22} = \psi_{t22}$) changes across cohorts. Model H improves fit: the contrast of χ^2 to model 5 is 41.29 for blacks and 202.29 for whites with 3 *df* less. The *bic* statistics are -1179.59 and -958.80. In model I, we test the hypothesis that the global effect of the common factor on educational attainment changes across cohorts, with unmeasured between-family and within-family variances unchanged. That is, we permit $\beta_{p21} = \beta_{t21}$ to vary across cohorts. Model I fits better than model 5 among blacks, when bic is used as a criterion: χ^2 = 389.11 with 158 *df* and *bic* is -1181.81. However, model I does not improve the fit for whites relative to model 5: χ^2 = 968.80 with 159 *df* and *bic* is -793.36. That is, there is some evidence of a global change in the effect of measured family background among blacks, but not among whites. Finally, model J says that the specific effect of each family background variable changes over cohorts. Model J does not improve fit substantially among blacks or among whites: χ^2 = 374.10 for blacks with 141 *df*, and χ^2 = 955.23 for whites with 140 *df* and the *bic* statistics are -1018.98 and -752.37. Thus, while a single parameter for change in the effect of family background variables has some merit (among blacks, at least), changes are not large enough to appear significant when all of the background coefficients are permitted to vary freely.

We report the final stage of model selection in the third panel of Tables 5 and 6. The new baseline is model 5 with constraints relaxed as in models D, E, F, G, and H; that is, we maintain inter-cohort equality constraints on the loadings of the education of each brother on the common family factor (Λ^y) and on the effects of measured family background (Γ and B), and we release the inter-cohort constraints on all variances (Ψ and Θ^ε). The new baseline model says that the relationships between the effects of the common family factor on education of respondents and their brothers and the effects of measured social background on schooling are constant across cohorts, while the unmeasured variance components, within- or between-families, vary across cohorts. Model A1 fits very well in comparison with model 5: χ^2 = 280.44 with 150 *df* for blacks and χ^2 = 829.85 with 146 *df* for whites. The *bic* statistics are -1201.56 and -951.35. Conditional on this new baseline model, we test hypotheses I and J again. As before, hypothesis I cannot be rejected for blacks only, and hypothesis J cannot be accepted either for blacks or whites. That is, the global effects of family background change over cohorts for blacks only. However, for whites, the fit of hypothesis J relative to model A1 is better than relative to model 5. It is possible that the effects of some, but not all exogenous variables change over time among whites.

In the left panel of table 7, we report the parameter estimates of model A3 for white brothers. While effects of household head's and mother's education do not change over cohorts, the effects of household head's occupation in the first two cohorts are smaller than in the last two cohorts. The negative effects of total number of siblings slightly increase over cohorts. However, the negative effects of farm background, non-intact family and southern origin decline regularly across cohorts. In model A4, we test the hypothesis of intercohort variation in the effects of just these three family background variables, i.e., farm background, non-intact family and Southern origin. When contrasted with model A1, model A4 does not improve fit significantly for black brothers, but it does improve significantly for white brothers. That is, the effects of family background do not change globally over cohorts for whites, but there is significant intercohort change in the effects of farm background, non-intact family, and Southern origin.

Model A4 is the preferred final model for whites; model A2 is the preferred final model for blacks. For both blacks and whites, the variances of unmeasured factors, within families and between families, change over cohorts. Given that the effects of the common family factor on education of each brother vary within each sibship, for whites, while the effects of socio-economic background, such as household head's occupation and education and mother's education, do not vary over cohorts, the effects of farm background, southern origin and family structure change over cohorts. However, among blacks, the effect of each measured family background variable on the education of brothers changes consistently across cohorts.

Estimates of Parameters

The parameter estimates of our preferred final model are reported in the right panels of Table 7 and Table 8. They show that, for white brothers, the negative effects of farm, non-intact family and Southern origin decline over cohorts. Growing up in a farm has almost no effect at all for the most recent cohort of white brothers. The disadvantage of growing up in a farm, living in non-intact family or growing up in the South is greater for white men born at the first two cohorts than those born later. For blacks, with the declining importance of measured family background variables, the contrast appears primarily to lie between the most recent cohort and the other three cohorts.

We cannot compare the coefficients of socio-economic background of blacks directly with those of whites, because the global effects on white brothers are constrained to equal one. However, by inspecting the coefficients reported for model A3, we do not see too much difference between them. Of other family background variables, first, the disadvantage of growing up in a farm or in the South is larger for black men than for white men. The differences between black and white men are increasing over time even though the global effects of family background on education for black men are declining. Second, living in a non-intact family seems more disadvantageous for white men than for black men in every cohort; thus, although non-intact families are an obstacle to school completion, the findings would appear to offer partial support for claims of resilience among black, non-intact families (also, see Hauser and Phang 1993). The negative effect of number of siblings is significantly larger among white men than among black men.

Comparing our findings to Hauser and Featherman (1976), the coefficients of model A3 are slightly different from coefficients of regression analyses in Hauser and Featherman (1976, Table 11). Generally, our effects on education of respondents are lower than theirs, except for those of household head's occupation and farm background. A distinct difference between our analyses and theirs occurs in the effect of Southern origin for black men. In Hauser and Featherman (1976), the effects of Southern origin decline rapidly for black men, from -1.783 to -.163; in this study, the effects of Southern origin on education for black men only decline from -.878 to -.443 in model A3 and from -.585 to -.362 in the preferred final model.

Our samples, as well as our models, are different from those estimated by Hauser and Featherman (1976). The model includes mother's education, which is not used by Hauser and Featherman. We also exclude cases in which mother's education is missing, so our sample is smaller. Moreover, respondents who are only children are not in our sample, but are in their sample. All the same, we think that the differing effects of Southern origin deserve further attention.

Variances

Declining family effects can also be seen in the unmeasured between-family variances and within-family variances. Table 9 and Table 10 report the estimated variances. We decompose the total variances for each brother into three components: measured between-family variance, unmeasured between-family variance and within-family variance. The measured between-family variance is the variance explained by the exogenous variables. We are particularly interested in two comparisons: comparisons of variances and percentages across cohorts for respondents, and comparisons of variances between the oldest and youngest brothers in triads.

All variances decline regularly across time, for both respondents and their brothers. That is, every component of educational inequality, within or between families, has declined among black and white men born in the first half of this century. Among white respondents, the share of common family variance due to measured causes does not vary much, but that of the unmeasured between-family variance declines over time while that of within-family variance increases over time. The same pattern can also be seen among white brothers. For black respondents, the relative importance of common measured variance also declines, but the shares both of unmeasured between- and within-family variance increase. Those trends in variances and percentages are more or less determined by the change in the global effect of the family factor, but that by no means minimizes the import of the trend.

Conclusion

Except for some small points, we generally confirm findings from earlier studies. Among white and black men, the effects of measured social background on educational attainments of brothers can be specified as working through a single, common family factor. However, those effects differ by relative birth order; they are larger for older than for younger brothers. Family type -- intact vs. non-intact -- is not associated with substantial differences in the effects of measured background variables on men's educational attainments.

The schooling of brothers is affected strongly by measured social background characterstics: paternal and maternal schooling, occupational status, numbers of siblings, intact family, farm origin, and Southern birth.[14] These effects are not markedly different among blacks and whites, but non-intact families and larger sibships appear to handicap whites more than blacks. Effects of social background on schooling have declined among blacks and whites, but the pattern of change appears different in the two groups. Among blacks, the data are consistent with a sharp, global decline in the effect of all background characteristics between cohorts born from 1935 to 1946 and all earlier cohorts, while, among whites, there appears to have been a gradual decline in the effects of just three specific background characteristics, farm background, southern origin and family structure.

As anticipated by Hauser and Featherman, we find that a modest vector of measured social background characteristics accounts for about one quarter of the variance in men's schooling, but for about one half the common family variance. Roughly one half of the variance in men's schooling occurs within families, that is, it is not explained either by measured or unmeasured common family characteristics. The observations hold equally for blacks, except the within-family variance appears to be somewhat smaller as a share of the total.

This decomposition of the components of variation in educational attainment has been rather stable across cohorts, even as the total variance in schooling has declined rapidly. Thus, the equalization of schooling among American men born in the first half of the century appears to have been effected by reduced inequality from every source. This need not have been the case. One might imagine that, for

[14] We note that, even though we are working with a subsample of the cases analyzed by Hauser and Featherman (1976), our estimated standard errors are much smaller. In part, this is due to our failure to use a correction for the CPS sample design effect, which was applied by Hauser and Featherman. More important, by constraining estimates of the effects of social background variables across multiple observations of schooling, that is, those of respondents and their brothers, we have substantially reduced the sampling variability of the elements of Γ, relative to the regression coefficients estimated by Hauser and Featherman.

example, that the effects of measured background characteristics could have remained constant, thus increasing the share of educational inequality created by family inequality. Likewise, one could imagine that family resources and practices, other than the purely socioeconomic, need not have become more equal. Thus, as orderly as the data appear to be, they point to a remarkable pattern of reduced inequality in the distribution of schooling. We hope that it may soon be possible to extend this statistical history to more recent cohorts as well as to women.

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Table 1. Identifiability of Birth Order for OCG Brothers

	Brother(s) Reported			
	Oldest	Both	Youngest	Total
A. OCG White Respondents and Their Brothers				
Non-Identifiable	656	3396	128	4180
	14.92	43.84	20.68	32.75
Identifiable	3741	4350	491	8582
	85.08	56.16	79.32	67.25
Total	4397	7746	619	12762
	100.00	100.00	100.00	100.00
B. OCG Black Respondents and Their Brothers				
Non-Identifiable	89	759	17	865
	19.39	52.78	32.69	44.38
Identifiable	370	679	35	1084
	80.61	47.22	67.32	55.62
Total	459	1438	52	1949
	100.00	100.00	100.00	100.00

Table 2. Exogenous Variables, Family Structure and Household Head

Variables		Intact Family	Non-Intact Family				
			Father	Mother	Male Relatives	Female Relatives	Missing
A. OCG White Respondents							
Sample Size		18836	592	885	526	101	49
Head's education	Mean	7.92	7.03	8.17	8.46	8.34	8.00
	SD	3.93	3.79	3.95	4.00	4.22	4.27
Head's occ SEI	Mean	28.40	25.55	26.63	28.29	28.48	23.14
	SD	15.63	12.54	14.80	14.23	15.56	9.56
Mother's education	Mean	8.59	7.88	8.26	8.12	8.39	8.84
	SD	3.68	3.93	3.89	3.94	4.52	3.60
Farm origin	Mean	.278	.336	.185	.282	.234	.290
	SD	.448	.473	.389	.451	.427	.461
Number of siblings	Mean	4.52	5.10	4.13	4.90	4.75	5.03
	SD	2.52	2.52	2.44	2.68	2.80	2.39
Southern origin	Mean	.207	.219	.229	.247	.262	.065
	SD	.405	.414	.420	.432	.444	.250

B. OCG Black Respondents

Sample Size		1616	88	145	112	55	9
Head's education	Mean	5.52	5.56	6.59	4.81	5.03	3.78
	SD	3.70	3.93	4.19	4.19	3.45	3.60
Head's occ SEI	Mean	20.95	21.72	18.33	21.16	18.58	18.42
	SD	8.85	12.53	8.18	9.39	7.22	3.62
Mother's education	Mean	6.87	6.20	6.65	5.95	5.95	3.67
	SD	3.67	3.80	3.75	4.24	3.46	3.74
Farm origin	Mean	.510	.500	.326	.602	.463	.444
	SD	.500	.503	.471	.492	.503	.527
Number of siblings	Mean	6.34	6.12	5.07	5.46	5.25	5.67
	SD	2.49	2.74	.648	2.48	2.37	2.65
Southern origin	Mean	.655	.612	.698	.652	.746	.667
	SD	.475	.490	.461	.479	.440	.500

Table 3. Family Structure and Birth Order, OCG White and Black Brothers

	Whites				Blacks			
	Baseline	Family Structure		Birth Order	Baseline	Family Structure		Birth Order
		Intact	Non-Intact			Intact	Non-Intact	
Effect on the Family Factor (Γ)								
Head's Education (γ_{11})	.127 (.009)	.122 (.009)	.161 (.029)	.111 (.008)	.183 (030)	.197 (.034)	.111 (.071)	.136 (.026)
Head's occ SEI (γ_{12})	.029 (.002)	.030 (.002)	.012 (.006)	.027 (.002)	.012 (.010)	.007 (.011)	.034 (.022)	.015 (.008)
Mom's Education (γ_{13})	.257 (.009)	.250 (.009)	.309 (.029)	.226 (.008)	.328 (.030)	.312 (.033)	.419 (.071)	.287 (.026)
Farm Origin (γ_{14})	-.867 (.055)	-.841 (.057)	-1.081 (.201)	-.753 (.049)	-1.729 (.190)	-1.708 (.208)	-1.914 (.462)	-1.550 (.165)
Number of Sibs (γ_{16})	-.224 (.009)	-.234 (.010)	-.149 (.031)	-.193 (.008)	-.029 (.033)	-.020 (.037)	-.082 (.074)	.002 (.029)
Souther Origin (γ_{17})	-.375 (.056)	-.362 (.058)	-.472 (.191)	-.357 (.050)	-.318 (.182)	-.296 (.200)	-.392 (.443)	-.237 (.160)
Global Effects on (Λ^y)								
Older (λ^y_{21})	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
Younger (λ^y_{31})	0.774 (.012)	.784 (.012)	.688 (.039)	1.000	0.690 (.038)	0.690 (.042)	0.675 (.079)	1.000
Variances, (Ψ, Θ^ϵ)								
between-family (ψ_{11})	3.514 (.106)	3.417 (.108)	4.341 (.446)	2.663 (.071)	5.252 (.513)	5.026 (.558)	6.314 (1.305)	3.489 (.296)
within-family (Θ^ϵ)								
Oldest(θ^ϵ_{11})	3.814 (.108)	3.817 (.110)	3.740 (.436)	4.726 (.087)	4.265 (.497)	4.340 (.546)	3.621 (1.210)	6.194 (.363)
Youngest (θ^ϵ_{22})	5.173 (.088)	5.040 (.091)	6.365 (.327)	4.665 (.086)	6.956 (.361)	6.909 (.397)	7.309 (.868)	6.043 (.358)
χ^2	112.49/20	57.52/10		405.19/21	33.64/20	28.38/10		83.06/21
bic	-117.11	-57.28		164.11	-149.56	-63.22		-108.98

Table 4. Descriptive Statistics: OCG Brother Sample, 1973

	Sibling Pair								Three Siblings							
	Cohort 1		Cohort 2		Cohort 3		Cohort 4		Cohort 1		Cohort 2		Cohort 3		Cohort 4	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Blacks (N)	185		133		115		93		398		379		403		291	
Education (R)	11.280	3.285	10.378	3.644	9.467	4.447	7.086	3.839	11.262	2.443	9.850	3.519	8.073	3.888	6.584	4.158
Education (S1)	11.064	2.600	10.129	3.264	9.121	4.010	7.783	3.255	10.651	2.810	9.559	3.428	7.962	3.806	6.748	3.815
Education (S2)	--	--	--	--	--	--	--	--	11.082	2.568	10.389	3.198	9.448	3.591	8.079	3.849
Head's Education	7.234	3.800	6.641	3.981	5.465	4.242	4.486	3.820	6.518	3.435	5.878	3.442	4.678	3.676	4.527	3.642
Head's occ SEI	21.428	11.748	21.017	11.173	20.911	9.292	19.666	6.928	20.370	8.118	20.816	7.930	20.691	8.207	21.179	9.427
Mom's Education	8.481	3.625	7.879	3.885	6.490	3.751	5.714	3.651	7.927	3.120	7.235	3.452	5.998	3.655	5.082	3.675
Farm Origin	0.268	0.444	0.395	0.491	0.444	0.499	0.615	0.489	0.370	0.483	0.464	0.499	0.644	0.479	0.668	0.472
Non-Int..t Family	1.230	0.422	1.195	0.397	1.322	0.469	1.257	0.440	1.166	0.372	1.191	0.394	1.193	0.395	1.198	0.399
Number of Sibs	4.759	2.707	4.103	2.628	3.741	2.211	4.253	2.384	6.811	2.185	6.806	2.191	6.730	2.195	6.882	2.201
Southern Origin	0.620	0.487	0.578	0.496	0.724	0.449	0.542	0.501	0.591	0.492	0.663	0.472	0.691	0.463	0.716	0.452
Whites (N)	2904		2062		1893		1312		2910		3090		3609		2829	
Education (R)	13.059	2.960	12.725	3.255	12.122	3.411	11.043	3.682	11.973	3.123	11.315	3.461	10.890	3.617	9.910	3.751
Education (S1)	13.067	2.818	12.497	2.935	11.986	3.082	10.945	3.413	11.817	3.079	11.025	3.249	10.736	3.314	9.896	3.372
Education (S2)	--	--	--	--	--	--	--	--	11.863	2.963	11.532	2.996	11.309	2.962	10.572	3.304
Head's Education	10.105	3.659	8.707	3.831	8.321	3.923	7.543	3.780	8.614	3.766	7.498	3.643	6.955	3.680	6.663	3.714
Head's occ SEI	33.212	18.573	30.742	17.121	30.143	16.220	29.004	15.996	28.357	15.450	26.099	13.516	26.025	13.581	25.856	13.119
Mom's Education	10.762	3.028	9.557	3.398	8.753	3.716	7.840	3.579	9.575	3.296	8.334	3.512	7.529	3.571	6.942	3.481
Farm Origin	0.127	0.333	0.182	0.386	0.226	0.418	0.293	0.455	0.214	0.410	0.300	0.458	0.357	0.479	0.425	0.494
Non-Intact Family	1.089	0.283	1.092	0.289	1.120	0.325	1.118	0.323	1.101	0.301	1.102	0.303	1.098	0.297	1.110	0.313
Number of Sibs	2.832	1.821	2.857	1.926	2.950	1.880	3.355	2.152	5.105	2.340	5.439	2.372	5.566	2.303	5.958	2.270
Southern Origin	0.193	0.395	0.198	0.399	0.160	0.367	0.181	0.385	0.231	0.421	0.248	0.432	0.234	0.424	0.208	0.406

Table 5. Model Selection, OCG White Brothers

Model		χ^2	*df*	Contrast		*bic*
Specific Hypotheses on Data						
0.	Baseline Model	1163.65	158	n.a.	--	-763.95
1.	0 + eq. consts on the effect of fb on resp. (λ^y_{k12})	1165.18	159	1.53	1	-774.62
2.	1 + eq. consts on the within-family variances (θ^ϵ_{k11})	1220.81	160	55.63	1	-731.19
3.	1 + eq. consts on the between-family unmeasured variance (ψ_{k22})	1168.19	160	3.01	1	-783.81
4.	3 + eq. consts on the effect of fb on the oldest (λ^y_{k22})	1203.62	161	35.43	1	-760.58
5.	3 + eq. consts on the within-family variances (θ^ϵ_{k22})	1171.09	161	2.90	1	-793.11
Specific Hypotheses on Cohorts						
A.	5 - eq. consts on the effect of fb on the oldest (λ^y_{k22}) in 3-brohter gps	1150.56	158	20.53	3	-777.04
B.	5 - eq. consts on the effect of fb on the oldest (λ^y_{k22}) in 2-brother gps	1144.35	158	26.74	3	-783.25
C.	5 - eq. consts on the effect of fb on the youngest (λ^y_{k32}) in 3-brother gps	1150.53	158	20.56	3	-777.07
D.	5 - eq. consts on within-family variances of resp. (θ^ϵ_{k11}) in 3-brother	1112.83	158	58.26	3	-814.77
E.	5 - eq. consts on within-family variances of resp. (θ^ϵ_{k11}) in 2-brother	1110.84	158	60.25	3	-816.76
F.	5 - eq. consts on within-family variances of the oldest (θ^ϵ_{k22})	1138.50	158	32.59	3	-789.10
G.	5 - eq. consts on within-family variances of the youngest (θ^ϵ_{k33})	1108.40	158	62.69	3	-819.20
H.	5 - eq. consts on unmeasured between-family variances (ψ_{22})	968.80	158	202.29	3	-958.80
I.	5 - eq. consts on the global effect of fb (β_{21})	1134.24	158	36.85	3	-793.36
J.	5 - eq. consts on the effects of fb variables (Γ)	955.23	140	215.86	21	-752.37
Final Model Selection						
A1.	D+E+F+G+H	829.85	146	342.24	15	-951.35
A2.	A1 + I	794.51	143	35.34	3	-950.09
A3.	A1 + J	595.42	125	234.43	21	-929.58
A4.	A1 - eq. consts on the effects of farm, non-intact family and south	656.99	137	172.86	9	-1014.41
A5.	A4 + I	646.88	134	10.11	3	-987.92

Table 6. Model Selection, OCG Black Brothers

Model		χ^2	*df*	Contrast		*bic*
Specific Hypotheses on Data						
0.	Baseline Model	430.68	158	n.a.	–	-1130.36
1.	0 + eq. consts on the effect of fb on resp. (λ^y_{k12})	430.70	159	0.02	1	-1140.22
2.	1 + eq. consts on the within-family variances (θ^ϵ_{k11})	431.27	160	0.57	1	-1149.53
3.	2 + eq. consts on the between-family unmeasured variance (ψ_{k22})	432.52	161	1.25	1	-1158.16
4.	3 + eq. consts on the effect of fb on the oldest (λ^y_{k22})	440.65	162	8.13	1	-1159.91
5.	3 + eq. consts on the within-family variances (θ^ϵ_{k22})	432.62	162	0.10	1	-1167.94
Specific Hypotheses on Cohorts						
A.	5 - eq. consts on the effect of fb on the oldest (λ^y_{k22}) in 3-brohter gps	424.48	159	7.14	3	-1146.44
B.	5 - eq. consts on the effect of fb on the oldest (λ^y_{k22}) in 2-brohter gps	421.02	159	11.60	3	-1149.90
C.	5 - eq. consts on the effect of fb on the youngest (λ^y_{k32}) in 3-brother gps	419.50	159	13.12	3	-1151.42
D.	5 - eq. consts on within-family variances of resp. (θ^ϵ_{k11})	354.12	159	78.50	3	-1216.80
F.	5 - eq. consts on within-family variances of the oldest (θ^ϵ_{k22})	389.87	159	42.75	3	-1181.05
G.	5 - eq. consts on within-family variances of the youngest (θ^ϵ_{k33})	392.34	159	40.28	3	-1178.58
H.	5 - eq. consts on unmeasured between-family variances (ψ_{22})	391.33	159	41.29	3	-1179.59
I.	5 - eq. consts on the global effect of fb (β_{21})	389.11	159	43.51	3	-1181.81
J.	5 - eq. consts on the effects of fb variables	374.10	141	58.52	21	-1018.98
Final Model Selection						
A1.	D+F+G+H	280.44	150	152.18	12	-1201.56
A2.	A1 + I	229.34	147	51.10	3	-1223.02
A3.	A1 + J	213.45	129	66.99	21	-1061.07
A4.	A1 - eq. consts on the effects of farm, non-intact family and south	261.00	141	19.44	9	-1132.08
A5.	A4 + I	220.90	138	40.10	3	-1142.54

Table 7. Parameter Estimates, Model A3 and Preferred Final Model, OCG White Brothers

	Model A3				Preferred Final Model			
	Cohort 1	Cohort 2	Cohort 3	Cohort 4	Cohort 1	Cohort 2	Cohort 3	Cohort 4
Global Effect (β_{21})	1.0	1.0	1.0	1.0	1.0	1.0	1.0	1.0
Coefficients of Family Background Variables on the Common Family Factor (Γ)								
Head's Education (γ_{11})	0.108 (.010)	0.146 (.011)	0.142 (.012)	0.120 (.015)	0.129 (.006)	"	"	"
Head's Occ SEI (γ_{12})	0.250 (.019)	0.264 (.024)	0.343 (.024)	0.328 (.032)	0.285 (.012)	"	"	"
Mom's Education (γ_{13})	0.247 (.011)	0.203 (.012)	0.176 (.011)	0.243 (.015)	0.216 (.006)	"	"	"
Farm Origin(γ_{14})	-.097 (.073)	-.710 (.075)	-1.011 (.069)	-1.055 (.084)	-.089 (.071)	-.707 (.073)	-1.014 (.067)	-1.085 (.080)
Non-Intact Family (γ_{15})	-.444 (.090)	-.601 (.104)	-.948 (.098)	-.877 (.120)	-.466 (.090)	-.585 (.104)	-.945 (.099)	-.895 (.120)
Number of Sibs (γ_{16})	-.243 (.013)	-.204 (.015)	-.163 (.014)	-.178 (.018)	-.202 (.008)	"	"	"
Southern Origin (γ_{17})	-.150 (.066)	-.390 (.075)	-.627 (.075)	-.601 (.096)	-.172 (.065)	-.395 (.074)	-.596 (.075)	-.597 (.096)
Global Effects on (Λ^y)								
Oldest 3-b (λ^y_{22})	1.033 (.013)	"	"	"	1.033 (.013)	"	"	"
2-b (λ^y_{22})	0.938 (.015)	"	"	"	0.938 (.015)	"	"	"
Youngest (λ^y_{32})	0.859 (.012)	"	"	"	0.858 (.012)	"	"	"
Variances, (Ψ, Θ^ϵ)								
between-family (ψ_{22})	2.067 (.084)	3.044 (.110)	3.271 (.111)	3.828 (.146)	2.084 (.085)	3.049 (.110)	3.299 (.111)	3.844 (.146)
within-family (Θ^ϵ)								
Resp.(θ^ϵ_{11}) 3b	5.194 (.168)	5.848 (.185)	6.393 (.178)	6.926 (.229)	5.201 (.169)	5.844 (.185)	6.382 (.178)	6.924 (.229)
2b	4.273 (.157)	4.979 (.217)	5.383 (.244)	5.683 (.312)	4.273 (.157)	4.971 (.217)	5.378 (.244)	5.669 (.312)
Oldest(θ^ϵ_{22})	4.135 (.107)	3.741 (.114)	4.137 (.116)	4.230 (.147)	4.135 (.107)	3.745 (.114)	4.122 (.116)	4.228 (.147)
Youngest (θ^ϵ_{33})	4.991 (.153)	4.597 (.144)	4.148 (.121)	5.618 (.182)	4.988 (.153)	4.597 (.143)	4.175 (.121)	5.628 (.183)

Table 8. Parameter Estimates, Model A3 and Preferred Final Model, OCG Black Brothers

	Model A3				Preferred Final Model			
	Cohort 1	Cohort 2	Cohort 3	Cohort 4	Cohort 1	Cohort 2	Cohort 3	Cohort 4
Global Effect (β_{21})	1.0	1.0	1.0	1.0	1.0	1.506 (.140)	1.790 (.161)	1.616 (.165)
Coefficients of Family Background Variables on the Common Family Factor (Γ)								
Head's Education (γ_{11})	0.100 (.031)	0.164 (.040)	0.238 (.042)	0.215 (.051)	0.122 (.016)	"	"	"
Head's Occ SEI (γ_{12})	0.164 (.096)	0.171 (.127)	0.344 (.145)	0.436 (.169)	0.182 (.047)	"	"	"
Mom's Education (γ_{13})	0.161 (.033)	0.301 (.039)	0.306 (.042)	0.210 (.049)	0.168 (.018)	"	"	"
Farm Origin (γ_{14})	-.938 (.201)	-1.130 (.236)	-1.086 (.259)	-1.412 (.321)	-.758 (.098)	"	"	"
Non-Intact Family (γ_{15})	-.026 (.226)	-.579 (.272)	-.853 (.290)	-.572 (.363)	-.346 (.099)	"	"	"
Number of Sibs (γ_{16})	-.102 (.038)	-.020 (.048)	-.111 (.055)	-.055 (.065)	-.050 (.017)	"	"	"
Southern Origin (γ_{17})	-.443 (.179)	-.438 (.229)	-.475 (.255)	-.878 (.318)	-.362 (.085)	"	"	"
Global Effects on, (Λ^y)								
Oldest 3-b (λ^y_{22})	1.025 (.037)	"	"	"	1.029 (.037)	"	"	"
2-b (λ^y_{22})	0.881 (.048)	"	"	"	0.881 (.048)	"	"	"
Youngest (λ^y_{32})	0.831 (.034)	"	"	"	0.833 (.034)	"	"	"
Variances, (Ψ, Θ^ϵ)								
between-family (ψ_{22})	2.918 (.284)	3.827 (.402)	4.345 (.474)	5.133 (.606)	2.951 (.286)	3.852 (.403)	4.363 (.475)	5.179 (.610)
within-family (Θ^ϵ)								
Resp.(θ^ϵ_{11})	3.175 (.285)	5.127 (.449)	7.204 (.595)	6.581 (.670)	3.185 (.286)	5.144 (.450)	7.182 (.593)	6.706 (.678)
Oldest(θ^ϵ_{22})	3.388 (.288)	4.154 (.395)	4.905 (.476)	5.553 (.604)	3.388 (.289)	4.134 (.396)	4.920 (.478)	5.472 (.601)
Youngest (θ^ϵ_{33})	3.895 (.332)	5.421 (.472)	7.226 (.602)	7.313 (.723)	3.887 (.332)	5.429 (.473)	7.228 (.603)	7.294 (.723)

Table 9. Variances, OCG White Brothers, Preferred Final Model

	Between-Family		Within-Family	Total	Percentage		
	Measred	UnMsred			Measrd	UnMsrd	W/in
Three-Brother Sibship							
Respondent C1	2.643	2.084	5.201	9.928	26.62	20.99	52.39
Respondent C2	2.937	3.049	5.844	11.830	24.83	25.77	49.40
Respondent C3	3.252	3.299	6.382	12.933	25.14	25.51	49.35
Respondent C4	3.203	3.844	6.924	13.971	22.93	27.51	49.56
Two-Brother Sibship							
Respondent C1	2.356	2.084	4.273	8.713	27.04	23.92	49.04
Respondent C2	2.943	3.049	4.971	10.963	26.84	27.81	45.34
Respondent C3	3.400	3.299	5.378	12.077	28.15	27.32	44.53
Respondent C4	3.377	3.844	5.669	12.890	26.20	29.82	43.98
Three-Brother Sibship							
Oldest C1	2.820	2.224	4.135	9.179	30.72	24.23	45.05
Oldest C2	3.134	3.254	3.745	10.133	30.93	32.11	36.96
Oldest C3	3.470	3.520	4.122	11.112	31.23	31.68	37.10
Oldest C4	3.417	4.102	4.228	11.747	29.09	34.92	35.99
Youngest C1	1.946	1.534	4.988	8.468	22.98	18.12	58.90
Youngest C2	2.162	2.245	4.597	9.004	24.01	24.93	51.06
Youngest C3	2.394	2.429	4.175	8.998	26.61	26.99	46.40
Youngest C4	2.358	2.830	5.628	10.816	21.80	26.16	52.03
Two-Brother Sibship							
Oldest C1	2.073	1.834	4.135	8.042	25.78	22.81	51.42
Oldest C2	2.589	2.683	3.745	9.017	28.71	29.75	41.53
Oldest C3	2.991	2.903	4.122	10.016	29.86	28.98	41.15
Oldest C4	2.971	3.382	4.228	10.581	28.08	31.96	39.96

Table 10. Variances, OCG Black Brothers, Preferred Final Model

	Between-Family		Within-Family	Total	Percent		
	Measrd	UnMsrd			Measrd	UnMsrd	W/in
Three-Brother Sibship							
Respondents C1	1.290	2.951	3.185	7.426	17.37	39.74	42.89
Respondents C2	3.239	3.852	5.144	12.235	26.47	31.48	42.04
Respondents C3	4.707	4.363	7.182	16.252	28.96	26.85	44.19
Respondents C4	4.257	4.179	6.706	15.142	28.11	27.60	44.29
Two-Brother Sibship							
Respondents C1	1.612	2.951	3.185	7.748	20.81	38.09	41.11
Respondents C2	4.677	3.852	5.144	13.673	34.21	28.17	37.62
Respondents C3	6.623	4.363	7.182	18.168	36.45	24.01	39.53
Respondents C4	3.309	5.179	6.706	15.194	21.78	34.09	44.14
Three-Brother Sibship							
Oldest C1	1.366	3.125	3.388	7.879	17.34	39.66	43.00
Oldest C2	3.430	4.079	4.134	11.643	29.46	35.03	35.51
Oldest C3	4.984	4.620	4.920	14.524	34.32	31.81	33.87
Oldest C4	4.507	5.484	5.472	15.463	29.15	35.47	35.39
Youngest C1	0.895	2.048	3.887	6.830	13.10	29.99	56.91
Youngest C2	2.248	2.673	5.429	10.350	21.72	25.83	52.45
Youngest C3	3.266	3.027	7.228	13.521	24.16	22.39	53.46
Youngest C4	2.954	3.594	7.294	13.842	21.34	25.96	52.69
Two-Brother Sibship							
Oldest C1	1.251	2.290	3.388	6.929	18.05	33.05	48.90
Oldest C2	3.630	2.990	4.134	10.754	33.75	27.80	38.44
Oldest C3	5.141	3.386	4.920	13.447	38.23	25.18	36.59
Oldest C4	2.568	4.020	5.472	12.060	21.29	33.33	45.37

Figure 1. MIMIC Model of Sibling Resemblance

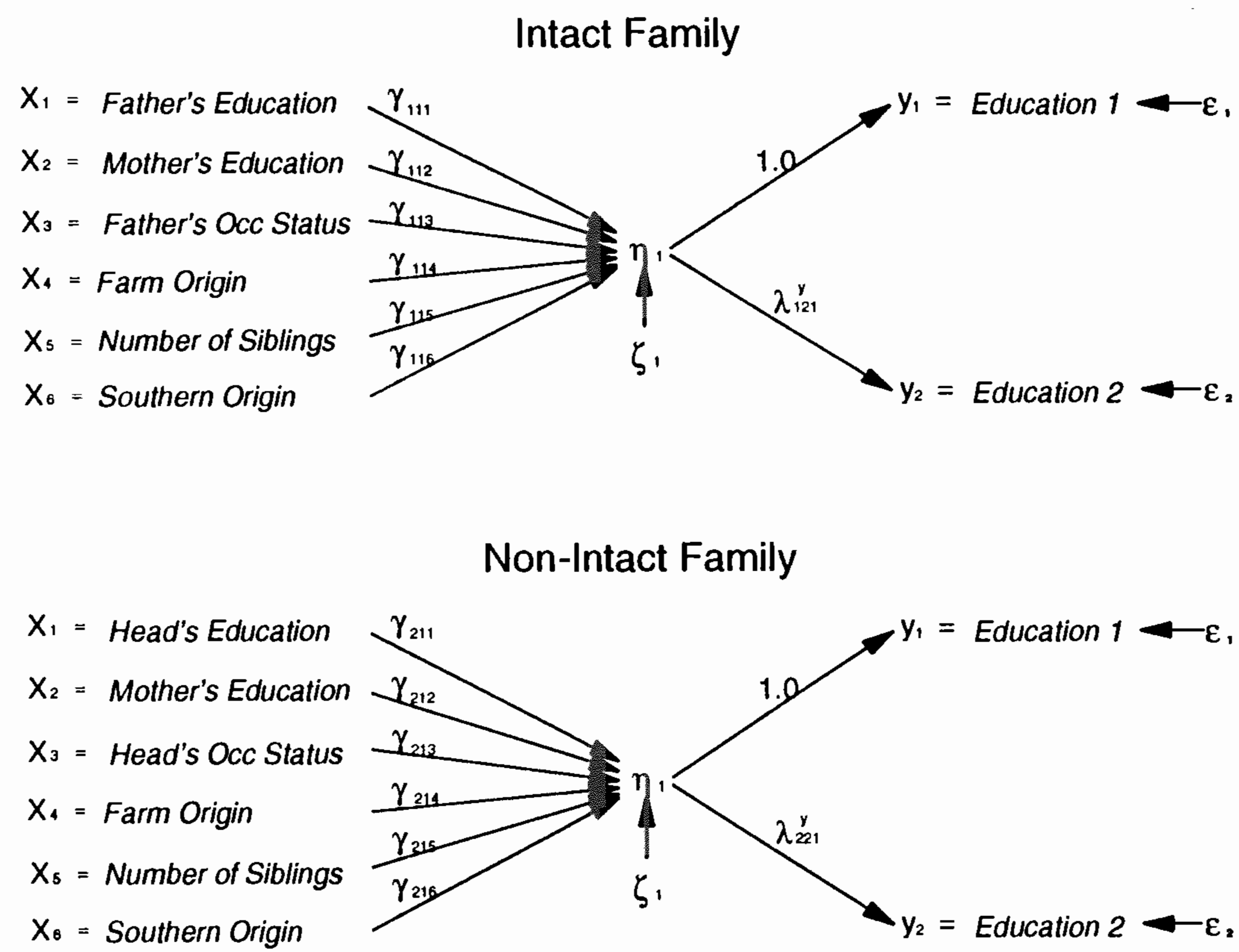
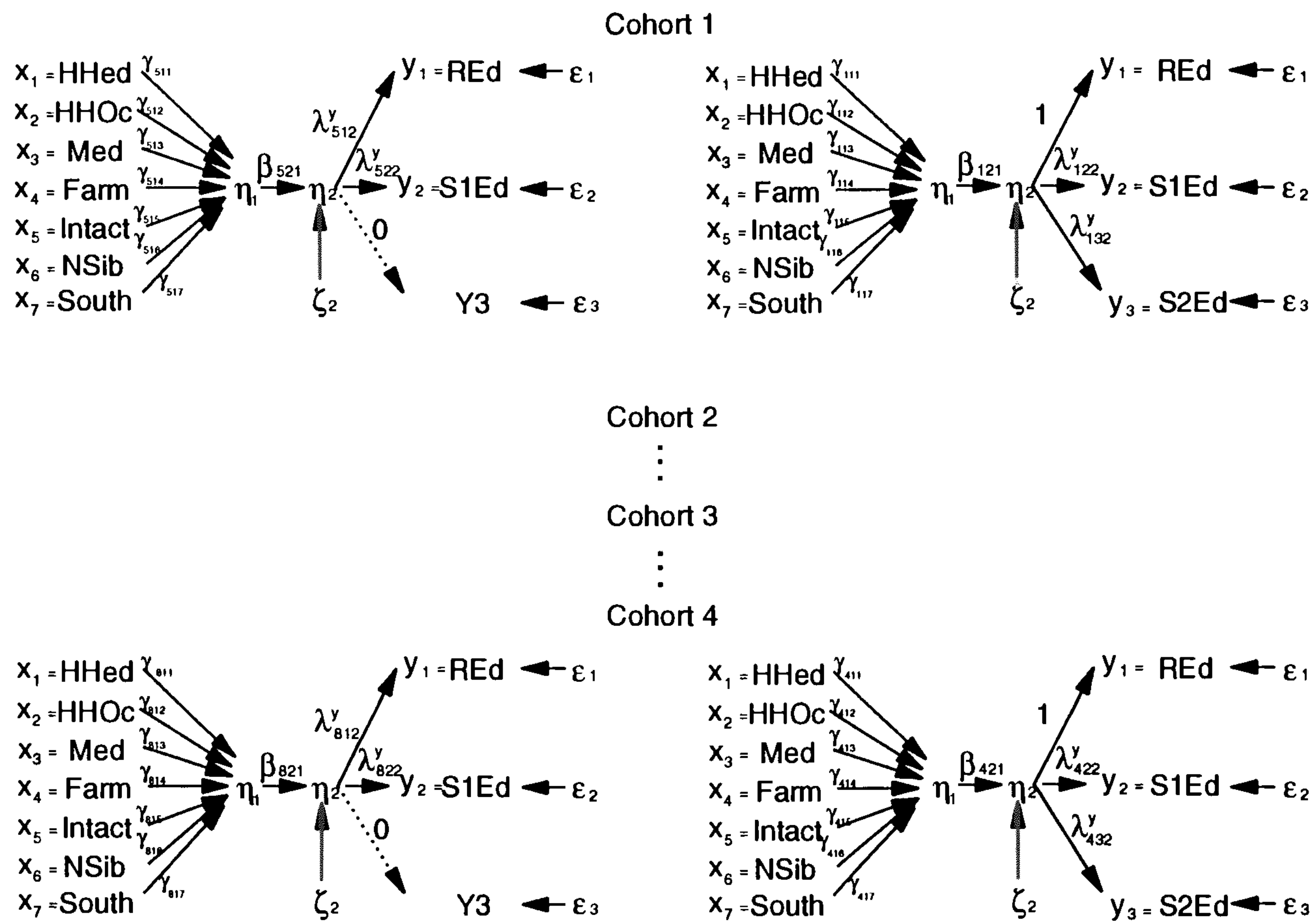


Figure 2. OCG Sibling Model: Brother Pairs and Triads in Four Cohorts



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