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Abstract

Biases due to measurement errors in structural equation models of the intergenerational transmission of socioeconomic status were assessed by estimating unobserved variable models with data from the remeasurement program of the 1973 Occupational Changes in a Generation-II survey. We found persuasive evidence that reports of social background and achievement variables by nonblack males are subject to strictly random errors, while reports of black males appear subject to significant nonrandom error. When measurement errors are ignored for nonblacks, occupational returns to schooling are underestimated by about 15 percent, the effects of some background variables are underestimated by as much as 22 percent, and variation in socioeconomic achievements not attributable to education or social origins is underestimated by as much as 27 percent. Biases appear to be substantially greater for nonblacks. Consequently, ignoring measurement error exaggerates racial differences in returns to schooling and occupational inequality not attributable to social origins.

RESPONSE ERRORS OF BLACK AND NONBLACK MALES IN MODELS OF STATUS INHERITANCE AND MOBILITY

Structural equation models have provided the foundation for research in social stratification for nearly a decade [Blau and Duncan, 1967; Duncan, Featherman and Duncan, 1972; Sewell and Hauser, 1975]. These models specify socioeconomic statuses as functions of social origins and intervening events and achievements. With the cumulation of data and findings, researchers have become increasingly concerned with precision and validity in measurement and parameter estimation. Some types of measurement error have been incorporated into substantive analyses of the achievement process using structural equation models that include unobserved variables [Siegel and Hodge, 1968; Jencks et al., 1972; Bowles, 1972; Bowles and Nelson, 1974; Bowles and Gintis, 1976; Mason et al., 1976; Treiman and Hauser, 1976].

Precision is not the central issue in the treatment of measurement error and data quality in socioeconomic achievement models. Incorrect specification of measurement error (e.g., ignoring it) can result in systematic bias in parameter estimates. The size and importance of such biases remain points of controversy. Jencks et al. conclude that "random measurement error is of relatively little importance in research of the kind described here" [1972:336]. Bowles [1972:S222] asserts that "social class background is considerably more important as a determinant of both educational attainment and economic success than has been indicated in recent analogous statistical treatments by Duncan and others." Bowles argues that retrospective reports of parental statuses are much less reliable than respondents' reports of their own attainments and that the effects of origin variables are consequently underestimated.

Patterns of response error have been built into models of the achievement process by obtaining multiple indicators of background and achievement variables and specifying models in which the covariation among the indicators is generated by unobserved "true scores." Figure 1 presents a path diagram of such a model with two measures of each of four variables. The model specifies that the jth measure of the ith variable, x_{ij}, is generated by the true score of that variable, T_i, plus a response error, e_{ij}, that is independent of T_i. That is, the measurement structure is

$$x_{ij} = \lambda_{ij}T_i + e_{ij}, \ (i = 1,\ldots,4;\ j = 1,2). \tag{1.1}$$

The model also specifies a fully recursive causal structure among the true scores:

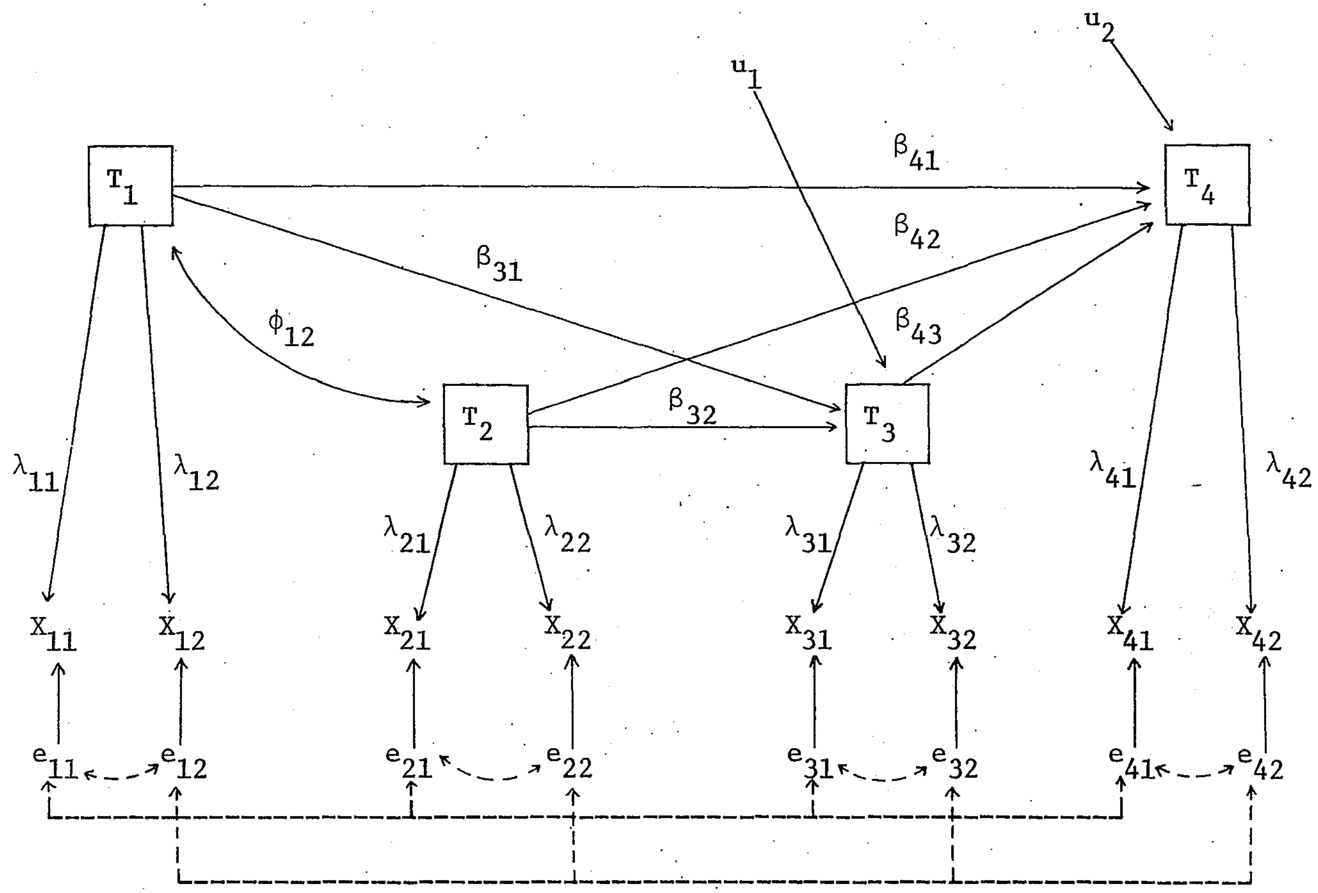
$$T_3 = \beta_{31}T_1 + \beta_{32}T_2 \qquad\qquad + u_1 \ , \tag{2.1}$$

$$T_4 = \beta_{41}T_1 + \beta_{42}T_2 + \beta_{43}T_3 + u_2 \ . \tag{2.2}$$

The method most often used to estimate the parameters of such models has been first, to estimate (or borrow) the parameters of the error structure, second, to estimate the covariance matrix of true scores, and then to estimate the structural coefficients relating the true scores.

To complete the model, the pattern of covariation among response errors must be specified. When multiple responses are obtained from the same individuals, three types of covariation among response errors appear particularly plausible. First, response errors in the report of a variable may covary with the respondent's true score on that variable. For example, individuals of high status may tend to understate their status while those of low status overstate their status. The implication for the measurement structure would be a nonunit slope of the population

FIGURE 1 -- A fully recursive structural equation model with measurement errors.



regression relating the observed measure, x_{ij}, to the true score, T_i. This type of correlated error is captured by the slope coefficient, λ_{ij}, while maintaining the lack of correlation between T_i and e_{ij}. A second source of covariation in response error would be a tendency for respondents to overstate the consistency between different variables ascertained on a single occasion. This "within-occasion/between-variable correlated error" is represented in Figure 1 by the dotted lines showing correlations among the e_{i1} and e_{i2}, for $i = 1, \ldots, 4$. A third source of correlated response error would be contamination of the respondent's second report of a given variable by his recollection of the earlier report of that variable. This "within-variable/between-occasion correlated error" is represented in Figure 1 by correlations among pairs of response errors, e_{i1} and e_{i2}, for $i = 1, \ldots, 4$.

Unfortunately, attempts to apply models like that in Figure 1 to the achievement process have been limited by a lack of appropriate data, by inadequate specifications, and by crude estimation procedures. Siegel and Hodge [1968], Jencks et al. [1972], Bowles and Nelson [1974], and Treiman and Hauser [1976] relied on between-occasion correlations of educational attainment, occupational status, and income computed from census tabulations. To these data, Bowles [1972; Bowles and Nelson, 1974] added findings from matched census and retrospective reports, which were obtained for part of the Chicago pretest sample of the 1962 Occupational Changes in a Generation (OCG) Survey [Blau and Duncan, 1967:457-462]. However, none of these data included covariances of measures of different variables ascertained on different occasions, i.e., no correlations between x_{ij} and $x_{i'j'}$, where

$i \neq i'$ and $j \neq j'$, were obtained. This lack of complete covariance information precluded estimation of correlated errors, and thus the resulting estimates were dependent upon untestable assumptions. Further, these researchers had to rely on tenuous assumptions about relationships between reporting errors in censuses and in other social surveys.

Bowles [1972] specified within-variable correlated error in his models, but assumed an arbitrary value for these correlations, e.g., $\rho_{e_{i1}e_{i2}} = .5$, rather than estimating them. The size of the error correlations is important, because ignoring positive within-variable correlated errors decreases estimated true score correlations while positive within-occasion correlated errors have the opposite effect. Bowles did not have enough information to identify either within-variable or within-occasion correlated error--it seems arbitrary that he specified a high level of correlation among errors between measurement occasions, but no such correlations within a single occasion. That is, Bowles' assumptions guaranteed he would obtain upper-bound estimates of intergenerational true score correlations.

The specification of models with variables in standard deviation units rather than in their natural metric has resulted in additional problems in the research of Bowles, Treiman and Hauser, Jencks et al., and Siegel and Hodge. Data quality assumptions stated in terms of error variances by Bowles and by Siegel and Hodge have been implemented in terms of standardized parameters. Yet these assumptions are not invariant to standardization. Moreover, the identifying information implied by unit slope coefficients in the measurement equations is lost under standardization. In addition, standardized measurement parameters (reliability coefficients) have been applied to heterogeneous populations [Bowles, 1972; Kalleberg, 1974;

Treiman and Hauser, 1976; Jencks et al., 1972; Featherman, 1973; Kelley, 1973] but the unstandardized parameters (error variances) are more likely to be invariant [Wiley and Wiley, 1970]. Finally, measurement parameters have been applied across studies where measurement techniques as well as populations differ. For example, Siegel and Hodge recognized differences in the quality of census and CPS (Current Population Survey) measurement procedures, but such differences have not always been considered in the "borrowing" of reliability coefficients.

In summary, while strong statements about the effects of measurement error can be found in the existing literature, these statements have been based on inadequate data and models. The issues have been well stated. Failure to incorporate response error structures into models of the achievement process may lead to underestimates of the effects of social background on schooling and achievement, or to overestimates of the effects of schooling on later achievements. Without estimates based upon more comprehensive data and a less restricted specification of error structures, we can accept neither the positions of Jencks et al. [1972] and Siegel and Hodge [1968] that the biases are negligible, nor the position of Bowles [1972] that they are substantial.

1973 OCG Data

Data from the remeasurement program of the 1973 Occupational Changes in a Generation-II study allow us to estimate and test less restrictive models of response error and to assess the effects of plausible error structures on parameters of the achievement process. The 1973 OCG study [Featherman and Hauser, 1975] was designed to achieve a strict replication of the 1962 study conducted by Blau and Duncan [1967]. The 1973 survey, executed in conjunction with the March 1973 Current Population Survey, represents approximately 53 million males in the civilian noninstitutional population between the ages of 20 and 65 in March 1973. Educational and labor-force data were obtained from the March 1973 CPS household interviews. In about three-fourths of the cases the CPS respondent was the spouse of the designated male. These data were supplemented in the fall of 1973 with social background and occupational career data from the mailout-mailback OCG questionnaire (OCGQ). In about three-fourths of these cases the OCGQ respondent was the designated male. Responses to OCGQ were obtained from this questionnaire or subsequent telephone or personal follow-ups for more than 27,000 members of the experienced civilian labor force. The overall response rate was greater than 88 percent. A random subsample of about 1,000 OCGQ respondents (600 nonblacks and 400 blacks) was selected for inclusion in the OCG remeasurement program (OCGR). Approximately three weeks after the mail return of their OCG questionnaires, telephone (and in a few cases personal) interviews were conducted with these respondents to obtain a second report of selected items on the OCG questionnaire.

Table 1 shows which variables were measured on each of the three occasions--CPS, OCGQ, and OCGR. Educational attainment (x_{43}), current (March) occupation (x_{63}), and age of the designated male (AGE) were ascertained in the March CPS interview. Reports of the three social background variables--father's (or other head of household's) occupation (x_{11}), father's (or other head of household's) educational attainment (x_{21}), and parental family income (x_{31})--were obtained from the fall OCG questionnaire. Also, the fall questionnaire ascertained a man's first full-time, civilian job after completing schooling (x_{51}) and a second measurement of

TABLE 1 -- Timing of measurements in the 1973 CPS and OCG surveys.

Variable	Measurement		
	March 1973 CPS household interview (CPS)	Fall 1973 OCG questionnaire (OCGQ)	Fall 1973 OCG remeasurement interview (OCGR)
1. Father's occupational status (FO)	-	x_{11}	x_{12}
2. Father's educational attainment (FE)	-	x_{21}	x_{22}
3. Parental income (PI)	-	x_{31}	x_{32}
4. Educational attainment (ED)	x_{43}	x_{41}	x_{42}
5. Occupational status of first job after completing schooling (O1)	-	x_{51}	x_{52}
6. Current occupational status (March or fall) (OC)	x_{63}	-	x_{62}
7. Age	AGE, AGE2	-	-

educational attainment (x_{41}). Thus, the CPS and OCGQ measurements provide two reports of educational attainment and one report of six other variables for each male in the full CPS-OCGQ sample. (The second measurement of ED was not intended to supplant the CPS item, but rather to improve the respondent's recall of the timing of schooling and labor force entry.) Within the OCGR subsample, each of the variables except age was remeasured. For technical reasons we were not able to ascertain March 1973 occupation in the OCGR interviews, therefore, we obtained a report of current (Fall 1973) occupation (x_{62}). While some job mobility occurred between the spring spring and fall surveys, we disregard it here on the argument that occupational status changes were negligible over the six- or seven-month period. Consequently, our estimates of unreliability in the reporting of current occupational status include effects of job mobility as well as response error. In summary, for OCGR respondents we have two measures of each of the social background variables (FO, FE, and PI), three reports of educational attainment (ED), two reports of both first and current occupation (O1 and OC), and a single report of age (AGE).

Each of the occupation reports was scaled using Duncan SEI scores for detailed 1960 Census occupation, industry, and class of worker categories [Duncan, 1961]. Thus, our estimates of the quality of occupation reports do not pertain to a description of occupations per se, but rather to a particular transformation of detailed job descriptions into a status metric [Featherman and Hauser, 1973]. Educational attainment is coded in exact years of schooling completed, and parental income is coded as the logarithm of price adjusted dollars.[1] Age is expressed in years divided by ten, and a quadratic age variable, AGE2, is defined as $(\text{years}-40)^2/10$.

Model Specification

Our strategy is to specify and estimate measurement models separately for the 578 nonblacks and 348 blacks of the remeasurement (OCGR) subsamples and then apply the estimated measurement models to the full CPS-OCGQ samples of 25,223 nonblacks and 2,020 blacks. In this way we estimate substantive parameters in the full samples that have been corrected for response error. It is instructive to compare the corrected estimates with naive estimates for the full samples, i.e., estimates assuming perfect measurement. After examining the biases in the naive estimates due to measurement error for nonblacks and blacks, we assess the implications of these biases for detecting racial differences in the stratification process.

Our structural model is presented in the path diagram of Figure 2.[2] The variables enclosed in boxes, FO, FE, PI, ED, 01, and OC are unobserved true scores. Linear and quadratic age terms, AGE and AGE2 are assumed to be measured without error in the CPS interviews. The term x_{ij}, refers to the jth report of the ith variable, as indicated in Table 1.

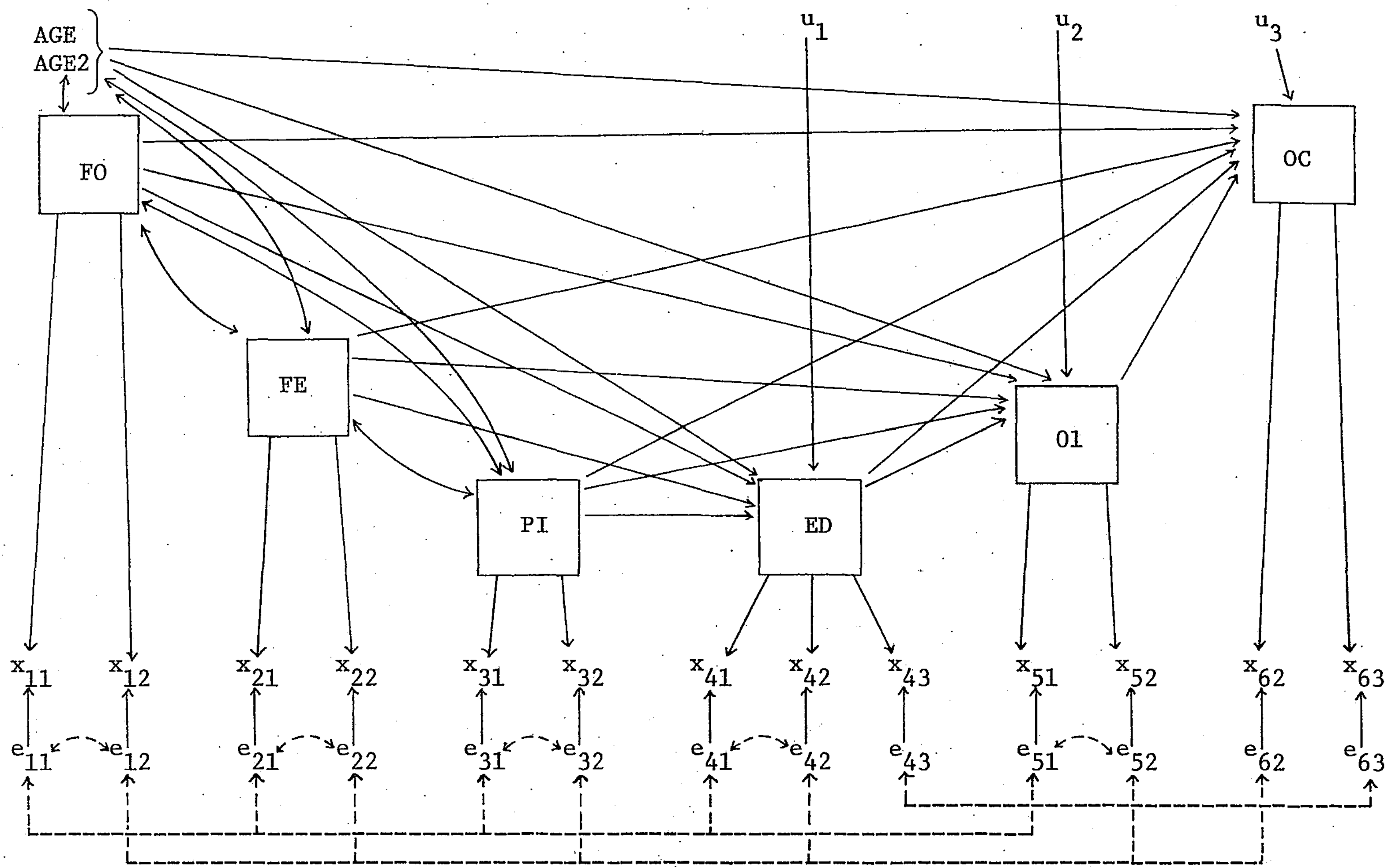
The substantive portion of Figure 2 is a fully recursive model among true scores, represented by the following structural equations:

$$\begin{aligned} ED = \alpha_1 + \beta_1(AGE) + \beta_2(AGE2) + \beta_3(FO) + \beta_4(FE) + \\ \beta_5(PI) \qquad + u_1 , \end{aligned} \tag{3.1}$$

$$\begin{aligned} 01 = \alpha_2 + \beta_6(AGE) + \beta_7(AGE2) + \beta_8(FO) + \beta_9(FE) + \\ \beta_{10}(PI) + \beta_{11}(ED) \qquad + u_2 , \end{aligned} \tag{3.2}$$

$$\begin{aligned} OC = \alpha_3 + \beta_{12}(AGE) + \beta_{13}(AGE2) + \beta_{14}(FO) + \\ \beta_{15}(FE) + \beta_{16}(PI) + \beta_{17}(ED) + \beta_{18}(01) \qquad + u_3 , \end{aligned} \tag{3.3}$$

FIGURE 2 -- A structural equation model of the stratification process with measurement errors



NOTE: Variables are defined in Table 1.

where the disturbances are independent of each other and of the explanatory variables in their respective equations. These substantive equations will be just-identified in terms of the true score variances and covariances. Thus, the fully recursive structure does not constrain estimates of parameters of the measurement model.

In algebraic form, the measurement portion of Figure 2 is

$$x_{11} = \lambda_{11}(FO) + e_{11}\,, \qquad (4.1a)$$

$$x_{12} = \lambda_{12}(FO) + e_{12}\,, \qquad (4.1b)$$

$$x_{21} = \lambda_{21}(FE) + e_{21}\,, \qquad (4.2a)$$

$$x_{22} = \lambda_{22}(FE) + e_{22}\,, \qquad (4.2b)$$

$$x_{31} = \lambda_{31}(PI) + e_{31}\,, \qquad (4.3a)$$

$$x_{32} = \lambda_{32}(PI) + e_{32}\,, \qquad (4.3b)$$

$$x_{41} = \lambda_{41}(ED) + e_{41}\,, \qquad (4.4a)$$

$$x_{42} = \lambda_{42}(ED) + e_{42}\,, \qquad (4.4b)$$

$$x_{43} = \lambda_{43}(ED) + e_{43}\,, \qquad (4.4c)$$

$$x_{51} = \lambda_{51}(01) + e_{51}\,, \qquad (4.5a)$$

$$x_{52} = \lambda_{52}(01) + e_{52}\,, \qquad (4.5b)$$

$$x_{62} = \lambda_{62}(OC) + e_{62}\,, \qquad (4.6a)$$

$$x_{63} = \lambda_{63}(OC) + e_{63}\,. \qquad (4.6b)$$

The model allows both within-occasion and within-variable correlated response error. Response errors of reports obtained from the fall OCG questionnaire, e_{11}, e_{21}, e_{31}, e_{41}, e_{52} and e_{51} may be intercorrelated, as may be errors of reports obtained from the fall OCG telephone remeasurement interview, e_{12}, e_{22}, e_{32}, e_{42}, e_{52} and e_{62} and the errors of the two reports obtained from the March CPS household interview, e_{43} and e_{63}. We allow within-variable correlated errors in the reports of variables obtained from the fall OCG questionnaire and the fall OCG telephone remeasurement interview, that is, correlations between e_{i1} and e_{i2} for $i = 1, \ldots, 5$. It seems plausible that recall contamination might occur in these responses, obtained an average of 24 days apart. However, we assume that such contamination does not occur between the March CPS reports and the fall OCG reports of educational attainment and occupational status. These were obtained more than five months apart, and from different respondents in about 70 percent of the cases.

We establish a metric for the true scores by fixing $\lambda_{11} = \lambda_{21} = \lambda_{31} = \lambda_{43} = \lambda_{51} = \lambda_{63} = 1.0$. That is, we fix the metric of the true scores to be the same as that of the observed reports that are used in models for the full CPS-OCGQ sample. The metrics of FO, FE, PI, and 01 are identical to those of the corresponding OCGQ reports, and the CPS reports define the metrics for ED and OC. A normalization of this kind is necessary because the metric of an unobserved variable is arbitrary, and consequently the slope coefficients with respect to indicators are identifiable only relative to each other. For example, given our normalization, a coefficient, λ_{i2}, greater (or smaller) than unity, indicates a conditional expectation slope of the OCGR report on the true score which is steeper (flatter) than the slope of the OCGQ report on the true score. However, the absolute values of the two slopes are indeterminate.[3] This normalization is imposed upon all of our models.

Our measurement models are all based on equations 4 and differ only in the specification of the covariances among the e_{ij} and the restrictions imposed upon the λ_{ij}. Our most restrictive specification, Model A, (see Table 4) permits only random measurement errors, so the e_{ij} are assumed to be mutually uncorrelated. It corresponds to the random measurement error models of Siegel and Hodge [1968:51-52], Jencks et al. [1972:330-336], Treiman and Hauser [1976], and the one implicitly used by other researchers applying "corrections for attenuation" [cf., Bohrnstedt, 1970]. Thus, in Model A the 91 variances and covariances among the thirteen reports (ignoring age) are to be reproduced by 41 free parameters: 7 slope coefficients, 13 error variances, 6 true score variances, and 15 true score covariances.

After assessing Model A, we consider more complex measurement models. Model B corresponds to the model specified by Bowles [1972]. It differs from Model A only in that within-variable error correlations ($\rho_{e_{i1},e_{i2}}$ for i = 1, . . . , 5) are fixed to be 0.5 instead of fixed to be zero. Model C allows both within-variable and within-occasion correlations. To identify these additional parameters, we must impose some other constraints. Within-occasion correlated errors are constrained to be equal when they involve the same pair of variables. That is, we have 10 constraints of the form

$$\rho_{e_{i1}e_{k1}} = \rho_{e_{i2}e_{k2}} \quad (i, k = 1, \ldots, 5;\ i \neq k),$$

and also,

$$\rho_{e_{43}e_{63}} = \rho_{e_{42}e_{62}}.$$

The other four within-occasion correlated errors, $\rho_{e_{i2}e_{62}}$ (i = 1, 2, 3, 5) are constrained. The availability of a third (CPS) measure of education, x_{43}, with an error component, e_{43}, uncorrelated with the error components of the OCGQ and OCGR measures identifies the within-variable error correlation, $\rho_{e_{41}e_{42}}$. We shall assume that within-variable error correlation between OCGQ and OCGR reports of other variables exists to the same degree that it can be detected in the education reports. That is, we constrain the within-variable error correlations to be equal across the five variables measured both in the OCG questionnaire and the remeasurement interviews, and show,

$$\rho_{e_{11}e_{12}} = \rho_{e_{21}e_{22}} = \ldots = \rho_{e_{51}e_{52}}.$$

Model C adds 16 free parameters for the measurement error correlations--one for the within-variable correlation, and 15 for the within-occasion correlations.

We estimate other models but these are variations of Models A, B, and C. Then we take the most appropriate or best fitting model, and reestimate it after eliminating statistically and substantively insignificant coefficients and constraining to unity those estimated slope coefficients that appear statistically indistinguishable from 1.0.

The measurement model parameter estimates for the nonblack and black OCGR subsamples provide true score variance-covariance matrices from which we could solve for the substantive parameters of equations 3. However, we can obtain more stable estimates of the substantive parameters by using the measurement error variances and error correlations from the OCGR subsamples to correct the observed variance-covariance matrices for the full CPS-OCGQ samples. In doing so, we assume that our OCGR-based estimates of equations 4.1a, 4.2a, 4.3a, 4.4c, 4.5a, and 4.6b apply to the CPS reports of ED and OC, and apply to the OCGQ reports of FO, FE, PI, and 01 in the full CPS-OCGQ samples of nonblacks and blacks.[4] We can then compare, for each racial group, substantive parameters estimated from the corrected and uncorrected full sample variance-covariance matrices.[5]

Estimation of Measurement Models

Assuming the joint distribution of the thirteen reports of status variables is multivariate normal, we obtain maximum likelihood estimates of parameters of the 13-equation measurement model using Jöreskog's [1970] "general method for the analysis of covariance structures." The estimates have been computed from pair-wise present correlations for nonblack and black males 20 to 65 years old in the experienced civilian labor force in March 1973.[6] The correlations among the thirteen reports are given in Tables 2 and 3 and means and standard deviations appear in the first two columns of Tables 5 and 6. It appears that there is a slight tendency for respondents to report higher statuses in the remeasurement telephone interviews. While this may indicate a social desirability effect in the interview situation that is not elicited by the questionnaire [Couch and Keniston, 1960; Campbell, Siegman, and Rees, 1967] it may also be due in part to lower-response rates for some items among lower-status persons in the telephone interview. There is a more pronounced tendency for the OCGR items to vary less than the same OCGQ items. Thus, we might expect to find smaller error variances in the OCGR items.

Goodness-of-fit tests for the various measurement models are reported in Table 4. The likelihood-ratio test statistic contrasts the null hypothesis that constraints on the observed variance-covariance matrix are satisfied in the population with the alternative that the variance-covariance matrix is unrestricted. In large samples, this statistic has a chi-square distribution with degrees of freedom equal to the difference between the number of variances and covariances and the number of independent parameters estimated under the hypothesized model. Moreover, when two measurement models

TABLE 2 -- Observed correlations among status variables: OCGR subsample of nonblack males in the experienced civilian labor force, March 1973 (N = 578)

Variable			(1)		(2)		(3)		(4)			(5)		(6)	
			x_{11}	x_{12}	x_{21}	x_{22}	x_{31}	x_{32}	x_{41}	x_{42}	x_{43}	x_{51}	x_{52}	x_{62}	x_{63}
1.	FO	x_{11}	--												
		x_{12}	.869	--											
2.	FE	x_{21}	.585	.589	--										
		x_{22}	.597	.599	.939	--									
3.	PI	x_{31}	.422	.437	.477	.467	--								
		x_{32}	.426	.450	.486	.478	.913	--							
4.	ED	x_{41}	.428	.430	.448	.445	.426	.439	--						
		x_{42}	.445	.443	.483	.492	.485	.502	.838	--					
		x_{43}	.419	.419	.467	.467	.486	.501	.801	.921	--				
5.	O1	x_{51}	.398	.410	.290	.300	.370	.358	.581	.644	.637	--			
		x_{52}	.409	.409	.325	.322	.363	.348	.578	.642	.631	.847	--		
6.	OC	x_{62}	.340	.369	.280	.284	.291	.296	.504	.563	.534	.585	.599	--	
		x_{63}	.364	.390	.291	.308	.307	.301	.519	.603	.566	.618	.620	.797	--

NOTE: See Table 1 for definitions of variables.

TABLE 3 -- Observed correlations among status variables: OCGR subsample of black males in the experienced civilian labor force, March 1973 (N = 348)

Variable			(1) x_{11}	x_{12}	(2) x_{21}	x_{22}	(3) x_{31}	x_{32}	(4) x_{41}	x_{42}	x_{43}	(5) x_{51}	x_{52}	(6) x_{62}	x_{63}
1.	FO	x_{11}	--												
		x_{12}	.639	--											
2.	FE	x_{21}	.442	.508	--										
		x_{22}	.437	.531	.916	--									
3.	PI	x_{31}	.207	.266	.320	.353	--								
		x_{32}	.271	.367	.361	.363	.841	--							
4.	ED	x_{41}	.137	.238	.398	.384	.419	.450	--						
		x_{42}	.159	.247	.398	.401	.374	.414	.914	--					
		x_{43}	.168	.239	.393	.371	.390	.369	.815	.870	--				
5.	O1	x_{51}	.295	.271	.281	.262	.267	.280	.481	.475	.476	--			
		x_{52}	.182	.265	.269	.254	.252	.328	.454	.498	.464	.771	--		
6.	OC	x_{62}	.230	.297	.321	.309	.281	.297	.491	.511	.510	.500	.537	--	
		x_{63}	.169	.327	.335	.342	.269	.316	.520	.540	.516	.517	.537	.724	--

NOTE: See Table 1 for definitions of variables.

TABLE 4 -- Chi-square goodness-of-fit tests for measurement models: nonblack and black males in the experienced civilian labor force, March 1973

	Nonblacks (N=578)			Blacks (N=348)		
Model	χ^2	df	p	χ^2	df	p
A. Random measurement error -- no constrained slopes	43.82	50	.718	130.64	50	.000
B. "Bowles" Model -- Within variable correlated error fixed at 0.5	81.61	50	.003	129.36	50	.000
C. Within-occasion and within-variable correlated error	31.06	34	.612	70.92	34	.000
D. Within-occasion correlated error	31.95	35	.616	74.43	35	.000
E. Within-variable correlated error	43.28	49	.703	128.32	49	.000
F. Random measurement error -- constrained slopes (final nonblack model)	45.27	55	.822	--	--	--
G. Some within-occasion and fixed within-variable correlated error	--	--	--	83.56	46	.001
H. Some within-occasion, fixed within-variable correlated error and constrained slopes (final black model)	--	--	--	84.25	48	.001

NOTE: Maximum likelihood estimates were computed with the ACOVSF program described in Jöreskog, Gruvaeus and van Thillo [1970].

are "nested," that is, when one model can be obtained by constraining the parameters of a more general model, the difference in chi-square values provides a likelihood-ratio test of the constrained parameters.

Measurement Models: Nonblacks

Goodness-of-fit tests of measurement models for nonblacks appear in the first three columns of Table 4. Model A, the random measurement error model, fits remarkably well ($p = .718$). In contrast, the "Bowles" model, Model B, differing only in that within-variable correlated error is fixed at 0.5 instead of zero, fits poorly ($p = .003$). Model C adds the 16 parameters for within-occasion and within-variable correlated error to the random measurement error model, but the fit does not significantly improve over Model A. The difference in chi-square values of 12.8 with 16 degrees of freedom is not statistically significant (compare lines A and C).

Lines D and E of Table 4, respectively, pertain to models with within-occasion correlated error, but no within-variable correlated error, and vice versa. Contrasting line D with line C, we see that the chi-square value for the within-variable correlated error parameter is not statistically significant. Comparing lines E and C, the chi-square value for the within-occasion correlated error parameters is 12.22 with 15 degrees of freedom, which is again less than its expected value on the null hypothesis. The point estimate of within-variable correlated error is 0.1 with an approximate standard error of 0.1 (not shown in the table). The largest point estimate of within-occasion correlated error is 0.07 with an approximate standard error of 0.07. Thus, neither in a global test, in separate tests for within-occasion and within-variable error correlations, nor in our examination of the several estimated within-occasion error correlations, do we find substantial evidence of correlated error.

The evidence that reporting errors are random for nonblack men is almost, but not quite, complete. Model F, the final measurement model, was constructed by imposing unit slopes on those free λ_{ij} that were within approximately one standard error of 1.0. Under Model A there were seven free slope parameters (λ_{ij}), but only the estimates of λ_{62}, λ_{41}, and λ_{42} were significantly different from 1.0. Further, the latter two estimated did not differ significantly from one another. Thus, in Model F we estimate only two free nonunit slope parameters, $\lambda_{41} = \lambda_{42}$ and λ_{62}. The five additional constraints in Model F raise chi-square by only 1.45 relative to Model A, and thus the 36 free parameters of Model F (2 slope coefficients, 13 error variances, 6 true-score variances, 15 true score covariances) provide a quite good representation of the 91 variances and covariances of the observed reports ($\chi^2 = 45.27$ with 55 df; $p = .822$).

Parameter estimates for this final measurement model for nonblacks appear in columns 3 through 5 of Table 5. Several features of these estimates are noteworthy. The OCGR interview reports, uniformly have smaller error variances than the OCGQ questionnaire reports. The three variables measured in the Duncan SEI metric FO, 01, and OC have error standard deviations ranging from 8 to 12, with those for FO and 01 somewhat smaller than those for OC. The reason may be that the retrospective reports are less detailed, or respondents may be ignoring transient components of their fathers', and their own first occupations which are not ignored in describing their own current occupations. The error standard deviation of

TABLE 5 -- Observed moments and measurement model parameter estimates: nonblack males in the experienced civilian labor force, March 1973 (N = 578)

	Variable		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	True	Observed	Mean	Observed Std. Dev.	Std. Dev. of Error[a]	Std. Dev. of True Score	Relative Slope[a]	Reliability Coefficient[b]	Test-Retest Correlations	Coding Reliability	Percent of Cases with Data Present
	T_i	x_{ij}	μ_{ij}	$\sigma_{x_{ij}}$	$\sigma_{e_{ij}}$	σ_{T_i}	λ_{ij}	$(\sigma^2_{T_i}/\sigma^2_{x_{ij}})\lambda^2_{ij}$	$\rho_{x_{i1},x_{i2}}$	$\rho_{x_{i1},x_{i1'}}$	
1.	FO	x_{11}	32.96	24.27	9.37 (.54)	22.37	1.00	.85	.87	.94	96
		x_{12}	33.62	23.73	7.97 (.59)		1.00	.89			95
2.	FE	x_{21}	8.97	4.19	1.12 (.09)	4.04	1.00	.93	.94	.99	95
		x_{22}	8.96	4.14	0.93 (.10)		1.00	.95			94
3.	PI	x_{31}	3.78	0.41	0.14 (.01)	0.38	1.00	.86	.91	.99	89
		x_{32}	3.81	0.39	0.09 (.01)		1.00	.95			90
4.	ED	x_{41}	11.98	3.42	1.78 (.06)		1.06 (.02)	.70	.84[c]	.95	93
		x_{42}	12.12	2.93	0.61 (.06)	2.71	1.06 (.02)	.96			94
		x_{43}	12.18	2.87	0.97 (.04)		1.00	.89			100[e]
5.	OI	x_{51}	34.61	24.71	9.86 (.52)	22.47	1.00	.87	.85	.94	89
		x_{52}	32.10	24.15	9.26 (.54)		1.00	.87			94
6.	OC	x_{62}	39.57	24.81	12.25 (.65)	23.11	0.93 (.04)	.76	.80[d]	---	100[e]
		x_{63}	41.34	25.21	10.08 (.80)		1.00	.84			100[e]

[a] Standard errors of parameter estimates appear in parentheses.

[b] These coefficients are squared "validity coefficients." They have approximate standard errors on the order of 0.03.

[c] $\rho_{x_{41},x_{43}} = .80$, $\rho_{x_{42},x_{43}} = .92$.

[d] This quantity is $\rho_{x_{62},x_{63}}$, the correlation between SEI scores of reports of March 1973 occupation and Fall 1973 occupation.

[e] Missing values have been allocated for NA cases.

the OCGQ report of Educational Attainment is anomalously large, nearly three times that obtained with the same item in the OCGR telephone interview. The two interview reports of education, OCGR and CPS are clearly superior to the questionnaire report.

As noted above, only two slope coefficients depart from the normalized value of 1.0. The CPS household interview report of educational attainment has a flatter slope than the other two reports, while the CPS report of occupational status has a steeper slope than the OCGR telephone interview report. Reliability coefficients (the squared true score-observed score correlations estimated from the measurement model) appear in column 6. It is striking that retrospective reports of social background variables are no less reliable than contemporaneous reports of status variables.

Correlations between the first and second reports of each of the variables appear in column 7. These observed "test-retest" correlations correspond to the reliability coefficients that would be obtained under a classical test theory model with congeneric forms in the measurement of each variable. For most variables these correlations are close to the mean of the estimated reliability coefficients of the indicators presented in column 6.

Column 8 presents external evidence of data quality for nonblacks: correlations between two independent codings of the OCGQ questionnaire responses for the variables FO, FE, PI, ED and O1. (The Bureau of the Census recoded OCG questionnaire responses after they were transcribed to telephone interview forms. Telephone interviewers used the transcribed responses to reconcile discrepancies after a second report was obtained.) These correlations reflect unreliability due to transcription, coding and keypunching error, but are free of unreliability due to response error. Thus, they provide an upper bound to the reliabilities attainable from the OCG questionnaire. We find very little coding unreliability in the precoded FE and PI variables. The coding reliability is .94 for FO and O1, which were coded into detailed Census codes from questions on occupation, industry, and class of worker and then transformed into the status metric. The correlation between codings of the education item in the OCG questionnaire is an unusually low .95. Thus, the relatively high error variance of the OCG questionnaire report on education may be due to unusually high coding or keypunch errors for that item.

Measurement Models: Blacks

Examining the fit of measurement models for blacks in Table 4, we encounter a notable lack of fit, compared to models estimated for nonblacks. Indeed, at conventional levels of statistical significance, we can reject all of our measurement models. Nevertheless, we can compare the fit of other models relative to the random measurement error model. Model B, the "Bowles" model, provides a negligibly better fit than the random error model. However, Model C adds 16 free correlated error parameters to the random error model, and reduces the chi-square value by about 45 percent, from 130.64 to 70.92. Furthermore, most of this improvement is attributable to the within-occasion correlated error, seen by comparing lines A and D. It is difficult to choose between Model D and Model C. Statistically, the improvement in fit from adding the within-variable error correlations to the within-occasion error correlations is minimal (χ^2 = 74.34 - 70.92 = 3.51 with 1 df, $0.05 < p < 0.10$). Substantively, the estimated within-variable error correlation is quite large, 0.44. In the absence of within-variable correlated errors, the largest within-occasion correlated errors are estimated to be about 0.2. In the presence of within-variable correlated errors the within-occasion error correlations fall to about 0.1.

Because there is no detectable within-variable correlated error in the nonblack models, and the parameter in the black models is of marginal statistical significance, we are reluctant to accept an estimate as high as 0.4. Our solution is to assume that within-variable error correlation (contamination that occurs across measurement occasion) is no larger than the largest within-occasion error correlation (contamination that occurs at a single occasion). Consequently, in Model G and Model H we fix the within-variable error correlation at 0.2.

In Model G we also eliminate the statistically and substantively insignificant within-occasion correlated errors. What remain are within-occasion correlated errors involving four pairs of variables (see Table 7). Response errors among OCGQ reports of FE and ED and errors among OCGR reports of the same two variables are estimated to be correlated at 0.09. A correlation of 0.12 is estimated among errors in PI and 01 in both the OCGQ and OCGR instruments, and a correlation of 0.15 is estimated among errors in ED and 01 reports in those instruments. Finally, after examining residuals from the correlations implied by the model and experimenting with different error correlations, we estimated a correlation of 0.29 among errors in the OCGQ reports of FO and 01, but <u>not</u> in the OCGR reports. That is, to the degree that Model G accurately represents the pattern of response errors of black respondents, it suggests a tendency for blacks to overstate the consistency between their parental income and first job status, between their educational attainment and first job status, and between their father's and their own educational attainment in both the OCGQ questionnaire and the OCGR telephone reinterview. The model also suggests a tendency for blacks to overstate the consistency of their father's job status and their own first job status in the OCGQ questionnaire, but not in the OCGR interview.

The λ_{ij} slope coefficients are more likely to depart from 1.0 in the models estimated for blacks. Under Model G, only λ_{22} and λ_{52} are estimated to be within one standard error of 1.0. In Model H, these two slopes are constrained to equal 1.0, increasing the chi-square value by only 0.69. Estimates of within-occasion error correlations are essentially the same as those estimated from Model G and are presented in Table 7. While Model H, our final measurement model for blacks, provides a statistically better representation of the pattern of response error than the random error model, the fit is rather poor compared to the successful fit we were able to obtain for nonblacks.[7] Consequently, our interpretations should be considered less definitive than those of the model for nonblacks due to the likelihood of substantial misspecification of our measurement model for blacks.

Estimates of the measurement error parameters for Model H, the final model for blacks, appear in columns 3 through 5 of Table 6 and in Table 7. As with the nonblack model, error standard deviations of the remeasurement interview reports are uniformly smaller than those of the OCG questionnaire reports (column 3 of Table 6). Again, error standard deviations for variables measured in the Duncan SEI metric, FO, O1, and OC, are near 10.0, showing some stability across variables and populations. Since blacks exhibit less total variation on these variables, the same amount of error

TABLE 6 -- Observed moments and measurement model parameter estimates: black males in the experienced civilian labor force, March 1973 (N = 348)

	Variable		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	True	Observed	Mean	Observed Std. Dev.	Std. Dev. of Error[a]	Std. Dev. of True Score	Relative Slope[a]	Reliability Coefficient[b]	Test-Retest Correlations	Coding Reliability	Percent of Cases with Data Present
	T_i	x_{ij}	μ_{ij}	$\sigma_{x_{ij}}$	$\sigma_{e_{ij}}$	σ_{T_i}	λ_{ij}	$(\sigma^2_{T_i}/\sigma^2_{x_{ij}})\lambda^2_{ij}$	$\rho_{x_{i1},x_{i2}}$	$\rho_{x_{i1},x_{i1'}}$	
1.	FO	x_{11}	16.62	13.45	9.97 (.46)	9.02	1.00	.45	.64	.88	93
		x_{12}	17.39	14.75	8.38 (.79)		1.34 (.12)	.68			92
2.	FE	x_{21}	6.65	4.03	1.44 (.10)	3.74	1.00	.86	.92	.98	90
		x_{22}	6.75	3.89	1.10 (.14)		1.00	.92			88
3.	PI	x_{31}	3.42	0.43	0.23 (.02)	0.37	1.00	.74	.84	.98	89
		x_{32}	3.45	0.43	0.13 (.04)		1.12 (.07)	.93			88
4.	ED	x_{41}	10.40	3.69	1.44 (.07)	3.00	1.13 (.04)	.85	.91[c]	.98	94
		x_{42}	10.56	3.32	0.79 (.09)		1.08 (.04)	.95			96
		x_{43}	10.50	3.35	1.50 (.07)		1.00	.80			100[e]
5.	OI	x_{51}	21.14	18.78	10.20 (.60)	16.16	1.00	.74	.77	.93	89
		x_{52}	21.22	19.19	10.09 (.59)		1.00	.71			94
6.	OC	x_{62}	25.77	19.37	10.68 (.69)	18.00	0.90 (.06)	.70	.72[d]	--	100[e]
		x_{63}	26.15	20.74	10.30 (.82)		1.00	.75			100[e]

[a] Approximate standard errors of parameter estimates appear in parentheses.

[b] These coefficients are squared "validity coefficients." They have approximate standard errors on the order of 0.05.

[c] $\rho_{x_{41},x_{43}} = .82$, $\rho_{x_{42},x_{43}} = .87$.

[d] This quantity is $\rho_{x_{62},x_{63}}$, the correlation between SEI scores of reports of March 1973 occupation and Fall 1973 occupation.

[e] Missing values have been allocated for NA cases.

TABLE 7 -- Estimates of nonzero correlations among measurement errors: OCGR subsample of black males in the experienced civilian labor force, March 1973 (N = 348)

Error term		(1) e_{11}	(1) e_{12}	(2) e_{21}	(2) e_{22}	(3) e_{31}	(3) e_{32}	(4) e_{41}	(4) e_{42}	(4) e_{43}	(5) e_{51}	(5) e_{52}	(6) e_{62}	(6) e_{63}
1. FO	e_{11}	--												
	e_{12}	0.20[a]	--											
2. FE	e_{21}	--	--	--										
	e_{22}	--	--	0.20[a]	--									
3. PI	e_{31}	--	--	--	--	--								
	e_{32}	--	--	--	--	0.20[a]	--							
4. ED	e_{41}	--	--	0.09	--	--	--	--						
	e_{42}	--	--	--	0.09	--	--	0.20[a]	--					
	e_{43}	--	--	--	--	--	--	--	--	--				
5. 01	e_{51}	0.29	--	--	--	0.12	--	0.15	--	--	--			
	e_{52}	--	--	--	--	--	0.12	--	0.15	--	0.20[a]	--		
6. OC	e_{62}	--	--	--	--	--	--	--	--	--	--	--	--	
	e_{63}	--	--	--	--	--	--	--	--	--	--	--	--	--

Note: [a]These correlations are specified to be fixed at 0.20.

variation results in lower reliability coefficients. Indeed, blacks exhibit less true variation (column 4) than nonblacks on all variables except educational attainment (ED), and this, together with somewhat higher error variation, results in substantially lower reliabilities for blacks on most reports (compare columns 3, 4, and 6 in Tables 5 and 6).

Different reports of the same variables are more likely to differ in slope coefficients for blacks as compared to nonblacks. OCGR remeasurement interview reports of FO and PI have steeper slopes than the OCGQ questionnaire reports, while the remeasurement interview report of ED is less steep than the questionnaire report, and the CPS report of ED has an even flatter slope. Finally, the remeasurement interview report of current occupational status has a flatter slope than the CPS interview report.

Coding reliability correlations (column 8 of Table 6) are slightly lower on the average for blacks (except for ED). This is probably due to restricted variance among blacks, but for variables in the Duncan SEI metric it may indicate that blacks tend to be in occupations and industries that are more difficult to code or that blacks tend to provide less detail in their responses to the occupation and industry questions.

We have evidence that the structure of response error among blacks is more complex than that for nonblacks in a number of ways. First, while a simple random error structure is adequate to account for nonblack responses, we have been less successful in fitting a structure to the pattern of black responses. Our best-fitting model suggests that there is correlation of response errors among blacks both within and between measurement occasions, and that the variation attributable to measurement errors is larger among blacks. Relative slopes of observed reports on true scores are also more likely to differ across instruments for blacks. Clearly these findings suggest caution in interpreting models of achievement processes among blacks, especially when those estimates take no account of response error. In the following sections we provide some indication of the biases encountered when measurement error is ignored.

Incorporating the Structure of Measurement Error into a Basic Model of the Intergenerational Transmission of Status

In this section we assess the effects of measurement error on the substantive portion of the model for nonblacks and blacks in the full CPS-OCGQ basic file sample. Tables 8 and 9 present observed (uncorrelated) and corrected correlations, means, and standard deviations for 25,223 nonblacks in the full sample; Tables 10 and 11 present the corresponding figure for 2,020 blacks. Corrected moments are obtained by applying measurement model parameters (Model F for nonblacks, Model H for blacks) estimated from the remeasurement samples to the observed moments from the full CPS-OCGQ samples. Comparisons of observed means and standard deviations for the full sample (Tables 8 and 10) with the corresponding quantities in the remeasurement program subsample (Tables 2 and 3) for each racial group reveal no large or systematic biases in the composition of the remeasurement subsample.[8]

Tables 12 and 13 present corrected and uncorrected estimates of structural equations (lines 1, 3, and 6 of each table) and reduced-form equations (lines 1, 2, 4, and 5) for nonblacks; Tables 14 and 15 present corresponding estimates for blacks. Coefficients are presented in both metric (unstandardized) and standardized form. We shall assume that the population values of a standardized coefficient of a background variable (FO, FE, or PI) does not differ enough from zero to be substantively interesting if it is estimated to be less than 0.100.[9]

TABLE 8 -- Uncorrected correlations, means, and standard deviations: CPS-OCG basic file nonblack males in the experienced civilian labor force, March 1973
(N = 25,223)

Variable	1	2	3	4	5	6	7	8
1. x_{11}	--							
2. x_{21}	.537	--						
3. x_{31}	.400	.466	--					
4. x_{43}	.411	.470	.483	--				
5. x_{51}	.392	.330	.293	.636	--			
6. x_{63}	.326	.275	.257	.571	.617	--		
7. AGE	-.174	-.297	-.248	-.210	-.067	.025	--	
8. AGE2	.014	.026	-.027	-.095	-.114	-.142	.144	--
Mean	31.09	8.78	3.77	12.07	33.81	41.11	3.97	16.04
Std.dev.	22.90	4.04	0.42	3.07	24.55	24.91	1.25	14.63

NOTE: See Table 1 for definitions of variables.

TABLE 9 -- Corrected correlations, means, and standard deviations: CPS-OCG basic file nonblack males in the experienced civilian labor force, March 1973
(N = 25,223)

Variable	1	2	3	4	5	6	7	8
1. FO	--							
2. FE	.612	--						
3. PI	.464	.514	--					
4. ED	.475	.516	.539	--				
5. O1	.469	.375	.339	.732	--			
6. OC	.391	.313	.298	.658	.737	--		
7. AGE	-.191	-.309	-.264	-.221	-.073	.027	--	
8. AGE2	.015	.003	-.028	-.100	-.124	-.155	.144	--
Mean	31.09	8.78	3.77	12.07	33.81	41.11	3.97	16.04
Std.dev.	20.90	3.88	0.40	2.91	22.48	22.78	1.25	14.63

NOTE: See Table 1 for definitions of variables. Correlations and standard deviations have been corrected with measurement model parameters estimated from a subsample of 578 observations.

TABLE 10 -- Uncorrected correlations, means, and standard deviations: CPS-OCG basic file black males in the experienced civilian labor force, March 1973
(N = 2020)

Variable	1	2	3	4	5	6	7	8
1. x_{11}	--							
2. x_{21}	.433	--						
3. x_{31}	.302	.384	--					
4. x_{43}	.244	.416	.409	--				
5. x_{51}	.252	.279	.277	.490	--			
6. x_{63}	.225	.284	.278	.500	.546	--		
7. AGE	-.143	-.324	-.230	-.412	-.145	-.109	--	
8. AGE2	.036	.033	-.042	-.077	-.042	-.103	.026	--
Mean	16.92	6.80	3.43	10.42	21.32	25.33	3.81	16.06
Std.dev.	14.53	4.02	0.45	3.37	18.53	20.06	1.25	14.72

NOTE: See Table 1 for definitions of variables.

TABLE 11. -- Corrected correlations, means, and standard deviations: CPS-OCG basic file black males in the experienced civilian labor force, March 1973
(N = 2020)

Variable	1	2	3	4	5	6	7	8
1. FO	--							
2. FE	.638	--						
3. PI	.482	.477	--					
4. ED	.374	.497	.530	--				
5. O1	.228	.358	.339	.655	--			
6. OC	.360	.354	.376	.651	.762	--		
7. AGE	-.196	-.347	-.268	-.460	-.174	-.127	--	
8. AGE2	.049	.035	-.049	-.086	-.050	-.120	.026	--
Mean	16.92	6.80	3.43	10.42	21.32	25.33	3.81	16.06
Std.dev.	10.57	3.75	0.39	3.02	15.47	17.21	1.25	14.72

NOTE: See Table 1 for definitions of variables. Correlations and standard deviations have been corrected with measurement model parameters estimated from a subsample of 348 observations.

TABLE 12 -- Corrected estimates of parameters of the stratification process: nonblack males in the experienced civilian labor force, March 1973
(N = 25,223)

Dependent Variable	Predetermined Variables							R^2	Components of Variation*		
	AGE	AGE2	FO	FE	PI	ED	01		Residual σ_u	Explained $\sigma_{\hat{t}}$	Total σ_t
1. ED	-.034 (-.014)	-.018 (-.092)	.025 (.178)	.175 (.233)	2.42 (.330)	--	--	.395	2.27	1.83	2.91
2. 01	1.54 (.086)	-.212 (-.138)	.381 (.354)	.675 (.117)	7.56 (.134)	--	--	.266	19.26	11.59	22.48
3. 01	1.73 (.096)	-.110 (-.072)	.243 (.226)	-.301 (-.052)	-5.94 (-.105)	5.57 (.722)	--	.581	14.55	17.14	22.48
4. OC	3.35 (.184)	-.283 (-.182)	.314 (.288)	.695 (.118)	8.42 (.147)	--	--	.227	20.03	10.85	22.78
5. OC	3.52 (.193)	-.188 (-.121)	.185 (.170)	-.218 (-.037)	-4.21 (-.073)	5.21 (.667)	--	.496	16.17	16.04	22.78
6. OC	2.65 (.146)	-.132 (-.085)	.063 (.058)	-.067 (-.011)	-1.23 (-.022)	2.42 (.309)	.502 (.495)	.598	14.44	17.62	22.78

NOTE: Standardized coefficients appear in parentheses. Estimates of measurement error variances are based on a subsample of 578 observations.

*Components are expressed as standard deviations. The additive decomposition is $\sigma_t^2 = \sigma_{\hat{t}}^2 + \sigma_u^2$.

TABLE 13 -- Uncorrected estimates of parameters of the stratification process: nonblack males in the experienced civilian labor force, March 1973
(N = 25,223)

Dependent Variable	Predetermined Variables: AGE	AGE2	FO	FE	PI	ED	01	R^2	Components of Variation*: Residual σ_u	Explained $\sigma_{\hat{t}}$	Total σ_t
1. ED	-.058 (-.024)	-.019 (-.092)	.021 (.160)	.183 (.241)	2.18 (.299)	--	--	.337	2.50	1.78	3.07
2. 01	1.48 (.075)	-.217 (-.129)	.296 (.276)	.895 (.147)	7.53 (.129)	--	--	.204	21.90	11.09	24.55
3. 01	1.75 (.089)	-.125 (-.074)	.194 (.181)	.026 (.004)	-2.83 (-.049)	4.76 (.595)	--	.439	18.39	16.27	24.55
4. OC	3.29 (.165)	-.288 (-.169)	.245 (.225)	.888 (.144)	8.06 (.136)	--	--	.176	22.61	10.45	24.91
5. OC	3.55 (.178)	-.202 (-.119)	.150 (.138)	.075 (.012)	-1.63 (-.028)	4.45 (.548)	--	.375	19.69	15.25	24.91
6. OC	2.86 (.143)	-.153 (-.090)	.074 (.068)	.065 (.011)	-0.52 (-.009)	2.58 (.318)	.392 (.387)	.459	18.32	16.88	24.91

NOTE: Standardized coefficients appear in parentheses.

*Components are expressed as standard deviations. The additive decomposition is $\sigma_t^2 = \sigma_{\hat{t}}^2 + \sigma_u^2$.

TABLE 14 -- Corrected estimates of parameters of the stratification process: black males in the experienced civilian labor force, March 1973
(N = 2020)

Dependent Variable	Predetermined Variables: AGE	AGE2	FO	FE	PI	ED	01	R^2	Components of Variation*: Residual σ_u	Explained $\sigma_{\hat{t}}$	Total σ_t
1. ED	-.689 (-.285)	-.015 (-.071)	.003 (.012)	.188 (.234)	2.57 (.333)	--	--	.435	2.27	1.99	3.02
2. 01	-0.32 (-.026)	-.047 (-.045)	-.095 (-.065)	1.17 (.284)	8.92 (.225)	--	--	.170	14.09	6.38	15.47
3. 01	2.19 (.177)	.006 (.005)	-.107 (-.073)	.485 (.118)	-0.45 (-.011)	3.65 (.712)	--	.457	11.40	10.46	15.47
4. OC	0.30 (.022)	-.144 (-.123)	.267 (.164)	.710 (.155)	9.81 (.223)	--	--	.210	15.30	7.89	17.21
5. OC	3.04 (.221)	-.086 (-.074)	.254 (.156)	-.038 (-.008)	-0.39 (-.009)	3.97 (.697)	--	.484	12.36	11.97	17.21
6. OC	1.65 (.120)	-.089 (-.077)	.322 (.198)	-.347 (-.076)	-0.11 (-.002)	1.65 (.290)	.636 (.572)	.662	10.01	14.00	17.21

NOTE: Standardized coefficients appear in parentheses. Estimates of measurement error variances are based on a subsample of 348 observations.

*Components are expressed as standard deviations. The additive decomposition is $\sigma_t^2 = \sigma_{\hat{t}}^2 + \sigma_u^2$.

TABLE 15 -- Uncorrected estimates of parameters of the stratification process: black males in the experienced civilian labor force, March 1973
(N = 2020)

Dependent Variable	Predetermined Variables: AGE	AGE2	FO	FE	PI	ED	01	R^2	Components of Variation*: Residual σ_u	Explained $\sigma_{\hat{t}}$	Total σ_t
1. ED	-.748	-.016	.009	.182	1.84	--	--	.320	2.78	1.91	3.37
	(-.278)	(-.068)	(.038)	(.217)	(.248)						
2. 01	-0.57	-.055	.171	.666	6.95	--	--	.129	17.29	6.66	18.53
	(-.039)	(-.043)	(.134)	(.144)	(.170)						
3. 01	1.29	-.016	.149	.213	2.37	2.49	--	.268	15.85	9.59	18.53
	(.087)	(-.013)	(.117)	(.046)	(.058)	(.454)					
4. OC	0.10	-.143	.137	.893	7.80	--	--	.132	18.69	7.29	20.06
	(.064)	(-.105)	(.099)	(.179)	(.176)						
5. OC	2.23	-.099	.111	.378	2.59	2.84	--	.287	16.93	10.75	20.06
	(.139)	(-.073)	(.081)	(.076)	(.058)	(.476)					
6. OC	1.71	-.093	.052	.292	1.63	1.83	.402	.388	15.69	12.50	20.06
	(.106)	(-.068)	(.037)	(.059)	(.037)	(.308)	(.372)				

NOTE: Standardized coefficients appear in parentheses.

*Components are expressed as standard deviations. The additive decomposition is $\sigma_t^2 = \sigma_{\hat{t}}^2 + \sigma_u^2$.

First we shall examine the corrected estimates for nonblacks in Table 12, obtained by applying least-squares regression to the corrected moments in Table 9. The reduced-form equations (lines 1, 2, and 4) reveal that the background variables FO, FE, and PI affect each aspect of socioeconomic achievement. Together with the age variables, they account for about two-fifths of the variance in educational attainment and about one-fourth of the variance in statuses of first and current occupations of nonblacks. The standardized reduced-form coefficients reveal that parental income (PI) has the strongest relative impact on educational attainment (ED), while father's occupational status (FO) has the largest effect on the two occupational statuses (01 and OC). It appears that the OCG questionnaire item assessing parental income is indeed capturing a dimension of socioeconomic background that contributes to variation in socioeconomic achievements net of the more conventional measures of social origins.

Educational attainment (ED) completely mediates net advantages in occupational status due to FE and PI (compare lines 2 with 3, and lines 4 with 5). That is, educational advantages (or disadvantages) account for the influence of father's education and parental income on a man's occupational standing. In contrast, the effect of father's occupational status on schooling accounts for less than one-half of its influence on the status of son's first or current occupation. The direct influence of father's occupational status (FO) on son's status is about one-fourth of an SEI point for each point of FO in the 01 equation (3) and about one-sixth of a point for each point of FO in the OC equation (5). The effects of a year of schooling are about 5.6 SEI points in status of first job and about 5.2 SEI points in status of 1973 job. Adding educational attainment more than doubles the proportion of variance explained (R^2) in both the 01 and OC equations.

Entering status of first job into the equation for current occupational status reduces the effect of educational attainment on current occupational status by a factor of more than one-half (compare lines 5 and 6). That is, more than one-half of the effect of schooling on current occupational standing reflects the payoff to schooling in selection of the first job, but schooling also directly affects one's standing later in the occupational career. The stability of occupational status is about one-half of one SEI point of current status for each SEI point of first job status. None of the social background factors appears to affect current occupational standing except by way of schooling and first jobs. Overall, background and educational attainment account for about 60 percent of the variance in status of first job and about 50 percent of the variance in status of current job.

Table 13 presents an analogous set of estimated coefficients, which are based on direct application of least squares to the observed full CPS-OCGQ sample moments of Table 8, ignoring response error. First we compare the variation in each dependent variable in Tables 12 and 13. The confounding of measurement error with true variation results in a 5 percent overstatement of the total variation, σ_t, in educational attainment and a 9 percent overstatement of the variation in first and current job status. Residual variation, σ_u, which includes measurement errors in the dependent variables in Table 13, is overestimated by 10 percent in the ED equation and by 13 to 27 percent in the 01 and OC equations. Explained variation in the dependent variables, $\sigma_{\hat{t}}$, is underestimated by 3 to 8 percent in each equation in Table 13. Thus, if we ignore measurement error, we slightly overstate the total amount of socioeconomic inequality and we slightly understate the inequality that is attributable to variation in socioeconomic background and educational attainment. The naive estimates substantially overestimate the amount of unexplained, or conditional, socioeconomic inequality. In all there is a 15 percent underestimate of the proportion of variance explained (R^2) in ED, and a 20 to 24 percent underestimate of the proportion of variance explained in 01 and OC.

The estimated effects of paternal education (FE) are nearly unaffected by correction for measurement error (the uncorrected estimates overstate its reduced-form effects), but there appear to be substantial downward biases in the estimated reduced-form coefficients of the other social background variables. The reduced-form effects of father's occupational status (FO) are underestimated by 16 to 22 percent and those of parental income (PI) are underestimated by about 10 percent in the ED reduced-form equation. Father's occupational status is the only social background variable to have nontrivial effects on first and current job status net of education (lines 3 and 5), and the uncorrected estimates of these effects are about 20 percent lower than the corrected estimates (but the bias disappears when zero restrictions are imposed on the FE and PI coefficients in equations 3 and 5; see appendix Tables 5 and 6).

The uncorrected estimates understate the effect of one year of schooling (ED) on status of first job (01) by 15 percent. The schooling coefficient is biased by about the same amount in the case of current occupational status (line 5 in Tables 12 and 13). In equation 6, the effect of status of first job on current occupational status is underestimated by 22 percent, while the effect of schooling is overestimated by 7 percent.

To summarize our results for nonblack males, ignoring measurement errors result in modest biases (10 to 20 percent) in the reduced-form effects of two of the three background variables--father's occupational status and parental family income. That is, we understate the effects of these two variables on educational attainment and their effects on first and current job status as transmitted by years of schooling.

Though not to the same degree, measurement error also reduces estimated returns to schooling net of social background. Note that the downward bias in the schooling coefficient contributes to the downward bias in the reduced-form effects of background variables. The largest single difference between the corrected and uncorrected structural coefficients involves neither status inheritance nor return to schooling, but is a substantial (22 percent) downward bias in stability of occupational status within the son's career. The other major difference between the corrected and uncorrected models is the overstatement in the latter model of the degree to which variation in socio-economic achievements is <u>not</u> determined by social background and education. After the effects of schooling and social background are taken into account, about one-quarter of the remaining variation in occupational status, which is sometimes ascribed to luck or chance, is actually random response error.

Table 14 gives our corrected estimates of structural coefficients in the stratification model for the full CPS-OCGQ sample of black males, obtained by applying least-squares regression to the corrected moments in Table 11. These results are more tentative than those for nonblacks because of the questionable fit of the measurement model. Furthermore, the full sample estimates for blacks are based upon substantially fewer cases than those for blacks, and consequently they are more susceptible to sampling errors. However, we shall discuss some of the larger and more interesting differences between the structural coefficients for blacks and those for nonblacks (reported in Table 12). First, there is essentially no direct transmission of advantage due to father's occupational status (FO) in the case of educational attainment (ED) or status of first job (O1) among blacks. However, net of education, father's occupational status has more influence upon the black respondent's current occupational status (OC) than upon white respondents occupational status (.254 versus .185 in equation 5 and .322 versus .063 in equation 6).[10] The effect of father's education on status of son's first job is greater among blacks than whites, and this difference persists when the influence of father's on son's schooling is controlled (lines 2 and 3 in Tables 14 and 12). In the case of educational attainment and current occupational status there is greater similarity between the races in the effects of father's education. There is substantial similarity between the races in the effect of parental income on each measure of achievement.

Blacks obtain first jobs whose status is 3.65 SEI points higher for each year of schooling and current jobs whose status is 3.97 points higher for each year of schooling. The effect of educational attainment on status of the first job is 66 percent as large among black as among white men, and the effect of schooling on current occupational status is 76 percent as large (lines 3 and 5 of Tables 14 and 12). At the same time, the stability of occupational status from first to current jobs is 27 percent greater among blacks than among whites. If blacks are more likely to persist in jobs of the same status, they are less likely than whites to gain or lose status after the first job as a result of their schooling. Net of background and the status of first job and the effect of schooling on current occupational status is 68 percent larger among whites than among blacks (line 6).

In the corrected data there is only a small difference in the variability in schooling among black and white men. The estimate of residual variation, σ_u, is the same, 2.27 years, however, the variability in schooling attributable to social background is 9 percent greater among black than among nonblack men, and this is reflected in σ_t, the total variation of schooling. At the same time, none of the components of status of the first or current occupations of black men is as large as 80 percent of the corresponding component of variation among nonblack men. That is, there is substantially less variability in the occupational status of black men than in the status of white men that can be attributed to social background or schooling, and there is substantially less variability in the occupational status of black men conditional on social background or schooling. For example, the variation in status of first job among black men that is explained by social background is 6.38 points on the Duncan scale, or only 55 percent of the corresponding component of variation among nonblack men (see $\sigma_{\hat{t}}$ in line 2 of Tables 14 and 12). Similarly, the variation in first job status that is explained by social background and schooling is only 61 percent as large among black as among nonblack men. These are the two most extreme comparisons between the races, and in other cases the components of variation are 70 to 75 percent as large among black as among nonblack men.

While there is less variation in occupational status among black than among white men, and while black occupational attainments are less dependent upon social background than are the attainments of whites, black men are also less able to translate the advantages of additional schooling into higher occupational attainments. Relative to whites, black men live under a perverse regime of equality of opportunity and of results in the world of work. The constraining influence of social background is not as great among blacks as among whites, but neither are educational attainments as easily translated into occupational status, and the range of job opportunities for men of equal background and schooling is less in the black than in the nonblack population.

Table 15 gives uncorrected estimates of parameters of the achievement process in the OCG sample of black men in the experienced civilian labor force. The consequences of ignoring measurement error appear to be greater in the case of black than in the case of nonblack men. For example, there is a downward bias of about 30 percent in the effect of schooling on the status of first and of current occupation (compare line 3 and line 5 of Table 14 with the corresponding lines in Table 15). Intragenerational stability of occupational status is underestimated by 37 percent in Table 15 (line 6).

In the three reduced-form equations (lines 1, 2, and 4) the uncorrected effects of parental income are about 20 to 30 percent lower than the corrected estimates. There is essentially no difference in the effect of father's education on son's education in the corrected and uncorrected equations, however, the effect of father's education on the status of first job is substantially understated in the uncorrected equations, and the effect of father's education is substantially overstated in the uncorrected equations for current occupational status. The pattern is the opposite in the case of father's occupational status. The corrected and uncorrected effects of father's occupational status on son's educational attainment are both virtually zero, but the uncorrected estimates overstate the influence of father's occupational standing on son's first occupation and understate its influence on the status of son's current occupation. These sharp changes are attributable to within-occasion correlated error in the measurement model for black men.

Measurement error variation is larger relative to true variation among black men. Consequently, the uncorrected measures of variation substantially overstate the amount of inequality in the dependent variables, and especially the component of variation that is conditional upon social background or schooling. For example, in the structural equations of the model (lines 1, 3, and 6 of Tables 14 and 15), the residual variation, σ_u, in the uncorrected data is overestimated by 22 percent in the case of educational attainment, 39 percent for status of first job, and 57 percent for status of current occupation. In the uncorrected model, we underestimate the explained variation, $\sigma_{\hat{t}}$, in each dependent measure by 4 to 10 percent (except in the reduced-form equation for status of first occupation). As a consequence of the upward bias in the residual variation and the downward bias in the explained variation when measurement errors are ignored, in the black sample the proportions of variance explained (R^2) are substantially lower in the uncorrected than in the corrected estimates.

It is not necessary to describe in detail uncorrected comparisons between the black and nonblack models of the stratification process, since these comparisons are implicit in the preceding discussion. Since the biases in structural and reduced-form coefficients are larger among black than among nonblack men, the uncorrected racial comparisons show unrealistically large differences between the races in the effects of social background and schooling. At the same time, the larger error variation among black responses leads to an understatement of racial differences in total and conditional variation in occupational attainment.

To summarize our results for black males, the pattern of apparent biases is similar to that of nonblacks, but the magnitude of biases are substantially greater. Uncorrected estimates of several reduced-form effects of background variables are 22 to 49 percent lower than the corrected estimates. Apparent biases in the transmission of occupational status from father to son, net of educational attainment, are even greater. Uncorrected estimates of occupational returns to schooling are about 30 percent of the corrected estimates. As we found for nonblacks, residual variation in achievement variables, inequality not attributable to variation in background characteristics, is consistently overestimated when measurement error is ignored, by 22 to 57 percent for blacks. Because biases are greater among blacks, ignoring measurement error exaggerates the advantages of nonblacks in converting educational attainments into occupational achievements and underestimates the degree to which there is less variation among blacks in occupational attainments independent of social origins than among nonblacks.[11]

Conclusions: Measurement Errors in Models of the Intergenerational Transmission of Socioeconomic Status

Several sociologists and economists have noted possible biases in effects of social background and schooling when intergenerational models of the stratification process are based on retrospective survey reports of status variables. The prevailing view has been that effects of social background are biased downward by errors in retrospective reports. Consequently, effects of schooling are biased upward, at least relative to those of social background. But research on these biases has been inconclusive because appropriate data and statistical models have not been available. Using data from the remeasurement program of the 1973 Occupational Changes in a Generation-II Survey, we have overcome some of these shortcomings by estimating and testing comprehensive structural models that incorporate both random and nonrandom response errors.

We think there is persuasive evidence that reports of social background and achievement variables by nonblacks are subject only to random response error. Moreover, we find no evidence that social background variables are measured substantially less reliably than contemporaneous achievement variables among nonblack men. Contrary to some previous expectations, response error leads to downward biases in estimated returns to schooling, and for nonblack men downward biases in estimated effects of social background variables are neither pervasive nor very large. Ignoring response error, we underestimate occupational returns of nonblack men by about 15 percent and the effects of father's occupational status and parental income on son's status by as much as 22 percent. Yet downward biases in estimated effects of father's educational attainment are negligible. Measurement error does have a substantial effect on estimates of status persistence within the occupational career. Also, by ignoring response errors among nonblack men, we overstate the total amount of variation in achievement variables that is independent of social background by 10 to 27 percent.

Among black men there are substantial departures from randomness in errors of reports about status variables. While we are not convinced that our final measurement model for black men is correct, we do find evidence suggesting contamination in the responses of blacks both within and across measurement occasions; moreover, error variation in responses of black men is estimated to be greater than among nonblacks. Consequently, when we compare corrected and uncorrected estimates of stratification models among black men, we find biases that are substantially larger than those for non-black men. Because of the questionable fit of our final measurement model for blacks, our assessment of these biases must be regarded as tentative. Occupational returns to schooling appear to be biased downward by about 30 percent, and bias appears to be even larger in the uncorrected estimate of intragenerational stability of occupational status among blacks. Because of the differing structures of response error among black and nonblack men, ignoring those structures leads to an exaggeration of black-nonblack differences in occupational returns to schooling and to an understatement of racial differences in total and conditional inequality of occupational attainment.

What do our results suggest about the intergenerational transmission of socioeconomic inequality in the United States? They demonstrate that by ignoring measurement error we have been systematically <u>underestimating</u> the degree to which schooling is converted into occupational successes, by about 15 percent for nonblacks, and probably by much more than that for blacks. However, there are two social forces generating the distribution of schooling: circumstances of birth and "meritocratic" sources independent of social origins. In our models that ignore measurement error, we have been overestimating the contribution of the second force by at least as much as we have been underestimating the contribution of the first source. While previous writers in the debate about the intergenerational transmission of socioeconomic status and the impact of measurement error bias have been somewhat negligent in specifying exactly which parameters of the stratification process are important and how much bias in these parameters can be called "substantial," it appears that our results lend conclusive evidence neither to those who have argued that the effects of response errors are trivial, nor to those who have argued that the effects are substantial. If nothing else, our results have removed the debate from the realm of speculation and hypothetical data toward the realm of empirical evidence.

Finally, we have--especially for nonblacks--made available for the first time a set of parameters that characterize the measurement of six socio-economic variables when specific measuring instruments are applied to specific populations. However, a cautionary note is in order. Our data were collected as part of a carefully designed and instrumented study that uses the resources, personnel, and procedures of the U.S. Bureau of the Census. It may be inappropriate to apply our estimates of measurement parameters to data obtained using instruments and procedures that differ from those of the OCG-II Survey. Indeed, within this survey and for a given population, nonblack males ages 20 to 65 in the experienced civilian labor force of March 1973, we have estimated reliability coefficients for our three measures of educational attainment, (OCGQ, CPS, and OCGR) as varied as .70, .89, and .96. The coefficients for educational attainment estimated by Siegel and Hodge [1968] have certainly been applied to data sets employing instruments to measure education, which are considerably more diverse than the three instruments used in the OCG-II Survey. We hope that our results make clear the need for careful consideration and restraint in the "borrowing" of measurement model parameters.

APPENDIX

ALTERNATIVE ESTIMATES OF SUBSTANTIVE PARAMETERS

TABLE A1 -- Corrected estimates of parameters of the stratification process: nonblack males in the experienced civilian labor force, March 1973
(Remeasurement Subsample, N = 578)

Dependent Variable	Predetermined Variables							R^2	Components of Variation*		
	AGE	AGE2	FO	FE	PI	ED	O1		Residual σ_u	Explained $\sigma_{\hat{t}}$	Total σ_t
1. ED	.008 (.004)	-.016 (-.082)	.021 (.174)	.168 (.250)	2.25 (.317)	--	--	.387	2.12	1.69	2.71
2. O1	1.86 (.099)	-.262 (-.160)	.366 (.365)	.215 (.038)	13.1 (.222)	--	--	.292	18.91	12.14	22.47
3. O1	1.81 (.096)	-.173 (-.105)	.249 (.248)	-.724 (-.130)	0.53 (.009)	5.59 (.673)	--	.570	14.73	16.96	22.47
4. OC	3.96 (.204)	-.268 (-.159)	.348 (.337)	.549 (.096)	10.8 (.178)	--	--	.272	19.72	12.05	23.11
5. OC	3.92 (.201)	-.182 (-.107)	.236 (.228)	-.352 (-.061)	-1.24 (-.021)	5.37 (.627)	--	.513	16.12	16.55	23.11
6. OC	3.04 (.156)	-.098 (-.058)	.115 (.111)	-.002 (-.000)	-1.50 (-.025)	2.66 (.311)	.484 (.470)	.608	14.47	18.02	23.11

NOTE: Standardized coefficients appear in parentheses.

*Components are expressed as standard deviations. The additive decomposition is $\sigma_t^2 = \sigma_{\hat{t}}^2 + \sigma_u^2$

TABLE A2 -- Uncorrected estimates of parameters of the stratification process: nonblack males in the experienced civilian labor force, March 1973
(Remeasurement Subsample, N = 578)

Dependent Variable	Predetermined Variables							R^2	Components of Variation*		
	AGE	AGE2	FO	FE	PI	ED	01		Residual σ_u	Explained $\sigma_{\hat{t}}$	Total σ_t
1. ED	-.027	-.016	.018	.164	2.13	--	--	.329	2.35	1.65	2.87
	(-.011)	(-.077)	(.152)	(.239)	(.304)						
2. 01	1.75	-.232	.294	.224	14.9	--	--	.228	21.71	11.80	24.71
	(.084)	(-.129)	(.289)	(.038)	(.248)						
3. 01	1.89	-.152	.206	-.586	4.40	4.95	--	.450	18.33	16.58	24.71
	(.091)	(-.084)	(.202)	(-.099)	(.073)	(.575)					
4. OC	3.83	-.281	.261	.710	11.4	--	--	.208	22.43	11.50	25.21
	(.181)	(-.153)	(.251)	(.118)	(.185)						
5. OC	3.95	-.209	.181	-.017	1.93	4.45	--	.380	19.85	15.54	25.21
	(.186)	(-.114)	(.175)	(-.003)	(.031)	(.506)					
6. OC	3.21	-.150	.101	.213	0.20	2.50	.393	.462	18.49	17.14	25.21
	(.151)	(-.081)	(.097)	(.035)	(.003)	(.285)	(.385)				

NOTE: Standardized coefficients appear in parentheses.

*Components are expressed as standard deviations. The additive decomposition is $\sigma_t^2 = \sigma_{\hat{t}}^2 + \sigma_t^2$.

TABLE A3 -- Corrected estimates of parameters of the stratification process: black males in the experienced civilian labor force, March 1973
(Remeasurement Subsample, N = 348)

Dependent Variable	Predetermined Variables								Components of Variation*		
	AGE	AGE2	FO	FE	PI	ED	01	R^2	Residual σ_u	Explained $\sigma_{\hat{t}}$	Total σ_t
1. ED	-.532 (-.236)	.003 (.016)	-.032 (-.097)	.242 (.302)	2.66 (.328)	--	--	.339	2.44	1.75	3.00
2. 01	0.61 (.050)	-.021 (-.020)	.353 (.197)	.490 (.113)	10.2 (.234)	--	--	.190	14.54	7.04	16.16
3. 01	2.31 (.190)	-.031 (-.030)	.456 (.255)	-.279 (-.065)	1.75 (.040)	3.18 (.591)	--	.421	12.30	10.49	16.16
4. OC	1.00 (.074)	-.156 (-.137)	.446 (.223)	.943 (.196)	9.73 (.200)	--	--	.261	15.47	9.20	18.00
5. OC	3.03 (.224)	-.168 (-.147)	.569 (.285)	.020 (.004)	-0.41 (-.008)	3.81 (.636)	--	.528	12.37	13.08	18.00
6. OC	1.86 (.137)	-.152 (-.133)	.337 (.169)	.162 (.034)	-1.30 (-.027)	2.19 (.365)	.510 (.458)	.649	10.66	14.50	18.00

NOTE: Standardized coefficients appear in parentheses.

*Components are expressed as standard deviations. The additive decomposition is $\sigma_t^2 = \sigma_{\hat{t}}^2 + \sigma_u^2$.

TABLE A4 -- Uncorrected estimates of parameters of the stratification process: black males in the experienced civilian labor force, March 1973
(Remeasurement Subsample, N = 348)

Dependent Variable	Predetermined Variables								Components of Variation*		
	AGE	AGE2	FO	FE	PI	ED	O1	R^2	Residual σ_u	Explained $\sigma_{\hat{t}}$	Total σ_t
1. ED	-.553	.001	-.007	.216	2.12	--	--	.277	2.85	1.76	3.35
	(-.219)	(.005)	(-.029)	(.260)	(.272)						
2. O1	0.55	-.023	.277	.673	8.10	--	--	.147	17.34	7.20	18.78
	(.039)	(-.019)	(.198)	(.144)	(.185)						
3. O1	1.95	-.026	.295	.127	2.76	2.52	--	.293	15.79	10.17	18.78
	(.138)	(-.021)	(.211)	(.027)	(.063)	(.450)					
4. OC	0.84	-.152	.028	1.47	8.39	--	--	.155	19.07	8.16	20.74
	(.054)	(-.116)	(.018)	(.286)	(.174)						
5. OC	2.47	-.156	.049	.832	2.15	2.95	--	.319	17.12	11.71	20.74
	(.159)	(-.118)	(.032)	(.162)	(.045)	(.476)					
6. OC	1.77	-.147	-.057	.787	1.16	2.04	3.61	.395	16.13	13.03	20.74
	(.114)	(-.111)	(-.037)	(.153)	(.024)	(.329)	(.327)				

NOTE: Standardized coefficients appear in parentheses.

*Components are expressed as standard deviations. The additive decomposition is $\sigma_t^2 = \sigma_{\hat{t}}^2 + \sigma_u^2$.

TABLE A5 -- Corrected estimates of parameters of the stratification process (subject to zero restrictions): nonblack males in the experienced civilian labor force, March 1973

(N = 25,223)

Dependent Variable	Predetermined Variables							R^2	Components of Variation*		
	AGE	AGE2	FO	FE	PI	ED	01		Residual σ_u	Explained $\sigma_{\hat{t}}$	Total σ_t
1. ED	-.034 (-.014)	-.018 (-.092)	.025 (.178)	.175 (.233)	2.42 (.330)	--	--	.395	2.27	1.83	2.91
2. 01	1.96 (.110)	-.211 (-.138)	.318 (.295)	.901 (.155)	12.5 (.220)	--	--	.303	18.77	12.37	22.48
3. 01	2.14 (.119)	-.118 (-.077)	.189 (.176)	--	--	5.15 (.667)	--	.572	14.71	17.00	22.48
4. OC	3.65 (.201)	-.282 (-.182)	.270 (.247)	.859 (.146)	11.9 (.207)	--	--	.253	19.69	11.46	22.78
5. OC	3.82 (.209)	-.194 (-.124)	.147 (.135)	--	--	4.91 (.628)	--	.491	16.25	15.96	22.78
6. OC	2.73 (.150)	-.134 (-.086)	.051 (.047)	--	--	2.30 (.294)	.507 (.500)	.598	14.44	17.62	22.78

NOTE: Standardized coefficients appear in parentheses. Estimates of measurement error variances are based on a subsample of 578 observations.

*Components are expressed as standard deviations. The additive decomposition is $\sigma_t^2 = \sigma_{\hat{t}}^2 + \sigma_u^2$.

TABLE A6 -- Uncorrected estimates of parameters of the stratification process (subject to zero restrictions): nonblack males in the experienced civilian labor force, March 1973

(N = 25,223)

Dependent Variable	Predetermined Variables: AGE	AGE2	FO	FE	PI	ED	01	R^2	Components of Variation*: Residual σ_u	Explained $\sigma_{\hat{t}}$	Total σ_t
1. ED	-.058	-.019	.021	.183	2.18	--	--	.337	2.50	1.78	3.07
	(-.024)	(-.092)	(.160)	(.241)	(.299)						
2. 01	1.60	-.214	.281	.845	10.1	--	--	.216	21.74	11.41	24.55
	(.081)	(-.128)	(.264)	(.139)	(.173)						
3. 01	1.87	-.126	.184	--	--	4.62	--	.437	18.42	16.23	24.55
	(.095)	(-.075)	(.172)			(.578)					
4. OC	3.32	-.285	.241	.804	9.58	--	--	.180	22.56	10.57	24.91
	(.167)	(-.168)	(.223)	(.131)	(.162)						
5. OC	3.58	-.201	.148	--	--	4.39	--	.375	19.69	15.25	24.91
	(.179)	(-.118)	(.136)			(.541)					
6. OC	2.84	-.152	.076	--	--	2.58	.393	.459	18.32	16.88	24.91
	(.143)	(-.089)	(.070)			(.318)	(.387)				

NOTE: Standardized coefficients appear in parentheses.

*Components are expressed as standard deviations. The additive decomposition is $\sigma_t^2 = \sigma_{\hat{t}}^2 + \sigma_u^2$.

TABLE A7 -- Corrected estimates of parameters of the stratification process (subject to zero restrictions): black males in the experienced civilian labor force, March 1973

(N = 2020)

Dependent Variable	Predetermined Variables								Components of Variation*		
	AGE	AGE2	FO	FE	PI	ED	01	R^2	Residual σ_u	Explained $\sigma_{\hat{t}}$	Total σ_t
1. ED	-.688 (-.285)	-.014 (-.071)	--	.194 (.241)	2.59 (.335)	--	--	.434	2.27	1.99	3.02
2. 01	-.340 (-.027)	-.047 (-.047)	--	.984 (.239)	9.30 (.234)	--	--	.176	14.04	6.49	15.47
3. 01	2.13 (.172)	.003 (.003)	--	.288 (.070)	--	3.59 (.700)	--	.453	11.44	10.41	15.47
4. OC	.506 (.037)	-.145 (-.125)	.257 (.158)	.923 (.201)	9.91 (.225)	--	--	.230	15.10	8.25	17.21
5. OC	3.14 (.228)	-.091 (-.077)	.257 (.158)	.181 (.040)	--	3.82 (.671)	--	.485	12.35	11.99	17.21
6. OC	1.80 (.131)	-.093 (-.079)	.257 (.158)	---	--	1.57 (.276)	.628 (.564)	.659	10.05	13.97	17.21

NOTE: Standardized coefficients appear in parentheses. Estimates of measurement error variances are based on a subsample of 348 observations.

*Components are expressed as standard deviations. The additive decomposition is $\sigma_t^2 = \sigma_{\hat{t}}^2 + \sigma_u^2$.

TABLE A8 -- Uncorrected estimates of parameters of the stratification process (subject to zero restrictions): black males in the experienced civilian labor force, March 1973

(N = 2020)

Dependent Variable	Predetermined Variables							R^2	Components of Variation*		
	AGE	AGE2	FO	FE	PI	ED	01		Residual σ_u	Explained $\sigma_{\hat{t}}$	Total σ_t
1. ED	-.747	-.015	--	.194	1.88	--	--	.319	2.78	1.90	3.37
	(-.277)	(-.067)		(.231)	(.254)						
2. 01	-0.87	-.055	.185	.518	5.02	--	--	.104	17.54	5.98	18.53
	(-.058)	(-.044)	(.145)	(.112)	(.123)						
3. 01	1.12	-.015	.185	--	--	2.67	--	.264	15.90	9.52	18.53
	(.076)	(-.012)	(.145)			(.485)					
4. OC	-0.08	-.144	.076	.990	5.62	--	--	.107	18.96	6.56	20.06
	(-.005)	(-.106)	(.055)	(.198)	(.128)						
5. OC	2.15	-.099	.076	.410	--	2.99	--	.279	17.03	10.60	20.06
	(.134)	(-.072)	(.055)	(.082)		(.502)					
6. OC	1.69	-.093	--	.410	--	1.89	.411	.385	15.73	12.45	20.06
	(.105)	(-.068)		(.082)		(.318)	(.380)				

NOTE: Standardized coefficients appear in parentheses.

*Components are expressed as standard deviations. The additive decomposition is $\sigma_t^2 = \sigma_{\hat{t}}^2 + \sigma_u^2$.

NOTES

[1]The OCG parental income item states: "When you were about 16 years old, what was your family's annual income?" The fourteen possible responses were:

No income (or loss),
$1-499,
$500-999,
$1,000-1,999,
$2,000-2,999,
$3,000-3,999,
$4,000-4,999,
$5,000-5,999,
$6,000-6,999,
$7,000-7,999,
$8,000-8,999,
$9,000-9,999,
$10,000-14,999,
$15,000 or more.

After examining plots of occupational status of first and current job and educational attainment by parental income category we determined that a logarithmic function of parental income was the appropriate functional form relating it to the achievement variables. The first two categories were collapsed and midpoints of intervals were used. A value of $19,750 was assigned to the open-ended category on the basis of a canonical analysis with ED, 01 and OC as criterion variables. Responses to pretest probes and plots of achievement variables by parental income categories by ten-year age cohorts clearly indicated that respondents tended not to adjust their responses to current dollars. Therefore, the dollar midpoint responses were adjusted by a four-year moving average of the Consumer Price Index, with the four years weighted to reflect the uncertainty in determining the exact year of birth from age in March 1973. The final scale was computed as the logarithm (base 10) of the price adjusted dollar category midpoints. Our scaling procedure explicitly attempted to maximize correlations between parental income and statuses of the respondent. As a consequence, intergenerational (father-son) correlations between PI and ED are larger than intragenerational (father's generation) correlations between PI and both FO and FE (Tables 8 through 11).

[2]Figure 2 shows the most general (least restricted) model that we estimated for each racial group. Ultimately, we eliminated some of the correlations among reporting errors.

[3]Another way of stating this normalization is that only the ratio of the slopes is identifiable. A more common normalization is to assume unit variances of true scores. However, this normalization does not allow the computation of metric coefficients relating unobservables. Error

variances and reliabilities (squared true score-observed correlations) are invariant with respect to normalization, although true score variances (and structural coefficients) do depend on which λ_{ij} are fixed to unity.

[4]Again we have an indeterminancy in the slope of the conditional expectation function of the observed score given true score, and we assume that the measures included in the full sample models define the true score metrics. That is, in our models for the full CPS-OCGQ samples we assume all such slopes to be unity. Since all of our metrics, except perhaps that of educational attainment, are to some degree arbitrary, it seems reasonable to normalize by taking the observed metrics as the standard. While our findings do suggest some relative differences in slope coefficients correspond to "true" metrics.

[5]Since the mean vector is not restricted by the model, the sample means provide the maximum likelihood estimates of the true score means.

[6]The Bureau of the Census uses the "hot deck" technique to allocate nonresponses in CPS reports of education and occupation, and we treat these allocations as responses. Allocated nonresponses are assigned the observed value of the last case processed with the same age, sex, and race. Thus, allocated responses have both systematic and random components. Elsewhere, of course, we assume that the pairwise correlations accurately represent the correlations that would have been obtained were complete data available. While this is an untestable assumption, the alternatives are more problematic. Replacement with means restricts variances and would result in underestimates of error variances. Random allocation would reduce the ability to detect nonrandom response error structure, while systematic allocation would have the opposite effect. Omitting all cases that have missing data would reduce the sample size by about 40 percent and probably eliminate many of the cases with less accurate responses. Models of the achievement process are almost always estimated from pairwise present correlations, and it is the response error structure in these analyses that we are attempting to assess.

[7]There are factors mitigating the lack of fit among blacks in our further application of Model H. First, the OCG samples are less efficient than simple random samples, but we have treated (weighted) observations as if we had a simple random sample. The appropriate design factor may be as small as .75, in which case we would not reject Model H at the .05 level. Second, when correlations are computed among blacks for whom data are present on all thirteen measured variables, the fit of measurement models improves substantially. Model A, the random error model, fits quite well for the "listwise" black sample ($\chi^2 = 43.97$ with 50 df; $p = .713$). Nevertheless, the proportionate reduction in chi-square upon entering within-occasion and within-instrument correlated error (Model D) is nearly the same as for the "pairwise" black sample ($\chi^2 = 23.88$ with 34 df; $p = .902$), and restricting the black sample to cases with no missing data reduces the number of cases by 46 percent.

[8] In the black remeasurement subsamples, variances of two of the socioeconomic background variables, FO and PI, are restricted relative to corresponding variances in the full (basic file) sample of black respondents. While this may suggest that selection of black remeasurement cases is biased toward those subject to less error, comparisons of correlations between the black subsample and full sample suggest just the opposite. Correlations involving background variables are generally lower in the remeasurement subsample. The apparent complexity of the measurement error structure for blacks precludes a more definitive assessment of selection bias in the remeasurement subsample.

[9] Standard errors of the corrected estimates cannot be computed, because estimates are based upon both the full CPS-OCGQ sample and the OCGR subsample. The standard errors computed by least-squares regressions for the uncorrected estimates are inappropriate because of the misspecification of the uncorrected models. For the nonblack model, we have been able to use the LISREL program of Jöreskog and van Thillo [1972] to estimate structural and measurement parameters within the OCGR subsample. Statistically, we do not reject the null hypothesis that the negligible coefficients are all zero (constraining to zero the four coefficients for FE and PI in the O1 and OC structural equations increases the chi-square value by 7.8; $p > .05$). Unfortunately, the more complex error structure in the model for blacks precluded computation of a similar statistical test for that model.

Corrected and uncorrected estimates based entirely upon the remeasurement program subsamples of nonblacks ($N = 578$) and blacks ($N = 348$) appear in appendix Tables A1 through A4. Comparing estimates from these subsample tables to those from corresponding CPS-OCGQ full sample Tables 12 through 15 reveal few differences. For nonblacks (Tables A1, A2, 12 and 13), the apparent biases due to measurement error are nearly identical in the two samples. The few large negative effects of background variables estimated in the full CPS-OCGQ sample (e.g., the effect of PI in line 3 of Table 12), are not evident in the subsample estimates, and conversely, the large negative effects of background variables estimated in the subsample (e.g., the effect of FE in line 3 of Table A1) are not evident in the larger sample, supporting our assumption that such negative effects are not substantially different from zero. The subsample and full sample estimates for blacks (Tables A3, A4, 14, and 15) are based upon fewer cases and are therefore more subject to sampling variability. In the corrected estimates for the black subsample we detect effects of father's occupational status upon status of first job that do not appear in the full sample estimates (lines 2 and 3 in Table A3 and 15). Also, apparent biases due to measurement error in the education coefficients and in the residual variation of ED and O1 for blacks are slightly larger in the full sample computations than in the subsample.

Corrected and uncorrected estimates with negligible effects of background variables constrained to equal zero appear in appendix Tables A5 and A6 for nonblacks, A7 and A8 for blacks (based upon the full CPS-OCGQ samples). Estimates of the structural equations were obtained from least-squares regression applied to the uncorrected and corrected moments; reduced for coefficients were obtained algebraically from structural equations. Imposing the constraints has little effect on the estimates except to reduce the apparent bias due to measurement error in the education coefficients (from 15 percent bias to 10 or 11 percent bias for nonblacks, from about 30 percent bias to 21 to 26 percent bias for blacks). The constrained estimates for nonblacks are discussed in detail by Bielby, Hauser, and Featherman [1976]. The estimates subject to zero restrictions, are not discussed in the text, since doing so might confound black-nonblack comparisons in the stratification process with the different zero restrictions imposed for the two racial groups.

[10]It should be recalled that we estimated a substantial correlation (about 0.3) between response errors in OCG reports of FO and O1 among black men, suggesting a tendency of respondents to overstate the consistency of the status of first job and of father's occupation. Correcting for this tendency causes the (uncorrected) effect of FO on O1 to disappear and also accounts for the persisting effect of FO on OC when O1 is introduced into the corrected OC equation. However, the observed correlation between father's occupational status and first job status among blacks is 20 percent higher in the remeasurement subsample than in the full CPS-OCGQ sample (.295 versus .252). We may be overestimating the amount of error correlation in the full sample, and consequently underestimating the net effect of FO on O1. Note that within the black remeasurement subsample (appendix Tables A3 and A4), FO has substantial net effects of O1 in both the corrected and uncorrected models. It should also be noted that the full black CPS-OCGQ basic file sample is less than one-tenth the size of the nonblack sample, consequently, there is considerable sampling error in the estimates discussed here.

[11]Components of mean racial differences in socioeconomic achievements are often analyzed with the technique of indirect standardization where means for blacks on predetermined variables are substituted into the equations for nonblacks [Duncan, 1969; Featherman and Hauser, 1974]. While there are conceptual reasons for standardizing this way instead of substituting nonblack means into the black equations, our results suggest a methodological reason as well: The coefficients of the nonblack equations are probably less subject to biases due to measurement error.

REFERENCES

Bielby, William T., Hauser, Robert M. and Featherman, David L.
1976 "Response errors of nonblack males in models of the stratification process." Forthcoming in Latent Variables in Socioeconomic Models, edited by Dennis J. Aigner and Arthur S. Goldberger. Amsterdam: North-Holland.

Blau, Peter M., and Duncan, Otis Dudley.
1967 The American Occupational Structure. New York: Wiley.

Bohrnstedt, George W.
1970 "Reliability and validity assessment in attitude measurement." In Attitude Measurement, edited by Gene F. Summers, pp. 80-99. Chicago: Rand McNally.

Bowles, Samuel.
1972 "Schooling and inequality from generation to generation." Journal of Political Economy 80 (May/June): S219-S251.

Bowles, Samuel, and Gintis, Herbert.
1976 Schooling in Capitalist America. New York: Basic Books.

Bowles, Samuel, and Nelson, Valerie I.
1974 "The 'inheritance of IQ' and the intergenerational reproduction of economic inequality." Review of Economics and Statistics 56 (February): 39-51.

Campbell, Donald T., Seigman, Carole R., and Rees, Matilda B.
1967 "Direction-of-wording effects in the relationships between scales." Psychological Bulletin 68 (November): 292-303.

Couch, Arthur, and Keniston, Kenneth.
1960 "Yeasayers and naysayers: agreeing response set as a personality variable." Journal of Abnormal and Social Psychology 60 (March): 151-174.

Duncan, Otis Dudley,
1961 "A socioeconomic index for all occupations." In Occupations and Social Status, edited by Albert J. Reiss, Jr., pp. 109-138. New York: Free Press.

1969 "Inheritance of poverty or inheritance of race?" In On Understanding Poverty, edited by Daniel P. Moynihan, pp. 85-110. New York: Basic Books.

Duncan, Otis Dudley, Featherman, David L. and Duncan, Beverly.
1972 Socioeconomic Background and Achievement. New York: Seminar Press.

Featherman, David L.
1973 "Comments on models for the socioeconomic career." American Sociological Review 38 (December): 785-791.

Featherman, David L., and Hauser, Robert M.
1973 "On the measurement of occupation in social surveys." Sociological Methods and Research 2 (November): 239-251.

1974 "Socioeconomic achievements of U.S. men, 1962 to 1972." Science 185 (July): 325-331.

1975 "Design for a replicate study of social mobility in the United States." In Social Indicator Models, edited by Kenneth C. Land and Seymour Spilerman, pp. 219-252. New York: Russell Sage.

Jencks, Christopher; Smith, Marshall; Acland, Henry; Bane, Mary Jo; Cohen, David; Gintis, Herbert; Heyns, Barbara; and Michelson, Stephan.
1972 Inequality: A Reassessment of the Effect of Family and Schooling in America. New York: Basic Books.

Jöreskog, Karl G.
1970 "A general method for the analysis of covariance structures." Biometrika 57 (August): 230-251.

Jöreskog, Karl G., Gruvaeus, Gunnar T; and van Thillo, Marielle.
1970 "ACOVS: a general computer program for analysis of covariance structures." Research Bulletin 70-15, Educational Testing Service, Princeton, New Jersey.

1972 "LISREL: a general computer program for estimating a linear structural equation system involving multiple indicators of unmeasured variables." Research Bulletin 72-56, Educational Testing Service, Princeton, New Jersey.

Kalleberg, Arne L.
1974 "A causal approach to the measurement of job satisfaction." Social Science Research 3 (December): 299-322.

Kelley, Jonathan.
1973 "Causal chain models for the socioeconomic career." American Sociological Review 38 (August): 481-493.

Mason, William M.; Hauser, Robert M; Kerckhoff, Alan C; Poss, Sharon Sanomirsky; and Manton, Kenneth.
1976 "Models of response error in student reports of parental socio-economic characteristics." In Schooling and Achievement in American Society, edited by William H. Sewell, Robert M. Hauser, and David L. Featherman, pp. 443-494. New York: Academic Press.

Sewell, William H., and Hauser, Robert M.
1975 Education, Occupation and Earnings: Achievement in the Early Career. New York: Academic Press.

Siegel, Paul M., and Hodge, Robert W.
1968 "A causal approach to the study of measurement error." In Methodology in Social Research, edited by Hubert M. Blalock and Ann B. Blalock, pp. 28-59. New York: McGraw-Hill.

Treiman, Donald J., and Hauser, Robert M.
1976 "The intergenerational transmission of income." Forthcoming in The Process of Stratification: Trends and Analyses, edited by Robert M. Hauser and David L. Featherman. New York: Academic Press.

Wiley, David E., and Wiley, James A.
1970 "The estimation of measurement error in panel data." American Sociological Review 35 (February): 112-117.