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# Schooling and Inequality from Generation to Generation

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## I. Introduction

For at least half a century schooling has been the chosen instrument of American social reformers. More and better schooling has been seen as an antidote to the brutalization of industrial life. More equal access to schooling has been sought as a powerful vehicle for the equalization of economic opportunity, the redistribution of income, and the elimination of poverty.

Until recently, the choice of education as the instrument of those who sought greater equality in the United States has not been based on any direct evidence of its efficacy in bringing higher incomes to the children of the poor. Rather, the popularity of educational reform among liberals and progressives stemmed from more political considerations: educational equalization seemed to offer a strategy for achieving the greater social equality that was politically viable. More equal education, it was confidently asserted, could achieve significantly greater equality of economic opportunity and incomes without challenging the basic economic institutions of society and without requiring any major redistribution of capital.

Yet over the past decade, important empirical support has been forthcoming for those who see education as—to quote Horace Mann—“the great equalizer.” First, the *possibility* of more equal schooling achieving a more equal distribution of income seemed to be confirmed by studies of the determinants of individual earnings.<sup>1</sup> The earnings functions estimated in these studies demonstrated a strong relationship between years of school-

This paper represents part of a larger work on alienation, class, and schooling that I am undertaking jointly with Herbert Gintis. I gratefully acknowledge his help and that of Valerie Nelson and Janice Weiss. The work presented here has been financially supported by the Social and Rehabilitation Service of the U.S. Department of Health, Education, and Welfare. An earlier draft of this paper was presented at the Far Eastern Meetings of the Econometric Society in Tokyo in June 1970.

<sup>1</sup> Hanoch (1967); Mincer (in press) is the most comprehensive and recent study.

ing and earnings. For white males 25–34 years old, for example, each additional year of schooling appeared to result in around \$350 additional annual income.<sup>2</sup> It seemed obvious that if more schooling could be given to the children of the poor, a significant increase in their incomes would result.

Second, the *actual role* of schooling as an equalizer seemed to be confirmed by recent studies of the intergenerational transmission of economic status. The social background of individuals was shown to exercise a relatively minor impact upon the number of years of schooling attained. Measures of parents' occupational and educational status appeared to explain only between a quarter and a third of the variance of years of schooling attained (Blau and Duncan 1967; Duncan 1968; Duncan, Featherman, and Duncan 1968; Hauser 1969, 1970). Further, the same studies seemed to show that additional schooling exerts a major effect upon earnings or occupational status *independent of the social class background of the individual*. This last finding is of central importance, for it seems to lay to rest a common objection to the early earnings functions, namely, that the apparently large impact of schooling upon earnings might be a statistical artifact resulting from the positive correlation between the social class background and the level of schooling of individuals.

These recent studies ran contrary to a large and venerable body of sociological literature on the relation between social class and schooling. Earlier studies had asserted that social class was a major determinant of the amount, quality, and vocational orientation of the educational experience of individuals (for example, Warner, Havighurst, and Loeb 1944; Hollingshead 1949). Yet the more recent work of the Duncans and their associates was based on far larger, more comprehensive samples, and used more rigorous and systematic statistical techniques. The older view of social class and schooling has gradually been discredited.

The confidence in education inspired by the demonstration that schooling is a major determinant of earnings underlay the basic strategy of the U.S. government's effort to combat poverty. In addition, the earnings-functions studies gave strong support to what has been called the human capital interpretation of inequality (Mincer 1958; Becker 1964). Education and training programs consumed the lion's share of War on Poverty funds. Underlying this allocation of resources was a new view of poverty. It explains the poverty of the poor by their low productivity, and this, in turn, is attributed to their low levels of schooling and training (Schultz 1966). Inadequate education is seen as the problem, and more education as the solution.

The empirical basis for this position, strong as it seemed at first, is no

<sup>2</sup> The figure is an average over all years of schooling (Hanoch 1967) and refers to the nonsouthern region.

longer uncontested. Evidence has begun to accumulate challenging the efficacy of schooling as an equalizer of incomes. Hanoch (1967) found that the internal rate of return to increased schooling (except for graduate studies) was considerably lower for blacks than for whites. Extending Hanoch's work, Weiss (1970) estimated earnings functions for black workers having 12 or fewer years of schooling. He found no statistically significant monetary return to additional schooling except for workers in the 35–44-year age group. Harrison's work (1969) with more recent data has done nothing to overturn the impression of negligible monetary returns to education for blacks. But low returns to schooling evidently are not confined to blacks. Using data on draft rejects—a group not atypical of the poverty population—Hansen, Weisbrod, and Scanlon (1970) estimated that the difference in annual earnings associated with an additional year of schooling was a paltry \$62.<sup>3</sup>

To evaluate the conflicting evidence on the role of schooling as a vehicle for the equalization of opportunity and income I have estimated a model of individual earnings determination and intergenerational transmission of economic status. Specifically, I estimated a recursive model similar to that used by Duncan and his associates:

$$\text{EDUC} = ax, \quad (1)$$

$$\text{INC} = b_1\text{EDUC} + bx, \quad (2)$$

as well as the reduced form equation:

$$\text{INC} = cx. \quad (3)$$

The following notation is used: EDUC = respondent's years of schooling; INC = respondent's annual income;  $x$  = a vector of variables measuring the respondent's socioeconomic origins. All variables are expressed in normalized form. In each equation there is a stochastic disturbance term which I have omitted here for simplicity of presentation.

I will consider three measures of the relationship among social class, schooling, and income. The first is the fraction of variance of EDUC explained by the socioeconomic background variables—a measure of the extent to which an individual's years of schooling is predetermined by his social and economic origins. The second is  $b_1$ , the regression coefficient of years of schooling in the earnings function. The third is the difference in the fraction of variance explained in equations (2) and (3), or the increment in the  $R^2$  associated with the introduction of the years-of-schooling variable. Because of the unambiguous causal ordering of the variables, this third measure is a legitimate indicator of the extent to which the dispersion of years of schooling is associated with the dispersion of income

<sup>3</sup> The figure is the regression coefficient of years of schooling in an equation predicting income. Other variables in the equation are age and race.

independently of the causally prior dispersion of socioeconomic background.<sup>4</sup>

While the estimated model presented below must be seen as provisional and based to some extent on conjecture, two conclusions of general importance seem warranted:

First, failure to adequately measure the social class background of individuals has led a number of researchers to premature and probably seriously misleading conclusions concerning the role of schooling in the process of income determination. The inadequate measurement of social background has arisen both from excluding important dimensions of social class and from erroneously measuring aspects of class actually incorporated in the models. I will deal with these two problems in Sections II and III, respectively. The results presented in Section IV—even bearing in mind their conjectural nature—are sufficient to recommend considerable skepticism concerning the credibility of estimates that take inadequate account of the twin problems of specification bias and erroneous measurement.

Second, my estimates—both of the model itself and of the remaining biases due to errors of measurement and specification—suggest 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. Likewise, my results seriously question the putative efficacy of schooling as an instrument for the equalization of incomes. The economic return to additional schooling for individuals of a given social class background will be shown to be considerably less than the simple relationship between schooling and earnings would suggest.

## II. The Specification of Socioeconomic Background

I will begin with an obvious point: assessment of the role of schooling in the intergenerational transfer of economic status requires a rather complete understanding of the ways in which the transfer of status takes place. More specifically, estimation of equations (1)–(3) is likely to yield biased results unless we are able to specify fully those characteristics associated with a person's family background that might have some plausible direct or indirect causal relationship to the likelihood of the person earning high income in his adult life.

Equations (1)–(3) can be estimated as a recursive system only because the relationship represented by equation (1)—the determination of years of schooling attained—is postulated as causally prior to that represented by equation (2)—the determination of income. While this seems a reasonable assumption, the unbiased estimation of equation (2) requires the

<sup>4</sup> The three measures  $M_1$ ,  $M_2$ , and  $M_3$  are not independent:  $M_3 = M_2^2 (1 - M_1)$ .

highly unlikely complementary assumption that the error term in the first equation is uncorrelated with that in the second. The most obvious source of bias is that some dimensions of the social background of the individual exert a positive direct effect upon both the level of schooling attained by the individual and his later income. If these aspects of the social background are not measured by the socioeconomic background variables, the level of schooling achieved (EDUC) will serve as a proxy for these unmeasured variables in equation (2). For this reason the coefficient of EDUC will be overestimated. For example, if parental wealth has a direct positive effect both upon the level of schooling attained and the later income of the individual, the coefficient of the schooling variable in the income-determination equation will represent not only the direct effect of education upon income, but part of the direct effect of unmeasured parental wealth upon the respondent's income as well. Thus unless we are able to completely specify and accurately measure all of the relevant background variables,  $x$ , all three of our measures of the role of schooling in income determination will be biased: The proportion of variance of years of schooling attained explained by social background will be underestimated, the regression coefficient of the years of schooling variable will be overestimated, and the increment in the explained variance of income associated with the introduction of the schooling variable will be overestimated.

Because available data do not allow the complete specification of the relevant social background of individuals, existing estimates of the role of schooling in the intergenerational transfer of economic status are biased, as I will show in this section. More concretely, I will argue that the absence of measures of family income, parental wealth, and the position of the parents in the hierarchy of work relations has systematically biased the resulting estimates in the direction of showing schooling to have a powerful effect upon income independent of the socioeconomic background of the individual. In the next section, I will show that even those variables that *are* frequently included in studies of income determination—father's occupational status score and father's educational level—are measured with a substantial degree of error, thus exacerbating the above biases due to specification errors.

Family income and wealth are obvious candidates for inclusion in the equation predicting years of schooling attained. Both measure the ability of parents to finance the direct and indirect costs of their children's education. Both probably are associated with dimensions of the value orientations and aspiration levels of the home not fully captured in the available socioeconomic status variables measuring father's occupation and education. Likewise, no compelling argument can be adduced for their exclusion from the income-determination equations. Quite the contrary. A direct relation of parental wealth to individual income—operating through inheritance—would seem an obvious aspect of the process of the intergenerational trans-

mission of economic status. Moreover, plausible conjectures suggest that the relationship between parental wealth and respondent's income may be of considerable magnitude.

Consider the following model, with all variables expressed in normalized form:  $NW = f_0 \text{ PNW}$ ;  $INC = f_1 NW + f_2 \text{ EDUC} + fx$ ; and the reduced form:  $INC = f'_0 \text{ PNW} + f'x$ , where  $NW$  = respondent's net worth, and  $\text{PNW}$  = parents' net worth. All other variables are as defined in the Introduction. Notice that the model postulates no direct relation between parents' net worth and the income of the respondent; the influence of  $\text{PNW}$  upon  $INC$  is totally mediated by  $NW$ ; likewise, the linear additive form precludes interaction between  $\text{PNW}$  and  $\text{EDUC}$  and  $x$ . Last, I have not postulated a direct partial relation between  $\text{PNW}$  and  $\text{EDUC}$ . If we now assume that the variance of  $\text{PNW}$  explains 25 percent of respondent's net worth,  $f_0 = .5$ . Further assume an average rate of return,  $i$ , of 7 percent, and that the standard deviation of  $NW$  is roughly three times as large as that for income. (This latter assumption is based on data from the 1967 Survey of Economic Opportunity [see Cromwell, in preparation].) Then  $f_1 = i (S_{NW}/S_{INC}) = .21$ , and the normalized regression coefficient of  $\text{PNW}$  in a reduced-form income equation is  $\partial_{INC}/\partial_{PNW} = f'_0 = f_0 f_1 \approx .10$ . Incorporation of even a small direct partial relation between parental net worth and respondent's years of schooling would raise  $f'_0$  to above .15, or only slightly less than the normalized partial regression coefficient of father's occupational status on son's income (estimated in Section IV).

The biases resulting from the exclusion of parental income and wealth measures would be minor if these variables were highly correlated with the parental occupation and education variables on which data are ordinarily available. But this is not the case. In the sample used here and originally studied by Duncan and his associates, the respondents' occupational status score and educational level together explain only 32 percent of the variance of income of non-Negro males of nonfarm background, aged 35–44 years.<sup>5</sup>

The statistical association between net worth and the socioeconomic status variables ordinarily used is similarly weak, as the correlation coefficients in table 1 indicate, for occupational status and years of schooling together explain only 5 percent of the variance of net worth.<sup>6</sup>

A third excluded dimension of the socioeconomic background of respondents is the parents' position in the hierarchy of work relations. The importance of this dimension can best be understood in the framework of the following model of intergenerational status transmission:<sup>7</sup>

<sup>5</sup> Correlations from Duncan et al. (1968) have been adjusted upward to account for errors in measurement of all three variables. This particular age group seems most relevant to our concerns here as it is the group most likely to have young children in the home.

<sup>6</sup> These figures would be somewhat higher if the measurement errors in the variables were taken into account.

<sup>7</sup> The model outlined here is developed in more detail in Bowles (in press).

TABLE 1  
ZERO-ORDER CORRELATIONS AMONG MEASURES OF SOCIOECONOMIC STATUS,  
NORTHEAST CENSUS REGION, NONFARM HOUSEHOLDS  
HEADED BY MALES, AGED 25-44

Variable	(1)	(2)	(3)	(4)
1. Income .....	1.0	...	...	...
2. Years of school .....	.375	1.0	...	...
3. Occupational SES .....	.384	.556	1.0	...
4. Net worth .....	.314	.184	.177	1.0

SOURCE.—Cromwell (in preparation). Data based on the 1967 Survey of Economic Opportunity tapes made available by the U.S. Office of Economic Opportunity. Households headed by males in the military service or enrolled in school are excluded. The sample size is 596. I am grateful to Cromwell for making these data available to me.

The social relations of production characteristic of advanced capitalist societies (and many socialist societies) are most clearly illustrated in the bureaucracy and hierarchy of the modern corporation.<sup>8</sup> Occupational roles in the capitalist economy may be grouped according to the degree of independence and control exercised by the person holding the job. The degree of occupational self-direction is positively associated with, but by no means completely determined by, the worker's position in the hierarchy of work relations. The personality attributes associated with the adequate performance of jobs in occupational categories defined in this broad way differ considerably, some apparently requiring independence and internal discipline, and others emphasizing such traits as obedience, predictability, and willingness to subject oneself to external controls.

These personality attributes are developed at a young age in the family and, to a lesser extent, in secondary socialization institutions such as schools. They are reinforced in the day-to-day experience of adults. Because people tend to marry within their own class, both parents are likely to have a similar set of these fundamental personality traits. Thus children of parents occupying a given position in the occupational hierarchy grow up in homes where child-rearing methods and perhaps even the physical surroundings tend to develop personality characteristics appropriate to adequate job performance in the occupational roles of the parents.<sup>9</sup> The

<sup>8</sup> Max Weber referred (1946) to bureaucracy as the "most rational offspring" (p. 254) of discipline, and remarked that "military discipline is the ideal model for the modern capitalist factory" (p. 261).

<sup>9</sup> Much of the evidence for this assertion is from three major studies of occupational self-direction and parental values by Melvin Kohn which he summarized in 1969. He concludes: "Whether consciously or not, parents tend to impart to their children lessons derived from the conditions of life of their own class—and thus help to prepare their children for a similar class position . . . . The conformist values and orientation of lower- and working-class parents . . . are inappropriate for training children to deal with the problems of middle class and professional life. . . . The family, then, functions as a mechanism for perpetuating inequality" (p. 200). On class differences in child rearing with respect to the importance of obedience, see Dolger and Ginandes (1946) and Kohn (1964). See also the study of differences in child-rearing practices in families



children of managers and professionals are taught self-reliance within a broad set of constraints (Winterbottom 1953; Kohn 1963); the children of production-line workers are taught obedience.

While this relation between parents' class position and child's personality attributes operates primarily in the home, it is reinforced by schools and other social institutions. Thus, for example, the authoritarian social relations of working-class high schools complement the discipline-oriented early socialization patterns experienced by working-class children. The relatively greater freedom of wealthy suburban schools extends and formalizes the early independence training characteristic of upper-class families.

In this interpretation, the educational system serves less to change the results of the primary socialization in the home than to ratify them and render them in adult form. The complementary relationship between family socialization and schools serves to reproduce social class differences in personality development from generation to generation.

The operation of the labor market translates these differences into income inequalities and occupational hierarchies. The personality traits, values, and expectations characteristic of different class cultures play a major role in determining an individual's success in gaining a high income. The apparent contribution of schooling to higher income, far from being the result of the independent role of schooling in the development of cognitive capacities, seems to be explained primarily by the personality characteristics of those who have higher educational attainments.<sup>10</sup>

In this view, the hierarchy of work relations is replicated in social class differences in values and child rearing. Because of the important role of affective characteristics as determinants of success both in school and on the job, class differences in parental values and child-rearing practices play an important role in replicating the class structure in the next generation. If this interpretation is correct, the failure of studies of income determination to adequately measure the parents' position in the work hierarchy results in an underestimate of the impact of socioeconomic background upon adult economic success and an overestimate of the effects of schooling.

Once again, the exclusion of a measure of the parents' position in the hierarchy of work relations would be of little import if this dimension of family background were highly correlated with the socioeconomic status score of the parents' occupation or with the parents' education. But this does not appear to be the case. Melvin Kohn has shown (1969, p. 166) that the relation between occupational self-direction and parental values (self-

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headed by bureaucrats as opposed to entrepreneurs by Maccoby, Gibbs, et al. (1954) and Miller and Swanson (1958). While the existence of class differences in child rearing is supported by most of the available data (but see Lewis 1965), the stability of these differences over time has been questioned by Bronfenbrenner (1963).

<sup>10</sup> This view is elaborated in Gintis (1971). For other studies stressing the noncognitive dimensions of the schooling experience, see Parsons (1959) and Dreeben (1968).

direction vs. conformity) is statistically significant even when the occupational status score of the parents is controlled.

The existence of biases in estimates of the schooling-income-socioeconomic-background relationship arising from the failure to consider relevant dimensions of socioeconomic background does not by itself require that the estimates be rejected. But on the other hand, one cannot place much confidence in the existing estimates unless a compelling argument can be made that these biases are small. No such case has been made—or even seriously attempted—by the main contributors to the literature on intergenerational status transmission and individual income determination.

I have attempted to eliminate one source of specification bias by hypothesizing a parents' income variable and developing estimates of the relevant row and column in the product-moment matrix on the basis of data from a variety of sources. The methods used are described in the Appendix. The crucial assumption used in the estimation is that the zero-order correlation between parents' income, on the one hand, and respondent's income, on the other, is equal to the correlation between father's occupation and respondent's occupational status.<sup>11</sup>

Available data do not allow use of variables even hypothetically measuring parental wealth and the position of the parents in the hierarchy of work relations.

I turn now to the problem arising from erroneous measurement of those socioeconomic background variables ordinarily available.

### III. Measurement Errors

The above biases are due to the incomplete specification of the model and would arise even if the variables I am forced to use were accurately measured. Unfortunately, the available data contain serious inaccuracies.

The presence of errors in measurement in the variables used lowers the explanatory power of the equations, and ordinarily biases downward the regression coefficients of the erroneously measured variables. Because the degree of error in the measurement of the father's occupation and education variables greatly exceeds that in the respondent's own years-of-schooling variable, we are led to expect that failure to take systematic account of these errors will lead to an underestimation of the importance of socioeconomic background as an influence upon educational attainments and later income. The relative importance of schooling as a determinant of income will be correspondingly overestimated.

<sup>11</sup> The resulting raw (uncorrected) correlation is .202, or slightly lower than the value of .258 found for this correlation (from the matrix used by Conlisk [1968]) in the small sample used by the Berkeley Guidance Study. That study somewhat supports the figures used in this essay, nonetheless the conjectural nature of this aspect of the exercise should be obvious.

In this section I attempt to estimate the magnitudes of the error components in each variable and to develop a method of estimating equations (1) and (2) that will reduce the biases due to the errors-in-variables problem.

I will use these estimates in Section IV to eliminate some of the measurement-error bias from a model of intergenerational status transmission and income determination.

The data used in this essay are from a U.S. Census survey of slightly more than 20,000 males 20–64 years of age in 1962.<sup>12</sup> Respondents were asked to report their own occupation and level of educational attainment (in years) as well as the occupation and educational level of their father or family head. Additional data collected included the number of siblings of the respondent and his income in the year previous to the survey. Because the importance of family size has been stressed by many students of mobility, I have included the number of siblings reported by the respondent as a measure of social background. The occupation of the respondents' fathers was scaled according to the Duncan socioeconomic status index. An index of years of schooling is the sole measure of educational attainment.

These data were collected by surveys and often required the respondent to provide retrospective information such as his father's occupation when the respondent was 16 years old. Quite apart from errors in responses likely in these cases, some of the data do not correspond exactly to the models which I seek to estimate. This errors-in-variables problem is to be distinguished from the problems associated with the inadequate specification of equations in the model due to the above-mentioned incomplete measurement of the social class of the respondent. Confining attention to the incomplete set of variables on which we do have data, we find that the data available often do not measure what they purport to measure, and further, that the measure itself, even if accurately observed, does not correspond to the variable in our model. For example, in a model of the effect of education upon economic success, we would like to measure respondent's permanent income, yet our observations purport to measure only annual income. We may generalize the problem as follows: For each variable,  $x$ , and for any individual observation,  $i$ , we have

$$x'_i = x_i + u_i, \quad (4)$$

where  $x_i$  = the true value of the variable,  $x'_i$  = the observed value of the variable, and  $u_i$  = the error in measurement. We know that errors of this type will bias the least-squares estimates of the regression coefficients as well as the coefficient of determination.

<sup>12</sup> Blau and Duncan (1967, pp. 10–19) give a more complete description of the sample properties. I will discuss estimates only for the group 25–34 years of age.

In order to eliminate the biases arising from the discrepancies between the observed and true values of the variables used, I will estimate the zero-order correlation coefficients among the true variables and use these corrected correlation coefficients to estimate the model of mobility.<sup>13</sup> If we assume that the errors,  $u_i$ , are uncorrelated with the true values,  $x_i$ , then it follows that<sup>14</sup>

$$\text{var}(x') = \text{var}(x) + \text{var}(u). \quad (5)$$

Now define  $r_j$ , the correlation of the true value of  $x_j$ , with its observed value, as

$$r_j = \sqrt{\frac{\text{var } x_j}{\text{var } x'_j}}, \quad (6)$$

or the square root of the fraction of the variance of  $x'_j$ , the observed measure, which is accounted for by the variance of  $x_j$ , the true measure. Then the observed correlation between any pair of variables  $x_k$  and  $x_j$ ,  $r'_{kj}$ , may be written as a function of the true correlation,  $r_{kj}$ , the correlations between the true and observed variables,  $r_k$  and  $r_j$ , and the correlation of the errors in the two observed variables,  $r_{ukj}$ :

$$r'_{kj} = r_{kj} r_k r_j + r_{ukj} \sqrt{1 - r_k^2} \sqrt{1 - r_j^2}. \quad (7)$$

The corrected correlation coefficients,  $r_{kj}$ , will be used as the normalized  $X'X$  matrix to estimate the model of class immobility. (See Appendix for corrected and uncorrected correlation matrices.)

For each variable I attempt to introduce independent data concerning the degree of error in the measures that I have used in my regression equations. While the information used to estimate the accuracy of the measures is itself subject to serious question arising from differences in samples, ages of respondents, and variable definitions, I believe that the errors arising from erroneous estimates of reliability are considerably less serious than those which would result if I were simply to use the uncorrected data.

I will consider the error in each variable in turn, and then deal with those pairs of variables for which the errors are likely to be correlated. (The somewhat complicated processes of estimating these values are described in more detail in the Appendix.)

I turn first to problems concerning the definition and measurement of income. Abstracting from inaccuracies in the respondents' reported income, I have already noted that annual income is not the correct variable

<sup>13</sup> This method is formally equivalent to that suggested by Johnston (1963) and others (see Appendix).

<sup>14</sup> To adopt a more realistic assumption would greatly complicate the task of calculating corrected correlation coefficients and would require unavailable data.

to use in a model of the intergenerational transfer of economic attainment. Most available studies do not allow us to distinguish between the variance in annual income due to year-to-year transient variations, on the one hand, and simple reporting errors, on the other. However, there are a number of estimates of the fraction of the variance of observed income that is accounted for by both reporting errors and the transient component in annual income. The estimate most consistent with the available data implies that only 70 percent of the variance of observed income is due to the variance of permanent income.<sup>15</sup> The square root of the figure, .84, is the estimate of the correlation of permanent and observed income which appears in column 3 of table 2.

I turn now to questions concerning the accuracy of respondent's reports of their own educational attainments. Immediately following the 1950 census, a postenumeration survey was conducted to check the accuracy of census responses (Bureau of the Census 1960). A comparison of the respondents' reports to both the census and the Post-Enumeration Survey allows an estimate of the correlation of the true and reported values. I have calculated a number of values of this correlation based on alternative assumptions concerning both the relative accuracy of the census and the

TABLE 2  
ESTIMATED ERRORS IN VARIABLES MEASURING SOCIOECONOMIC  
BACKGROUND, INCOME, AND EDUCATIONAL LEVEL

Variable Required (1)	Measure Used (2)	Estimated Correlation of Observed Measure with True Value of Variable Required (3)
1. Respondent's permanent income . . . . .	Respondent's annual income	.84
2. Respondent's educational attainment	Respondent's years of school attained (index)	.91
3. Occupational status of the father or family head of respondent . . . . .	Duncan's status score for the occupation of father or family head	.80
4. Educational attainment of father or family head of respondent . . . . .	Years of school attained (index) by father or family head	.80
5. Parent's permanent income . . . . .	Parents' annual income	.84
6. Family size . . . . .	Number of siblings	.96

SOURCE.—See Section III and Appendix.

<sup>15</sup> The choice of this figure is explained in the Appendix, where a series of alternative estimates are given.

Post-Enumeration Survey and the correlation of errors in reporting to the two surveys (see Appendix for method of calculation and the alternative estimates). The most plausible assumptions yield a correlation of .91 between the true and observed values of the index of educational attainments. These correlations are reported in column 3 of table 2.<sup>16</sup>

Note that while I have estimated the degree or error in reporting one's educational attainments (col. 3 of table 2), I have assumed that years of schooling is an accurate measure of the level of educational attainment. Years of schooling attained should not be construed as an accurate measure of the total school resources devoted to a respondent's schooling. While the amount of resources "enjoyed" per year is associated with the years of school eventually attained, the correlation is far from perfect. Whatever bias arises due to this discrepancy operates—though not necessarily with equal force—for both the respondent's schooling and that of his parents.<sup>17</sup>

Consider now the accuracy of the respondents' reports of their parents' occupation and education. The data used here are from a survey in which respondents were asked to report the highest level of schooling attained by the father or family head, as well as the occupation held by the father or family head at the time the respondent was a teenager.

As part of their survey of intergenerational mobility, Blau, Duncan, and their associates administered a survey of 570 males in Chicago; the usual questions concerning parents' status were asked, along with an item eliciting the respondent's address when he was 16 years old (Blau and Duncan 1967, pp. 457–62). The decennial censuses nearest to the respondent's sixteenth birthday were then searched to extract the census report of the respondent's father's occupation.<sup>18</sup>

The zero-order correlation between the occupational status as reported by father and by son was found to be .74. There is a downward bias in this measure, as the census years from which the father's own reports were taken did not correspond exactly to the sixteenth year of age of the respondent. On the other hand, an upward bias is implicit in the method by which the sample to be studied was selected. The study automatically excluded respondents who could not correctly recall another retrospective

<sup>16</sup> The correlation between the true and observed values of respondents' occupational status score calculated in the same manner was .92.

<sup>17</sup> See the next section for a discussion of biases arising from the inadequate measurement of schooling.

<sup>18</sup> Of the original 570, only 137 cases could be used in the study. Inclusion required that the respondent had correctly recalled his address and had responded to the question concerning father's occupation, and also that his father had responded to the census question on occupation. A study of the matched responses then compared occupation of the father as reported by the respondent with that reported (presumably by the father or mother of the respondent) to the census. Then those reporting farm occupations to both surveys were eliminated, reducing the total to 115, and the occupations were scaled by the Duncan status score.

fact (their address at age 16), as well as those who did not answer the question concerning father's occupation and those whose father also had not replied to that question when asked by the census enumerator. While these considerations would seem to point on balance to an upward bias in the estimated accuracy of the responses, use of the figure is consistent with the data on cohorts collected by the 1962 Occupational Changes in a Generation Survey, as well as with other census data (see Appendix). I have, therefore, used this figure, and, as I did not have independent evidence on the accuracy of reports of parents' educational attainments, I applied the figure to that variable as well.

Use of this reliability estimate in conjunction with independent census data implies a negligible correlation of errors in reporting occupational and educational attainments (see Appendix).

The estimate of the accuracy of reports of number of siblings is based on the following reasoning: Duncan reports a correlation of .96 between mothers' census and reinterview reports of number of children ever born. Making the extreme assumption that the reinterview was totally accurate and that adults report number of siblings as inaccurately as they report number of children ever born, the correlation between the reported and actual number of siblings is .96.<sup>19</sup>

The evidence of this and the previous section points unambiguously to the existence of shortcomings in the specification of socioeconomic background, for errors in the measurement of some variables appear to be significant. Furthermore, as was indicated in the previous section, important socioeconomic background variables are omitted altogether.

A test proposed by Duncan et al. (1968) provides independent evidence on these shortcomings of measurement and specification. Because brothers ordinarily share a common socioeconomic background, estimates of the equation predicting years of schooling also allow the prediction of the degree of correlation between the educational attainments of brothers.

If we rewrite equation (1) for the respondent, with variables in normalized form  $EDUC = ax$ , and assume that siblings share a common socioeconomic background and similar relations between educational attainment and socioeconomic background, then the predicted correlation between brothers' years of schooling,  $\hat{r}^*$ , will be:

$$\hat{r}^* = \sum_{i=1}^n a_i r_{ieduc}, \quad (8)$$

where  $a_i$  is the normalized regression coefficient of the  $i$ th socioeconomic background variable from the equation predicting  $EDUC$ , and  $r_{ieduc}$  is the

<sup>19</sup> If the reinterview mentioned by Duncan was as inaccurate as the census and the errors are uncorrelated, a figure of .98 would be more appropriate.

zero-order correlation between the brother's years of schooling and variable  $i$  (assumed equal to that for the respondent himself).<sup>20</sup>

If the predicted correlation  $\hat{r}^*$  falls short of the actual correlation,  $r^*$ , we must conclude that the factors influencing the respondent's level of schooling have been inadequately specified or erroneously measured, or both. The observed correlation for the sample of males, aged 25–34 years, studied here is .54 (Duncan et al. 1968, table B.5). After correcting this observed coefficient for errors in reporting,<sup>21</sup> this coefficient rises to .65. Using the equation as estimated by Duncan et al. (1968, p. 54), uncorrected for errors in variables and including only father's occupational status and educational level and respondent's number of siblings as measures of socioeconomic background, the predicted coefficient is .26, less than half of the observed coefficient. A discrepancy this large suggests not merely that biases exist, but that they are likely to be significant.<sup>22</sup> In the next section I will subject my own estimates to a similar test.

#### IV. An Empirical Model of Schooling and Income Determination

The estimates of reliability of the socioeconomic status, education, and income variables have been used to calculate a correlation matrix among the true values of the variables. (This correlation matrix, along with recalculated standard deviations of these variables, appears in the Appendix.) I have used the corrected correlation matrix to estimate equations (1), (2), and (3). The resulting estimates can tell us little about the magnitude of the biases arising from inadequate specification of the socioeconomic background of individuals. The only adjustment for this problem is the hypothetical introduction of a variable representing parents' income. Yet the estimated equations, as presented in table 3, do cast some light on the seriousness of the errors-in-variables problem.

The following characteristics of the results should be noted: First, the measures of family background explain 52 percent of the variance of the

<sup>20</sup> The value of  $\hat{r}^*$  is thus equal to the fraction of variance in sons explained by equation (1).

<sup>21</sup> I am assuming, conservatively, that the respondent's reports of brother's educational attainment are no more error prone than reports of his own educational attainment.

<sup>22</sup> Of course, the omitted variables may not measure dimensions of the socioeconomic background of the individual. The only other potentially important common aspect of background which may be conceived of as exogenous in this model seems to be the common (but not identical) genetic inheritance of brothers. Duncan et al. (1968) explicitly measured the effect on educational attainments of differences in childhood IQ, and thus were able to extend the above calculations to take some account of the common genetic inheritance of siblings. The resulting value of  $r^*$  for non-Negro native men 25–64 years of age was .34, still far short of .573, the observed correlation for this sample.



TABLE 3  
REGRESSION EQUATIONS FOR MODEL OF SOCIOECONOMIC BACKGROUND, SCHOOLING, AND INCOME, U.S. NON-NEGRO MALES  
WITH NONFARM BACKGROUNDS, 25-34 YEARS OLD

DEPENDENT VARIABLE	NORMALIZED REGRESSION COEFFICIENTS OF:*					R <sup>2</sup>
	Father's Occupational Status	Father's Years of Schooling	Parents' Income	Respondent's Number of Siblings	Respondent's Years of Schooling	
1. Respondent's years of schooling .....	.155 (58.8)	.204 (108.6)	.412 (832.3)	-.181 (185.6)	...	.517
2. Respondent's earnings .....	.112 (29.5)	†	.112 (27.7)	-.059 (10.9)	.203 (76.2)	.151
3. Respondent's earnings .....	.173 (77.3)	†	.197 (105.7)	-.102 (33.8)	...	.130

NOTE.—Refers to experienced civilian labor force, 1962, as represented by the occupational changes in a generation survey sample.  
\* *F*-statistics in parentheses.  
† In a prior estimate this coefficient was insignificantly different from zero. The equation presented here was estimated without this variable.

years of schooling obtained by the respondent.<sup>23</sup> Second, years of schooling attained appears to be a significant determinant of the earnings of the respondent. The estimated increment in annual income associated with an additional year of schooling is \$265. Yet this partial relationship of schooling to income net of socioeconomic background is less than 60 percent as large as the gross return indicated by the simple relationship between the two variables.<sup>24</sup> This finding suggests that much of the apparent economic return to schooling is in fact a return to socioeconomic background.<sup>25</sup>

Third, the variance of earnings explained by the social background variables alone is only slightly less than that explained by these variables along with the educational attainments of the respondent. The social background variables alone explain 13.0 percent of the variance of earnings.<sup>26</sup> The additional variance explained by years of schooling is only 2.1 percent. I infer from this result that years of schooling attained exerts a comparatively minor independent influence on earnings independent of social background. Most of the impact of years of schooling on earnings appears to be a direct transmission of economic status from one generation to the next.<sup>27</sup>

As an internal check on the plausibility of the correction for errors in measurement, I can now calculate the correlation between brothers' years of schooling predicted by my estimate of equation (1).<sup>28</sup> The predicted correlation, .52, still falls short of the true (corrected) .65. Nonetheless this calculation suggests that the correction for errors in variables has resulted in a considerable improvement in prediction.<sup>29</sup>

Returning to table 3, it might be argued that the explanatory power of the schooling variable would be increased if I had a measure of the quality as well as the duration of schooling. While this is undoubtedly true, the

<sup>23</sup> This is almost twice the fraction of variance explained by the uncorrected variables in Duncan et al. (1968).

<sup>24</sup> This calculation is based on the normalized regression coefficient of years of schooling in equation (2), the corrected zero-order correlation coefficient between schooling and income, and the corrected standard deviation of the two variables.

<sup>25</sup> It is shown below, in an extension of this work, that introduction of a measure of early IQ reduces the apparent net effect of schooling still further (see Bowles and Gintis 1972).

<sup>26</sup> The social background variables here include number of siblings. If that variable is excluded, the  $R^2$  falls to .12. Using the uncorrected data from Duncan et al. (1968), the  $R^2$  is .054 with number of siblings in the equation and .045 without.

<sup>27</sup> This does not appear to be the case when the respondent's occupational status is used as the dependent variable. The results suggest that while education exerts a major independent influence on occupational attainment, this influence does not translate into a major independent influence on earnings. The discrepancy between these results may be explained by the wide dispersion of earnings within occupational categories.

<sup>28</sup> The predicted correlation is calculated as described in the text using correlation coefficients from table A3 and normalized regression coefficients from table 3.

<sup>29</sup> Gintis and I show elsewhere (1972) that the introduction of a hypothetical variable measuring childhood IQ in this model raises the predicted correlation to a value slightly below the actual (corrected) correlation.

importance of this point is easily exaggerated. It is difficult to conceive of operational measures of school quality which will be at the same time important in their influence on adult economic success and not highly correlated with years of schooling attained and the social class of the respondent. The most commonly suggested—average school resources per year of schooling—would seem to be highly correlated with both years of schooling and social class background. A second measure—scholastic achievement—appears to fail on all counts. To the extent that we can explain the variance among individual students in scholastic achievement, the social class of the student seems to be the main explanatory variable. The increment in the explained variance of scholastic achievement scores associated with the introduction of school policy and resource variables in an equation already including crude measures of the social background of the student is ordinarily very small.<sup>30</sup> Better measures of school policies and resources would undoubtedly alter the picture somewhat, but even a substantial change would not be of great import as long as the scholastic achievement scores themselves are highly correlated with the number of years of schooling attained. If the Armed Forces Qualification Test (AFQT) may be taken as a proxy for a scholastic achievement test, we arrive at an estimate of .68 for this correlation,<sup>31</sup> suggesting that a substantial part of the variance of scholastic achievement is already measured by the years-of-schooling variable.

The more serious problem involved in the use of the achievement score (or its relatives such as the AFQT score) is that there is very little evidence that the effect of schooling upon economic success operates to any significant degree through the effect of schooling upon the types of cognitive development measured in these tests. If the sole medium through which schooling operated was cognitive development as measured by achievement tests, then we would expect to find that the addition of an individual's test score to an equation using years of schooling to predict individual earnings would result in the coefficient of schooling falling to zero. On the other hand, if education contributed to earnings entirely independently of its effect on cognitive development, the coefficient of years of schooling would be reduced by a relatively minor amount.<sup>32</sup>

<sup>30</sup> See Coleman et al. (1966) and Bowles (1970). This statement does not imply that variations in school policy or resources have no effect on scholastic outcomes, but rather that school policy and resources as conventionally measured exercise very little influence not already measured statistically by the social class variables. For a discussion of this problem, see Bowles and Levin (1968*a*, 1968*b*).

<sup>31</sup> The correlation refers to U.S. males 25–34 years old and is corrected for errors in both variables. The test-retest reliability of the AFQT is about .95. The uncorrected correlation is from Duncan (1968). Other sources report a considerably higher correlation between AFQT score and years of schooling completed. See Personnel Research Division, Adjutant General's Office (1945) for evidence on this correlation as well as on the reliability of the AFQT.

<sup>32</sup> The reduction in the coefficient of education in this case would be explained by

Evidence from unpublished studies by Cutright, Conlisk, and Griliches, summarized by Gintis (1971), indicates that the regression coefficient of years of schooling is only slightly reduced and remains highly significant upon the introduction of a measure of cognitive development to a function predicting individual earnings or income. Though preliminary, these results suggest that schooling exerts an influence upon earnings largely independent of its contribution to scholastic achievement.<sup>33</sup>

In summary, the importance of biases arising from the exclusion of school quality variables cannot be adequately estimated, in part because of our inability adequately to specify what we mean by school quality. We cannot exclude the possibility that these biases may be large. Nonetheless, the above considerations lead me to doubt that important modifications in the results would follow the successful inclusion of a school quality variable in the model.

## V. Conclusion

Given the available data and restrictive assumptions required in the analysis, my own estimates are unavoidably subject to considerable error. Two possible sources of error in my treatment of the errors-in-variables problem seem to be particularly important: the assumption that errors in the main variables are uncorrelated with the true values, and the fragmentary nature of some of the evidence concerning the reliability of the variables and the correlation among errors in the variables. Moreover, while the direction of the specification biases seems reasonably certain, its real magnitude cannot be inferred with any confidence from the estimated coefficients of the hypothetical parents' income variable. We must await real data on this variable and on the other relevant aspects of family background before much of a quantitative nature can be said about this problem.

Nonetheless, the above estimates of correction for errors in variables, along with the earlier discussion of specification bias, leave little doubt that the estimation of models of social mobility and income determination which confine the measurement of socioeconomic background to respondents' reports of their parents' occupation and education level will result in significant biases. It is equally clear that these biases systematically understate the importance of social class in the determination of income and educational attainment.

Yet even having corrected for errors in measurement of the limited range of variables on which data are available, and having eliminated at least

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the assumed positive correlation between the level of cognitive development and the noncognitive variables that, in this model, are hypothesized as the medium through which education affects earnings (see Gintis 1970).

<sup>33</sup> This argument is based on Gintis (1969).

some part of the specification bias, I have estimated a set of equations which, while explaining a large portion of the variance of years of schooling, explains relatively little of the variance of income. Of course to some extent the limited explanatory power of equations (2) and (3) may be explained by the remaining shortcomings in the specification of the model—the omission of measures of parental wealth, the parents' position in the hierarchy of work relations, and perhaps other dimensions of social class as well as the quality of schooling. Furthermore, Mincer's work (in press) suggests that a substantial increment in explained variance of income would result from the introduction of measures of work experience and on-the-job learning.

Yet the low  $R^2$  in equations (2) and (3) may be the result of a mechanism in the individual income-determination process, the implications of which have, to my knowledge, been entirely ignored by economists and sociologists interested in income inequality and stratification. One's social class and educational level do not determine one's income; rather, they determine (presumably subject to some random influences) one's opportunity. Opportunity takes the form of a choice among jobs, each offering a different configuration of monetary and nonmonetary rewards. The income received by an individual is thus the result of a choice—a choice constrained by what could be called the occupational opportunity set. Only if preferences for various attributes of jobs are independent of socioeconomic background and level of schooling will the estimation of equations (1)–(3) yield unbiased estimates of the relations between social class, schooling, and real income, or equivalently, the *opportunity* to earn high money incomes. Yet this does not appear to be the case. There is considerable evidence that rich, high-status parents place a larger value on the nonmonetary aspects of work and a lower value on monetary returns than poorer, lower-status parents. Further evidence indicates that the job preferences of teen-age children show a similar relationship to socioeconomic status (Hyman 1966). The biases which arises in this case are illustrated in figure 1. Each job is characterized by an expected money income and a level of nonmonetary benefits (or costs) of the associated work.<sup>34</sup> The occupational opportunity set  $oab$  indicates the jobs available to an individual from a family of low socioeconomic status who has attained relatively few years of schooling. The opportunity set  $oa'b'$  refers to the job opportunities of an individual from a high-status family who has attained relatively many years of schooling.<sup>35</sup> Now if preferences for the monetary and nonmonetary aspects of work are associated with social class or with years of schooling, the occupations chosen may be illustrated by points  $x$  and  $x'$ . The money-income difference between the two individuals

<sup>34</sup> Nothing is lost by assuming (unrealistically) a single dimension for the non-monetary aspects of each job.

<sup>35</sup> The opportunity loci  $ab$  and  $a'b'$  need not, and in general will not, be parallel.

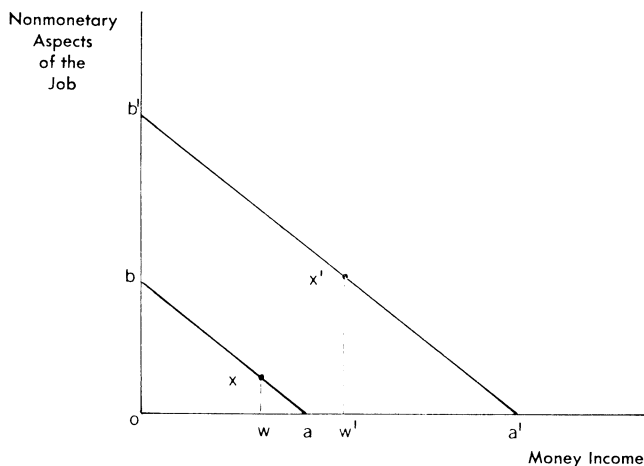


FIG. 1.—Occupational opportunities with endogenous preferences

$(w' - w)$  thus understates the differences in real income or opportunity, which may be measured on the money income axis as  $(a' - a)$ . As a result, the power of equations (2) and (3) to explain money income may greatly understate the relationship between real income or job opportunity on the one hand, and social class and schooling on the other. If the main determinant of job preferences is socioeconomic background, then the relative importance of the background variables as an influence on economic opportunity will be underestimated. Similarly, if education is the main determinant of job preferences, the coefficient of the years-of-schooling variable will be underestimated.<sup>36</sup>

Considering the biases arising from the problem of endogenous occupational preferences, along with the biases due to incomplete specification of family background, the results reported in table 3 demonstrate a surprisingly strong relationship between socioeconomic background on the one hand, and educational attainments and income on the other.

The substantial impact of even the limited measurement of socioeconomic status possible in this study can be illustrated as follows: Define a composite social class index,  $S$ , as the equally weighted sum of the three socioeconomic status variables.<sup>37</sup> Now we may ask: What is the expected difference in educational attainments and income for two hypothetical white males, ages 25–34, one originating from a high-status family and

<sup>36</sup> The greater explanatory power both of schooling and parental socioeconomic status in predicting the occupational status of the respondent (results calculated in a manner similar to equations [1]–[3] are not reported here; see also Duncan et al. [1968]) may be due in part to this association among job preferences, social class, and schooling. See also n. 27 above.

<sup>37</sup> That is:  $S = 1/3 \text{ FED} + 1/3 \text{ FOCC} + 1/3 \text{ PARinc}$ .

the other from a low-status family? Let us define high- and low-status families as those scoring exactly one standard deviation above or below the mean on the three family-status scores.<sup>38</sup> By this definition, the high-status individual's origins—at the eighty-ninth percentile of the composite social class (*S*) distribution—are translated into an advantage of 4.6 years of schooling over the respondent whose low-status origins place him at the eleventh percentile.<sup>39</sup> Further, even in the evidently unlikely event that both individuals attained the same years of schooling, the individual of high-status origins could expect to earn \$1,630 more annually over the ages 25–34.

Calculated in terms of the probability of receiving either a relatively high or low income, the impact of social class on income determination is even more striking. Taking account of the direct impact of socioeconomic background on income as well as its indirect effect via the relationship between social class and years of schooling attained, the high-status individual has 2.8 times the probability of receiving an income over \$10,000 annually than the low-status individual. Analogously, the low-status individual is 2.8 times more likely to receive less than \$1,200 annually.

While these calculations must be taken as only illustrative of orders of magnitude, they do suggest that substantial inequality of economic opportunity exists in the United States and that the educational system is a major vehicle for the transmission of economic status from one generation to the next.

## Appendix

### Methods Used to Estimate the Corrected Correlation Matrix

#### 1. *Equivalence to the More Familiar Errors-in-Variables Approach*

I will first show that the method used is equivalent to the generalized errors-in-variables approach as described by Johnston (1963). First write the normal equations in the form

$$\sum_{i=1}^n \hat{b}_i M'_{x_i x_k} = M'_{x_k x_j} \quad (k = 1 \dots n \text{ equations}), \quad (\text{A1})$$

<sup>38</sup> Assuming a normal distribution of *S* in the population under study, and utilizing the correlation coefficients in table A3, an individual whose family scored a standard deviation above the mean on all three status scores will be 1.21 SD above the mean on *S*, or at the eighty-ninth percentile of the social class distribution. Analogous calculations for an individual whose family scored a standard deviation below the mean place that person at the eleventh percentile.

<sup>39</sup> The impact (in standard deviation units) of a 2 SD difference in each socioeconomic background variable is calculated from equations identical with those in table 3, except that the number-of-siblings variable has been omitted so as to isolate the direct and indirect (via family size) effect of social class upon educational attainments and income. The equations omitting the NSIBS variable are virtually identical with those reported in table 3, and may readily be calculated from the corrected zero-order correlation matrix. The corrected standard deviations of years of schooling and income used to translate the normalized impact into raw figures appear in table A5.

where  $\hat{b}_i$  is the estimated regression coefficient of variable  $i$  in an equation predicting variable  $j$ , and

$$M'_{x_i x_k} \text{ and } M'_{x_k x_j}$$

are the observed second-order sample moments. If the variables are measured with error (uncorrelated with the true variables), we may estimate unbiased regression coefficients using  $n$  equations which appear in the form

$$\begin{aligned} \hat{b}_1 (M'_{x_1 x_1} - \text{var } u_1) + \sum_{i=2}^n \hat{b}_i (M'_{x_1 x_i} - \text{cov } [u_1, u_i]) \\ = M'_{x_j x_1} - \text{cov } (u_j, u_1), \quad (\text{A2}) \end{aligned}$$

etc., where  $\hat{b}_i$  are the unbiased regression coefficients and  $u_i$  is the error term in the observed variable  $x'_i$ . Expressing equation (A2) in standard deviation units of the true variables yields the system

$$\begin{bmatrix} 1 & M_{12} & \cdot & \cdot & M_{1n} \\ \cdot & & & & \cdot \\ \cdot & & & & \cdot \\ \cdot & & & & \cdot \\ M_{n1} & \cdot & \cdot & \cdot & 1 \end{bmatrix} \begin{bmatrix} \beta_1 \\ \cdot \\ \cdot \\ \cdot \\ \hat{\beta}_n \end{bmatrix} = \begin{bmatrix} M_{1j} \\ \cdot \\ \cdot \\ \cdot \\ M_{nj} \end{bmatrix}, \quad (\text{A3})$$

where  $\beta_i$  is the unbiased estimate of the normalized regression coefficient of variable  $i$ , and where

$$M_{ij} = \frac{M'_{x_i x_j} - \text{cov } (u_i, u_j)}{\sigma_{x_i} \sigma_{x_j}}. \quad (\text{A4})$$

Rearranging, we have

$$M_{ij} = \frac{M'_{x_i x_j}}{\sigma'_{x_i} \sigma'_{x_j}} \cdot \frac{\sigma'_{x_i} \sigma'_{x_j}}{\sigma_{x_i} \sigma_{x_j}} - \frac{\text{cov } u_i u_j}{\sigma_{x_i} \sigma_{x_j}}. \quad (\text{A5})$$

Using the notations introduced in the text, and noting that

$$(\sigma'_{x_i} \sigma'_{x_j}) / \sigma_{x_i} \sigma_{x_j} = (r_{x_j})^{-1},$$

it follows from the fact that

$$\text{cov}(u_i, u_j) = r_{nij} \sigma_{ui} \sigma_{uj},$$

and,

$$1 - r_i^2 = \frac{\sigma_{u_i}^2}{\sigma_{x_i}^2}$$

that

$$M_{ij} = \frac{r'_{ij}}{r_i r_j} - \frac{\sqrt{1 - r_i^2} \sqrt{1 - r_j^2}}{r_i r_j} \cdot r_{nij}, \quad (\text{A6})$$

which is identical with equation (7).



## 2. *The Accuracy of Reported Annual Income as a Measure of Permanent Income*

Various estimates of the ratio of the variance of permanent income to observed annual income are available. The first is Friedman's estimate based on the elasticity of consumption with respect to income.<sup>40</sup> Friedman's estimates for nonfarm or urban families in 1935-36 and 1941 appear as lines 1 and 2 of table A1.

We arrive at a second estimate if we define permanent income as that measured by the weighted sum of the income in a number of years, and then inspect the fraction of variance in any individual year's income explained by the incomes of other years. Using 3 adjacent years' income for a sample of 24,788 whites, we arrive at the estimates which appear in lines 3-5 of table A1. Alternatively, we may use the correlation of incomes in adjacent years. Assuming that both

TABLE A1  
ANNUAL INCOME AS MEASURE OF PERMANENT INCOME

Sample	Method	Fraction of Observed Variance in Annual Income Due to Variance in Permanent Income
Nonfarm or urban families:		
1. 1935-36	Income elasticity of consumption	.82*
2. 1941		.87*
White veterans:		
3. 1962	Fraction of variance of each year's income explained by 3 other years' incomes	.73†
4. 1963		.71†
5. 1964		.77†
Urban spending units:		
6. 1947, 1948 . . . . .	Correlation of adjacent years' incomes	.83‡
White veterans:		
7. 1962, 1963	Correlation of adjacent years' incomes	.83†
8. 1963, 1964		.83†
All Wisconsin taxpayers:		
9. Average of six coefficients for all consecutive pairs of years, 1929-35 . . . . .	Correlation of adjacent years' incomes	.84‡
White veterans and Wisconsin taxpayers:		
10. Various years ...	Best estimate in explaining correlations among incomes at times separated by various numbers of years, assuming serially correlated errors	.70§

\* From Friedman (1957), p. 67.

† Calculated from correlations of yearly income found by Cutright (1969). The additional year was 1958.

‡ From Friedman (1957), p. 187.

§ Data from Friedman (1957) and Cutright (1969). Method described in text.

<sup>40</sup> Friedman (1957) shows that the elasticity of consumption with respect to income is an estimate of the fraction of the variance of observed income attributable to the variance of permanent income. This estimate is based on the assumption that the transient components in annual income are not serially correlated.

income figures are imperfect measures of the underlying permanent incomes, and that the errors in each are uncorrelated, this correlation coefficient is an estimate of the portion of variance of observed income which is due to variations in the permanent income. Various estimates on this basis appear as lines 6-9 in table A1.

It seems likely that the transient components in annual income will be serially correlated. This possibility is indicated by the fact that the correlation coefficients among annual incomes for given individuals decline as the number of years intervening between the years increases (Friedman 1957; Cutright 1969). Thus we may assume that the estimates in lines 1-9 represent overestimates of the fraction of the variance of observed annual income due to variations in permanent income.

Gintis has attempted to deal with the problem of serially correlated errors in a recent unpublished paper (1970). He used the following basic model: Let

$$y_t^i$$

be the observed annual income of the  $i$ th individual in year  $t$ ;

$$y_p^i$$

is the permanent income of the  $i$ th individual, and

$$e_t^i$$

is the deviation from the permanent income in year  $t$ . Then

$$y_t^i = y_p^i + e_t^i \quad (\text{A7})$$

and

$$e_{t+1}^i = \rho e_t^i + u_{t+1}^i \quad (\text{A8})$$

where the  $u^i$  are serially correlated. Gintis then showed that the correlation between incomes in  $h$  years apart,  $r_h$ , will be a function of the autoregressive pattern as described by  $\rho$ , and  $\text{var}(e)/\text{var}(y_p)$ , the ratio of the variance of  $e$  to the variance of permanent income. Thus,

$$r_h = \frac{1 + \rho^h [\text{var}(e)/\text{var}(y_p)]}{1 + [\text{var}(e)/\text{var}(y_p)]}. \quad (\text{A9})$$

Using this equation and the data from Friedman and Cutright, I have arrived at the following estimates:  $\text{var}(e)/\text{var}(y_p) = .43$  and  $\rho = .5$ . These estimates imply that 70 percent of the variance of observed income is attributable to the variance of permanent income.<sup>41</sup> The discrepancy between this estimate and those mentioned previously is explained by the fact that the lower estimate is based on the assumption that errors in observed income are serially correlated. Because the Gintis model fits the data so well,<sup>42</sup> I conclude that the assumption of serially correlated errors is appropriate. I will therefore use

<sup>41</sup> In arriving at this estimate I have eliminated observations based on correlations with an income reported very early in the work experience of the respondent. It seems likely that the error variance of this early year's reported income is atypically large, and that the average annual rate of increase of observed income over this period is quite high. Both of these characteristics of the pattern of earnings in the early years will cause estimates based on expression (A9) to overstate the size of the error component in annual income.

<sup>42</sup> Simple models in which, assuming serially uncorrelated errors, ( $\rho = 0$ ) would

$$.84 (= \sqrt{.70})$$

as the estimated correlation between observed and permanent income<sup>43</sup> (see col. 3 of table 2).

### 3. *Errors in Reporting Education and Occupation*

On the basis of the 1950 Post-Enumeration Survey, the Census Bureau published matrices recording responses to the original census enumerators and to the Post-Enumeration surveyors. I have scaled these responses using the Duncan occupational status scale, as well as his educational attainments scale, and correlated the two responses. The correlations were .86 for educational attainment and .83 for occupational status.<sup>44</sup>

In order to use these correlation coefficients to estimate the error variance as a fraction of the variance of observed occupational status and educational attainment, two basic assumptions must be made: one concerning the accuracy of the original census relative to the Post-Enumeration Survey, and the other concerning the degree to which errors in reporting to the original census are correlated with errors in reporting to the Post-Enumeration Survey. Because the *Current Population Survey* data upon which my mobility estimates are based were collected by highly trained census enumerators, I think it reasonable to assume that these data are about as accurate as the Post-Enumeration Survey which also was conducted by a well-trained staff, and that both are highly reliable by comparison with the general census data. Because the Post-Enumeration Survey took place very shortly after the census, it also seems reasonable to assume that the errors in both sources are positively correlated. I have estimated various measures of the accuracy of the Post-Enumeration Survey, based on alternative assumptions concerning the degree of correlation of errors, and the relative accuracy of the Post-Enumeration Survey and the census. These estimates are presented in table A2. I have chosen the middle assumptions for each as the basis for the estimate of the correlation between observed and true variables for educational attainment and for occupational status.<sup>45</sup> These figures appear in column 3 of table 2.

### 4. *Correlation of Errors and Internal Consistency of the Reliability Estimates*

The evidence from the Duncan and Blau census follow-up study (1967, pp. 457-

not predict the observed pattern of correlation coefficients, even as a rough approximation.

<sup>43</sup> Data from the 1950 Post-Enumeration Survey of the U.S. census suggests that the correlation of reported incomes in the same year to two separate surveys is .80. This figure presumably represents pure reporting error, as it does not contain transient year-to-year variations in income. If we assume that the error component in the census is twice as great as in the Post-Enumeration Survey, and further that the correlation between errors in reporting to these two surveys is .5, the estimated correlation between reported and actual annual income is .87. This calculation suggests that most of the "error" in reported annual income is due to erroneous reporting rather than to transience of annual income.

<sup>44</sup> Although the respondents' occupational status score is not used in the estimated equations, the estimated reliability of this variable will be used in calculating other reliabilities.

<sup>45</sup> Let  $r_p$  and  $r_c$  represent the correlations between the true measure and that measured by the Post-Enumeration Survey and the census, respectively. Further, define  $r_{pc}$  as the correlation between census and Post-Enumeration responses, and  $r_u$  as the correlation of errors in the two measures. Then (using equation [7]) we have

$$r_{pc} = r_p r_c + r_u \sqrt{1 - r_c^2} \sqrt{1 - r_p^2}.$$

TABLE A2  
ESTIMATED CORRELATION OF TRUE AND OBSERVED VALUES FOR RESPONDENTS' REPORTS  
OF OWN OCCUPATIONS AND YEARS OF SCHOOLING

ERROR VARIANCE OF CENSUS/ ERROR VARIANCE OF PES	CORRELATION OF ERRORS IN CENSUS AND PES REPORTS		
	.0	.5	.7
	Occupation		
1.0 .....	.935	.870	.770
1.0 .....	.959	.923	.882
2.0 .....	.970	.950	.930
	Education		
1.0 .....	.930	.850	.730
2.0 .....	.953	.910	.865
3.0 .....	.965	.940	.920

62) yielded the figure .8 as the estimate of the correlation between the observed and true occupational status of parents. It was further assumed that the figure of .8 also represents the correlation between reported and real educational attainments of parents. It remains to be shown, first that these estimates, together with other evidence on reliabilities and with the fragments of independent evidence available, imply a negligible correlation of errors in reporting occupational status and parents' income. Second, I will show that the reliability estimates themselves are consistent with the available data and with each other.

In order to do this I adopt the following procedure: First define  $r_{oc}$  as the correlation between the true occupational status and true educational attainment. Let a superscript  $p$  indicate that the correlation refers to the father or head of household of the respondent. Let the superscript prime ( $'$ ) indicate the observed correlations, and the subscript  $u$  indicate that the correlation refers to the correlation of error terms rather than the variables themselves.

Using equation (7), I can now write the observed correlations as a function of the true correlations, the correlations of the observed with the true variables, and the correlations of the error terms. Thus:

$$r_{oc}^{p'} = r_{oc}^p r_o^p r_c^p + r_{uoc}^p \sqrt{1 - r_o^{p2}} \sqrt{1 - r_c^{p2}} \tag{A10}$$

$$r_{oc}' = r_{oc} r_o r_c + r_{uoc} \sqrt{1 - r_o^2} \sqrt{1 - r_c^2} \tag{A11}$$

We have some independent information on the trend in the correlation between occupational status and educational achievement. This evidence, in turn, will allow some inferences about the relationship between

$$r_{oc}^p$$

and  $r_{oc}$ . The evidence of a cohort analysis of the occupational changes in a generation sample suggest no trend in the correlation between occupational status and educational attainment.

Because the respondent's own occupational status is reported for different points in the individual life cycle for different age cohorts, not much can be inferred from the correlations between respondent's own occupational status and educational attainment. However, respondents were asked to report their parents'

occupational status at a roughly similar time in the life cycle of the parents, namely, when the respondents were about 16. Thus the correlations among the parents' variables provide evidence largely independent of the position in the life cycle. For the four 10-year age cohorts from 25 to 64, the correlations are (from oldest to youngest): .5313, .4863, .5300, and .4885. There is no apparent secular trend, thus motivating the assumption that<sup>46</sup>

$$r_{oc} = r_{oc}^p.$$

Assuming that

$$r_{uoc} = r_{uoc}^p$$

it can be seen from equation (A10) and (A11) that our estimated reliabilities and the assumption that  $r_{oc} = r_{oc}^p$  imply that the correlations of errors

$$r_{uoe} \text{ and } r_{uor}^p$$

are negligible. It is assumed on this basis that the correlation of errors in reporting education and income is also zero. Assumption of a (perhaps more plausible) positive correlation of errors would lower the corrected correlation between schooling and income and would thus result in a lower estimate of the independent effect of schooling upon income.

The evidence that there has been no secular trend in the relationship between occupational status and educational attainment may be further checked in a manner which provides evidence on the consistency of the no-trend assumption with our estimated correlations of true and observed variables.

If there has been no trend in the relationship between occupational status and educational attainment, and if our estimate of the accuracy of the respondent's reports of his own and his parents' occupations and education are accurate, then the corrected correlation of educational attainment and occupational status for the 25-34-year-old respondents' parents should be roughly equal to the analogous correlation for the 35-44-year-old respondent's own occupation and education. (The latter age group is selected as that which is most likely to have 16-year-old children, and thus to correspond to the parents' status retrospectively reported by the 25-34-year-old respondents referring to the period roughly 9-19 years ago when they were 16.) The corrected correlation for the 35-44-year-olds is .7676, while that for the parents of 25-34-year-olds is .7633. It should be stressed that the striking similarity of these two correlations demonstrates only the consistency of our estimates. Other estimates might also be consistent, although a little experimentation will show that an alternative set of consistent estimates is not easy to come by.

Table A3 summarizes the corrections; see also tables A4 and A5.

<sup>46</sup> These are the observed correlations reported in Duncan et al. (1968, p. 51). Folger and Nam (1967) present evidence that the degree of association between educational attainment and occupational status declined over the period 1940-60. The Folger and Nam results must be seriously questioned, however. It may be seen from equations (A10) and (A11) that any significant decline in this relationship, namely,

$$r_{\hat{oc}} < r_{oc}^p$$

implies a large negative correlation of errors in reporting occupation and educational attainment. This seems to be highly unlikely.

TABLE A3  
CORRECTED ZERO-ORDER CORRELATIONS AMONG MAIN VARIABLES FOR NON-NEGRO MALES, AGED 25-34, WITH  
NONFARM BACKGROUNDS IN EXPERIENCED LABOR FORCE, 1962

Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)
1. Income .....	1.000	.4345	.3004	.3566	-.1889	.2283	.3004
2. Occupation .....	.840/.923	1.000	.4786	.7751	-.2793	.4632	.4786
3. Parents' income .....	*	*	1.000	.6031	-.1811	.4018	.4900
4. Years of schooling .....	.840/.910	.933/.91	.840/.91	1.000	-.3734	.5518	.5677
5. Number of siblings .....	.840/.960	.923/.96	.840/.960	.910/.96	1.000	-.3504	-.2982
6. Father's education .....	.840/.800	.923/.80	.840/.910	.910/.80	.96/.8	1.000	.7633
7. Father's occupation .....	.840/.800	.923/.80	.840/.923	.910/.80	.96/.8	.8/.8	1.000

NOTE.—Corrected coefficients appear above the diagonal. The first and second number in the cells below the diagonal are the estimated correlation between the true and observed variables denoted by the column and row headings, respectively (for sources see tables 2 and A4).

\* Corrected correlations for these cells are assumed to be equal to the respective corrected correlations of father's occupation with respondent's income and occupation, respectively.

TABLE A4  
ZERO-ORDER CORRELATIONS AMONG MAIN VARIABLES FOR NON-NEGRO MALES, AGED 25-34 WITH  
NONFARM BACKGROUNDS IN EXPERIENCED LABOR FORCE, 1962  
(ORIGINAL DATA AND SOURCES)

Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)
1. Income .....	1.000	.3369	...	.2726	-.1523	.1534	.2019
2. Occupation .....	*	1.000	...	.6510	-.2475	.3420	.3334
3. Parents' income .....	...	...	1.000	.4610	-.1460	.3071	.3799
4. Years of schooling .....	*	*	†	1.000	-.3262	.4017	.4133
5. Number of siblings .....	*	*	†	*	1.000	-.2691	-.2290
6. Father's education .....	*	*	†	*	*	1.000	.4885
7. Father's occupation .....	*	*	†	*	*	*	1.000

\* From Duncan et al. (1968), p. 51.

† Data provided by the California Guidance Study. I am grateful to Marjorie Honzik and the Institute of Human Development at the University of California (Berkeley) for making these data available to me, and to John Conlisk for drawing my attention to the data source and assistance in acquiring the data. These correlations are estimated for a sample of about seventy individuals and their families.

‡ From Duncan et al. (1968), p. 51. These figures are the actual correlations for 55-64-year-old respondents' own income, education, and occupation.

TABLE A5

MEANS AND STANDARD DEVIATIONS OF MAIN VARIABLES FOR NON-NEGRO MALES, 25-34,  
WITH NONFARM BACKGROUNDS IN EXPERIENCED LABOR FORCE, 1962

VARIABLE	STANDARD DEVIATION		
	MEAN*	Uncorrected*	Corrected†
Income .....	\$6,140	\$4,290	\$3,604
Years of schooling .....	12.38	3.04	2.77
Number of siblings .....	3.49	2.86	2.75
Father's education .....	9.17	3.53	2.82
Father's occupation .....	34.59	22.35	17.88

\* From Duncan et al. (1968), p. 51.

† Estimated standard deviation of the true variable, calculated as the product of the uncorrected standard deviation and the estimated correlation between the true variable and the observed variable (from table 2).

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