Regression Toward Mediocrity in Economic Stature

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This paper provides estimates of the correlation in lifetime earnings between fathers and sons. Intergenerational data from the National Longitudinal Survey are used. Earlier studies, conducted for the United States, report elasticities of children's earnings with respect to parent's earnings of 0.2 or less, suggesting extensive integenerational mobility. These estimates, however, are biased downward by error-contaminated measures of lifetime economic status. Estimates presented in this paper correct for the problem of measurement error and find the intergenerational correlation in income to be on the order of 0.4. This suggests considerably less intergenerational mobility than previously believed. (*JEL* J62)

"Few sons, indeed are like their fathers.
Generally they are worse;
but just a few are better."
[Homer, *Odyssey*]

"Like father, like son."
[Old Proverb]

Over a century has passed since Sir Francis Galton wrote his classic paper "Regression Towards Mediocrity in Hereditary Stature," in which he measured the relationship between the heights of children and those of their parents. Galton found that children of short parents tended to be shorter than average, while children of tall parents tended to be taller than average. However, the transmission of height across generations was imperfect. The child gained only two thirds of an inch for each inch the parents exceeded the average: "The Deviates of the Children are to those of the Mid-Parent as 2 to 3" (Galton, 1886 plate IX). Galton concluded that there was "regression towards mediocrity" (p. 546) in height. With the passage of time, heights would tend toward equality.

The idea that attributes of parents are transmitted to children transfers naturally to the economic domain. To what extent do individuals "inherit" their position in the income distribution? The answer to this question is fundamental to issues of equal opportunity and economic justice. Surprisingly, little evidence exists measuring the extent to which economic advantages and disadvantages are transmitted across generations. The few studies conducted for the United States have found the elasticity of children's earnings with respect to parent's earnings to be on the order of 0.2 or less, suggesting extensive economic mobility across generations.[1] Gary Becker (1988) in

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[1]Previous estimates of intergenerational mobility in the United States, using actual income for both fathers and sons, include W. H. Sewell and R. M. Hauser (1975), who report a father–son income elasticity of 0.15, and Jere Behrman and Paul Taubman (1985), who obtain an estimate of 0.18. J. R. Kearl and Clayne L. Pope (1986) report an estimate of 0.18 for a sample of late-19th-century Mormons. Gary Solon (1992), using repeated observations on incomes of fathers and sons from the Michigan Panel Study of Income Dynamics (PSID), obtains much higher estimates, indicating the intergenerational correlation to be at least 0.4. Behrman and Taubman (1990) report similar findings from the PSID. Joseph Altoji and Thomas A. Dunn (1991) report a 0.36 correlation using the National Longitudinal Survey. Lee Soltow's (1965) pioneering Norwegian study found a correlation of only 0.14 for father–son earnings. British estimates have been higher: see, for example, Anthony B. Atkinson et al. (1983) who report an estimate of 0.45. Sociological

his presidential address to the American Economic Association noted: "In every country with data that I have seen...earnings strongly regress to the mean between fathers and sons" (p. 10).

Previous estimates of intergenerational correlations, however, have been criticized on several grounds. Gary Solon (1989) has argued that existing estimates of intergenerational mobility are biased downward by a combination of measurement error and unrepresentative samples. Arthur Goldberger (1989) has suggested that focusing on correlations in earnings across generations might lead economists to underestimate the role of family background on inequality; integenerational links could be stronger for more broadly defined measures of socioeconomic status.

This paper uses data from the National Longitudinal Survey (NLS) to measure the amount of intergenerational economic mobility present in the United States. The NLS provides a sample of 876 independent father–son pairs, and reports earnings and occupational status for up to 15 years. Several estimation strategies are developed to correct for measurement error in reported economic status. The empirical findings contrast sharply with those of earlier research. The elasticity of child's earnings with respect to parent's earnings is found to be on the order of 0.4. This estimate suggests that mobility in the United States is considerably less than had previously been believed.

Section I of this paper will discuss the measurement of mobility and introduce some of the difficulties inherent in such an exercise. Section II specifies the econometric models that will be estimated. Section III describes the data. Section IV discusses the empirical findings, and Section V offers conclusions.

I. The Measurement of Intergenerational Mobility

Two basic approaches have been utilized in measuring the extent of intergenerational earnings mobility. The first approach posits a simple first-order Markov process relating father's and son's outcomes:

$$Y_i^{\text{son}} = \alpha + \beta Y_i^{\text{father}} + \varepsilon_i^{\text{son}}$$

where Y_i is a measure of permanent status, $\varepsilon_i^{\text{son}}$ is a white-noise error term, and the slope-coefficient β measures the degree of regression toward the mean in economic status.[2] Estimates of β close to unity are indicative of limited mobility, while values of β close to zero suggest rapid regression to the mean. This equation should be regarded as a reduced-form equation, with the coefficient β being determined by a multitude of factors, including those cultural and genetic endowments transmitted from parents to children.[3] Since the purpose of this paper is to estimate the projection of son's status on father's status alone, no attempt is made to decompose the coefficient β into its causal components.

A transition matrix provides an alternative way to depict intergenerational mobility. This matrix cross-classifies the income rank, say quartile, of fathers and their sons.

models of economic success have concentrated on occupational attainment (see e.g., Otis Dudley Duncan et al., 1972). The theory literature includes Alan S. Blinder (1976), John Conslisk (1977), Gary Becker and Nigel Tomes (1979, 1986), and Glenn Loury (1981). Paul Menchik (1979) has studied wealth correlations. Lee Lillard and Robert Willis (1978) and Mary Jo Bane and David Ellwood (1986) have conducted studies of intragenerational mobility.

[2]Permanent status is defined analogously to permanent income. If permanent status is measured in logarithms, then β is simply the elasticity of son's status with respect to father's status. If the underlying variances in father's and son's status are equal, then the regression parameter β is equal to the correlation between father's and son's status.

[3]Sociologists and geneticists have developed detailed structural models of the socioeconomic attainment process (see e.g., Duncan et al., 1972; L. L. Cavalli-Sforza and Marcus Feldman, 1973). See Behrman et al. (1980) for a discussion of the relative importance of nature versus nurture in economic outcomes. Arleen Leibowitz (1974) investigates the effect of early human-capital investments on economic attainments.

The column elements P_{ij} of this matrix indicate the probability that a son will attain status-quartile i given that his father has attained status-quartile j. Complete immobility would imply an identity transition matrix, with son's rank correlating perfectly with father's rank. Complete "equality of opportunity" implies that all of the matrix elements equal $1/n$, where n is the number of income groups ranked (i.e., $n = 4$ in the case of ranking by income quartiles). In this case, the son's income rank is independent of the father's rank.

The major empirical difficulty in applying either of these approaches abides in the fact that permanent status is not observed. Ideally, data would be available on the economic status of fathers and sons over their entire working lives. Instead, only a few annual "snapshots" of their economic status are typically observed. Given these observations, a researcher must try to construct an estimate of the father's and son's lifetime, or "permanent," status. This exercise is complicated by the fact that the fathers and sons are typically observed at different stages in their earnings life cycle. This "between-generation" age variation could bias estimates of β.[4] Beyond potential life-cycle biases, short-term proxies for lifetime economic status, such as annual earnings, are tainted by transitory fluctuations. This measurement error causes the variance of observed status to exceed the variance of permanent status, causing ordinary least-squares estimates of β to be biased downward.[5]

II. Econometric Models

In light of the above discussion, the problem is formulated using the following statistical model:

$$Y_i^{\text{son}} = \alpha + \beta Y_i^{\text{father}} + \varepsilon_i^{\text{son}} \tag{1}$$

where Y_i is a measure of permanent status and $\varepsilon_i^{\text{son}}$ is a white-noise error term.[6]

It is not permanent status, Y_i, that is observed, but rather, current status, Y_{it}, for individual i at time t. Current status is specified to be

$$Y_{it} = Y_i + \gamma Y_{it} + w_{it}. \tag{2}$$

This specification assumes observed status, Y_{it}, to be composed of three factors. The first is an individual fixed effect, Y_i, which may be referred to as "permanent status."[7] The second is a set of systematic factors, X_{it}, such as age or labor-market experience, which cause observed status to deviate from permanent status. The third component, w_{it}, is a transitory error component. Essentially, this specification assumes an earnings-experience profile for fathers and sons, with common slope γ and an individual-specific intercept.

The estimation of equation (1) requires an estimate of Y_i, the permanent component of observed status. The estimation strategy proceeds as follows. First, data for fathers and sons are pooled for all the years observed. Equation (2) is then estimated using ordinary least squares on the pooled data. This first-stage "sweeping regression" includes potential experience variables and experience-squared variables as well as year-indicator variables.[8] The residuals resulting from this regression will be referred to as "adjusted current status" and are given

[4]Stephen Jenkins (1987) demonstrates that no general conclusion can be reached on the direction of the bias.

[5]For a discussion of biases in previous estimates of intergenerational earnings correlations, see Solon (1989).

[6]While equation (1) specifies a linear relationship between father's status and son's status, it is possible that a nonlinear relationship is more appropriate, with intergenerational transmission being stronger in some parts of the income distribution than in others (see e.g., Stanley W. Siebert, 1989). Nonparametric-regression results as well as nonlinear specifications indicate that a linear model is appropriate for this sample.

[7]Permanent status is simply the person-fixed effect from the regression in (2) and may be regarded as the composite effect of time-invariant unobserved (or excluded) variables.

[8]The use of a quartic rather than a quadratic specification in experience has little effect on the results.

by

$$(3)\quad Y_{it} - \hat{\gamma}X_{it} = Y_i + (\gamma - \hat{\gamma})X_{it} + w_{it}$$
$$= Y_i + v_{it}$$

where Y_i is permanent status and v_{it} is the error term. This provides an adjusted status measure for both fathers and sons with

$$(4)\quad Y_{it}^{\text{son}} = Y_i^{\text{son}} + v_{it}^{\text{son}}$$

$$(5)\quad Y_{it}^{\text{father}} = Y_i^{\text{father}} + v_{it}^{\text{father}}.$$

The adjusted status measure provides an error-tainted estimate of permanent status. Equations (4) and (5) combined with (1) present a classic errors-in-variables problem with adjusted current status comprising a permanent component and a (possibly serially correlated) transitory error component. Estimation of (1) by regressing the adjusted data from (4) on those of (5) yields biased estimates of the intergenerational-mobility parameter β. Although errors in measuring the dependent variable, Y_i^{son}, would be incorporated into the error term of such a regression, the presence of the transitory error component, v_{it}^{father}, would cause the ordinary-least-squares (OLS) estimate of β to be biased downward with[9]

$$(6)\quad \text{plim}\,\beta_{\text{OLS}} = \beta\left[\frac{\sigma_Y^2}{\sigma_Y^2 + \sigma_v^2}\right].$$

The magnitude of this bias depends on the ratio of signal to total variance, $\sigma_Y^2/(\sigma_Y^2 + \sigma_v^2)$, where σ_Y^2 is the variance of father's permanent status, and σ_v^2 is the variance of father's transitory status.

[9]This assumes that the errors v_{it}^{son} and v_{it}^{father} are mutually uncorrelated and also uncorrelated with the permanent component of status. This specification assumes there is no mean-reversion in the errors (see John Bound and Alan B. Krueger, 1991). While the consistency of β is unaffected by measurement error in the dependent variable, the correlation coefficient, R, is biased by measurement error in either the dependent or independent variables.

Three different approaches to the estimation of (1) in the presence of measurement error are employed:

1. ordinary least squares using averaged father's earnings;
2. instrumenting fathers earnings; and
3. method-of-moments estimates using the covariance restrictions implied by the statistical model.

A. *Ordinary Least Squares Using Averaged Earnings*

The availability of repeated observations on father's earnings, wage, and occupational status makes it possible to proxy father's permanent status with various averages of these measures. Averaging father's status should reduce the ratio of signal to total variance, thus reducing the extent of the errors-in-variables bias. Using a T-period mean of father's status,

$$(7)\quad \bar{Y}_T^{\text{father}} = \sum_{t=1}^{T} \frac{Y_{it}^{\text{father}}}{T}$$

as the independent variable in the regression with the dependent variable from (4) and assuming the errors to be serially uncorrelated would yield an estimate of β with probability limit

$$(8)\quad \text{plim}\,\beta_{\text{avg}} = \beta\left[\frac{\sigma_Y^2}{\sigma_Y^2 + \dfrac{\sigma_v^2}{T}}\right].$$

This procedure reduces the bias in the estimate of β by averaging away the transitory component of the father's earnings. The extent of the bias reduction is an increasing function of the number of periods, T, over which the average is taken. This derivation, however, implausibly assumes that v_{it} is white noise. If, instead, the transitory component of fathers' earnings followed a first-order autoregressive AR(1) scheme,

$$(9)\quad v_{it}^{\text{father}} = \rho v_{it-1}^{\text{father}} + \xi_{it}$$

then the implied probability limit of the averaged estimator under stationarity would be

$$(10)\quad \operatorname{plim}\beta_{\text{avg}} = \beta\left[\frac{\sigma_Y^2}{\sigma_Y^2 + \dfrac{\sigma_\zeta^2}{T(1-\rho^2)}\left(\dfrac{1+2\rho\left[T-\dfrac{(1-\rho^T)}{1-\rho}\right]}{T(1-\rho)}\right)}\right]$$

B. *Instrumenting for Father's Status*

While employing an average of father's status should reduce the errors-in-variables bias by attenuating the "noise" component of measured status, an alternative approach is to employ an instrumental-variables estimator. An appropriate instrument would be correlated with father's permanent status, but uncorrelated with the transitory component of observed status, allowing a consistent estimate of β to be obtained. Two potential instruments are considered.

1. *Instrument Using Duncan Index.*—The first instrumental-variables estimator (IV1) considered uses a four-year average of the father's Duncan index of socioeconomic status to instrument for current earnings and wages in the first stage of a two-stage least-squares estimation procedure.[10] The Duncan index, being based on occupation characteristics rather than individual characteristics, would appear to satisfy the requirements of being correlated with father's long-run status but uncorrelated with transitory status fluctuations. In the present context, however, the Duncan index is unlikely to be an ideal instrument, as the index is itself commonly employed in structural models of status attainment. To illustrate the effect this has on the proposed estimator, suppose son's permanent status is generated by

$$(11)\quad Y_i^{\text{son}} = \alpha + \beta Y_i^{\text{father}} + \delta D_i^{\text{father}} + \varepsilon_i^{\text{son}}$$

where D_i^{father} is the ith father's average value of the Duncan index. This model proposes an independent Duncan-index effect on son's long-run status, after controlling for father's long-run status. This instrumental-variables estimator would have probability limit[11]

$$(12)\quad \operatorname{plim}\beta_{\text{IV1}} = \beta + \frac{\delta\sigma_{\text{D}}(1-\tau^2)}{\tau\sigma_Y}$$

where σ_{D} is the standard deviation of father's Duncan index and τ is the correlation between father's long-run status and father's Duncan index. Equation (12) implies that β_{IV1} is upward-inconsistent for $\delta > 0$ and $0 < \tau < 1$.[12] Estimates using this instrumental variable may consequently be thought of as providing an upper bound on estimates of the intergenerational-mobility parameter β.[13] Combining these estimates with the downward-biased OLS estimates would then provide bounds on the true value of β.

2. *Instrument Using Forward Quasi-Difference.*—If the transitory component of fa-

[10] The Duncan index is a widely used measure of socioeconomic status. It is measured on a scale from 0 to 96, with higher scores associated with more prestigious occupations. The index is generated by a two-step procedure. In the first step, the prestige rankings for a few occupations are regressed on a representative measure of education and earnings for that occupation. The parameters estimated are used to rank other occupations, given levels of education and earnings. The results of the predicted values are scaled to form and index. For a description, see Duncan (1961).

[11] For a derivation in a similar context, see Solon (1992).

[12] This assumes that the Duncan index has a positive ceteris paribus effect on son's earnings and is positively correlated with father's permanent earnings.

[13] It should be noted that D_i is assumed to be nonstochastic in this derivation. If D_i is measured with error and this error is positively correlated with the transitory component of earnings, then this effect would reduce the upward bias induced by the presence of D_i in the structural system determining earnings. This would be the case if transitory fluctuations are associated with changes in occupation. Thus, there is some ambiguity about the direction of inconsistency for this estimator.

ther's status can be parameterized, it is possible to find a consistent instrumental-variable estimator. A first-order autoregressive scheme is used where

$$(13)\quad v_{it}^{\text{father}} = \rho_{\text{father}}\left(v_{it-1}^{\text{father}}\right) + \xi_{it}^{\text{father}}.$$

This specification implies the orthogonality of a "forward quasi-differenced" instrument for father's status:

$$(14)\quad Y_{it+1}^{\text{father}} - \rho Y_{it}^{\text{father}}$$

since the transitory component of current status which depends on errors dated t or earlier,

$$(15)\quad Y_{it} = Y_i + \xi_{it} + \rho\xi_{it-1} + \rho^2\xi_{it-2}\cdots + \rho^k\xi_{it-k} + \cdots$$

is orthogonal to that in the forward quasi-difference,

$$(16)\quad Y_{it+1} - \rho Y_{it} = Y_i(1-\rho) + \xi_{it+1}.$$

Given a consistent estimate of ρ, and assuming that the AR(1) specification holds, this instrument permits a consistent two-stage least-squares estimate of β. An estimate of ρ may be obtained by estimating the following relation implied by equations (4) and (5):

$$(17)\quad \operatorname{Cov}\left(Y_{it}^{\text{father}}, Y_{is}^{\text{father}}\right) = \operatorname{Var}\left(Y_i^{\text{father}}\right) + \rho_{\text{father}}^{|t-s|}\operatorname{Var}\left(v_{it}^{\text{father}}\right) \qquad t > s.$$

This equation may be estimated by nonlinear least squares.[14]

[14]This specification is similar to that used by Lillard and Willis (1978) and seems warranted by the data used in this study. The estimation of equation (17) explains over 99.8 percent of the variation in the earnings and wage covariances, and the autocorrelation function declines geometrically.

C. *Complete-Model Estimation Using Covariance Restrictions*

A natural extension of the forward quasi-difference instrumental-variable estimator is to impose the covariance restrictions implied by the complete statistical model, (1)–(5) and (9), and to estimate the unknown parameters by minimizing the distance between the theoretical and empirical moments. The moments implied by the model are as follows:[15]

$$(18)\quad \operatorname{Var}(Y_{it}^{\text{son}}) = \operatorname{Var}(Y_i^{\text{son}}) + \operatorname{Var}(v_{it}^{\text{son}})$$

$$(19)\quad \operatorname{Var}(Y_{it}^{\text{son}}) = \operatorname{Var}\left(Y_i^{\text{father}}\right) + \operatorname{Var}\left(v_{it}^{\text{father}}\right)$$

$$(20)\quad \operatorname{Var}(v_{it}^{\text{son}}) = \operatorname{Var}(\xi_{it}^{\text{son}}) \times \left\{1 + \rho_{\text{son}}^2 + \rho_{\text{son}}^4 + \cdots + \rho_{\text{son}}^{2t-2}\right\} + \rho_{\text{son}}^{2t}\operatorname{Var}(v_{i0}^{\text{son}})$$

$$(21)\quad \operatorname{Var}\left(v_{it}^{\text{father}}\right) = \operatorname{Var}\left(\xi_{it}^{\text{father}}\right) \times \left\{1+\rho_{\text{father}}^2+\rho_{\text{father}}^4+\cdots+\rho_{\text{father}}^{2t-2}\right\} + \rho_{\text{father}}^{2t}\operatorname{Var}\left(v_{i0}^{\text{father}}\right)$$

$$(22)\quad \operatorname{Cov}(Y_{it}^{\text{son}}, Y_{is}^{\text{son}}) = \operatorname{Var}(Y_i^{\text{son}}) + \rho_{\text{son}}^{|t-s|}\operatorname{Var}(v_{it}^{\text{son}}) \qquad t > s$$

$$(23)\quad \operatorname{Cov}\left(Y_{it}^{\text{father}}, Y_{is}^{\text{father}}\right) = \operatorname{Var}\left(Y_i^{\text{father}}\right) + \rho_{\text{father}}^{|t-s|}\operatorname{Var}\left(v_{it}^{\text{father}}\right) \qquad t > s$$

$$(24)\quad \operatorname{Cov}\left(Y_i^{\text{son}}, Y_{is}^{\text{father}}\right) = \beta\operatorname{Var}\left(Y_i^{\text{father}}\right)$$

[15]This derivation also specifies father's and son's transitory income as following an AR(1) process. The errors of the autoregressive components are assumed to be uncorrelated between fathers and sons.

where the $\mathrm{Var}(v_{i0})$ represents the initial variance. Estimation requires fitting the following parameters:

$$M(q) = \{\mathrm{Var}(Y_i^{\mathrm{son}}), \mathrm{Var}(Y_i^{\mathrm{father}}), \mathrm{Var}(x_{it}^{\mathrm{son}}), \mathrm{Var}(x_{it}^{\mathrm{father}}), \rho_{\mathrm{son}}, \rho_{\mathrm{father}}, \mathrm{Var}(v_{i0}^{\mathrm{son}}), \mathrm{Var}(v_{i0}^{\mathrm{father}}), \beta\}.$$

Stacking the empirical moments into a vector **M**, an equally weighted minimum-distance estimator is employed (see Gary Chamberlain, 1982; John M. Abowd and David Card, 1989).[16] This estimator minimizes the sum of squared errors $[\mathbf{M}(\theta) - \mathbf{M}]'[\mathbf{M}(\theta) - \mathbf{M}]$ with respect to θ. Standard errors for the estimated parameters are derived under the assumption of data normality. In this case, the variances of the second moments are functions of the second moments themselves, allowing standard formulas (which require an estimate of the kurtosis matrix) to be used in forming the standard errors for the parameters.[17] In addition to providing an estimate of the intergenerational-mobility parameter β, the method-of-moments estimator provides an estimate of the ratio of signal to total variance for each year of the data on fathers and sons. This provides sufficient information to calculate the β implied under the OLS or averaging schemes by using the relations shown in equations (6) and (10). This provides an internal consistency check of the model, since the implied estimates may be compared to the actual OLS and averaged estimates, with large discrepancies between the actual and predicted values being indicative of misspecification.

III. The Data

The data analyzed below are from the National Longitudinal Survey (NLS). The original NLS was initiated in 1966 and comprised four cohorts, each with approximately 5,000 respondents.[18] A number of households in the survey yielded more than one respondent. Given household and relationship identifiers, it is possible to match related "pairs." For the purpose of this study 876 father–son pairs from the "mature men" and "young men" cohorts were employed using data through 1981. It is important to note that the number of observations used in the different analyses varies with the number of missing relevant values for observation. In selecting the sample, the eldest son was retained for families yielding more than one match. This preserves independence across observations and attempts to reduce the potential life-cycle bias by retaining the son farthest out on his earnings life cycle.[19] Furthermore, only observations for which both the father and son were fully employed are used.[20] This selection is designed to capture earnings measures as close as possible to average, or "permanent," earnings. In effect, this selection eliminates much of the transitory variation in the earnings data associated with the transitions of students to full-time work and of older workers to retirement. All monetary variables used were deflated into 1981

[16]An optimal weighting scheme is not employed because the data are unbalanced, allowing the matrix of fourth moments to be singular.

[17]C. Hendricks Brown (1983) shows that, under the assumption of data normality and randomly missing observations, the asymptotic covariance between pairwise covariances s_{ij} and s_{gh} is approximately $[N_{ijgh}/(N_{ij}N_{gh})](s_{ig}s_{jh} + s_{ih}s_{gh})$ where N_{ijgh} is the number of observations with complete information on the i, j, g, and h variables and N_{ij} and N_{gh} are defined analogously. Estimating the elements of **V** (the fourth-moment matrix) with this relation allows standard errors for the parameters to be estimated using: $(\mathbf{g}'\mathbf{g})^{-1}\mathbf{g}'\mathbf{V}\mathbf{g}(\mathbf{g}'\mathbf{g})^{-1}$ where $\gamma = d\mathbf{M}(\theta)/d\theta$.

[18]Original cohorts sampled were men aged 45–59, women aged 30–44, young men aged 14–24, and young women aged 14–24.

[19]This selection could overrepresent sons who stayed at home until a later age. Using all of the father–son pairs has an insignificant effect on the results.

[20]For the purposes of this study, an individual working on average 30 hours per week, at least 30 weeks per year, was defined as fully employed. Moderate adjustments to this definition have no noticeable effects on the results.

TABLE 1—CHARACTERISTICS OF FATHER–SON SAMPLE

Statistic	1965	1966	1967	1968	1969	1970	1971	1973	1975	1976	1978	1980	1981
Mean of father's log earnings	9.848	9.840	—	9.904	—	9.843	—	—	9.895	—	—	9.828	—
Mean of son's log earnings	—	9.348	9.373	9.303	9.541	9.501	9.540	9.683	9.746	9.815	9.913	9.893	9.857
Mean of father's log wage	—	6.664	6.707	—	6.758	—	6.773	—	—	6.775	6.730	6.729	6.555
Mean of son's log wage	—	6.427	6.461	6.532	6.590	6.585	6.597	6.713	6.746	6.819	6.850	6.814	6.827
Mean of father's Duncan index	—	35.4	35.5	—	36.5	—	35.5	37.2	37.2	37.4	39.1	38.4	39.2
Mean of son's Duncan index	—	29.3	29.8	32.0	35.0	36.9	38.6	40.4	43.9	46.3	46.9	46.7	46.9
Standard deviation of father's log earnings	0.544	0.527	—	0.554	—	0.574	—	—	0.577	—	—	0.549	—
Standard deviation of son's log earnings	—	0.386	0.393	0.532	0.421	0.501	0.483	0.499	0.493	0.498	0.438	0.441	0.529
Standard deviation of father's log wage	—	0.548	0.514	—	0.531	—	0.511	—	—	0.554	0.564	0.619	0.674
Standard deviation of son's log wage	—	0.350	0.360	0.371	0.366	0.426	0.410	0.390	0.402	0.419	0.449	0.411	0.442
Standard deviation of father's Duncan index	—	25.2	25.1	—	24.7	—	25.1	25.4	26.0	25.4	26.0	25.0	26.4
Standard deviation of son's Duncan index	—	18.9	20.5	23.6	23.7	24.8	24.6	25.2	25.8	26.8	25.7	25.1	25.8
Average father's age (1965)	49.7												—
Average son's age (1981)	—												33.8
Father's highest grade, average	10.2												—
Son's highest grade, average	—												14.2

dollars using the consumer price index. The analysis that follows utilizes three measures of economic success:

1. income from wages and salaries;
2. hourly wage; and
3. the Duncan index of socioeconomic status.

The Duncan index provides a more broadly based measure of economic status than earnings or wages. It is possible that such occupational indexes of socioeconomic status contain smaller transitory and life-cycle components of variation and thus provide a more accurate measure of permanent status. The inclusion of the Duncan index in the analysis represents an attempt to address this possibility. Summary statistics for the sample may be found in Table 1. These statistics relate to full-time workers as defined above and are unadjusted for life-cycle effects. It may be seen that the mean age for sons in 1981 is just under 34 years, while the mean age of the fathers is 49.7 years in 1965. This pairing represents the earliest observational date for fathers and the latest observational date for sons. This fact underscores the need for a life-cycle adjustment to the reported data. The young men's earnings are initially below those of the mature men but grow until they reach near equality with fathers' average earnings in the latest years of the sample.

IV. Empirical Results

This section presents estimates of the intergenerational-mobility parameter, β, using the three measures of economic status mentioned above. Transition-matrix results are also provided. The data used are adjusted for heterogeneity in labor-force experience. Estimation procedures, as discussed above, include ordinary least squares, ordinary least squares using various averages on father's economic status, instrumental-variables estimates, and finally an estima-

TABLE 2—SUMMARY OF RESULTS

Correction for measurement error	Unadjusted for life-cycle differences[a]			Adjusted life-cycle differences		
	Earnings	Wage	Duncan index	Earnings	Wage	Duncan index
No correction				Table 3	Table 4	Table 5
Mean β:	0.248	0.262	0.34	0.294	0.295	0.33
Averaging (four-year)				Table 6	Table 7	Table 8
Mean β (son in 1981):	0.464	0.367	0.371	0.538	0.391	0.33
Mean β (all years):[a]	—	—	—	0.346	0.337	—
Instrument (Duncan index)	—	—	—	Table 9	Table 10	—
Mean β:	—	—	—	0.417	0.485	—
Instrument (ρ difference)	—	—	—	Table 11	Table 12	—
Mean β:	—	—	—	0.36	0.379	—
Minimum distance	—	—	—	Tables 13, 14	Tables 13, 14	—
β:	—	—	—	0.402	0.376	—
Transition matrices	Table 15					

[a]Results for individual years are available from the author upon request.

tion of the complete model using a method-of-moments procedure. A catalogue of the estimations considered, along with a summary of the estimates, is presented in Table 2.

A. *Ordinary-Least-Squares Results*

Tables 3–5 present what might be called the "base case" set of results. The data used are adjusted for life-cycle effects, but no correction is made for the presence of measurement error. The results are presented in a matrix format with the elements of the matrix containing the following information:

1. an estimate of the intergenerational-mobility parameter β, and its standard error, σ_β;
2. the correlation between father's status and son's status;
3. the standard deviations for son's status and father's status; and
4. the sample size.

This information is provided for various cross-sectional pairwise combinations of father's status and son's status. For example, the first cell of Table 3 indicates that a regression of the logarithm of son's earnings from wages and salaries in 1981 on the logarithm of father's earnings in 1965 yields an estimate for β of 0.400 with a standard error of 0.059. The corresponding correlation coefficient is 0.378. Son's earnings and father's earnings have standard deviations of 0.516 and 0.486, respectively, and 277 father–son pairs are used in the estimation. Cross-sectional estimates of β found in Table 3 range from 0.195 to 0.428, with the average cross-sectional estimate being 0.294. Estimates using an hourly-wage measure are slightly higher, with a mean estimate of 0.295 and a range of 0.225–0.363 being found in Table 4. When the Duncan index is used as the measure of status, estimates increase further, with Table 5 having a mean of 0.330 and a range of 0.251–0.429. This could be regarded, prima facie, as evidence supporting Goldberger's (1989) speculation that general socioeconomic indexes might be more highly correlated across generations than purely pecuniary measures. Table

TABLE 3—LOG EARNINGS FROM WAGES AND SALARIES: OLS ADJUSTED FOR LIFE-CYCLE EFFECTS

Son	Statistics	Father 1965	1966	1968	1970
1981	β (σ_β)	0.400 (0.059)	0.428 (0.064)	0.400 (0.055)	0.357 (0.062)
	r	0.378	0.385	0.42	0.345
	SD_s, SD_f	0.516, 0.486	0.508, 0.456	0.499, 0.520	0.528, 0.512
	N	277	257	250	250
1980	β (σ_β)	0.331 (0.052)	0.290 (0.057)	0.269 (0.049)	0.201 (0.053)
	r	0.357	0.298	0.332	0.234
	SD_s, SD_f	0.433, 0.467	0.430, 0.443	0.407, 0.503	0.423, 0.492
	N	280	263	249	251
1978	β (σ_β)	0.308 (0.053)	0.304 (0.061)	0.276 (0.049)	0.272 (0.050)
	r	0.322	0.29	0.327	0.317
	SD_s, SD_f	0.451, 0.471	0.449, 0.428	0.421, 0.499	0.433, 0.505
	N	294	273	264	264
1976	β (σ_β)	0.276 (0.055)	0.326 (0.061)	0.253 (0.053)	0.195 (0.056)
	r	0.273	0.299	0.273	0.204
	SD_s, SD_f	0.508, 0.501	0.518, 0.475	0.494, 0.534	0.496, 0.521
	N	310	293	280	278
1975	β (σ_β)	0.382 (0.055)	0.360 (0.063)	0.317 (0.053)	0.310 (0.055)
	r	0.369	0.322	0.337	0.319
	SD_s, SD_f	0.526, 0.508	0.525, 0.469	0.510, 0.542	0.509, 0.525
	N	307	283	276	278
1973	β (σ_β)	0.332 (0.055)	0.296 (0.058)	0.231 (0.050)	0.212 (0.052)
	r	0.315	0.281	0.261	0.233
	SD_s, SD_f	0.509, 0.483	0.501, 0.477	0.467, 0.527	0.477, 0.524
	N	328	307	293	295
1971	β (σ_β)	0.267 (0.060)	0.210 (0.065)	0.219 (0.060)	0.219 (0.061)
	r	0.265	0.205	0.231	0.228
	SD_s, SD_f	0.491, 0.487	0.492, 0.479	0.478, 0.502	0.490, 0.510
	N	259	244	237	237

Notes: Each cell contains (i) estimates of β and its standard error (σ_β), (ii) the correlation coefficient (r) between father's and son's earnings, (iii) the standard deviations of son's income (SD_s) and father's income (SD_f), and (iv) the sample size (N).

2 notes the averages for the estimates unadjusted for life-cycle effects.[21] The unadjusted estimates are comparable to those of earlier studies. Adjusting for the life-cycle factors has a moderate positive effect on the estimates, resulting in an 18-percent average increase in β for earnings and an average 11-percent increase for wages. This finding contrasts with those of Atkinson (1981), who found an adjustment for age heterogeneity to have an insignificant impact on his findings.

B. *Ordinary-Least-Squares Results with Averaging*

Tables 6–8 direct attention toward the possibility of measurement error in the annual proxies for permanent status. Averag-

[21]Not reported in the tables. These and other results summarized in Table 2 but not tabulated are available from the author upon request.

TABLE 4—LOG HOURLY WAGE: OLS ADJUSTED FOR LIFE-CYCLE EFFECTS

		Father			
Son	Statistic	1966	1967	1969	1971
1981	β (σ_β)	0.339 (0.052)	0.288 (0.054)	0.318 (0.055)	0.340 (0.056)
	r	0.381	0.314	0.35	0.372
	SD_s, SD_f	0.421, 0.473	0.428, 0.466	0.418, 0.461	0.423, 0.463
	N	247	264	240	234
1980	β (σ_β)	0.363 (0.052)	0.335 (0.052)	0.345 (0.054)	0.334 (0.056)
	r	0.411	0.373	0.388	0.365
	SD_s, SD_f	0.407, 0.461	0.416, 0.463	0.405, 0.454	0.406, 0.443
	N	241	254	233	233
1978	β (σ_β)	0.263 (0.061)	0.242 (0.063)	0.319 (0.059)	0.349 (0.060)
	r	0.262	0.227	0.328	0.356
	SD_s, SD_f	0.468, 0.467	0.474, 0.444	0.435, 0.447	0.427, 0.436
	N	249	269	241	237
1976	β (σ_β)	0.277 (0.047)	0.296 (0.050)	0.277 (0.050)	0.267 (0.052)
	r	0.328	0.324	0.319	0.299
	SD_s, SD_f	0.420, 0.498	0.436, 0.478	0.434, 0.499	0.430, 0.481
	N	283	296	270	270
1975	β (σ_β)	0.361 (0.047)	0.303 (0.049)	0.270 (0.047)	0.283 (0.049)
	r	0.42	0.342	0.334	0.333
	SD_s, SD_f	0.437, 0.508	0.434, 0.490	0.423, 0.523	0.413, 0.487
	N	274	284	266	265
1973	β (σ_β)	0.294 (0.046)	0.225 (0.047)	0.230 (0.049)	0.230 (0.051)
	r	0.346	0.26	0.274	0.264
	SD_s, SD_f	0.422, 0.496	0.409, 0.472	0.411, 0.489	0.399, 0.458
	N	301	316	278	278
1971	β (σ_β)	0.271 (0.050)	0.285 (0.052)	0.265 (0.054)	0.291 (0.059)
	r	0.327	0.322	0.301	0.308
	SD_s, SD_f	0.426, 0.514	0.428, 0.485	0.424, 0.483	0.428, 0.453
	N	250	260	238	231

Notes: Each cell contains (i) estimates of β and its standard error (σ_β), (ii) the correlation coefficient (r) between father's and son's wages, (iii) the standard deviations of son's wage (SD_s) and father's wage (SD_f), and (iv) the sample size (N).

ing father's status should improve the ratio of signal to total variance, thus reducing the extent of the errors-in-variables bias. Results are presented using son's status in 1981 and two-, three-, and four-year averages of father's status. The estimates corresponding to a four-year averaging of father's income are 0.538 for earnings, 0.391 for wages, and 0.330 for the Duncan index. While the estimate using sons data in 1981 is higher than average, the mean for the averaging estimator when estimated for all years of sons data is 0.346 for earnings and 0.337 for wages. These results indicate considerably less mobility than was previously thought to exist. The Duncan-index results are the least affected by the averaging procedure. This suggests that the Duncan index might provide a better proxy for permanent status than year-to-year measures of earnings or wages.

C. *Instrumental-Variables Results*

An instrumental-variable estimator provides an alternative method to reduce the

TABLE 5—DUNCAN INDEX: OLS ADJUSTED FOR LIFE-CYCLE EFFECTS

		Father			
Son	Statistic	1966	1967	1969	1971
1981	β (σ_β)	0.316 (0.068)	0.381 (0.058)	0.327 (0.063)	0.312 (0.063)
	r	0.278	0.338	0.286	0.269
	SD_s, SD_f	26, 23	26, 23	26, 23	26, 23
	N	262	334	306	314
1980	β (σ_β)	0.295 (0.068)	0.336 (0.058)	0.312 (0.061)	0.295 (0.060)
	r	0.258	0.302	0.279	0.266
	SD_s, SD_f	26, 23	26, 23	25, 23	26, 23
	N	266	337	310	318
1978	β (σ_β)	0.251 (0.066)	0.312 (0.057)	0.265 (0.063)	0.274 (0.060)
	r	0.221	0.279	0.229	0.241
	SD_s, SD_f	26, 23	26, 24	26, 23	26, 23
	N	278	355	321	335
1976	β (σ_β)	0.309 (0.066)	0.347 (0.058)	0.313 (0.065)	0.296 (0.061)
	r	0.261	0.296	0.253	0.247
	SD_s, SD_f	28, 23	28, 24	28, 22	28, 23
	N	299	378	340	358
1975	β (σ_β)	0.357 (0.066)	0.338 (0.057)	0.288 (0.062)	0.276 (0.061)
	r	0.301	0.289	0.242	0.233
	SD_s, SD_f	27, 23	27, 23	27, 22	27, 23
	N	294	381	350	362
1973	β (σ_β)	0.413 (0.060)	0.378 (0.054)	0.345 (0.060)	0.335 (0.056)
	r	0.358	0.333	0.289	0.292
	SD_s, SD_f	26, 22	26, 23	26, 22	26, 23
	N	318	400	365	385
1971	β (σ_β)	0.429 (0.069)	0.422 (0.058)	0.358 (0.064)	0.339 (0.061)
	r	0.361	0.374	0.306	0.299
	SD_s, SD_f	26, 22	25, 22	26, 22	26, 23
	N	260	330	305	319

Notes: Each cell contains (i) estimates of β and its standard error (σ_β), (ii) the correlation coefficient (r) between father's and son's Duncan index, (iii) the standard deviations of son's Duncan index (SD_s) and father's Duncan index (SD_f), and (iv) the sample size (N).

downward bias in least-squares of β. An appropriate instrument would be correlated with father's permanent status but uncorrelated with the transitory component of observed status, allowing a consistent estimate of β to be obtained. Two instruments are considered, and adjusted data are used throughout.

1. *Instrument Using the Duncan Index.*—As discussed above, the estimates generated by instrumenting for father's status with his average Duncan index would yield upward-inconsistent estimates of β, with the magnitude of the bias depending on the size of the ceteris paribus effect of the Duncan index on status and on the degree of correlation between the Duncan index and father's long-run status. Table 9 contains the results for earnings. It may be seen that these estimates, in addition to being more spread out, are also much higher. Estimates range from 0.265 to 0.677, while the mean estimate is 0.417. Table 10 presents the estimates for wages. Again, the

TABLE 6—LOG OF EARNINGS FROM WAGES AND SALARY: OLS ADJUSTED, AVERAGES

First year in average of father's log earnings	Statistic	Father's earnings: Single-year measure	Two-year average	Three-year average	Four-year average
1965	β (σ_β)	0.400 (0.059)	0.446 (0.067)	0.521 (0.070)	0.538 (0.078)
	r	0.378	0.393	0.458	0.448
	SD_s, SD_f	0.516, 0.486	0.498, 0.440	0.492, 0.433	0.502, 0.418
	N	277	240	209	192
1966	β (σ_β)	0.428 (0.064)	0.516 (0.066)	0.531 (0.075)	
	r	0.385	0.467	0.447	
	SD_s, SD_f	0.508, 0.456	0.490, 0.443	0.498, 0.419	
	N	257	219	200	
1968	β (σ_β)	0.400 (0.055)	0.453 (0.066)		
	r	0.416	0.420		
	SD_s, SD_f	0.499, 0.520	0.508, 0.470		
	N	250	224		
1970	β (σ_β)	0.357 (0.062)			
	r	0.345			
	SD_s, SD_f	0.528, 0.512			
	N	250			

Notes: The dependent variable is son's earnings in 1981. Each cell contains (i) estimates of β and its standard error (σ_β), (ii) the correlation coefficient between father's and son's income (r), (iii) the standard deviations of son's income (SD_s) and father's income (SD_f), and (iv) the sample size (N).

TABLE 7—LOG OF HOURLY WAGE: OLS ADJUSTED, AVERAGES

First year in average of father's log wage	Statistic	Father's wage: Single-year measure	Two-year average	Three-year average	Four-year average
1966	β (σ_β)	0.339 (0.052)	0.330 (0.056)	0.374 (0.062)	0.391 (0.066)
	r	0.381	0.355	0.391	0.396
	SD_s, SD_f	0.421, 0.473	0.423, 0.456	0.411, 0.429	0.406, 0.412
	N	247	240	206	188
1967	β (σ_β)	0.288 (0.054)	0.360 (0.062)	0.378 (0.066)	
	r	0.314	0.362	0.376	
	SD_s, SD_f	0.428, 0.466	0.418, 0.420	0.413, 0.410	
	N	264	225	203	
1969	β (σ_β)	0.318 (0.055)	0.374 (0.063)		
	r	0.350	0.383		
	SD_s, SD_f	0.418, 0.461	0.414, 0.424		
	N	240	209		
1971	β (σ_β)	0.340 (0.056)			
	r	0.372			
	SD_s, SD_f	0.423, 0.463			
	N	234			

Notes: The dependent variable is son's wages in 1981. Each cell contains (i) estimates of β and its standard error (σ_β), (ii) the correlation coefficient between father's and son's wage (r), (iii) the standard deviations of son's wage (SD_s) and father's wage (SD_f), and (iv) the sample size (N).

TABLE 8—DUNCAN INDEX: OLS ADJUSTED, AVERAGES

First year in average of father's Duncan index	Statistic	Father's Duncan index: Single-year measure	Two-year average	Three-year average	Four-year average
1966	β (σ_β)	0.316 (0.068)	0.339 (0.068)	0.336 (0.075)	0.330 (0.080)
	r	0.278	0.298	0.285	0.271
	SD_s, SD_f	26, 23	26, 23	26, 22	26, 21
	N	262	256	227	216
1967	β (σ_β)	0.381 (0.058)	0.376 (0.065)	0.383 (0.069)	
	r	0.338	0.320	0.316	
	SD_s, SD_f	26, 23	26, 22	26, 22	
	N	334	297	280	
1969	β (σ_β)	0.327 (0.063)	0.349 (0.068)		
	r	0.286	0.291		
	SD_s, SD_f	26, 23	26, 22		
	N	306	286		
1971	β (σ_β)	0.312 (0.063)			
	r	0.269			
	SD_s, SD_f	26, 23			
	N	314			

Notes: The dependent variable is son's Duncan index in 1981. Each cell contains (i) estimates of β and its standard error (σ_β), (ii) the correlation coefficient between father's and son's Duncan index (r), (iii) the standard deviations of son's Duncan index (SD_s) and father's Duncan index (SD_f), and (iv) the sample size (N).

TABLE 9—LOG EARNINGS FROM WAGES AND SALARIES: INSTRUMENTAL-VARIABLE ESTIMATION USING DUNCAN INDEX FOR INSTRUMENT

Son	Statistic	Father: 1966	1968	1970
1981	β (σ_β)	0.677 (0.127)	0.581 (0.108)	0.648 (0.144)
	N	212	198	199
1980	β (σ_β)	0.443 (0.113)	0.349 (0.101)	0.392 (0.128)
	N	217	202	204
1978	β (σ_β)	0.518 (0.120)	0.397 (0.109)	0.468 (0.133)
	N	223	211	208
1976	β (σ_β)	0.356 (0.112)	0.281 (0.105)	0.342 (0.130)
	N	241	229	226
1975	β (σ_β)	0.526 (0.129)	0.422 (0.115)	0.512 (0.138)
	N	226	215	214
1973	β (σ_β)	0.356 (0.113)	0.268 (0.106)	0.315 (0.122)
	N	247	231	230
1971	β (σ_β)	0.345 (0.147)	0.265 (0.130)	0.292 (0.162)
	N	202	193	189

Notes: Each cell contains (i) estimates of β and its standard error (σ_β) and (ii) the sample size (N).

TABLE 10—LOG WAGE: INSTRUMENTAL-VARIABLE ESTIMATION USING DUNCAN INDEX FOR INSTRUMENT

		Father			
Son	Statistic	1966	1967	1968	1971
1981	β (σ_β)	0.567 (0.121)	0.620 (0.143)	0.619 (0.131)	0.670 (0.151)
	N	203	200	195	193
1980	β (σ_β)	0.517 (0.121)	0.558 (0.135)	0.517 (0.128)	0.589 (0.164)
	N	199	197	193	192
1978	β (σ_β)	0.597 (0.137)	0.651 (0.150)	0.680 (0.155)	0.711 (0.172)
	N	200	200	198	192
1976	β (σ_β)	0.369 (0.099)	0.399 (0.107)	0.390 (0.115)	0.435 (0.127)
	N	229	228	224	221
1975	β (σ_β)	0.433 (0.117)	0.460 (0.125)	0.394 (0.120)	0.483 (0.136)
	N	215	211	209	264
1973	β (σ_β)	0.369 (0.103)	0.398 (0.116)	0.410 (0.121)	0.398 (0.138)
	N	237	238	226	224
1971	β (σ_β)	0.304 (0.123)	0.310 (0.135)	0.338 (0.156)	0.383 (0.215)
	N	196	194	188	184

Notes: Each cell contains (i) estimates of β and its standard error (σ_β) and (ii) the sample size (N).

estimates are noticeably higher, with a mean of 0.485 and a range of 0.304–0.711.

2. *Instrument Using Forward Quasi-Difference.*—The implementation of the forward quasi-difference estimator, using the instrumental variable, $Y_{it-1} - \rho Y_{it}$, requires an estimate of the autoregression parameter ρ. This estimate is generated by fitting the nonlinear covariance restriction given in equation (17). This equation is estimated using nonlinear least squares.[22] The autoregressive parameter, ρ is estimated to be 0.45 for earnings and 0.57 for wages.

Results from this estimator are higher than previous (consistent) estimates but are also considerably more dispersed. Table 11 contains the results for earnings mobility using the ρ-differenced estimator. The mean estimate of β is 0.360, with estimates ranging from 0.243 to 0.609. Table 12 presents the results for wages. The mean estimate is 0.379, with estimates ranging from 0.181 to 0.670. These estimates, like those instrumenting with the Duncan index, exhibit considerable variation across cross sections. As expected, these estimates are typically less than the estimates based on the upward-inconsistent Duncan estimator.

D. *Complete-Model Estimation Results Using Covariance Restrictions*

As mentioned above, the covariance restrictions given by equations (18)–(24) provide both an estimate of the parameter β and sufficient information to calculate the ratio of signal to total variance for each year of fathers' and sons' earnings and wages.

[22] The nonlinear least-squares results are:

$$\text{Cov}(Y_{it}^{\text{father}}, Y_{is}^{\text{father}}) = 0.163 + 0.45^{|t-s|}(0.92)$$

for log earnings and

$$\text{Cov}(Y_{it}^{\text{father}}, Y_{is}^{\text{father}}) = 0.155 + 0.57^{|t-s|}(0.86)$$

for log wages. All of the parameters are significant at the 5-percent level.

TABLE 11—LOG EARNINGS FROM WAGES AND SALARIES: INSTRUMENTAL-VARIABLE ESTIMATION USING FORWARD QUASI-DIFFERENCE INSTRUMENT

		Father		
Son	Statistic	1965	1966	1968
1981	β (σ_β)	0.504 (0.091)	0.609 (0.085)	0.497 (0.086)
	N	240	219	224
1980	β (σ_β)	0.299 (0.079)	0.331 (0.076)	0.249 (0.074)
	N	246	221	226
1978	β (σ_β)	0.379 (0.086)	0.380 (0.086)	0.372 (0.076)
	N	259	233	237
1976	β (σ_β)	0.313 (0.086)	0.332 (0.081)	0.244 (0.083)
	N	275	251	255
1975	β (σ_β)	0.410 (0.087)	0.497 (0.086)	0.431 (0.083)
	N	265	242	251
1973	β (σ_β)	0.295 (0.086)	0.330 (0.086)	0.246 (0.078)
	N	286	259	260
1971	β (σ_β)	0.259 (0.098)	0.349 (0.101)	0.243 (0.102)
	N	225	211	212

Notes: Each cell contains (i) estimates of β and its standard error (σ_β) and (ii) the sample size (N).

TABLE 12—LOG WAGE: INSTRUMENTAL-VARIABLE ESTIMATION USING FORWARD QUASI-DIFFERENCE INSTRUMENT

		Father		
Son	Statistic	1966	1967	1969
1981	β (σ_β)	0.212 (0.143)	0.670 (0.136)	0.388 (0.089)
	N	240	225	209
1980	β (σ_β)	0.412 (0.134)	0.685 (0.135)	0.332 (0.091)
	N	233	219	207
1978	β (σ_β)	0.322 (0.158)	0.521 (0.110)	0.452 (0.098)
	N	240	227	211
1976	β (σ_β)	0.399 (0.105)	0.323 (0.079)	0.322 (0.074)
	N	271	252	243
1975	β (σ_β)	0.330 (0.094)	0.403 (0.078)	0.290 (0.071)
	N	259	243	227
1973	β (σ_β)	0.181 (0.105)	0.294 (0.072)	0.265 (0.074)
	N	287	259	244
1971	β (σ_β)	0.440 (0.122)	0.383 (0.091)	0.336 (0.092)
	N	238	218	207

Notes: Each cell contains (i) estimates of β and its standard error (σ_β) and (ii) the sample size (N).

TABLE 13—MINIMUM-DISTANCE ESTIMATES

Parameter	Log earnings	Log wages
$\mathrm{Var}(Y_i^{son})$	0.140	0.116
	(0.013)	(0.011)
$\mathrm{Var}(Y_i^{father})$	0.169	0.164
	(0.010)	(0.010)
$\mathrm{Var}(\xi_{it}^{son})$	0.085	0.056
	(0.007)	(0.005)
$\mathrm{Var}(\xi_{it}^{father})$	0.070	0.048
	(0.006)	(0.006)
ρ_{son}	0.475	0.476
	(0.090)	(0.106)
ρ_{father}	0.464	0.461
	(0.091)	(0.119)
$\mathrm{Var}(v_0^{son})$	0.102	0.066
	(0.019)	(0.014)
$\mathrm{Var}(v_0^{father})$	0.051	0.043
	(0.024)	(0.019)
β	0.402	0.376
	(0.048)	(0.042)
Value of objective function:	0.029	0.015

Notes: Estimates reported for log earnings use sons data for 1981, 1980, 1978, 1976, 1975, 1973, and 1971 and fathers data for 1975, 1970, 1968, 1966, and 1965. Estimates reported for log wages use sons data for 1981, 1980, 1978, 1976, 1975, 1973, and 1971 and fathers data for 1976, 1971, 1969, 1967, and 1966. Standard errors are given in parentheses.

This makes it possible to use the estimates generated by the method-of-moments estimation procedure to calculate the implied β under OLS and the averaging scheme. This provides an internal consistency check on the model. Estimates for the parameters are reported in Table 13. Adjusted data were used in the calculations. Estimates of 0.402 for earnings and 0.376 for wages are obtained. Table 14 reports the estimated ratio of signal to total variance for earnings. It is found to be 0.655 for fathers and 0.560 for sons under the assumption of stationarity in the variance of v. This in turn implies an OLS estimate of 0.263 and a four-year averaged estimate of 0.336. The actual values for the estimates were 0.294 for OLS and 0.346 for averaging. This suggest an overall coherency to the estimates. Table 14 also reports the estimates for wages. Here, the ratio of signal to total variance (under stationarity) is found to be 0.729 for fathers and 0.617 for sons. The implied OLS and averaged estimates are 0.274 and 0.332, respectively, while the actual estimates were 0.295 and 0.337. Again, the predicted outcomes correspond closely to those predicted by the full statistical model.

E. *Transition-Matrix Results*

Table 15 provides an alternative depiction of the evidence. Using the transition-matrix approach, the data for father's and son's status are allocated into quartile groups. Individuals are placed into groups corresponding to their observed status. The quartiles are then cross-tabulated for fathers and sons. This allows measurement of the probability of a child attaining a given quartile conditional upon his father's status-ranking. This approach allows an investigation into possible asymmetries in mobility across the status distribution. It should be noted that these results are not adjusted for measurement error, and as shown above, this could seriously alter the groupings that are reported. These results are reported to allow comparison with estimates provided by Atkinson et al. (1983) and should be interpreted with caution. The first row for each quartile presents the results for earnings, the second row presents results for wages, and the third row presents results for the Duncan index. In the case of perfect equality of opportunity, all of the cell entries would equal 0.25. A chi-square test strongly rejects this hypothesis for all three measures. Indeed, a consideration of sons' earnings data in 1981 and fathers' data in 1965 reveals that 40 percent of children whose fathers fall into the lowest earnings quartile find themselves in the lowest income group, while 12 percent attain the top quartile. A full 69 percent do not rise above the second earnings quartile. At the top end of the earnings distribution, 41 percent of sons whose fathers are in the top earnings

TABLE 14—DECOMPOSITION OF LOG-EARNINGS AND LOG-WAGE VARIANCE INTO PERMANENT AND TRANSITORY COMPONENTS

Variance	Permanent component: "signal"	Transitory component "noise"	Signal to total variance ratio	Implied β_{OLS} (average actual β_{OLS})	Implied β: four-year average (average actual β_{AVG})
Father:					
Log earnings	0.169	0.089	0.655	0.263 (0.294)	0.336 (0.346)
Log wage	0.164	0.061	0.729	0.274 (0.295)	0.332 (0.337)
Son:					
Log earnings	0.14	0.11	0.56		
Log wage	0.116	0.072	0.617		

Notes: The transitory components reported are those under stationarity. Year-by-year estimates can be calculated using the results in Table 13.

TABLE 15—ESTIMATED QUARTILE TRANSITION PROBABILITIES: EARNINGS, WAGES, AND DUNCAN INDEX

Son's quartile (1981)	Measure	Father's quartile: Top	Second	Third	Bottom
Top	log earnings	0.41	0.25	0.17	0.12
	log wage	0.52	0.15	0.25	0.11
	Duncan index	0.32	0.37	0.18	0.10
Second	log earnings	0.33	0.27	0.22	0.19
	log wage	0.31	0.28	0.18	0.24
	Duncan index	0.37	0.18	0.29	0.21
Third	log earnings	0.17	0.27	0.31	0.29
	log wage	0.11	0.30	0.36	0.35
	Duncan index	0.20	0.23	0.26	0.34
Bottom	log earnings	0.09	0.21	0.30	0.40
	log wage	0.06	0.28	0.21	0.30
	Duncan index	0.12	0.22	0.26	0.35

Notes: The first row for each quartile corresponds to father's log earnings in 1965 (sample size = 278, $X^2 = 33.5$). The second row for each quartile corresponds to father's log wage in 1966 (sample size = 248, $X^2 = 44.15$). The third row for each quartile corresponds to father's Duncan index in 1966 (sample size = 263, $X^2 = 28.03$).

quartile are themselves in the top quartile while 9 percent fall into the lowest group. Within the middle two quartiles, departures from cell equality are less marked, with movements up or down being more symmetric. These estimates resemble those reported by Atkinson et al. (1983) using British data in that there is somewhat more upward mobility from the bottom than downward mobility from the top. The results are similar for wages, with 65 percent of the sons whose fathers are in the lowest quartile finding themselves in the lower half of the wage distribution. A stronger effect is again found at the top of the income distribution, with 52 percent of sons with fathers in the top quartile being located in the top quartile themselves and with only 6 percent falling to the lowest quartile. A full 83 percent of sons whose fathers are in the top quartile of the wage distribution are themselves located in the upper two quartiles of the distribution. Again, the hypothesis of equal cell entries is easily rejected. Results for the Duncan index are similar to those for earnings and wages. Cell equality is again rejected. Bearing in mind the fact that these numbers are not corrected for measurement error, these estimates reinforce the suspicion aroused by the regression results that status mobility is less than previously believed.

V. Conclusions

This paper has used data from the National Longitudinal Survey to measure the degree of intergenerational earnings mobility present in the United States. In particular, it has examined the extent to which the economic outcomes of sons resemble those of their fathers. Previous estimates have found the elasticity of children's earnings with respect to parent's earnings to be on the order of 0.2 or less. These estimates, however, have been based on error-contaminated measures of lifetime economic status. This has induced a downward bias in the reported estimates. I use a variety of estimators in this paper to address the problem of measurement error. These include a simple averaging scheme, instrumental-variables estimates, and a generalized-method-of-moments estimate based on the covariance restrictions implied by the statistical model employed. The principal findings may be summarized as follows.

1. Intergenerational mobility is almost certainly less, and perhaps much less, than was previously thought. An estimate on the order of 0.4 or higher for the elasticity of child's earnings with respect to parent's earnings seems warranted by the evidence.
2. Correcting for measurement error dramatically changes the results, often greatly reducing estimates of mobility. Ratios of signal to total variance are estimated to be 0.655 and 0.729 for fathers' earnings and wages, respectively.
3. Correcting the data for experience heterogeneity has a moderate effect on the results, causing estimated mobility to be reduced.
4. More broadly defined measures of socioeconomic status may provide a better proxy for permanent status than year-to-year measures of earnings or wages.
5. Intergenerational mobility is similar for wages, earnings, and an index of socioeconomic status after corrections for measurement error have been made.
6. Transition-matrix results provide additional support for the regression-based findings.

In considering these results, certain caveats should be borne in mind. First, it is entirely possible that the intergenerational transmission of status is stronger in some parts of the income distribution than in others. It seems plausible that rigidities are strongest at the extremes of the status distribution. Unfortunately, data limitations preclude an adequate investigation of this hypothesis. The estimates provided in this study, which include only the working poor, may themselves be biased downward due to this selection. Secondly, it would be profitable to decompose the estimated correlation into its causal components. This would help to identify those factors that inhibit or promote mobility. Finally, while the size of this sample compares favorably with other studies, a richer data source would increase the range of hypotheses that could be considered.

In spite of these caveats, the weight of the evidence provides the grounds for a healthy skepticism toward the above-mentioned view that the elasticity of children's earnings with respect to parent's earnings is "0.2 or less." Indeed, mobility in the United

States may be considerably less than has previously been believed.

REFERENCES

Abowd, John M. and Card, David, "On the Covariance Structure of Earnings and Hours Changes," *Econometrica*, March 1989, *57*, 411–45.

Altonji, Joseph and Dunn, Thomas A., "Relationships Among the Family Incomes and Labor Market Outcomes of Relatives," National Bureau of Economic Research (Cambridge, MA) Working Paper No. 3724, June 1991.

Atkinson, Anthony B., "On Intergenerational Income Mobility in Britain," *Journal of Post Keynesian Economics*, Winter 1981, *3*, 194–218.

______, **Maynard, A. K. and Trinder, C. G.,** *Parents and Children: Incomes in Two Generations*, London: Heinemann, 1983.

Bane, Mary Jo and Ellwood, David, "Slipping Into and Out of Poverty: The Dynamics of Spells," *Journal of Human Resources*, Winter 1986, *21*, 1–23.

Becker, Gary S., "Family Economics and Macro Behavior," *American Economic Review*, March 1988, *78*, 1–13.

______ **and Tomes, Nigel,** "An Equilibrium Theory of the Distribution of Income and Intergenerational Inequality," *Journal of Political Economy*, December 1979, *87*, 1153–89.

______ **and** ______, "Human Capital and the Rise and Fall of Families," *Journal of Labor Economics*, July 1986, *43*, PS1–PS39.

Behrman, Jere and Taubman, Paul, "Intergenerational Earnings Mobility in the United States: Some Estimates and a Test of Becker's Intergenerational Endowments Model," *Review of Economics and Statistics*, February 1985, *67*, 144–51.

______ **and** ______, "The Intergenerational Correlation Between Children's Adult Earnings And Their Parents' Income: Results From The Michigan Panel Survey of Income Dynamics," *Review of Income and Wealth*, June 1990, *36*, 115–27.

______, **Hrubec, Z., Taubman, P. and Wales, T.,** *Socioeconomic Success: A Study of the Effects of Genetic Endowments, Family Environment, and Schooling*, New York: North-Holland, 1980.

Blinder, Alan S., "Inequality and Mobility in the Distribution of Wealth," *Kyklos*, 1976, *29*, 607–38.

Bound, John and Krueger, Alan B., "The Extent of Measurement Error in Longitudinal Earnings Data: Do Two Wrongs Make a Right?" *Journal of Labor Economics*, January 1991, *9*, 1–24.

Brown, Hendricks C., "Asymptotic Comparison of Missing Data Procedures for Estimating Factor Loadings," *Psychometrika*, June 1983, *48*, 269–91.

Cavalli-Sforza, L. L. and Feldman, Marcus, "Models for Cultural Inheritance I. Group Mean and Within-Group Variation," *Theoretical Population Biology*, March 1973, *4*, 42–55.

Chamberlain, Gary, "Multivariate Regression Models for Panel Data," *Journal of Econometrics*, January 1982, *18*, 5–46.

Conslisk, John, "An Exploratory Model of the Size Distribution of Income," *Economic Inquiry*, July 1977, *15*, 345–66.

Duncan, Otis Dudley, "A Socioeconomic Index for All Occupations," in Albert J. Reiss, Jr., ed., *Occupations and Social Status*, New York: Free Press of Glencoe, 1961, pp. 109–38.

______, **Featherman, D. L., and Duncan, B.,** *Socioeconomic Background and Achievement*, New York: Seminar Press, 1972.

Galton, Francis, "Regression Towards Mediocrity in Hereditary Stature," *Journal of the Anthropological Institute of Great Britain and Ireland*, 1886, *15*, 246–63.

Goldberger, A., "Economic and Mechanical Models of Intergenerational Transmission," *American Economic Review*, June 1989, *79*, 504–13.

Jenkins, Stephen, "Snapshots versus Movies: Lifecycle Biases and the Estimation of Intergenerational Earnings Inheritance," *European Economic Review*, July 1987, *31*, 1149–58.

Kearl, J. R. and Pope, Clayne L., "Unobservable Family and Individual Contributions to the Distributions of Income and Wealth," *Journal of Labor Economics*, July 1986, *4*, 548–82.

Leibowitz, Arleen, "Home Investments in Children" *Journal of Political Economy*, March 1974, *82*, S111–31.

Lillard, Lee and Willis, Robert, "Dynamic Aspects of Earnings Mobility," *Econometrica*, September 1978, *46*, 985–1012.

Loury, Glenn, "Intergenerational Transfers and the Distribution of Earnings," *Econometrica*, July 1981, *49*, 843–67.

Menchik, Paul, "Intergenerational Transmission of Inequality: An Empirical Study of Wealth Mobility," *Economica*, November 1979, *46*, 349–62.

Sewell, W. H. and Hauser, R. M., *Education, Occupation and Earnings*, New York: Academic Press, 1975.

Siebert, Stanley W., "Inequality of Opportunity: An Analysis Based on the Microeconomics of the Family," in R. Drago and Robert Perlman, eds., *Microeconomic Issues in Labor Economics*, New York: Harvester Wheatsheaf, 1989, pp. 177–97.

Solon, Gary, "Biases in the Estimation of Intergenerational Earnings Correlations," *Review of Economics and Statistics*, February 1989, *71*, 172–4.

______, "Intergenerational Income Mobility in the United States," *American Economic Review*, June 1992, *82*, 393–408.

Soltow, Lee, *Toward Income Equality in Norway*, Madison: University of Wisconsin Press, 1965.