

# Intergenerational Income Mobility Among Daughters

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The early empirical literature on intergenerational income mobility in the United States focused mainly on the association of earnings between fathers and sons. This literature suggested that the elasticity of son's earnings with respect to father's earnings is 0.2 or less.<sup>1</sup> Most of these studies, however, used single-year or other short-run measures of father's earnings. Presumably, we should be more interested in estimating the intergenerational association in *long-run* income status, in which case reliance on short-run measures induces a downward errors-in-variables bias. In addition, the peculiar homogeneity of some of the early studies' samples<sup>2</sup> aggravated the errors-in-variables bias by diminishing the "signal" proportion of the sample variation in measured earnings.

A more recent wave of studies has set out to reduce this bias by using intergenerational data from two longitudinal surveys, the Panel Study of Income Dynamics and the National Longitudinal Surveys of labor-market experience.<sup>3</sup> Because these data pertain to national probability samples, they avoid the homogeneity of the

earlier samples. Furthermore, the longitudinal nature of the data has enabled exploration of the empirical importance of using long-run instead of short-run income measures. Most of the evidence from the new studies suggests that the elasticity between the permanent components of son's and father's earnings is about 0.4.

Unfortunately, both the old and new literatures focused very disproportionately on sons.<sup>4</sup> Presumably, this neglect of daughters has stemmed partly from unconscious sexism and partly from a recognition that, in a society in which married women's labor-force participation rates are lower than men's, women's earnings may often be an unreliable indicator of their economic status. Of the few studies that have presented evidence on daughters' intergenerational mobility, some have excluded from their samples the many women with zero earnings, and most have given little attention to the role of husbands' earnings.<sup>5</sup>

This paper presents new evidence on daughters' intergenerational mobility in a framework that encompasses women not in the labor force, uses broader measures of economic status than just the women's own earnings, and highlights the role of husbands' earnings. For purposes of comparison, we perform a parallel analysis of

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<sup>1</sup> Gary S. Becker and Nigel Tomes (1986).

<sup>2</sup> For example, the fathers in Jere R. Behrman and Paul Taubman's (1985) study were drawn from a sample of white male twins in which both members of each twin-pair had served in the armed forces and then cooperated with a succession of surveys.

<sup>3</sup> See Solon (1999) for a detailed survey of the new studies and a discussion of the biases in the earlier literature.

<sup>4</sup> The senior author of the present paper is at least as guilty of this lapse as anyone.

<sup>5</sup> The two studies most like our own are Alexandra L. Minicozzi (1997) and John Shea (1997). Minicozzi estimates a 0.41 coefficient in the regression of the log of a two-year average of the daughter's annual earnings (when ages 28 and 29) on the log of an estimate of the present discounted value of her parents' lifetime earnings. Daughters with zero earnings at age 28 or 29 are omitted from the analysis, and other sources of the daughter's family income, such as husband's earnings, are not considered. Shea estimates a 0.54 coefficient in the regression of the log of a multiyear average of the daughter's annual earnings on the log of a multiyear average of her parents' family income. He also estimates a 0.39 coefficient in the regression in which the dependent variable is instead the log of a multiyear average of the daughter's family income. Shea does not look specifically at the role of husband's earnings in the latter estimate. See Solon (1999) for a more complete summary of the literature.

sons' mobility, which similarly contributes to the sons literature by focusing on *family* income and on the role of wives' earnings. We find that intergenerational transmission of income status may be somewhat weaker for daughters than for sons, but is still quite substantial. We also find that assortative mating is an important element in the intergenerational transmission process.

The next section of the paper presents a simple model that illustrates the potential importance of assortative mating in the intergenerational transmission of economic status. Section II describes our data, Section III lays out our econometric framework, Section IV presents our empirical results, and Section V summarizes and discusses our findings.

### I. Assortative Mating and Intergenerational Mobility

Broadly speaking, the expression "assortative mating" refers to any nonrandomness in the process of who mates with whom. The glossary appended to Elizabeth Epstein and Ruth Guttman's (1984) often-cited survey article defines assortative mating as "character-specific mate selection which would not be expected to occur by chance." The reasons for such systematic mate selection are discussed in economic terms in the theoretical analyses by Becker (1991) and David Lam (1988). The mostly noneconomic empirical literature surveyed by Epstein and Guttman documents positive correlations between spouses with respect to age, physical size, intelligence test scores, religion, ethnicity, and certain values and personality traits. Empirical research by economists has focused mainly on educational attainment and earnings. Michael Kremer (1997), for example, reports that the spouse correlation in years of schooling in the United States is a little above 0.6. Steven J. Haider (1998) reports similar results for schooling and also estimates that the spouse correlation in hourly wage rates is above 0.3. He explains that the spouse correlation in annual earnings is much smaller because the spouse correlation in annual hours of work is slightly negative.

We are interested in the role that assortative mating plays in the intergenerational persistence of economic status. We express that role formally with a rudimentary ver-

sion of the model developed by Lam and Robert F. Schoeni (1993, 1994). For the moment, we simplify by assuming that all daughters marry and participate in the labor force. We characterize the intergenerational determination of daughters' earnings with the regression equation

$$(1) \quad \log E_{wi} = \alpha_w + \beta_w y_{0i} + \varepsilon_{wi}$$

where  $\log E_{wi}$  denotes the permanent component of log earnings for a daughter from family  $i$ ,  $y_{0i}$  denotes the permanent component of her parents' log family income, the error term  $\varepsilon_{wi}$  reflects the combined effects on daughter's earnings of factors orthogonal to parental income, and the slope coefficient  $\beta_w$  is the intergenerational elasticity of daughter's long-run earnings with respect to her parents' long-run income. This elasticity is positive if, as suggested by previous research and our own evidence below, daughters from well-off families tend to earn more.

Following Lam and Schoeni, we assume that assortative mating can be summarized by a correlation  $\gamma$  between the daughter's log earnings and her husband's log earnings:

$$(2) \quad \gamma = \text{Corr}(\log E_{wi}, \log E_{hi})$$

where  $\log E_{hi}$  is the permanent component of the husband's log earnings. Like Lam and Schoeni, we assume that marital sorting on earnings depends only on total earnings and does not depend differentially on the separate components shown on the right side of equation (1). Of course, assortative mating in the real world is much more complex than this. Furthermore, this model does not address the family labor-supply behavior that accounts for both the frequency with which married women do not participate in the labor force and Haider's finding that the correlation between spouses in their earnings is less than the correlation in their hourly wage rates. Nevertheless, the model is sufficient to illustrate some key lessons for the empirical analysis of intergenerational mobility.

First, Lam and Schoeni (1994) show that the regression of the daughter's husband's log earn-

ings on her parents' log income can be written as

$$(3) \quad \log E_{hi} = \alpha_h + \beta_h y_{oi} + \varepsilon_{hi}$$

where the elasticity  $\beta_h$  of the daughter's husband's earnings with respect to her parents' income is

$$(4) \quad \beta_h = \beta_w \gamma \sqrt{\text{Var}(\log E_h) / \text{Var}(\log E_w)}.$$

Thus, if there is no assortative mating on earnings ( $\gamma = 0$ ), this elasticity is zero. On the other hand, with positive assortative mating on earnings, this elasticity is positive. The possibility that this elasticity is substantial in the United States is suggested by Joseph G. Altonji and Thomas A. Dunn's (1991) finding of a 0.26 sample correlation between multiyear averages of the age-adjusted log earnings of the daughter's husband and her father. Lam and Schoeni emphasize that the husband's earnings may be just as elastic as the daughter's own earnings with respect to her parents' economic status. As equation (4) shows, this is possible if assortative mating is strong and husbands' log earnings are more dispersed than wives' log earnings. Indeed, using a sample from York, England, A. B. Atkinson et al. (1983) estimate that the elasticity of the daughter's husband's earnings with respect to her father's earnings is just as great as the elasticity of a son's earnings with respect to his own father's earnings. Lee A. Lillard and M. Rebecca Kilburn (1995) report a similar finding for Malaysia.

Second, the model also has strong implications for the connection between the daughter's family income and that of her parents. Suppose that her family income is comprised entirely of her own earnings and her husband's, and let  $S$  denote her husband's share of their combined earnings. Then the elasticity of her family income with respect to that of her parents is

$$(5) \quad \beta = S\beta_h + (1 - S)\beta_w,$$

the share-weighted average of the separate elasticities of the daughter's own earnings and her husband's. If there is no assortative mating, so

that  $\beta_h = 0$ , and if the husband's earnings are greater than the wife's (as they emphatically are for most of our sample), then the daughter's family income is much less elastic with respect to her parents' income than her own earnings are. On the other hand, suppose that assortative mating is very positive, and  $\beta_h$  is just as large as  $\beta_w$ . Then in the typical family, in which  $S$  is much more than half, the association between the daughter's family income and that of her parents is *mostly* accounted for by her husband's earnings.

In the remainder of this paper, we report on an empirical investigation of daughters' intergenerational income mobility that gives special attention to the role of assortative mating. We also present a parallel investigation for sons. Unlike the simple model in this section, our empirical analysis considers individuals that do not marry or do not participate in the labor force.

## II. Data

Our empirical analysis is based on data from the Panel Study of Income Dynamics (PSID), a longitudinal survey conducted by the University of Michigan's Survey Research Center. The PSID began in 1968 with a national probability sample of about 5,000 families, and it has conducted annual reinterviews each year since. For purposes of intergenerational mobility research, the crucial feature of the survey is that it has followed children from the original families as they have grown into adulthood and formed their own households. As a result, it is now possible to relate the children's income status as adults to the status of their parents as annually reported by the parents themselves since the outset of the survey.

Our daughters sample consists of daughters from the original 1968 sample who also participated in the 1992 survey.<sup>6</sup> Their income reports in that survey pertain to the 1991 calendar year.

<sup>6</sup> We therefore exclude daughters who had disappeared from the survey by 1992 because of death, refusal to cooperate, or inability of the Survey Research Center to locate them. Solon's (1992) discussion of sample attrition notes a tendency for greater attrition among low-income individuals and explains why this probably results in an attenuation bias in the estimation of intergenerational elasticities.

TABLE 1—CHARACTERISTICS OF DAUGHTERS SAMPLE

Variable	Mean	Standard Deviation	Minimum	Maximum
Daughter's age in 1991	33.57	4.76	25.00	40.00
Daughter's log family income in 1991	10.51	0.83	6.91	13.21
1968 household head's age in 1967	39.00	9.11	19.00	91.00
1968 head's log family income in 1967	9.11	0.63	5.30	11.09
1968 head's average of 1967–1971 log family incomes	9.31	0.56	7.24	11.36
Sample size	533			

We use only the Survey Research Center component of the sample, i.e., we exclude the Survey of Economic Opportunity component (the so-called “poverty sample”). We also restrict our analysis to the cohort born between 1951 and 1966. Daughters born before 1951, who were older than 17 at the 1968 interview, are excluded to avoid overrepresenting daughters who left home at late ages. The 1966 birth-year restriction assures that the daughters’ 1991 income measures are observed at ages of at least 25 years. Income observations at younger ages would be particularly noisy measures of long-run status. By the same token, where more than one daughter from the same family meets all of our other sample restrictions, only the oldest is retained in our analysis, because her 1991 status is likely to be a more accurate indicator of her long-run status. Because the PSID’s information on individual earnings pertains to heads of household and wives (including female cohabitators), we restrict our daughters sample to heads and wives.<sup>7</sup> We exclude cases in which family income is nonpositive or the individual earnings variables are imputed by “major assignments.”

Following much of the recent literature, we reduce the errors-in-variables problem from noisy measurement of parents’ long-run income by averaging parental income over multiple years. In particular, we use family income for the years 1967–1971 (as reported in the 1968–1972 interviews) for the 1968 head of house-

hold. We exclude cases in which any of these income observations are missing, nonpositive, or based on major assignments of individual earnings.

As shown in Table 1, the resulting sample contains 533 daughters. The sample’s mean age in 1991<sup>8</sup> is 33.6, and its mean log family income in 1991 is 10.5, which implies a geometric mean of \$36,500 for the level of family income. In the daughters’ families of origin, the mean age in 1967 of the 1968 household heads is 39.0, and the geometric mean of 1967 family income is \$9,062. When we use a five-year average of parental log income for 1967–1971 to smooth out transitory fluctuations, as expected the sample variance declines to 80 percent (the square of 0.56/0.63) of the sample variance for the single-year 1967 measure.

### III. Econometric Framework

Let  $y_{1i}$  denote the permanent component of log family income for a daughter from family  $i$ , and let  $y_{0i}$  denote the same variable for her parents. Following most of the empirical literature on intergenerational mobility, we will express the intergenerational persistence of income status with the regression equation

$$(6) \quad y_{1i} = \alpha + \rho y_{0i} + \varepsilon_i$$

<sup>8</sup> We measure daughter’s age as 1991 minus birth year as recorded in the 1992 interview file. We measure the 1968 household head’s age on the basis of his or her birth year as recorded in the 1983 interview. When that information is missing, we base it instead on the age variable recorded in the 1968 interview.

<sup>7</sup> Our analysis of married couples includes couples that have cohabited at least a year, i.e., since the previous year’s interview. Our sample of married daughters includes 11 cohabitators.

where the error term  $\varepsilon_i$  reflects the combined effects on daughter's income of factors orthogonal to parental income and the slope coefficient  $\rho$  is the intergenerational elasticity of long-run income.

Because the available longitudinal surveys do not track either daughters or their parents long enough to enable direct measurement of permanent income, we model the daughter's log family income in year  $t$  as

$$(7) \quad y_{1it} = y_{1i} + \delta_1 + \gamma_1 A_{1it} + \lambda_1 A_{1it}^2 + \nu_{1it}$$

where  $A_{1it}$  is her age in year  $t$  and  $\nu_{1it}$  is a transitory fluctuation around her long-run income-age profile due to both actual transitory movement and random measurement error. Similarly, we model the parents' log family income in year  $s$  as

$$(8) \quad y_{0is} = y_{0i} + \delta_0 + \gamma_0 A_{0is} + \lambda_0 A_{0is}^2 + \nu_{0is}$$

where  $A_{0is}$  is the age of the parental household head in year  $s$ . The quadratic specifications for the age profiles are less restrictive than they may seem at first, because different quadratics are allowed for the daughter's and parents' generations, which are observed over different age ranges.

The implied relationship between the daughter's log income in year  $t$  and the parents' log income in year  $s$  is

$$(9) \quad y_{1it} = (\alpha + \delta_1 - \rho\delta_0) + \rho y_{0is} + \gamma_1 A_{1it} + \lambda_1 A_{1it}^2 - \rho\gamma_0 A_{0is} - \rho\lambda_0 A_{0is}^2 + \varepsilon_i + \nu_{1it} - \rho\nu_{0is}.$$

If least-squares estimation is performed for this regression of the daughter's log income in year  $t$  on the parents' log income in year  $s$  and age controls for both generations, the correlation between the key regressor  $y_{0is}$  and the error component  $\nu_{0is}$  induces an errors-in-variables bias in the estimation of the intergenerational elasticity  $\rho$ . In particular, if all the error components are uncorrelated with each other, parental permanent income, and both generations'

ages, then the least-squares estimator of  $\rho$  in this regression is subject to the classical errors-in-variables inconsistency

$$(10) \quad \text{plim } \hat{\rho} = \rho\sigma_y^2 / (\sigma_y^2 + \sigma_\nu^2) < \rho$$

where  $\sigma_y^2$  denotes the population variance in parents' permanent status  $y_{0i}$  and  $\sigma_\nu^2$  is the variance of the measurement noise  $\nu_{0is}$ .

Like many recent studies of intergenerational mobility, we will reduce this errors-in-variables bias by measuring parental status with a multiyear average of parental log income. Specifically, we will apply least squares to the regression

$$(11) \quad y_{1it} = (\alpha + \delta_1 - \rho\delta_0) + \rho\bar{y}_{0i} + \gamma_1 A_{1it} + \lambda_1 A_{1it}^2 - \rho\gamma_0 \bar{A}_{0i} - \rho\lambda_0 \bar{A}_{0i}^2 + \varepsilon_i + \nu_{1it} - \rho\bar{\nu}_{0i}$$

where  $\bar{y}_{0i} = \sum_{s=1967}^{1971} y_{0is} / 5$  is the five-year average of the 1968 household head's log family incomes for 1967 through 1971,  $\bar{A}_{0i}$  is his or her average age over those years (which, of course, is his or her age in 1969),  $\bar{A}_{0i}^2$  denotes the average of his or her squared age over those years (which is just two plus the square of age in 1969), and  $\bar{\nu}_{0i}$  averages the measurement noise over the five years. When we apply least squares to this regression, the probability limit of the resulting  $\hat{\rho}$  is the same as in equation (10) except that  $\sigma_\nu^2$ , the single-year noise variance, is replaced by the variance of the averaged noise  $\bar{\nu}_{0i}$ . Under a broad range of assumptions, the latter variance is smaller, which presumably is the main reason that, in Table 1, the sample variance of the five-year average of parental log income is only 80 percent of that of the 1967 value. Accordingly, the errors-in-variables bias in estimating the intergenerational elasticity will be reduced, though not eliminated.<sup>9</sup>

Once we have estimated the intergenerational income elasticity in this way, we will perform a

<sup>9</sup> As shown below, the resulting estimate of  $\rho$  is 0.43. If instead we use single-year measures of parental log income, as in equation (9), we estimate  $\rho$  as 0.32 using 1967 income, 0.34 using 1968 income, 0.38 using 1969 income, 0.41 using 1970 income, and 0.38 using 1971 income.

TABLE 2—ESTIMATED INTERGENERATIONAL ELASTICITIES FOR DAUGHTERS

Dependent Variable	Full Daughters Sample	Married Daughters	Married Daughters Whose Husbands Have Positive Earnings
Log family income	0.429 (0.063)	0.408 (0.055)	0.387 (0.055)
Log of couple's combined earnings		0.386 (0.065)	0.348 (0.063)
Log of husband's earnings			0.360 (0.079)
Log of husband's share of combined earnings			0.012 (0.052)
Sample size	533	372	365

Note: Numbers in parentheses are estimated standard errors.

more detailed analysis of the 70 percent of our daughters sample that is married. For that subsample, in addition to estimating equation (11), we will reestimate the equation with a new dependent variable, the log of the sum of the daughter's earnings ( $E_{wit}$ ) and her husband's earnings ( $E_{hit}$ ).<sup>10</sup> As shown in equation (5), we can write the elasticity  $\beta$  of the couple's combined earnings with respect to the daughter's parents' income as  $\beta = S\beta_h + (1 - S)\beta_w$  where  $\beta_w$  is the elasticity of the daughter's own earnings with respect to her parents' income,  $\beta_h$  is the elasticity of her husband's earnings with respect to her parents' income, and  $S = E_h / (E_w + E_h)$  is the typical share of husband's earnings in combined earnings.

For purposes of inferring the  $\beta$ 's on the right side of equation (5), it is useful to write the log of the couple's combined earnings as

$$(12) \quad \log(E_{wit} + E_{hit}) = \log(E_{hit}) - \log(S_{it})$$

where  $S_{it}$  is the share of husband's earnings in couple  $i$ 's combined earnings in year  $t$ . Then in addition to estimating  $\beta$  by reestimating equation (11) with  $\log(E_{wit} + E_{hit})$  as the dependent variable, we also can reestimate the equation with  $\log(E_{hit})$  as the dependent variable and with  $\log(S_{it})$  as the dependent variable. The difference between the coefficient vectors in these last two regressions is identically equal to the coefficient

vector in the regression with  $\log(E_{wit} + E_{hit})$  as the dependent variable. Estimating the regression with  $\log(E_{hit})$  as the dependent variable, as was done in the English and Malaysian studies mentioned in Section I, produces an estimate of  $\beta_h$ , the elasticity of husband's earnings with respect to the daughter's parents' income. Estimating the regression with  $\log(S_{it})$  as the dependent variable produces an estimate of  $\beta_S$ , the elasticity of the husband's share with respect to the daughter's parents' income. Since  $\beta_S = (1 - S)(\beta_h - \beta_w)$ ,  $\beta_S$  will be close to zero if  $\beta_w \cong \beta_h$ , that is, if the elasticities of the daughter's earnings and her husband's earnings with respect to her parents' income are nearly the same.<sup>11</sup>

#### IV. Results

All of our empirical analyses use least squares to estimate equation (11), with various choices of dependent variable and sample. We begin by estimating the elasticity of daughter's family income with respect to her parents' family income for our full sample of 533 daughters. As shown in the first column of Table 2, the estimated elasticity is 0.43. This estimate is similar to Shea's (1997) estimate of 0.39, previously mentioned in footnote 5. It also is sim-

<sup>10</sup> We measure both earnings variables with the PSID's "total labor income" measures.

<sup>11</sup> Of course, one might try to estimate  $\beta_w$  more directly by reestimating equation (11) with daughter's log earnings as the dependent variable. That approach is awkward, however, because of the frequency with which daughter's earnings are zero. This also is why it is not very useful to write equation (12) instead as  $\log(E_{wit} + E_{hit}) = \log(E_{wit}) - \log(1 - S_{it})$ .

ilar to most recent estimates of the elasticity of son's earnings with respect to father's earnings, but, as we will see below, sons show larger elasticity estimates when parents' family income is the measure of parental status.

In an analysis not shown in the table, we crudely adjust each generation's family income measure for family size and composition by dividing income by the official poverty line for a family of that type. The elasticity of the daughter's "income-to-needs ratio" with respect to that of her parents is estimated at 0.49 (with estimated standard error 0.05), even higher than the elasticity estimate for unadjusted income. This result is suggestive of substantial intergenerational persistence in family structure.<sup>12</sup>

One of the main contributions of our study is to explore the role of assortative mating in intergenerational mobility of married daughters. Toward that end, we now exclude the female household heads who comprise 30 percent of our original sample, and we focus on the remaining 372 married daughters.<sup>13</sup> As shown in the second column of Table 2, when we reestimate the intergenerational elasticity in family income for this subsample, the estimate falls

<sup>12</sup> Two other analyses explore nonlinearity in the intergenerational relation and the impact of age restrictions. To explore nonlinearity, we include the square of  $\bar{y}_{0i}$  as an additional regressor in equation (11). The estimated coefficient of the squared term is only  $-0.004$  (with estimated standard error 0.063). We also redo the analysis in the first column of Table 2 with the sample restricted to daughters at least 30 years old (instead of just 25) in 1991. This analysis is motivated by Robert T. Reville's (1995) finding that the estimated elasticity of son's earnings with respect to father's earnings becomes much higher when the sons are observed in their 30's instead of their 20's. Changing the age floor for our daughters sample from 25 to 30 reduces the sample size to 395 and raises the estimated intergenerational income elasticity from 0.429 to 0.443 (0.070). Performing the same exercise for our sons sample in the first column of Table 3 reduces the sample size to 380 and raises the estimated intergenerational income elasticity from 0.535 to 0.584 (0.062).

<sup>13</sup> This subsample also excludes one married daughter whose own earnings and husband's earnings in 1991 were both zero. Including this case in the subsample raises the estimated intergenerational elasticity in family income from 0.408 to 0.409. We could boost the size of our married-daughters subsample by replacing some of the excluded unmarried daughters with their married younger sisters. Doing so results in estimates similar to those reported in Table 2. This same point applies to our later analysis of married sons.

slightly to 0.41.<sup>14</sup> To begin exploring the roles of each spouse's earnings, we also reestimate the intergenerational elasticity with  $\log(E_{wit} + E_{hit})$ , the log of the sum of the daughter's earnings and her husband's earnings, as the dependent variable. Doing so reduces the estimated intergenerational elasticity slightly further to 0.39.

Because the analysis in the last column of Table 2 will require taking the log of husband's earnings, we are forced to drop seven cases in which husband's 1991 earnings are zero. When we reestimate the intergenerational elasticities with the remaining sample of 365 daughters, the estimates fall to 0.39 when the dependent variable is log family income and 0.35 when it is the log of the couple's combined earnings. Next we follow the procedure discussed in Section III for decomposing the latter estimate into the parts associated with the daughter's earnings and her husband's earnings.

First, we estimate the elasticity of the daughter's husband's earnings with respect to her parents' income. The estimate of 0.36 is very close to the 0.35 estimate for the couple's combined earnings. The fourth row of the table, which reports the estimated elasticity of the daughter's husband's share of their combined earnings with respect to her parents' income, makes explicit that the discrepancy is only 0.01 and is insignificantly different from zero.<sup>15</sup> As explained in Section III, this result suggests that  $\beta_w \cong \beta_h$ , i.e., the elasticities of the daughter's earnings and her husband's earnings with respect to her parents' income are nearly the same.

As shown in equation (5), the elasticity of the couple's combined earnings with respect to her parents' income is just a weighted average of these two elasticities. The two elasticities contribute similarly in the sense

<sup>14</sup> A parallel reestimation for the subsample of 160 unmarried daughters produces an estimate of 0.435 (with estimated standard error 0.112).

<sup>15</sup> Our linear model for  $\log(S_{it})$  is convenient for the accounting framework developed in Section III, but it ignores that  $\log(S_{it})$  is a limited dependent variable with an upper bound at zero. When we perform maximum likelihood estimation of a Tobit model for  $\log(S_{it})$ , we duplicate our linear model result that the estimated coefficient of the parental income variable is small and statistically insignificant. The same goes for our later analysis of married sons.

TABLE 3—ESTIMATED INTERGENERATIONAL ELASTICITIES FOR SONS

Dependent Variable	Full Sons Sample	Married Sons	Married Sons with Positive Earnings
Log family income	0.535 (0.059)	0.541 (0.062)	0.508 (0.058)
Log of couple's combined earnings		0.585 (0.067)	0.552 (0.063)
Log of son's earnings			0.523 (0.077)
Log of son's share of combined earnings			-0.030 (0.046)
Sample size	501	340	338

Note: Numbers in parentheses are estimated standard errors.

that the two numbers averaged together are about the same as each other. In another sense, though, the elasticity of the husband's earnings contributes more importantly. The weight on the husband's elasticity is his share in their combined earnings and, for most couples, that is a majority share. In our regression sample of 365 married daughters, the mean value of the husband's earnings share is 0.71. The median also is 0.71, even the 25th percentile is more than half at 0.54, and the 75th percentile is 0.95. If we use a weight of  $S = 0.71$  in equation (5), we conclude that, if there were no assortative mating in the sense of a zero elasticity between the daughter's husband's earnings and her parents' income, the elasticity of the couple's combined earnings with respect to her parents' income would be only a little more than one-fourth of what it actually is.

So what have we learned so far? First, our various estimates of the intergenerational income elasticity for daughters range between 0.35 and 0.49. While these estimates suggest a considerable degree of intergenerational mobility, they also suggest considerable intergenerational persistence, considerably more than one might have guessed from the early literature on sons. Second, assortative mating appears to play a crucial role. For the typical married daughter, a major factor in the intergenerational transmission of income status is that the elasticity of her husband's

earnings with respect to her parents' income is just as great as the elasticity of her own earnings.

Finally, for comparison, in Table 3 we display the results of a parallel analysis for sons. For our full sample of 501 sons,<sup>16</sup> we estimate an intergenerational family income elasticity of 0.54, even higher than the corresponding 0.43 estimate for daughters. The  $t$ -ratio for the contrast between these two estimates is 1.23, so the contrast is not statistically significant at conventional significance levels. At first, the 0.54 estimate may seem surprisingly high relative to the typical 0.4 estimate in the recent literature on sons' intergenerational earnings mobility. A closer look at the literature, however, reveals that higher estimates are common when parental status is measured by family income rather than father's earnings. In Solon (1992), for example, the regression of son's 1984 log earnings on father's 1967 log earnings produces an elasticity estimate of 0.39, but the regression of son's 1984 log family income on father's 1967 log family income produces an estimate of 0.48.<sup>17</sup> Given that the present study reduces the errors-in-variables bias by using a multiyear average of parental income, it is unsurprising that the elasticity estimate becomes even a little larger. In fact, when we use single-year measures of parental log income in our present sample, the resulting elasticity estimates are 0.43 with 1967 income, 0.46 with 1968 income, 0.45 with 1969 income, 0.43 with 1970 income, and 0.43 with 1971 income.

Restricting the sample to the 340 married sons leaves the estimated intergenerational family income elasticity at 0.54 and also produces an estimate of 0.58 for the elasticity of the couple's combined earnings with respect to his parents' family income.<sup>18</sup> The  $t$ -ratio for the

<sup>16</sup> This sample excludes three sons whose 1991 incomes are recorded as \$1. Including these sons does not affect our estimates by much. For example, including them results in an estimated intergenerational family income elasticity of 0.52 instead of 0.54.

<sup>17</sup> The same pattern appears in Altonji and Dunn (1991), H. Elizabeth Peters (1992), Minicozzi (1997), and Casey B. Mulligan (1997).

<sup>18</sup> This subsample also excludes two cases in which the son's earnings and his wife's earnings in 1991 were both zero. Including these cases raises the estimated intergenerational elasticity in family income to 0.58.

contrast between the 0.54 estimate and the corresponding 0.41 estimate for daughters is 1.62, just short of significance at the 0.10 level. The *t*-ratio for the contrast between the 0.58 estimate and the corresponding 0.39 estimate for daughters is 2.13, which is significant at the 0.05 level. In the last column, where two sons with zero earnings in 1991 are dropped from the sample, the estimated intergenerational family income elasticity decreases to 0.51, and the estimated elasticity for the couple's combined earnings decreases to 0.55. The *t*-ratio of 1.51 for the contrast between the 0.51 estimate and the corresponding 0.39 for daughters is not statistically significant at conventional levels. The *t*-ratio of 2.29 for the contrast between the 0.55 estimate and the corresponding 0.35 estimate for daughters is statistically significant at the 0.05 level.

When the 0.55 elasticity estimate for combined earnings is decomposed to explore the roles of the son's earnings and his wife's earnings, the elasticity for his own earnings is estimated at 0.52. As shown in the last row, the discrepancy between this estimate and the 0.55 estimate for combined earnings is small and statistically insignificant. As found for daughters, it appears again that the elasticities for own earnings and spouse's earnings are similar. In other words, assortative mating clearly is at work in the intergenerational transmission of income status for sons as well as daughters. It is less important, though, in the sense that the wife's elasticity enters equation (5) for sons with a smaller weight than the husband's elasticity receives in the equation for daughters. In our regression sample of 338 married sons, the mean value of the wife's earnings share is 0.29, and the median also is 0.29.

### V. Conclusion

Using the Panel Study of Income Dynamics to estimate intergenerational income elasticities for daughters, we have obtained estimates ranging from 0.35 to 0.49. These estimates are smaller than our corresponding estimates for sons, though in some instances the contrasts between the estimates for daughters and sons fall short of statistical significance. In any case, the estimated intergenerational elasticities for

daughters are quite substantial. We also have found that assortative mating plays an important role in the intergenerational transmission process. Among married offspring, spouse's earnings appear to be just as elastic as the offspring's own earnings with respect to the parents' income.

One direction for future research is to explore whether these results can be replicated with other data sets. This can be checked for the United States with data from the National Longitudinal Surveys of labor-market experience, but analyses of data from other countries ultimately may be even more informative. Not only do data sets from other countries sometimes provide much larger sample sizes; international comparisons may generate valuable clues about the causal processes underlying the intergenerational associations we have measured. Lam and Schoeni (1994), for example, compare the Brazilian and U.S. associations of daughter's husband's wage rate with her father's years of schooling, and they interpret the contrast in terms of differences between Brazil and the United States in assortative mating. More generally, the accumulation of additional evidence on intergenerational mobility by gender should stimulate further theoretical and empirical analysis of how assortative mating, parental decisions to invest in their children, and other factors contribute to intergenerational mobility for both daughters and sons.

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