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"Ability" Biases in Schooling Returns and Twins: A Test and New Estimates

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Abstract

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Keywords

analysis of education, human capital formation, occupational choice, labor productivity

Disciplines

Demography, Population, and Ecology | Family, Life Course, and Society | Social and Behavioral Sciences | Sociology

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“Ability” Biases in Schooling Returns and Twins:

A Test and New Estimates

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Running Head: “Ability” Biases in Schooling Returns

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ABSTRACT

Identical twins long have been used to control for “ability” in efforts to obtain unbiased estimates of the earnings impact of schooling and of biases in estimates that do not control for earnings endowments. This study (1) presents new estimates of schooling returns and of “ability” bias using a new twins sample, (2) develops and applies a test of the significance of that bias, and (3) demonstrates that there may be “ability” bias even if the genetically-endowed component of ability does not affect schooling decisions directly as long as this component of ability is correlated with other family characteristics such as income that do affect schooling and that it is not possible to identify separately these individual components of “ability” bias. The basic empirical result is that, net of measurement error, upward “ability” bias is statistically significant in OLS estimates, causing an overestimate of the schooling impact of 12%.

JEL CLASSIFICATIONS: I21 Analysis of Education, J24 Human Capital Formation, Occupational Choice, Labor Productivity

I. INTRODUCTION

For over two decades – at least dating back to Behrman and Taubman (1976) – identical twins have been used to control for unmeasured earnings endowments or “ability” in efforts to obtain unbiased estimates of the earnings impact of schooling and of the extent of bias in estimates that do not control for such endowments. “Ability” has been used as the rubric for all unmeasured earnings endowments, which may include genetic endowments of ability, pre-school human capital investments, or motivation. The twins studies infer the existence and the extent of “ability” bias by comparing schooling coefficient estimates from individual estimates with those from within identical (monozygotic, MZ) twins estimates. The latter estimates control perfectly for genetic endowments at the time of conception.

One long-standing criticism of the use of estimates obtained by (essentially) differencing across siblings, including estimates based on differences across identical twins, is that such estimates exacerbate biases towards zero due to measurement error (Bishop 1976, Griliches 1979). The bias is exacerbated because differencing does not eliminate the measurement error in schooling but does eliminate the common schooling component, so the impact of measurement error relative to the difference in schooling is much larger than the impact of measurement error relative to the schooling of an individual. Recently there has been a revival of twins studies on the returns to schooling that contrast with the earlier efforts by controlling for measurement error in schooling.

Table 1 summarizes recent twins studies incorporating attention to measurement error in schooling. These have been based on three different twins samples: (1) The Twinsburg Ohio annual Twins Festival sample of female and male twin pairs, 147 MZ pairs from the 1991 round of which are used by Ashenfelter and Krueger (1994), 333 MZ pairs from the 1991-3 rounds of which are used by Ashenfelter and Rouse (1998) and 453 MZ pairs from the 1991-3, 1995 rounds of which are used by Rouse (1998); (2) the National Academy of Science-National Research Council sample with 1973 earnings for white male U.S. twins (who had been veterans), 141 MZ pairs from which are used by Behrman, Rosenzweig and Taubman

(1994);¹ and (3) the Australian Twins Register 1980/2-88/9 sample of female and male twins, 602 MZ pairs of which are used by Miller, Mulvey and Martin (1995).²

TABLE 1 PLACED HERE

There are a number of common features of these studies. First, all in their preferred estimates control for measurement error in schooling by using reports of other respondents. Those using the first and third twins samples use reports from the co-twin and find significant correlations in the errors of the reports of a twin on own schooling and on the co-twin's schooling. Second, the cross-sectional estimates and the within estimates for MZ twins (both controlling for measurement error) suggests the possibility of substantial "ability" bias for the second and third twins samples: Behrman, Rosenzweig and Taubman (1994) report an individual estimate that is over 200% of their MZ within estimate and Miller, Mulvey and Martin (1995) report an individual estimate that is over 160% of their MZ within estimate. Third, the Ashenfelter and Krueger (1994) estimates using the Twinsburg sample are much different in this respect, with the individual estimate *less* (not greater as for the other two samples) than the MZ within estimate, which is consistent with *negative* "ability" bias. However the Rouse (1998) replication with three times as large a sample (including the subsample used by Ashenfelter and Krueger) and the Ashenfelter and Rouse (1998) study using twice as large a sample suggest that this aspect of the Ashenfelter and Krueger results is an artifact of their sample, though in neither of these more recent studies are individual estimates presented with control for measurement error for direct comparison with the within-MZ estimates with control for measurement error. Fourth, only the Ashenfelter and Rouse study presents a formal statistical test for the existence of "ability" bias. However, as the authors point out the estimate of "ability" bias used in the test is itself biased and is based on an assumption that ability is not correlated with any other determinant of earnings than schooling, so the results with regard to the possible importance of "ability" bias are incomplete.³ It is thus not clear whether "ability" bias in estimates of the return to schooling based on twins studies is a statistically significant phenomenon or not.

Our contributions in this study are (1) to present new estimates of the returns to schooling and of “ability” bias based on a new sample of twins and (2) to develop and apply a test of the significance of that bias. We also demonstrate that there may be “ability” bias even if the genetically-endowed component of ability does not affect schooling decisions directly as long as this component of ability is correlated with other family characteristics such as income that do affect schooling and we show that it is not possible to identify separately these individual components of “ability” bias. For our estimates we use a larger sample than in the studies that are summarized above -- 720 female and male MZ twins pairs (plus 278 singletons from MZ pairs) for whom we have complete data. The sample is from the largest birth-certificate-based twins registry in the United States, the Minnesota Twins Registry, which is based on all multiple births in that state in 1936-1955.⁴ Our basic empirical result is that, net of measurement error, upward “ability” bias is statistically significant in OLS estimates, causing an overestimate of the schooling impact of 12%.

II. BASIC MODEL OF WAGES, SCHOOLING AND “ABILITY” BIAS

A. Individual Estimates and Causes of “Ability” Bias

Consider the following linear representation of an equation relating log wages W_{ij} for the i th member of family j to his or her schooling S_{ij} ⁵ and to two sets of unobserved variables representing (i) an endowment that we will call “ability” of the individual i from family j that directly affects earnings, represented by μ_{ij} , and (ii) a random earnings shock that is specific to i in j , inclusive of measurement errors in earnings, represented by v_{ij} :⁶

$$(1) \quad W_{ij} = a + \beta S_{ij} + \mu_{ij} + v_{ij},$$

where β is the effect of schooling. S_{ij} is itself a function of unobserved variables that pertain to the family and to the individual child in the family:

$$(2) \quad S_{ij} = d \mu_{ij} + f_j + u_{ij},$$

where d is the effect of the earnings endowment of the child μ on schooling investment; f_j is the joint influence of exogenous features of the family environment “family background” (e.g., prices, family

income, and parents' human capital characteristics) that affect schooling but do not directly affect earnings; and u_{ij} is a disturbance that affects S_{ij} but not W_{ij} except indirectly through S_{ij} .

As is well known, OLS estimates of (1) are biased if d is nonzero, as is likely to be the case if human capital investments and earnings respond directly to endowments as found in Behrman, Rosenzweig and Taubman (1994, 1996). The standard omitted variable bias result is that:

$$(3) \quad b_{OLS} = \beta + b_{\mu,S},$$

where $b_{\mu,S}$ is the regression coefficient of S in the auxiliary regression of μ on S , where, based on (2),

$$(4) \quad b_{\mu,S} = (ds_{\mu}^2 + s_{\mu,f})/s_s^2.$$

Expression (4) indicates that the bias term is more positive, for a given variance in schooling s_s^2 , the larger is the schooling response to ability d , the greater the variance in ability s_{μ}^2 , and the greater the covariance between ability and household determinants of schooling other than ability $s_{\mu,f}$. Note that the implication of this last covariance term, which does not seem to be well-recognized in this literature, is that in general so-called "ability" bias is due not only to the impact of endowed ability on schooling but also includes effects of correlated family background factors that directly affect schooling. If the correlation between earnings endowments of children and family variables that influence schooling is positive, as would be the case, for example, if some components of earnings ability is transferred genetically, what is called "ability" bias overstates true "ability" bias.⁸

The importance of the presence of the covariance term is that estimated "ability" bias in estimated returns to schooling depends on the extent to which parental resources are important in determining schooling attainment. School finance policies thus can affect "ability" bias, and if these change over time, estimates of "ability" bias drawn from different birth cohorts could yield different estimates of "ability" bias as well as returns to schooling. For example, to the extent that school and other policies reduce over time the impact of family financial constraints on schooling decisions, so-called "ability" bias may also fall. Put another way, OLS estimates of the "returns" to schooling may change over time due solely to changes

in “ability” bias caused by changes, say, in the effects of family background variables on schooling investment, even if there is no actual change in the true returns to schooling.

B. Use of MZ Within Estimates to Test for the Significance of “Ability” Bias.

Under certain assumptions, it is well known that the true return to schooling β can be identified using data on MZ twin-pairs as well as the variance of the unmeasured earnings endowment μ (for identical twins). It is not possible, however, to identify the parameter δ , the effects of the endowment on schooling. It is possible to identify the numerator of the bias term, the covariance of schooling and the endowment in (4) and the variance of schooling (net of measurement error) so that a statistical test for “ability” bias can be carried out. To see this note that equation (1) can be rewritten for a MZ twin pair:

$$(5) \quad W_{ij}^M = a + \beta S_{ij}^M + \mu_j^M + v_{ij}^M,$$

$$(6) \quad W_{kj}^M = a + \beta S_{kj}^M + \mu_j^M + v_{kj}^M,$$

where the superscript M refers to MZ twins. Within-MZ-twin estimators are obtained by differencing (5) and (6). With a within-MZ-twin estimator the common unobserved endowment component μ_j^M is swept out.⁹

For MZ twins, the within-family variances and covariances are:

$$(7) \quad \text{Var}(? W^M) = 2\beta^2 s_u^2 + 2s_v^2,$$

$$(8) \quad \text{Var}(? S^M) = 2s_u^2, \text{ and}$$

$$(9) \quad \text{Cov}(? W^M, ? S^M) = 2\beta s_u^2.$$

It can be seen that dividing (9) by (8), which is the expression for the within-MZ estimator b_{MZ}^w , yields β , and expressions (7) through (9) also provide information on the specific variances of u and v , which thus permit identification of the remaining earnings endowment variance term in (5) or (6), μ^2 . Thus with MZ twins estimators, we can identify the true effects of schooling net of endowments on earnings β and the variance of the earnings endowment.

The omitted variable bias term $b_{\mu,S}$ can also be identified, as the numerator of the bias term, the

covariance between S and μ , is given by:

$$(10) \quad ds_{\mu}^2 + s_{\mu,f} = Cov(W,S) - \beta s_s^2.$$

Thus, a test can be performed for the significance of the omitted variable bias. We estimate and test the significance of the total effect of a change in the earnings endowment on schooling that would be obtained by regressing schooling on the unmeasured earnings endowment, the behavioral relationship that gives rise to “ability” bias, with the unmeasured family factor f omitted:

$$(11) \quad S_{ij} = \gamma \mu_j + e_{ij},$$

The expression for γ has the same numerator as the bias term:

$$(12) \quad \gamma = (ds_{\mu}^2 + s_{\mu,f})/s_{\mu}^2 = d + b_{\mu,f},$$

where $b_{\mu,f}$ is the coefficient obtained by regressing the earnings endowment on the unmeasured family background factor in (2). The MZ twins data can thus be used to identify the compound relationship between the earnings endowment and schooling but not the individual effect of the endowment on schooling d or the relationship between family background and the earnings endowment¹⁰. Testing for the significance of γ is equivalent to testing for the significance of so-called “ability” bias, however!¹

C. Measurement Error in Schooling.

We have assumed that schooling is measured without error. It is well-known that random measurement error in a regressor variable, in this case schooling, biases regression coefficient estimates towards zero. Moreover, within-sibling (twin) estimates are likely to suffer more from measurement error than individual estimates (Bishop, 1976; Griliches, 1979). In the case in which true schooling S_{ij} is measured with random error w , such that

$$(13) \quad S'_{ij} = S_{ij} + w_{ij},$$

the within-MZ moment expressions based on true schooling, (7) through (9), become, with $s_w^2 \dots 0$:

$$(14) \quad Var(\gamma W^M) = 2\beta^2 s_u^2 + 2s_v^2,$$

$$(15) \quad Var(\gamma S^M) = 2s_u^2 + 2s_w^2,$$

$$(16) \quad \text{Cov}(W^M, S^M) = 2\beta s_u^2,$$

from which it can be seen that β is not identified.

Bishop and Griliches show that the bias in the estimated β , b^w , from sibling deviation estimators when the measurement error is not correlated across adult siblings¹² is:

$$(17) \quad \text{plim } b^w = \beta(1 - s_w^2/(s_s^2(1 - \rho_s))),$$

where ρ_s is the correlation across siblings in S . If ρ_s is positive as is the case for schooling, the measurement error bias towards zero in the sibling deviation estimates is greater than is that in the standard individual estimates (for which $\rho_s = 0$) and is even greater for estimates based on twins, for whom correlations in schooling are higher than for singulate siblings. For example, the cross-twin schooling correlations in the combined NAS-NRC Twins and the Minnesota Twin Registry samples that are used in Behrman, Rosenzweig and Taubman (1994) are 0.74 for MZ twins and 0.52 for single-sex DZ twins. If measurement error is about 10 percent of the true variance in schooling for everyone (i.e. $s_w^2/s_s^2 = 0.1$), as suggested by the findings of Bielby, Hauser and Featherman (1977) in their study of schooling reports from the 1973 CPS, the bias towards zero in estimates based on individuals would be nine percent but would be 35 percent for the within MZ estimator.¹³

The procedure that is followed in the studies that are summarized in Table 1 to control for measurement error is to use additional measures of schooling whose measurement errors are not correlated. It can be easily shown that with one additional measure of schooling for each twin and orthogonality across the measurement errors of all four schooling indicators, the true return to schooling¹⁴ is identified. Under this latter assumption, these "instrumental" variables add two equations to the model of the form:

$$(18) \quad S_{ij}' = \rho S_{ij} + w_{ij}', \quad n=i,k;$$

and two additional parameters (ρ and s_w^2) while enabling the computation of three additional independent (within-MZ) moments from the data. Correlated measurement errors across the different reports (i.e., own schooling, co-twin's schooling) also can be identified. The studies that are summarized in Table 1 all find

that control for measurement error increases the estimated return to schooling. The four studies that use the first and third sample all use schooling measures from co-twins and all find significantly correlated measurement errors for a given respondent across reports.

III. DATA

To estimate the effects of schooling on earnings and “ability” biases from using OLS requires data on schooling, actual work experience (so that the schooling effects are not confounded by labor force and job experience) and earnings for adult MZ twins pairs. We use information from a new survey of twins. The new twins data were obtained by resurveying a subset of the twins from the Minnesota Twin Registry (MTR) based on a survey instrument designed by us in collaboration with the Temple University Institute of Survey Research. The MTR is the largest birth-record-based twins registry in the United States, assembled over the 1983-90 period starting with birth records on all twins (both monozygotic, MZ, and dizygotic, DZ) born in Minnesota in 1936-55, with biographical data currently on about 8,400 of the 10,400 surviving intact twin pairs born in Minnesota in that period. Details of the sample and its characteristics are in Lykken *et al.* (1990). The MTR staff obtained from the State Health Department all birth certificates reporting multiple births. Then, through an extensive process, they located over 80% of the twins and sent them a four-page Biographical Questionnaire (BQ) with an introductory Newsletter describing the project and a letter signed by Minnesota's governor urging participation.⁴ 80% of the individuals contacted, 71% of concordant pairs, supplied information for the BQ. The staff of the MTR also has attempted to maintain the participating twins' interest by sending them newsletters about the project and computer-generated, personalized reports on scores on the personality tests administered by them.

Our survey instrument was mailed out in May 1994 to the 5862 members of same-sex pairs who had filled out the BQ and for whom the MTR had current addresses. An additional 776 members of same-sex pairs for whom updated addresses had been located between May and September 1994 were sent

questionnaires in November 1994. 3708 twins returned a completed questionnaire, for a response rate of surviving twins of over 60%.¹⁵ The item response on returned questionnaires is very high, exceeding that on recent Current Population Surveys and the 1990 Census. For example, only 9% of ever employed workers in our sample did not answer the questions on earnings or self employment income; on the CPS more than 20% do not.

The estimates in this paper are based on all of the returned questionnaires from intact female and male MZ twin pairs for which the key schooling, work experience, tenure with current employer and earnings variables are available. This totals 944 women (or 473 female MZ twin pairs) and 496 men (or 248 male MZ twins pairs). For an additional 278 MZ twins pairs (128 females, 150 males) we have usable responses from only one twin in the pair (our “singleton” sample). There are two features of the data that are particularly relevant to the analysis of the impact of schooling on earnings and the biases that occur due to the presence of unobserved earnings endowments. First, there is information on both total lifetime (full-time) work experience and tenure on the last job, which are of particular importance for women, many of whom have interrupted labor force careers. Second, to maximize the size of the sample of female twins that could be used, the earnings on the last job was elicited rather than only earnings in the year prior to the survey. A well-known problem in analyzing wages of women is that many women choose not to be in the labor force for some portion of their working lives and that such labor force participation may be selective (Gronau 1974, Heckman 1974). Only 82% of the women in the sample, for example, worked in 1993. But 97% of the sample members worked at some point in their lives, 91% in the past five years. Completed years of schooling are available for virtually all of the twins. To take into account the variability in work time during the year in which earnings are reported, the wage, earnings and time-worked information was used to construct full-time annual earnings based on either earnings in 1993 or on the last job, inflated in the latter case by the relevant CPI.

It is possible that this twins sample is not representative of all United States' females and males

belonging to the same birth cohorts. Clearly there may be selectivity associated with being born in Minnesota, both siblings reporting all relevant data, or being a MZ twin. However, the characteristics of the intact MZ twin pairs, the characteristics of the sample of MZ twins in which only one twin responded to the survey, and the population of individuals residing in Minnesota in 1990 are quite similar. Table 2 provides descriptive statistics on the individual characteristics of the 1440 twins in our MZ twins pair sample, the MZ twins in our MZ singleton sample, and of adult residents of Minnesota born in the same years represented in the 5% sample for the 1990 Census. Except for the MZ twins pair sample being more female, in most respects the means do not differ much among these three groups. The MZ twins pair sample is more female than the MZ singleton sample and than the Census data because of gender differentials in survey response.¹⁶ Because on average women have significantly lower earnings and slightly higher schooling levels than men, the mean average annualized income and full-time work experience are smaller for the MZ twins pair sample than for the MZ singleton sample, and the former is a little smaller than for the Census despite the fact that the mean schooling is slightly higher for the twins than for the Census sample.

TABLE 2 PLACED HERE

IV. ESTIMATES

The estimation method that we use to implement the model in Section 1 matches the theoretical moments of the system of equations, i.e., $cov(q_{ij}, q_{i'j'})$ for all i, i', j, j' , denoted by S , to the sample moments denoted by X . The objective function $Q = (x - s(p))' W (x - s(p))$ is minimized with respect to the parameter vector p ; x is the vector of elements obtained by stringing out the lower triangular elements of the matrix X , $s(p)$ is the corresponding vector obtained from S which depends on the set of model parameters p , and W is a weighting matrix. The weighting matrix is that which corresponds to the "optimal minimum distance" estimator discussed by Chamberlain (1982, 1984). We use an estimation method that

imposes no distributional assumption, in contrast, for example, to the normality assumption used by Ashenfelter and Krueger (1994) in their estimates based on moment estimators.

Table 3 presents five sets of estimates based on two specifications of the earnings relation (1). The first specification includes only age in addition to schooling, for comparison to the earlier estimates in the literature. The second specification includes instead of age full-time work experience and job tenure. Columns one and two present the estimated parameters for the two specifications of the earnings relation estimated under the OLS assumption that schooling (and experience and tenure for the second specification) is uncorrelated with earnings endowments. The next two columns report the earnings function estimates for the same two specifications from within-MZ estimates, which permit covariation between schooling, work experience, tenure and unmeasured earnings endowments⁴⁷. These columns also report the estimates of β for schooling and for the two experience variables. All columns also report the estimate of the unmeasured earnings endowment variance obtained from the model; these estimates are statistically significant in all specifications.

TABLE 3 PLACED HERE

All four of the sets of estimates are obtained based on the model in which there is potential measurement error in own and twin's schooling reports, with the measurement errors assumed uncorrelated for a twin. In all specifications, both of the measurement error variances are statistically significant, but the errors represent only 6.9% and 7.5% of the true variances in reported schooling for own and cross-twin reports, respectively. For the last column the specification is identical to that in the fourth column except that measurement errors across a respondent's reports of own and of co-twin schooling are allowed to covary. The estimate at the bottom of that column of this correlation ρ_w is not significantly nonzero, in contrast to the estimates reported in the four studies for the first and third samples in Table 1. Therefore we focus on the estimates in which there is no correlation assumed in schooling measurement errors (the point estimates hardly differ).

Our preferred specification is that with full-time work experience plus job tenure included, each of which has a statistically significant positive effect on earnings in each model (controlling only for age does not change the estimates much). The estimated schooling impact parameter from the within-MZ model is 0.104, which is 12% lower than that estimated from a model with the same specification (column 2) but that assumes that earnings endowments are uncorrelated with schooling, full-time work experience and job tenure. The direction of the difference suggests that there is positive “ability” bias in OLS estimates based on individuals, though the magnitudes of the schooling β s across models suggests a much smaller bias than those found in studies using the second and third twins samples reported in Table 1. Each of the individual estimated β s are statistically significant, and indicate that all three human capital variables are positively associated with unmeasured earnings endowments. A^2 test of the joint constraint that all of the β coefficients are zero strongly rejects the null hypothesis that endowments are not related to schooling or work experience ($\chi^2 = 11.34$, 3 d.o.f., significant at the 1% level). Therefore these estimates provide evidence that there is significant “ability” bias in OLS estimates.

Finally, we assessed whether the twins pairs in which both members responded to the questionnaire are different in some way that affects the estimates from those in which only one member of the pair responded (as we note above with regard to Table 2, the former are more female). While the singletons do not contribute to the identification of either the β s or the γ s, it is possible to test the assumption that the singleton sample moments come from the same structure as in the statistically-preferred model specification obtained from the intact twin-pair sample. A^2 test (6.549, for 4 degrees of freedom, not significant at 10% level) indicates that the estimates do not differ significantly for the singletons.

V. CONCLUSION

In this study we present estimates of the returns to schooling and of “ability” bias based on a new sample of twins that is larger than those used in previous studies that control for measurement error. We

also develop and apply for the first time a test of the significance of that bias. We further demonstrate that there may be "ability" bias even if endowed ability does not affect schooling decisions directly as long as ability is correlated with other family characteristics such as income that do affect schooling and we show that it is not possible to identify separately these individual components of "ability" bias. Our basic empirical result is that, net of measurement error, upward "ability" bias is statistically significant in OLS estimates, causing an overestimate of the schooling impact of 12%. We note the "ability" bias that we find is considerably smaller than found in some previous studies. Given that the "ability" bias depends not only on household schooling investment response to schooling, but also on correlated dimensions of family background and related school and school financing policies, it is not surprising that estimates of "ability" bias differ for different time periods and for different countries. "Ability" bias is not a fundamental underlying parameter and may differ importantly depending on the nature of markets and policies for schooling and for labor.

NOTES

1. This is a subsample (for which for both twins information existed to control for measurement error) of the sample of 1019 MZ twins pairs used in Behrman and Taubman (1976) and Behrman *et al* (1980) without control for measurement error.
2. This same sample also is used in Miller, Mulvey and Martin (1997).
3. Ashenfelter and Krueger (1994, p. 1171) also present a t test that rejects the equality of an OLS (with no control for measurement error) and their restricted maximum likelihood estimate of the returns to schooling. But this is a compound test of measurement error, “ability” bias, and other restrictions that are imposed in their maximum likelihood estimates. From this rejection it is not possible to infer whether or not “ability” bias is significant. Behrman, Rosenzweig and Taubman (1994) present in an appendix table the covariance estimates necessary to do the test that we present in Section II below, but do not comment on their implications for “ability” bias nor present the test *per se*.
4. We used the males in this sample for some of our estimates in Behrman, Rosenzweig and Taubman (1994), but that was before we had re-surveyed these twins so at that time we did not have actual earnings data (only nine occupational categories that we linked to census data to represent earnings) nor did we have a means of controlling for measurement error. We used females in this sample for our estimates of the impact of college characteristics in Behrman, Rosenzweig and Taubman (1996), but in that study we did not present estimates of standard wage functions such as are summarized in Table 1.
5. Many studies also include work experience and/or job tenure, as do we in our empirical estimates. But the problems in estimating the effects of experience and tenure are exactly parallel to those for estimating the effects of schooling. So, to keep the discussion as simple as possible, we here include only schooling among the observed variables. For the same reason we leave the discussion of measurement error in schooling to the end of this section. In our estimates we allow experience to be correlated with ability but we do not deal with potential measurement errors in experience due to lack of instruments. None of the cited prior twins studies allow for measurement error in variables other than schooling.
6. Bound et al. (1994) present results on measurement errors in earnings based on company records and self-reports in the PSID and find that measurement errors in earnings are not correlated with schooling in one year of the data but are in the other.
7. If other observed variables such as experience are included in the earnings relation, the auxiliary regression includes those variables among the right-side variables and there are parallel expressions for the biases in the OLS estimated coefficients of those variables.
8. Becker in his 1967 Woytinsky Lecture (1975) emphasizes the importance of the correlation between endowed ability and family background characteristics that affect the ability to finance schooling.
9. With sibling estimators (including fraternal or dizygotic(DZ) twins) the common family endowments, but not the individual-specific endowments, are swept out. Therefore sibling estimators do not eliminate the bias and in fact can result in greater biases than individual estimates as Griliches (1979) noted long ago. The use of partial measures of endowments such as IQ scores, even putting aside the question of whether they truly are exogenous, have the same problem as sibling estimators of only partially controlling for endowments.

10. Behrman, Rosenzweig and Taubman (1994, 1996) show that the household school allocation response to different individual specific endowments can be identified with data on both MZ and DZ twins pairs, and find that households use schooling investments to reinforce such endowment differentials (not to compensate for them).

11. A Hausman test that compared the cross-section and within-estimates of β , taking into account measurement error, also provides a test of “ability” bias. In our estimates, we use the moment restrictions of the model so that our estimates are more efficient than those obtained using regression techniques.

12. These expressions for the biases due to measurement error using within-sib estimators assume that the correlation across siblings in the measurement error (ρ_w) is zero. If this correlation is not zero, for sibling deviation estimates $\text{plim}\hat{\beta}_g = \beta_g(1 - f \cdot s_w^2/s_g^2)$, where $f = (1 - \rho_w)/(1 - \rho_s)$. Note that the bias due to measurement error in the sibling deviation estimates declines as ρ_w increases and is less than the bias in standard individual estimates if $\rho_w > \rho_s$. We are not aware of direct observations on f or on ρ_w . But if twins tend to attend the same schools and the cross-twin correlation is due to unobserved schooling quality, ρ_w may be fairly large. In any case, this measurement problem has not been dealt with in the literature making use of siblings or twins.

13. Bielby *et al.* (1977) used information on the schooling of a sub-set of non-black respondents from the 1973 CPS and reports of schooling from the same respondents in a telephone-based follow-up survey to estimate a measurement model. Their estimates indicated that the ratio of measurement error variance to the true schooling variance was .051 in the 1973 CPS and was .13 in the follow-up survey.

⁰ Among the first six birth cohorts studied (birth years of 1936, 1937, 1938, 1949, 1954, and 1955) 27% of the same-sex pairs identified through birth certificates are known to have been broken by the death of at least one of the members prior to our survey.

⁰ MTR sends custom-made MTR T-shirts to those who answered the survey as part of their ongoing effort to maintain high participation rates in surveys that they administer to the twins in their registry.

16. The Twinsburg sample that is used in three of the studies summarized in Table 1 also is more female than the larger population, with the overall sample for four rounds 59% female.

17. Estimates reported in columns one and two correspond to IV-GLS estimates, those in columns three through five are equivalent to IV-within-MZ estimates, with co-twin reports of schooling used as an instrument for schooling in all specifications, except that we impose the additional restrictions (not rejected by the data) that σ_u , measurement error variances and variances in earnings and schooling shocks and in μ are the same for both twins.

| Table 1. Summary of recent ln earnings relations for identical twins with control for measurement error in schooling | | | | |
|---|---|--|---------------------|--|
| Source | Sample size, year, gender and race composition, and mean age | Preferred estimates^a | | Other variables in IV within MZ estimates |
| | | IV Level | IV Within MZ | |
| 1. Twinsburg, Ohio annual twins festival sample | | | | |
| Ashenfelter and Krueger (1994, Tables 3 and 9) | 147 MZ twins pairs at 1991 Festival -- 54% female, 94% white, mean age 36.5 | 0.116 | 0.129 | |
| Rouse (1998, Table 4) | 453 MZ twins pairs at 1991-3, 1995 Festivals – 59% female, 92% white, mean age 37.9 | 0.111 ^b | 0.119 | tenure, union, married |
| Ashenfelter and Rouse (1998, Table 3) | 333 MZ twins pairs at 1991-3 Festivals - 59% female, 92% white, mean age 38.1 | .113 ^b | .100 | tenure, union, married |
| 2. United States' National Academy of Science-National Research Council twins sample of white veteran males born in 1917-27 | | | | |
| Behrman, Rosenzweig and Taubman (1994, Table 4) | 141 white male MZ pairs, 1973, mean age 51 | 0.101 | 0.050 | |
| 3. Australian Twin Register 1980-2 and 1988-9 | | | | |
| Miller, Mulvey and Martin (1995, Table 5) | 602 MZ twins pairs -- 51.4% female, mean age 36.8 | 0.073 | 0.045 | married |
| <p>^a All the estimates reported here use schooling reports for each twin by another respondent as an instrument to control for measurement error. Behrman, Rosenzweig and Taubman (1994) use the report by each twin's oldest child. The other studies use a report by the co-twin and find significant positive correlated measurement errors across reports by each respondent (i.e., for own and for co-twin's schooling).</p> <p>^b IV individual estimates are not presented. This is the GLS individual estimate. For the subsample used in Ashenfelter and Krueger (1994, Table 3), the GLS individual estimate is 0.088 and the IV individual estimate is 0.116 (32% greater). Rouse (1998, Table 2) also reports that, without instrumenting, for the Ashenfelter and Krueger (1994) subsample the OLS estimate is 0.084 and the within MZ estimate is 0.092 in comparison with her full-sample estimates of 0.105 for OLS and 0.075 for the within MZ estimate.</p> <p>^c Use average income of full-time workers in each two-digit occupational group from the 1986 Census of Population and Housing for earnings variable.</p> | | | | |

Table 2. Descriptive statistics

| Characteristic | MZ twins (Pairs sample) | MZ twins (Singleton sample) | 1990 Minnesota census population (5% sample) |
|--------------------------------|------------------------------------|--|---|
| Annualized income (1993 \$) | 29,754 (22,849) | 33,322 (31,539) | 31,878 (62,766) |
| Years of schooling | 14.1 (2.33) | 14.2 (2.50) | 13.4 (2.45) |
| Age | 46.8 (5.44) | 45.7 (5.45) | 48.0 (5.68) |
| Female | .656 (.475) | .460 (.500) | .480 (.500) |
| Full-time work experience | 20.2 (9.73) | 22.2 (9.09) | - |
| Job tenure | 10.9 (9.52) | 9.60 (7.94) | - |
| N | 1440 | 278 | 41,826 |

Table 3. Estimates: Dependent variable= \ln annualized earnings^a

| Estimation procedure/ coefficient estimate | GLS individual | GLS individual | MLMD “within” | MLMD “within” | MLMD “within” (correlated errors) |
|---|------------------|------------------|----------------|------------------|-----------------------------------|
| β , schooling coefficient | .113 (21.9) | .118 (23.5) | .105 (5.65) | .104 (6.13) | .104 (5.96) |
| Age | .00401 (1.89) | - | - | - | - |
| Full-time work experience | - | .0102 (7.87) | - | .00839 (3.21) | .00841 (3.22) |
| Job tenure | - | .00482 (3.85) | - | .00411 (1.78) | .00406 (1.75) |
| Female | -.483 (19.5) | -.364 (13.7) | - | - | - |
| $s^2(\mu)$ | .137 (9.14) | .0884 (6.99) | .141 (8.65) | .110 (6.85) | .109 (6.82) |
| $\gamma_s = d_s + b_{\mu,fs}$ | 0 | 0 | 1.19 (1.80) | 1.54 (1.99) | 1.53 (1.89) |
| $\gamma_x = d_x + b_{\mu,fx}$ | - | 0 | - | 9.66 (4.51) | 9.70 (4.50) |
| $\gamma_t = d_t + b_{\mu,ft}$ | - | 0 | - | 4.85 (2.98) | 4.89 (2.98) |
| γ_w | 0 | 0 | 0 | 0 | -.0542 (0.62) |
| N | 1440 | 1440 | 720 | 720 | 720 |

^a Data are from the Minnesota Twins Registry. All estimates use co-twin’s report of schooling to control for random measurement error in own schooling reports. Absolute values of asymptotic t-ratios are in parentheses.

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