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June 22, 1992

Professor Alan Krueger
Industrial Relations Section
Princeton University
Fax # 609-258-2907

Dear Alan,

The weight of the evidence and the intuition supports the notion that the person reporting one variable erroneously errs in the same direction on other variables of the same type he reports i.e. there is positive correlation across measurement errors for the same person that is more highly correlated than the measurement errors across persons. This neatly explains your Table 3 results and orders most of your other numbers (with minor exceptions that I note). Thus your final estimates (fixed effect i.v. or method of moments) are large and probably biased upward, and substantially so since twin's schooling is so highly correlated.

art is scribbly

Invoke exchangeability of twins. Then let

$$Y_1 = \gamma S_1^* + U_1$$

$$Y_2 = \gamma S_2^* + U_2$$

$$S_2^1 = S_2^* + \epsilon_2^1$$

$$S_2^2 = S_2^* + \epsilon_2^2$$

$$S_1^2 = S_1^* + \epsilon_1^2$$

$$S_1^1 = S_1^* + \epsilon_1^1$$

$$F_1 = F^* + \epsilon_F^1 \quad \text{(Father)}$$

$$F_2 = F^* + \epsilon_F^2$$

$$M_1 = M^* + \epsilon_M^1 \quad \text{(Mother)}$$

$$M_2 = M^* + \epsilon_M^2$$

$$y = \gamma s^* + u_5$$

$$y' = \gamma s^{*\prime} + u_5'$$

$$t = t^* + v_4$$

$$s' = s^{*\prime} + u_3'$$

$$t' = t^{*\prime} + v_4'$$

$$s = s^* + u_3$$

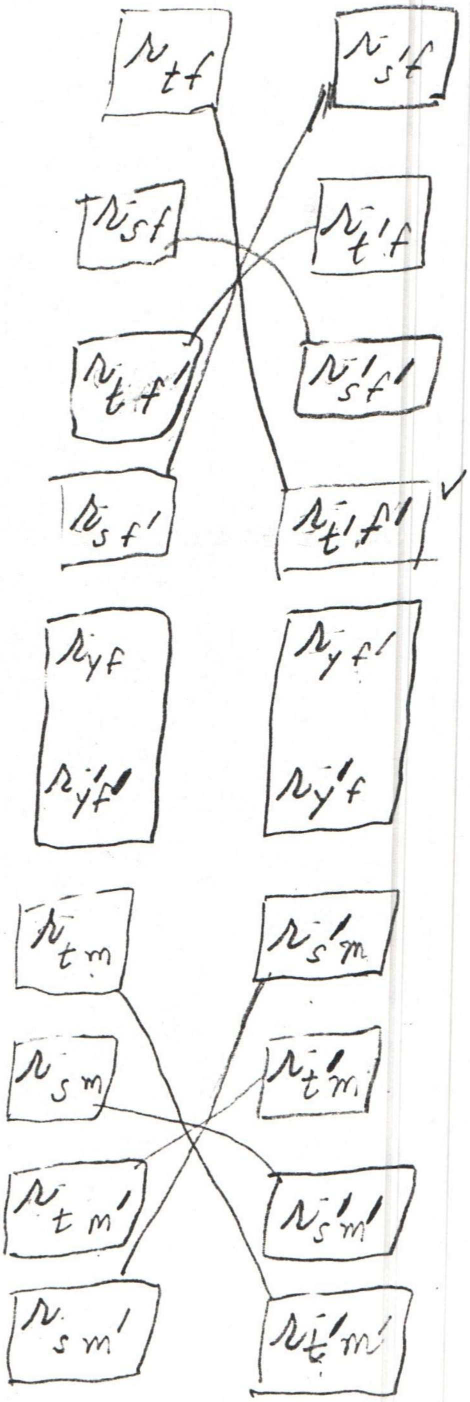
$$f = f^* + u_2$$

$$f' = f^* + u_2'$$

$$m = m^* + u_1$$

$$m' = m^* + u_1'$$

Assume each right hand side disturbance conditionally independent of right hand side "variable". Recall that the superscript records the reporter; subscript the reported variable. If measurement error is more alike for the reporter than across reporters (Cov $(\epsilon_x^i, \epsilon_y^i) >$ Cov $(\epsilon_x^j, \epsilon_y^i)$, $j \neq i$, with the roles of x and y interchangeable), then (numbers below are correlations which should obey the same ordering under exchangeability as covariates and are given in Table 2).



Cov $(S_2^1, F_1) >$ Cov (S_2^2, F_1)

(.416) (.361)

Cov $(S_1^1, F_1) >$ Cov (S_1^2, F_1)

(.345) (.266)

(*) Failure Cov $(S_2^1, F_2) <$ Cov (S_2^2, F_2)

(.389) (.320)

(*) Failure Cov $(S_1^1, F_2) <$ Cov (S_1^2, F_2)

(.357) (.278)

(allowing for measurement error in income) Cov $(Y_1, F_1) \geq$ Cov (Y_1, F_2)

(.155) (.159)

Cov $(Y_2, F_2) \geq$ Cov (Y_2, F_1)

(.091) (.088)

ignores Y, M data

Cov $(S_2^1, M_1) >$ Cov (S_2^2, M_1)

(.410) (.392)

Cov $(S_1^1, M_1) >$ Cov (S_1^2, M_1)

(.348) (.343)

(*) Failure Cov $(S_2^1, M_2) <$ Cov (S_2^2, M_2)

(.337) (.322)

Cov $(S_1^1, M_2) <$ Cov (S_1^2, M_2)

(.316) (.321)

(.316) (.321)

Assuming the error variance levels are equal for both brothers (but may differ across variables) we can use correlations and hence Table 2. I print numbers in parenthesis above. Except for the third and fourth rows, (the father's education and son's education), these correlations support my view. It would be better if you checked the covariances. Your tables are not detailed enough.

For fraternal twins the numbers are (in the format listed above)

.407 > .357
.332 > .408
.253 < .230
.259 < .392
.019 > .025
-.107 > .028
.244 > .224
.025 > .127
.180 < .109
.180 < .216.

Using covariances and appropriate std. errors, one could test jointly for this hypothesis. One sided inequality tests would be appropriate.

If true it implies that in random samples,

$$plim\ \hat{\gamma}_{OLS} < plim\ \hat{\gamma}_{IV} \qquad (For\ \gamma > 0)$$

Proof:

$$Y_1 = \gamma S_1^{*} + U_1$$

$$Y_1 = \gamma S_1^{1} + U_1 - \gamma \epsilon_1^{1}$$

$$plim\ \gamma_{OLS} = \gamma \left[\frac{V(S^*)}{V(S^*) + V(\epsilon_1^1)} \right]$$

Assumes $\sigma_{\epsilon S} = 0$.

where "V" denotes variance. $V(\epsilon_1^1) = V(\epsilon_2^2)$ by exchangeability.

$$plim\ \hat{\gamma}_{IV} = plim \frac{Cov(Y_1, S_1^2)}{Cov(S_1^1, S_1^2)}$$

$$= \gamma \frac{V(S^*)}{V(S^*) + Cov(\epsilon_1^1,\epsilon_1^2)}$$

By hypothesis Cov $V(\epsilon_1^1,\epsilon_2^1) < V(\epsilon_1^1)$ let $V(\epsilon_1^1,\epsilon_1^2)$ = Cov (A) where "A" denotes across brothers for future use). Thus

$$\text{plim } \hat{\gamma}_{OLS} < \text{plim } \hat{\gamma}_{iv}$$

In fact, the latter is consistent if a similar argument holds good for the fixed effects estimator. Thus

$$plim\ \hat{\gamma}_{fe} < plim\ \hat{\gamma}_{iv,fe} \qquad (assume\ \gamma > 0).$$

Proof:

$$Y_1 - Y_2 = \gamma(S_1^* - S_2^*) + U_1 - U_2$$

Thus

$$= \gamma(S_1^1 - S_2^2) + (U_1 - U_2) - \gamma(\epsilon_1^1 - \epsilon_2^2)$$

$$plim\ \hat{\gamma}_{fe} = \gamma\left[\frac{V(S^*)(1-\rho)}{V(S^*)(1-\rho) + V(\epsilon)(1-\rho_a)}\right] = \gamma\left[\frac{V(S^*)}{V(S^*) + V(\epsilon)\dfrac{1-\rho_a}{1-\rho}}\right].$$

where

$$\rho_a = \frac{Cov(\epsilon_1^1,\epsilon_2^2)}{[V(\epsilon_1^1)\ V(\epsilon_1^2)]^{1/2}} \qquad (correlation\ in\ errors\ across\ brothers)$$

and $V(\epsilon_1^1) = V(\epsilon_2^2) = V(\epsilon)$ by exchangeability and ρ = Correl $(\dot{S}_1, \dot{S}_2)$.

Provided $\rho > \rho_a$,

$$\hat{\gamma}_{fe} < \hat{\gamma}_{OLS} \quad (\text{assume } \gamma > 0).$$

If ρ_a is sufficiently big relative to ρ, can reverse. Thus if measurement error more correlated across twins than their schooling, the ordering is reversed. From Table 3 that does not appear to be the case.

Then, we notice using $(S_1^2 - S_2^1)$ as i.v above

$$plim \hat{\gamma}_{fe,iv} = plim \frac{Cov(S_1^2 - S_2^1, Y_1 - Y_2)}{Cov(S_1^2 - S_2^1, S_1^1 - S_2^2)}$$

$$= \gamma \left[\frac{V(S^*)}{V(S^*) + \dfrac{Cov(A) - Cov(W)}{1 - \rho}} \right]$$

$\epsilon_1^1 \neq u_3$
$\epsilon_1^2 = u_4'$
$\epsilon_2^2 = u_3'$
$\epsilon_2^2 = u_4$

where Cov(A) = Cov("Across Brothers") = $Cov(\epsilon_1^2, \epsilon_1^1) = Cov(\epsilon_2^1, \epsilon_2^2)$ by exchangeability

Cov(W) = Cov(Within) $= Cov(\epsilon_2^1, \epsilon_1^1) = Cov(\epsilon_1^2, \epsilon_2^2)$ by exchangeability.

As long as reporting errors are more correlated for errors made by the same person than for errors made by the same person than for errors made by different people

$$Cov(A) < Cov(W)$$

and

$$plim\ \hat{\gamma}_{fe,iv} > plim\ \hat{\gamma}_{fe}$$

Also observe that the higher ρ (the more correlated is true schooling across twins), the more pronounced this effect. This is consistent with Table 3.

Your Table 3 shows that

$$\hat{\gamma}_{OLS} < \hat{\gamma}_{fe} < \hat{\gamma}_{fe} < \hat{\gamma}_{fe,iv}$$

(The iv is not quite clear because brother's education enters some results in Table 3 - I have the correct expression for this case if you wish.)

The first inequality on the left is just a consequence of measurement error. The final inequality is a consequence of Cov(A) < Cov(W). The middle inequality implies that

i.e. $$Cov(W) > \rho\ Cov(A)$$

which is just an implication of the hypothesis "tested" at the beginning of this letter. Similar remarks

$$COV(A) > \frac{COV(A) - COV(W)}{1 - \rho}$$

apply to your method of moments estimator.

I didn't have time to develop this argument at the conference partly because I left my notes and scratch paper in my car when I rushed into the conference at the last minute.

Please confirm the hypothesis with covariances. Formal tests would be interesting. I see little support for your version Cov(A) > Cov(W).

Additional Comments

(1) Tests of the various coefficient estimates would be useful. See Durbin (1954) for details.

(2) The paper seems inconsistent. You test for family background bias using Chamberlain's π matrix assuming no measurement error. It is not a correct test if there is measurement error - a major claim of your paper.

(3) If, in fact, there are no family background effects, $\hat{\gamma}_{iv}$ is more efficient than $\hat{\gamma}_{fe,iv}$ and the estimated growth in returns to schooling is not so great.

(4) You never establish why MZ twins, or even twins, are needed. Given that you impose a common family effects structure, they are not. Even MZ twins have some freedom in schooling and income choices.

I have the expressions for the more complicated multivariate versions of these comparisons. Still, the discussion I give above is simpler and captures all of the essential ideas.

I enjoyed the chance to chat with you and to read your paper.

Sincerely,

Jim

James J. Heckman

cc: Orley Ashenfelter
Art Goldberger (Wisconsin)
Walter Oi (Rochester)
Tom MaCurdy (Stanford)
Derek Neal (Chicago)

26 June 1992

TO: Jim Heckman, Alan Krueger

FROM: Art Goldberger

RE: Twinsburg

Because the labelling of co-twins as 1 and 2 is arbitrary, I think that a symmetric treatment is required. This can be accomplished by using between-within components of variance and covariance. In the interim, we can average the correlations in Alan's Table 2 that correspond to the same population moment. (I do that below, after introducing a notation scheme). After all, the means and standard deviations in Alan's Table 1 <u>are</u> in effect averaged across 1's and 2's.

I like Jim's idea of introducing correlated errors to account for patterns in the empirical results. But many of the comparisons on page 3 of his fax concern arbitrarily labelled items, and hence are probably not informative. Also Jim ~~he~~ overlooks the Y,M correlations.

There are some typos on Jim's page 3 (and ~~one~~ on page 6) in which the role of sups and subs is reversed. I've highlighted those on the attached sheets. Also there seem to be some lines missing in the middle of page 5. The other scribbling on the attached pages are notes to myself.

Digression: Do you know that MZ's have been introduced into the presidential campaign? In Wednesday's news conference, Perot was discussing free trade agreements. He said something like "Suppose that you are a businessman planning a new factory in Texas. Your identical twin is considering opening a similar factory in Mexico, where wages and environmental regulations are weaker ... ".

My notation scheme: For an individual,

m^* = mother's true education, f^* = father's true education,

s^* = own true education, t^* = sib's true education,

m = reported mother's education, f = reported father's education,

s = reported own education, t = reported sib's education,

y = reported own logwage.

For the sib of the individual, primes are attached to each of those symbols. Because they have the same parents, $m^{*\prime} = m^*$ and $f^{*\prime} = f^*$. And because "sib" is reflexive (if that's the right word), $t^* = s^{*\prime}$ (and $t^{*\prime} = s^*$).

The model is

Individual	Sib
$m = m^* + u_1$	$m' = m^* + u_1{}'$
$f = f^* + u_2$	$f' = f^* + u_2{}'$
$s = s^* + u_3$	$s' = s^{*\prime} + u_3{}'$
$t = t^* + u_4$	$t' = t^{*\prime} + u_4{}'$
$y = \gamma\, s^* + u_5$	$y' = \gamma\, s^{*\prime} + u_5{}'$

-- with $t^* = s^{*\prime}$ (and $t^{*\prime} = s^*$). All starred variables are uncorrelated with all disturbances. But correlations among disturbances -- those within individual (same as those within sib) and those of individual crossed with sib -- are open for discussion.

Because of the inherent symmetry, population variances and covariances are invariant to joint switching of no-prime and prime. For examples,

$$\sigma_{11} = V(u_1) = V(u_1{}') = \sigma_{1'1'}$$

$$\sigma_{1'2} = C(u_1{}',u_2) = C(u_1,u_2{}') = \sigma_{12'}$$

$$\sigma_{mm} = V(m) = V(m') = \sigma_{m'm'}$$

$$\sigma_{mf'} = C(m,f') = C(m',f) = \sigma_{m'f} \quad .$$

So correlations have the same invariance.

With that in mind, I display the empirical correlation matrices (for identical twins) obtained by averaging those items in Alan's Table 2 which differed only because of the arbitrary labelling of 1 and 2. The display has the format I used on p. 314 in Kinometrics and in "Heritability" (Economica, Nov 1979). The numbers should be pretty much the same as those you'll get for correlations after manova on the original data. If memory serves, you can get the same results directly by double-entry correlations (in which each person appears twice, once as a 1 and once as a 2).

	SELF					CROSS				
	m	f	s	t	y	m'	f'	s'	t'	y'
m	1	.597	.335	.366	.094	.837	.574	.354	.340	.107
f		1	.332	.347	.123		.857	.359	.328	.124
s			1	.698	.327			.658	.898	.218
t				1	.194				.643	.311
y					1					.563

Note: Both matrices are symmetric, so lower portions are omitted.

Jim's suggestion is that the disturbances are more correlated within individuals than across them. For example, if $\sigma_{12} > \sigma_{12'}$, then $\sigma_{mf} > \sigma_{mf'}$.

Turning to the estimation of γ. If I understand it, Jim proposes various conditions on disturbance correlations in evaluating the (probability limits of) the various estimators:

OLS σ_{sy}/σ_{ss} He takes $\sigma_{35} = 0$.

IV $\sigma_{t'y}/\sigma_{t's}$ He takes $\sigma_{4'5} = 0$, but allows $Cov(A) = \sigma_{3'4} \neq 0$.

FE $\sigma_{\Delta s \Delta y}/\sigma_{\Delta s \Delta s}$ He takes $\sigma_{35} = 0$ and $\sigma_{3'5} = 0$, but allows $\sigma_{3'3} \neq 0$.

IVFE $\sigma_{\Delta t \Delta y}/\sigma_{\Delta t \Delta s}$ He takes $\sigma_{45} = 0$ and $\sigma_{4'5} = 0$, but allows $Cov(W) = \sigma_{34} \neq 0$, and $Cov(A) = \sigma_{3'4} \neq 0$.

Note that for the disturbances, some within-individual correlations are taken to be zero, while some cross correlations are allowed to be nonzero.

In my notation, what the penultimate line on his page 6 says is that $\sigma_{34}\, \sigma_{s^*s^*} > \sigma_{3'4}\, \sigma_{s^*s^{*\prime}}$.

It's not clear to me what this has to do with the "hypothesis tested" at the beginning of his letter.

No doubt I've made some errors in transcribing my notation here.

On substantive matters, I agree that it's crucial to get information on why the twins got different education. Was one doing much better in school, or did the parents (able to afford sending only one kid to college) toss a coin? In other words, to what future policy action is the rate of return estimate relevant?

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July 1, 1992

To: Alan Krueger and Art Goldberger,

The enclosed revision of my 6/22 letter fixes up a few typos and expands the argument. I am innocent of Art's charge that I ignore exchangeability (better - interchangeability). I just took the Ashenfelter-Krueger notation and used it in a correct interchangeable way.

There were a few typos but most are typos (see that the correct numbers are at the base of mistyped symbols). One section of text at the end got deleted.

I expand the argument about the ordering founding in A-K. It is consistent with my evidence (which is not subject to labelling bias as Art claims) that measurement error is more correlated across readings on the same person that across persons. The evidence is also consistent with the claim that variance of measurement error is greater for reports made about another person than about oneself.

Any reactions?

Best wishes to you both.

Jim

James J. Heckman

P.S. I just got Orley's fax. I disagree that what I have written is nitpicking. Since the evidence supports the iv estimator, it suggests only a 20-25% bias in returns to schooling - not 100% or more as you report. Don't miss the forest for the ecosystem!

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~~June 22, 1992~~ Date was clean by Jim.

Professor Alan Krueger
Industrial Relations Section
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Fax # 609-258-2907

Dear Alan,

The weight of the evidence and the intuition supports the notion that the person reporting one variable erroneously errs in the same direction on other variables of the same type he reports <u>i.e.</u> there is positive correlation across measurement errors for the same <u>person</u> that is more highly correlated than the measurement errors across persons. There is also indirect evidence that the error variance for reported own variables exceeds that of own reported sib variables. These observations explain your Table 3 results and orders most of your other numbers (with minor exceptions that I note). Thus your final estimates (fixed effect i.v. or method of moments) are large and probably biased upward, and substantially so since twin's schooling is so highly correlated.

Invoke exchangeability of twins. Then use your notational (respecting the arbitrariness of the labelling but recognizing that we are dealing with different persons in a twin pair).

$$Y_1 = \gamma S_1^* + U_1$$

$$Y_2 = \gamma S_2^* + U_2$$

$$S_2^1 = S_2^* + \epsilon_2^1$$

$$S_2^2 = S_2^* + \epsilon_2^2$$

$$S_1^2 = S_1^* + \epsilon_1^2$$

$$S_1^1 = S_1^* + \epsilon_1^1$$

$$F_1 = F^* + \epsilon_F^1 \quad \text{(Father)}$$

$$F_2 = F^* + \epsilon_F^2$$

$$M_1 = M^* + \epsilon_M^1 \quad \text{(Mother)}$$

$$M_2 = M^* + \epsilon_M^2$$

Assume each right hand side disturbance is mean zero and conditionally independent of all right hand side "variables" (within and across equations). Recall that the superscript records the reporter; subscript the reported variable. If measurement error is more alike for the reporter than across reporters (Cov $(\epsilon_x^i, \epsilon_y^i) > \text{Cov}(\epsilon_x^j, \epsilon_y^i, j \neq i)$, with the roles of x and y interchangeable), then (numbers below the inequalities are correlations which should obey the same ordering under exchangeability as the covariates and are given in Table 2): (Failure denoted by *)

	$\text{Cov}(S_2^1, F_1) > \text{Cov}(S_2^2, F_1)$	
	(.416)	(.361)
	$\text{Cov}(S_1^1, F_1) > \text{Cov}(S_1^2, F_1)$	
	(.345)	(.266)
(*) [Failure]	$\text{Cov}(S_2^1, F_2) < \text{Cov}(S_2^2, F_2)$	
	(.389)	(.320)
(*) Failure	$\text{Cov}(S_1^1, F_2) < \text{Cov}(S_1^2, F_2)$	
	(.357)	(.278)
(allowing for measurement error in income)	$\text{Cov}(Y_1, F_1) \geq \text{Cov}(Y_1, F_2)$	
	(.155)	(.159)
	$\text{Cov}(Y_2, F_2) \geq \text{Cov}(Y_2, F_1)$	
	(.091)	(.088)
(*) Failure	$\text{Cov}(Y_1, M_1) \geq \text{Cov}(Y_1, M_2)$	
	(.102)	(.126)
	$\text{Cov}(Y_2, M_2) \geq \text{Cov}(Y_2, M_1)$	
	(.087)	(.088)

$$\text{Cov}(S_2^1, M_1) > \text{Cov}(S_2^2, M_1)$$

(.410) (.392)

$$\text{Cov}(S_1^1, M_1) > \text{Cov}(S_1^2, M_1)$$

(.348) (.343)

(*) Failure

$$\text{Cov}(S_2^1, M_2) < \text{Cov}(S_2^2, M_2)$$

(.337) (.322)

$$\text{Cov}(S_1^1, M_2) < \text{Cov}(S_1^2, M_2)$$

(.316) (.321)

Assuming the error variance levels are equal for both brothers (but may differ across variables) we can use correlations and hence Table 2. I print numbers in parenthesis above. Except for the third and fourth rows, (the father's education and son's education), these correlations support my view. It would be better if you checked the covariances. Your tables are not detailed enough.

Exploiting the interchangeability of twins, we may write (using w as a weight)

$$w\,\text{Cov}(S_2^1, F_1) + (1 - w)\,\text{Cov}(S_1^2, F_2)$$
$$> w\,\text{Cov}(S_2^2, F_1) + (1 - w)\,\text{Cov}(S_1^1, F_2)$$

(.347) > (.358)

$$w\,\text{Cov}(S_1^1, F_1) + (1 - w)\,\text{Cov}(S_2^2, F_2)$$
$$> w\,\text{Cov}(S_1^2, F_1) + (1 - w)\,\text{Cov}(S_2^1, F_2)$$

(.333) > (.327)

$$w\,\text{Cov}(Y_1, F_1) + (1 - w)\,\text{Cov}(Y_2, F_2) \geq w\,\text{Cov}(Y_1, F_2) + (1 - w)\,\text{Cov}(Y_2, F_1)$$

(.123) ≥ (.123)

$$w\,\text{Cov}(Y_1, M_1) + (1 - w)\,\text{Cov}(Y_2, M_2) \geq w\,\text{Cov}(Y_1, M_2) + (1 - w)\,\text{Cov}(Y_2, M_1)$$

(.0945) ≥ (.107)

$$w\,Cov(S_2^1, M_1) + (1-w)\,Cov(S_1^2, M_2)$$

$$> w\,Cov(S_2^2, M_1) + (1-w)\,Cov(S_1^1, M_2)$$

$$(.3655) \qquad > (.354)$$

$$w\,Cov(S_2^2, M_2) + (1-w)\,Cov(S_1^1, M_1)$$

$$> w\,Cov(S_2^1, M_2) + (1-w)\,Cov(S_1^2, M_1)$$

$$(.353) \quad > \quad (.340)$$

using $w = ½$ produces numbers in parentheses. The agreement using exchangeable pairs seems closer.

For fraternal twins the numbers are (in the format listed above)

.407 > .357
.332 > .408
.253 < .230
.259 < .392
.019 > .025
-.107 > .028
.244 > .224
.025 > .127
.180 < .109
.180 < .216.

Using covariances and appropriate std. errors, one could test jointly for this hypothesis. One sided inequality tests would be appropriate. Better, one sided multiple comparisons considering all inequalities as a set.

If the hypothesis is true it implies that in random samples of twins,

$$plim\ \hat{\gamma}_{OLS} < plim\ \hat{\gamma}_{IV} \qquad (For\ \gamma > 0)$$

Proof:

$$Y_1 = \gamma S_1^* + U_1$$

$$Y_1 = \gamma S_1^1 + U_1 - \gamma \epsilon_1^1$$

$$plim\ \gamma_{OLS} = \gamma\left[\frac{V(S^*)}{V(S^*) + V(\epsilon_1^1)}\right]$$

where "V" denotes variance. $V(\epsilon_1^1) = V(\epsilon_2^2)$ by exchangeability. Note that I follow your logic throughout and assume that U_i is a structural disturbance (i.e. I assume no measurement error in Y or else I assume that the measurement error in Y is always uncorrelated with all other measurement errors).

$$plim\ \hat{\gamma}_{IV} = plim\ \frac{Cov(Y_1, S_1^2)}{Cov(S_1^1, S_1^2)}$$

$$= \gamma\ \frac{V(S^*)}{V(S^*) + Cov(\epsilon_1^1, \epsilon_1^2)}$$

By hypothesis Cov $V(\epsilon_1^1, \epsilon_2^1) < V(\epsilon_1^1)$. Let $Cov(\epsilon_1^1, \epsilon_1^2) = Cov\ (A)$ where "A" denotes across brothers for future use). Thus

$$plim\ \hat{\gamma}_{OLS} < plim\ \hat{\gamma}_{IV}$$

In fact, the latter is consistent if $Cov(\epsilon_1^1, \epsilon_1^2) = 0$. A similar argument holds good for the fixed effects estimator. Thus

$$plim\ \hat{\gamma}_{fe} < plim\ \hat{\gamma}_{IV,fe} \qquad (assume\ \gamma > 0).$$

Proof:

$$Y_1 - Y_2 = \gamma(S_1^* - S_2^*) + U_1 - U_2$$

Thus

$$= \gamma(S_1^1 - S_2^2) + (U_1 - U_2) - \gamma(\epsilon_1^1 - \epsilon_2^2)$$

$$plim\ \hat{\gamma}_{fe} = \gamma\left[\frac{V(S^*)(1 - \rho)}{V(S^*)(1 - \rho) + V(\epsilon)(1 - \rho_a)}\right] = \gamma\left[\frac{V(S^*)}{V(S^*) + V(\epsilon)\ \dfrac{1 - \rho_a}{1 - \rho}}\right].$$

where

$$\rho_a = \frac{Cov(\epsilon_1^1, \epsilon_2^2)}{[V(\epsilon_1^1)\; V(\epsilon_2^2)]^{1/2}} \qquad \textit{(correlation in errors across brothers)}$$

and $\quad V(\epsilon_1^1) = V(\epsilon_2^2) = V(\epsilon)$ by exchangeability and ρ = Correl $(\dot{S}_1, \dot{S}_2)$.

Provided $\quad \rho > \rho_a$,

$$\hat{\gamma}_{fe} < \hat{\gamma}_{OLS} \quad \text{(assume } \gamma > 0) \quad .$$

If ρ_a is sufficiently big relative to ρ, this inequality can reverse. Thus if measurement error is more correlated across twins than their schooling, the ordering is reversed. From Table 3 that does not appear to be the case.

Then, we notice using $(S_1^2 - S_2^1)$ as i.v above

$$plim\,\hat{\gamma}_{fe,iv} = plim\frac{Cov(S_1^2 - S_2^1, Y_1 - Y_2)}{Cov(S_1^2 - S_2^1, S_1^1 - S_2^2)}$$

$$= \gamma\left[\frac{V(S^*)}{V(S^*) + \dfrac{Cov(A) - Cov(W)}{1 - \rho}}\right]$$

where Cov(A) = Cov("Across Brothers") = $Cov(\epsilon_1^2, \epsilon_1^1) = Cov(\epsilon_2^1, \epsilon_2^2)$ by exchangeability

Cov(W) = Cov(Within) = $Cov(\epsilon_2^1, \epsilon_1^1) = Cov(\epsilon_1^2, \epsilon_2^2)$ by exchangeability.

As long as reporting errors are more correlated for errors made by the same person than for errors made by the same person than for errors made by different people

$$Cov(A) < Cov(W)$$

and

$$plim\ \hat{\gamma}_{fe,iv} > plim\ \hat{\gamma}_{fe}$$

Also observe that the higher ρ (the more correlated is true schooling across twins), the more pronounced this effect. This is consistent with Table 3.

Your Table 3 shows that

$$\hat{\gamma}_{OLS} < \hat{\gamma}_{fe} < \hat{\gamma}_{iv} < \hat{\gamma}_{fe,iv}$$

(The iv is not quite clear because brother's education enters some results in Table 3 - I have the correct expression for this case if you wish.)

The first inequality on the left is consistent with $\rho_s > \rho$. The final inequality is a consequence of Cov(A) < Cov(W) since the final equality implies that

i.e.
$$\text{Cov(W)} > \rho \text{ Cov(A)}$$

which is just an implication of the hypothesis "tested" at the beginning of this letter (within error covariances exceed between where within and between are within and between brothers).

The middle inequality ($\hat{\gamma}_{iv} < \hat{\gamma}_{fe}$) implies that

$$V(\epsilon_2^2) = V(\epsilon_1^1) > \text{Cov}(\epsilon_1^1, \epsilon_2^2) + (1 - \rho)\, \text{Cov}(\epsilon_1^1, \epsilon_1^2)$$
$$= \text{Cov}(\epsilon_2^2, \epsilon_1^1) + (1 - \rho)\, \text{Cov}(\epsilon_2^2, \epsilon_2^1)\,.$$

In the extreme case where $\epsilon_2^2 = \epsilon_1^2$ (person 2 makes exactly the same errors about his own schooling and the other sibs) or $\epsilon_1^1 = \epsilon_2^1$, the middle inequality implies that

$$V(\epsilon_2^2) = V(\epsilon_1^1) > \text{Cov}(\epsilon_2^2, \epsilon_1^1)\,(2 - \rho)$$

i.e.

$$1 > \rho_s(2 - \rho)$$

so we have an upper bound on ρ_s i.e. the more highly corelated true schooling, the more correlated can cross-brother errors be.

In the more general case, it seems plausible that $\text{Cov}(\epsilon_1^1, \epsilon_2^2) < \text{Cov}(\epsilon_1^1, \epsilon_1^2)$ because, in addition to the differences in response error across brothers, the target being aimed at is "different" i.e there is fresh noise, in the covariance on the left and this creates the possibility of less dependence. This pattern, of course is not "tested" above but is consistent with the "evidence".

Then

$$1 > \rho_a + (1 - \rho)\rho_A$$

where

$$\rho_A = Cov(\epsilon_2^2, \epsilon_1^2) / Var(\epsilon_1^1)$$

(the latter is meaningful only if $Var(\epsilon_2^1) = Var(\epsilon_2^2) = Var(\epsilon_1^1) = Var(\epsilon_1^2)$). More plausibly, $Var(\epsilon_1^2) > Var(\epsilon_1^1)$ in which case we write

$$\rho_A^* = \frac{Cov(\epsilon_2^2, \epsilon_2^1)}{\sqrt{Var(\epsilon_2^2), Var(\epsilon_2^1)}}$$

so the inequality becomes

$$1 > \rho_a + (1 - \rho)\, \rho_A^* \left[\frac{Var\ (\epsilon_2^1)}{Var\ (\epsilon_2^2)} \right]^{1/2} .$$

Thus the more highly correlated is true schooling (higher ρ), the more correlated can the measurement errors be.

Moreover since we found that $\rho_a > \rho$ seemed to describe one set of facts,

$$1 > \frac{1-\rho_a}{1-\rho} > \rho_A^* \left[\frac{Var\ (\epsilon_2^1)}{Var\ (\epsilon_2^2)} \right]^{1/2} .$$

Thus $Var(\epsilon_2^1)$ can be greater than $Var(\epsilon_2^2)$ but not more than

$$\left[\frac{1}{\rho_A^*} \right]^{1/2}$$

greater (in proportion).

The evidence seems to be consistent with this form of measurement error: (a) people are less accurate about others than themselves ($Var(\epsilon_2^1) = Var(\epsilon_1^2) > Var(\epsilon_1^1) = Var(\epsilon_2^2)$) and (b) mistakes are more correlated across the same person than across persons. The evidence in the front of this letter seems consistent with (b).

I didn't have time to develop this argument at the conference partly because I left my notes and scratch paper in my car when I rushed into the conference at the last minute.

Please confirm the hypothesis with covariances. Formal tests would be interesting. I see little support for your version Cov(A) > Cov(W).

Additional Comments

(1) Tests of the various coefficient estimates would be useful. See Durbin (1954) for details.

(2) The paper seems inconsistent. You test for family background bias using Chamberlain's π matrix assuming no measurement error. It is not a correct test if there is measurement error - a major claim of your paper.

(3) If, in fact, there are no family background effects, $\hat{\gamma}_{iv}$ is more efficient than $\hat{\gamma}_{fe,iv}$ and the estimated increase in the returns to schooling is not so great when the ordinary i.v. estimator is used.

(4) You never establish why MZ twins, or even twins, are needed. Given that you impose a common family effects structure, they are not. Even MZ twins have some freedom in schooling and income choices.

I have the expressions for the more complicated multivariate versions of these comparisons. Still, the discussion I give above is simpler and captures all of the essential ideas.

I enjoyed the chance to chat with you and to read your paper.

Sincerely,

Jim

James J. Heckman

cc: Orley Ashenfelter
Art Goldberger (Wisconsin)

This was also faxed to me.

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July 2, 1992

To: Art Goldberger

From: Orley Ashenfelter

Subject: The paper blitz

Since you have become a part of the daisy chain, and Jim Heckman's recent letter and revised letter, it seems to me you should see the FAX we sent Jim yesterday.

I think this is strictly a tempest in a teapot, but Jim obviously doesn't agree. Is this an example of econometrician's testosterone problems?

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July 1, 1992

Professor James Heckman
Department of Economics
University of Chicago
Chicago, IL 60657

FAX 312-702-8490

Dear Jim:

Many thanks for your FAX of June 22nd and the detailed comments on our paper using data on twins to measure the returns to schooling.

As usual, you have pointed out an entire class of additional models that might be applied to our problem in view of the unique data set we have collected. We think it is worth doing some further analysis along these lines, but it is pretty clear from the magnitudes of the discrepancies from the classical measurement error model that you have observed that none of the basic conclusions of our paper will be altered.

Just to make sure that we do not lose sight of the forest for the trees, let us reiterate our two main conclusions: **(1) The within-twin slope of the relationship between wage rates and schooling is no smaller, and may be larger, than the cross-sectional slope. (This is not what we expected nor what exists in the literature.) (2) There does seem to be evidence of measurement error in schooling levels that almost certainly biases downward the estimated schooling coefficient** (although, as you point out, one can argue about the magnitude of the bias.)

It is interesting that your reaction to the correlations in Table 3 led you to think of an alternative to the classical measurement error model. As we pointed out in the paper, we also found it fascinating that the classical measurement error model has many implications beyond those normally emphasized and that these are testable when enough data are available. Frankly, our reaction to Table 3 was different than yours: We were surprised at how accurate the classical model seemed to be based on our "eyeball tests"! For example, as you note, there are many correlations that should all be equal under the classical model--many of which you report in your letter. Unlike you, we were impressed by "eyeball" tests of how close many of these correlations seemed to be to each other, and we stated as much in the paper. (Indeed, even your hypothesis of person-specific measurement error orders the correlations correctly in only 7 of the 10 comparisons for identical twins, and it does less--4 out of 10 comparisons--well for fraternals.) Frankly, we think it would make sense to set out a complete model of measurement error and test and estimate the relevant parameters, but we also think this should probably be done in a separate paper.

Incidentally, we intend to interview again in Twinsburg this August, and this will no doubt provide us with some re-interview data on the same twins. This will enrich the class of measurement error models that it is feasible to estimate by quite a bit. Want to come along and do some interviewing? (We already have the local restaurant scene cased out and we will bring the wine!)

Sincerely,

Orley Ashenfelter   Alan Krueger

cc: Art Goldberger
Walter Oi
Tom MaCurdy
Derek Neal

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TO: Art Goldberger

FAX#: 608-262-04747

LOCATION: Dept of Economics U of Wisc

TELEPHONE#: ______

FROM: Orley Ashenfelter

TOTAL NUMBER OF PAGES (including cover page): ______

MESSAGE:

I think this new installment will amuse you too!

WE ARE TRANSMITTING ON ADLER-ROYAL FX915

OUR FAX #: 609-258-2907

July 2, 1992

Professor James Heckman
Department of Economics
University of Chicago
Chicago, IL 60637

Dear Jim:

Many thanks for the revised FAX, especially with the revised calculations on p. 3.

It seems to me that you have now shown conclusively that, at least with respect to the alternative you suggest, the classical measurement error model cannot be rejected in our data.

By averaging the correlations to produce the numbers on p. 3-4 of your letter you presumably reduce noise, and now the results strongly suggest classical measurement error. The pairs of numbers that should be equal, using your assumptions, are

our?

.347 should be equal to .358

.333 " " .327

.123 " " .123

.095 " " .107

.366 " " .354

.353 " " .340

It should be .335!

__Under the alternative scenario you suggest, all the numbers in the left column should be greater than (or equal to) the numbers in the right column. Three are greater, one is equal, and two are less. My god, what more evidence could there be for classical measurement error?__

Frankly, we are indebted to you for setting up an alternative--and especially for showing that the data do not support it!

Sincerely,

Orley Ashenfelter