The return to schooling: Estimates from a sample of young Australian twins

Paul Miller a,*, Charles Mulvey a, Nick Martin b

a Economics M251, The University of Western Australia, 35 Stirling Highway, Crawley WA 6009, Australia
b Queensland Institute of Medical Research, Australia

Received 1 December 2001; received in revised form 1 July 2004; accepted 20 October 2004
Available online 17 May 2005

Abstract

1. Introduction

The publication of Ashenfelter and Krueger’s study of the economic return to schooling from a new sample of twins in 1994 heralded a renewal of interest in using the natural experiment of twins in labour market analysis. On the basis of an analysis of the increments in earnings associated with extra schooling for identical twins in the US labour market, Ashenfelter and Krueger (1994) concluded that family and genetic effects make virtually no contribution to the returns to schooling. The novel feature of their sample was the multiple reports on the levels of schooling for each respondent that permitted adjustment for measurement error. This measurement error was shown to bias downwards the return to schooling estimated in conventional studies.

Ashenfelter and Krueger’s (1994) work for the US labour market has been replicated for other countries.1 Miller et al. (1995) use data from a sample of Australian twins. Their data set was more representative of the population than the data set used by Ashenfelter and Krueger (1994). Miller et al. (1995) report a modest role for ability and family background in the relationship between schooling and income such that the usual OLS estimates of the return to schooling in Australia are not biased upward by the omission of family and ability effects. This finding supports the major conclusion from the study by Ashenfelter and Krueger.

Isacsson (1999) conducted a similar analysis for Sweden. This study also reports that the measurement error adjusted estimate of the return to schooling obtained from a sample of identical twins was comparable to that found in studies based on samples of individuals. A feature of Isacsson’s (1999) study is the use of post-estimation adjustments for measurement error using the reliability ratio of the measured schooling variable. This compares with Ashenfelter and Krueger (1994) and Miller et al. (1995) where an IV estimator is used for this purpose. Isacsson (1999, p. 477) notes “The reason for not using a fixed effects by instrumental variables estimator (FE by IV). . . . is that the FE by IV results are very different for different age cohorts in the sample of monozygotic twins”.2

In a more recent study, Bonjour et al. (2003) conduct analyses of the return to education using female identical twins in the UK. They also find that measurement error biases

---

1 Other recent studies of the economic returns to schooling in the US labour market based on samples of twins include Ashenfelter and Rouse (1998), Behrman and Rosenzweig (1999) and Rouse (1999).

2 Ashenfelter and Krueger’s (1994) results are also different from those reported in Ashenfelter and Rouse (1998). This latter set of results, based on a larger sample, suggests that the conventional OLS estimates of the return to schooling are slightly upward biased. The difference in this regard may be due to sampling error (Ashenfelter and Rouse (1998, p. 264)).
estimated returns to education down, and omitted ability biases the conventional estimate up. The overall effect of these offsetting biases is such that the returns to education obtained with a within-twins study (of 7.7 percent) is very similar to that found in a conventional study based on a sample of individuals.

Ashenfelter and Krueger’s (1994) findings and approach have been questioned in a number of studies. Blackburn and Neumark’s (1995) analyses, based on the inclusion in an earnings function of test scores as ability controls, suggest that OLS estimates that ignore ability are biased upwards, by about 40 percent. They also argue, based on the results of a series of Hausman (1978) tests, that there is little support for their schooling variable being either endogenous or measured with error, or that ability is measured with error by the test scores available. Black et al. (2000) examine the consequence of a negative covariance between the measurement error in the schooling variable and the true value of that variable. By restricting the twins sample to pairs where the twins agree on the difference in education levels, they provide a lower bound to the estimate of the returns to education corrected for measurement error. Bound and Solon (1999) and Neumark (1999) suggest that Ashenfelter and Krueger’s (1994) findings may stem from ability differences between twins that are not removed in the fixed effects model, and from between-twins differences in schooling that are chosen endogenously. Any ability differences between twins that remain in the fixed effects model are shown by Neumark (1999) to be associated with greater biases in the within-twin IV model than in the standard within-twin estimator.

There has been some research that looks at ability or endowment differences within twin pairs. Martin et al. (1997) present an overview of factors that might make monozygotic twins less than fully identical. They argue that a wide range of antenatal genetic and environmental influences can cause genotypic divergence. Included are (p. 390, 392) “...endogenous accidents of development and differentiation, from perturbations in the gradients of developmental fields to somatic mutation, somatic recombination, differences in tissue-specific methylation patterns and the time of such events”. For example, Martin et al. (1997) note that there is only one optimal placental implantation site, and it is unlikely that both twins in a di-chorionic pair would be able to benefit from this site. However, the same authors note (p. 391) “Reconvergence may occur after birth because the twins do not passively undergo differing experiences; on the contrary, it now seems likely that they actively seek, select and perceive similar environments because of genetic similarities in brain physiology”.

Behrman et al. (1994) discuss birth weight differences within pairs of identical twins, and the links of these to schooling outcomes. They report (p. 1162) that an increase in birth weight by 4 ounces is associated with an increase in schooling of almost one-half a year; and among identical twin pairs, 69 percent had birth weights differing by at least this amount. This has been used as evidence that identical twins differ in endowments and ability (see, for example, Neumark (1999)). Similarly, Ashenfelter and Rouse (1998) examine a range of issues that could give rise to ability differences within twin pairs and hence lead to differences in levels of education. These include birth order and the possibility that twins are treated differently and so end up with different non-genetic abilities. They conclude, however, that (p. 279) “These results provide further evidence that the within-twin schooling differences are not solely determined by within-twin ability
differences and that within-twin estimates of the return to schooling contain less ability bias than cross-sectional estimates”. Similar findings are reported by Bonjour et al. (2003).

This paper presents analyses of the return to schooling based on a large sample of young Australian twins. The estimators proposed by Ashenfelter and Krueger (1994) are applied. There are three main reasons for conducting this research. First, the research effort that followed the publication of the important paper by Ashenfelter and Krueger (1994) has raised some concerns over the robustness of the findings obtained from the IV estimation of the fixed-effects model (see, for example, Isacsson (1999)). Study with alternative data sets and age groups is needed to examine this issue. Second, the research by Miller et al. (1995) for the Australian labour market was based on the mean earnings of the occupation of employment. Their study was able to capture the inter-occupational earnings effects of schooling, but not the intra-occupational earnings effects. In this study a more conventional measure of earnings is available. Examination of the economic returns to schooling in Australia within the context of a study of twins using this superior information on earnings will provide a check on the findings reported by Miller et al. (1995). Third, the analysis of these data provides an opportunity to examine the importance of some of the recent developments (e.g., Black et al.’s (2000) bounding procedure for the within-twins IV estimator) in this line of literature.

The structure of the paper is as follows. Section 2 outlines the data. Section 3 presents analyses of the data using the fixed-effects models proposed by Ashenfelter and Krueger (1994). Estimates from Ashenfelter and Krueger’s (1994) selection effects model are examined in Section 4 by way of a sensitivity analysis. Section 5 reports results from an application of Black et al.’s (2000) bounding procedure for the within-twins estimates. Section 6 contains a summary and conclusion.

2. The Australian twins survey

2.1. The sample

The sample of twins analysed in this study were members of the young adult cohort of the Australian Twin Register. They constitute a volunteer twin panel born between 1964 – 1971. Nearly all were first registered with the panel between 1980 and 1982 by their parents. A total of 4264 twin pairs were recruited at this time.

The data presented in this paper are derived from responses to a telephone interview conducted by lay interviewers during the period 1996 – 2000. The individual response rate for these telephone interviews was over 80 percent. The final sample included almost 1200 pairs of identical twins, and 1550 pairs of fraternal twins. Individuals were excluded from the study where they or their co-twin did not answer questions on any of the variables used in the analysis, or where either the respondent or their co-twin were not employed on either a full- or part-time basis. The final sample comprised 759 sets of identical twins, and 1031 sets of fraternal twins. As with the data analysed by Ashenfelter and Krueger (1994), Miller et al. (1995) and Ashenfelter and Rouse (1998), each twin provided reports on both their own level of education and that of their co-twin. This permits application of the IV estimators proposed by Ashenfelter and Krueger (1994) and Black et al. (2000).
Table 1 lists descriptive statistics for the sample of young Australian twins. Appendix A presents definitions of variables. The income data available have been reported using a prompt card that listed weekly, fortnightly and annual incomes against a letter code. The annual income equivalents of the letter codes have been used. The studies by Miller et al. (1995) and Isacsson (1999) were also based on annual incomes. As with Isacsson (1999), individuals with low incomes have been excluded. The exclusions in this regard are implemented according to the full- or part-time status of the workers. Hence full-time workers earning less than $10,000 per annum are omitted from all analyses. Pairs of twins are omitted if either twin meets this criterion. Only two percent of the sample (44 pairs of twins) are affected by this sample selection rule.

The twins included in the study range in age from 23 to 36 years. The mean age is 30 years, with the standard deviation being 2.4. On average the twins have almost 14 years of education. Most of the twins have completed high school, and close to 40 percent have post-school qualifications. As with the study of older twins by Miller et al. (1995) and that by Ashenfelter and Krueger (1994), the report of the respondent’s level of education by his or her co-twin is slightly lower than the self-reported measure of educational attainment. The sample is reasonably evenly balanced between males and females, and also between twins who are married and those who are not married. Around 80 percent of the twins were employed on a full-time basis. There is little difference between the gender and marital status mixes of the samples of identical and non-identical twins. Similarly, the mean incomes of the identical and non-identical twins are approximately the same.

It is noted that around 47 percent of twins report the same own level of education. For identical twins the percentage is 52 percent. This is similar to the data employed in the studies by Isacsson (1999) and Ashenfelter and Krueger (1994), and also in the study by Miller et al. (1994).

Table 1 lists descriptive statistics for the sample of young Australian twins. Appendix A presents definitions of variables. The income data available have been reported using a prompt card that listed weekly, fortnightly and annual incomes against a letter code. The annual income equivalents of the letter codes have been used. The studies by Miller et al. (1995) and Isacsson (1999) were also based on annual incomes. As with Isacsson (1999), individuals with low incomes have been excluded. The exclusions in this regard are implemented according to the full- or part-time status of the workers. Hence full-time workers earning less than $10,000 per annum are omitted from all analyses. Pairs of twins are omitted if either twin meets this criterion. Only two percent of the sample (44 pairs of twins) are affected by this sample selection rule.

The twins included in the study range in age from 23 to 36 years. The mean age is 30 years, with the standard deviation being 2.4. On average the twins have almost 14 years of education. Most of the twins have completed high school, and close to 40 percent have post-school qualifications. As with the study of older twins by Miller et al. (1995) and that by Ashenfelter and Krueger (1994), the report of the respondent’s level of education by his or her co-twin is slightly lower than the self-reported measure of educational attainment. The sample is reasonably evenly balanced between males and females, and also between twins who are married and those who are not married. Around 80 percent of the twins were employed on a full-time basis. There is little difference between the gender and marital status mixes of the samples of identical and non-identical twins. Similarly, the mean incomes of the identical and non-identical twins are approximately the same.

It is noted that around 47 percent of twins report the same own level of education. For identical twins the percentage is 52 percent. This is similar to the data employed in the studies by Isacsson (1999) and Ashenfelter and Krueger (1994), and also in the study by Miller et al. (1994).
Table 2 presents correlation coefficients between selected variables for the sample of young Australian twins. The first section of this table has been computed for identical twins; the second section for non-identical twins. The general patterns in these correlations mirror those reported in other studies of twins. The correlation between the educational attainments of members of sets of twins is .625 in the case of identical twins, and only .435 for non-identical twins. That is, identical twins are more alike in terms of their educational attainments than non-identical twins (see Baker et al. (1996) and Miller et al. (2001)). The correlation between the self-reported measure of educational attainment and the report on this educational attainment by the co-twin is around .85 for identical twins, and .78 for non-identical twins. The values reported for these correlations are remarkably consistent between the alternative choice of twins as **twin** and **co-twin**. Both measures are slightly lower than those reported for the sample of older twins by Miller et al. (1995). The within-twin pair correlation in incomes is .5 for identical twins, but only .14 for non-identical twins.

The main features of these data on young twins match reasonably closely data on older samples of twins (see Miller et al. (1995)). Finally, results from OLS regression of log annual earnings using these data were compared to analyses of the large samples of individuals from the Australian Census of Population and Housing. The results are presented in Appendix A. There it is shown that the return to education in the sample of young twins is remarkably consistent between the alternative choice of twins as **twin** and **co-twin**. Both measures are slightly lower than those reported for the sample of older twins by Miller et al. (1995). The within-twin pair correlation in incomes is .5 for identical twins, but only .14 for non-identical twins.

### Table 2
Correlation coefficients between selected variables: Australian twins survey

<table>
<thead>
<tr>
<th>Variable</th>
<th>EDUC₁¹</th>
<th>EDUC₁²</th>
<th>EDUC₂¹</th>
<th>EDUC₂²</th>
<th>Income 1</th>
<th>Income 2</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>A. Identical Twins (Sample Size = 759 Twin Pairs)</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>EDUC₁¹</td>
<td>1.000</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>EDUC₂¹</td>
<td>0.725</td>
<td>1.000</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>EDUC₂²</td>
<td>0.625</td>
<td>0.849</td>
<td>1.000</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>EDUC₁²</td>
<td>0.850</td>
<td>0.670</td>
<td>0.724</td>
<td>1.000</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Income 1</td>
<td>0.313</td>
<td>0.293</td>
<td>0.276</td>
<td>0.305</td>
<td>1.000</td>
<td></td>
</tr>
<tr>
<td>Income 2</td>
<td>0.228</td>
<td>0.286</td>
<td>0.287</td>
<td>0.209</td>
<td>0.501</td>
<td>1.000</td>
</tr>
<tr>
<td><strong>B. Non-identical Twins (Sample Size = 1031 Twin Pairs)</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>EDUC₁¹</td>
<td>1.000</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>EDUC₂¹</td>
<td>0.531</td>
<td>1.000</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>EDUC₂²</td>
<td>0.435</td>
<td>0.788</td>
<td>1.000</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>EDUC₁²</td>
<td>0.781</td>
<td>0.496</td>
<td>0.525</td>
<td>1.000</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Income 1</td>
<td>0.237</td>
<td>0.207</td>
<td>0.175</td>
<td>0.263</td>
<td>1.000</td>
<td></td>
</tr>
<tr>
<td>Income 2</td>
<td>0.180</td>
<td>0.244</td>
<td>0.263</td>
<td>0.174</td>
<td>0.135</td>
<td>1.000</td>
</tr>
</tbody>
</table>

Notes: EDUC₁¹ = twin 1’s report on own educational attainment; EDUC₁² = twin 1’s report on twin 2’s educational attainment; EDUC₂¹ = twin 2’s report on own educational attainment; EDUC₂² = twin 2’s report on twin 1’s educational attainment; Income 1 = twin 1’s income; Income 2 = twin 2’s income.
young Australian twins is 5.8 percent, while that obtained from the Census for a sample of 20 – 34 year olds was 6.4 percent. This small difference in estimated returns may be associated with the timing of the data: the twins data were collected at the end of the 1990s, whereas the Census data relate to 1991.

3. Traditional and IV analyses of sample of young twins

Table 3 presents estimates from the analyses of the incomes of individuals and for the within-twins analyses for identical twins. The first two columns of results are for the sample of identical twins treated as if they were a sample of individuals. The first set of results has been obtained using ordinary least squares (OLS), the second using instrumental variables (IV). For the OLS results, the model for the determination of earnings for twin i in family j may be expressed as:

\[ \ln Y_i = \alpha_0 + \alpha_1 \text{EDUC}_{ij} + \alpha_2 Z_{ij} + \beta X_j + \mu_j + \epsilon_{ij} \quad i = 1, 2, j = 1, \ldots, n \]  

(1)

where \( \text{EDUC}_{ij} \) is, using the notation of Table 2, individual i’s self-reported level of education, \( Z_{ij} \) is a vector of other characteristics that can vary across twins within a family, and which includes a dummy variable recording whether the individual is employed full-time, marital status and an interaction between marital status and gender, \( X_j \) is a vector of observed variables that vary by family but not across members of a twin pair (e.g., age), and \( \mu_j \) represents unobservable components that vary by family. When this specification of the model is applied to samples of individuals, the estimates will be subject to omitted variables bias owing to omission of controls for ability and the unobserved family component \( \mu_j \).

For the IV estimations, the co-twin’s report on educational attainment has been used as the instrument for the twin’s self-reported educational attainment. Using the notation of Table 2 once again, \( \text{EDUC}_{ij} \) has been used as an instrument for \( \text{EDUC}_{ij} \), where \( - i \) denotes the co-twin of individual i.

In this model, to allow for non-linearities, the age variable is entered in the Gompertz form that is argued to be superior for the study of young workers (Borland and Suen (1994), Griliches (1976)). A robust covariance matrix is computed to take account of the clustering in the data associated with the treatment of the twins as a sample of individuals.

The key features of the OLS estimates for individuals are the return to education of around 6 percent, the coefficient of \( - 0.169 \) on the female variable, the effects of marriage on earnings for males and females, and the coefficient of 0.767 on the full-time employment variable. The returns to education reported in the Australian literature are typically in the order of 6 to 7 percent, and hence the Table 3 estimate is at the lower end of this range. This level of return is lower than that generally reported for the US labour

\[ \hat{h} = -0.11 \text{Age} - 0.1 \text{Age} \times 100. \]

As \( \hat{h} \) is \( -6.907 \), evaluated at Age=30, this partial effect is 3.4 percent. Evaluated at Age=35, it equals 2.1 percent.

\[ \hat{h} = 0.1 \text{Age} - 0.1 \text{Age} \times 100. \]

6 Gender and ability will vary only by family for identical twins, but may vary across twins in the case of non-identical twins.

7 The partial effect of age on earnings is \( -0.11 \text{Age} - 0.1 \text{Age} \times 100. \) As \( \hat{h} \) is \( -6.907 \), evaluated at Age=30, this partial effect is 3.4 percent. Evaluated at Age=35, it equals 2.1 percent.
market, and this difference is usually attributed to the more centralized system of wage determination in Australia and the more egalitarian distribution of income that results from it. The 6 percent for this sample is similar to that obtained from an independent sample of 20 – 34 years olds, where the dependent variable is defined in the same way and the same specification of the estimating equation is employed (see Appendix A). Similarly, the female earnings differential of around 17 percent is comparable to the estimates reported in previous studies that have examined variations in annual earnings (see Miller et al. (1995)).

Marriage is associated with a wage premium among males of the order of six percent, though the marriage effect for females is not significant. This evidence mirrors the marriage earnings differentials in Australia reported by Eastough and Miller (2004). However, it departs from the conventional findings of a sizeable earnings advantage for married men and an earnings disadvantage among married women (e.g., Voon and Miller (in press)). There is no apparent reason for this difference, though it is noted that Bonjour et al. (2003) report that marital status was not a significant determinant of earnings of female identical twins in the UK. There is no evidence for Australia on the annual earnings differential between part-time and full-time workers. The evidence reported in Appendix A based on the 1991 Census suggests that the earnings differential is substantial, in the order of 100 percent. In comparison, the Table 3 results suggest the earnings of full-time workers are 2.2 times those of part-time workers. As the mean hours of work for part-time

Table 3
Estimates of models of log annual earnings: individuals and identical twins, Australian twins survey

<table>
<thead>
<tr>
<th>Variable</th>
<th>Identical twins as individuals&lt;sup&gt;(c)&lt;/sup&gt;</th>
<th>Identical twins: fixed effects models</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>OLS</td>
<td>IV</td>
</tr>
<tr>
<td>Constant</td>
<td>9.348 (99.38)</td>
<td>9.154 (88.71)</td>
</tr>
<tr>
<td>Own Education</td>
<td>0.060 (11.91)</td>
<td>0.075 (12.66)</td>
</tr>
<tr>
<td>Age (e&lt;sup&gt;−0.1580&lt;/sup&gt;)</td>
<td>−6.907 (6.67)</td>
<td>−6.722 (7.12)</td>
</tr>
<tr>
<td>Female</td>
<td>−0.169 (4.99)</td>
<td>−0.179 (5.38)</td>
</tr>
<tr>
<td>Married</td>
<td>0.061 (1.70)</td>
<td>0.066 (2.01)</td>
</tr>
<tr>
<td>Female× Married</td>
<td>−0.025 (0.50)</td>
<td>−0.028 (0.58)</td>
</tr>
<tr>
<td>Employed</td>
<td>0.767 (18.92)</td>
<td>0.755 (21.39)</td>
</tr>
<tr>
<td>Full-Time</td>
<td>0.4073</td>
<td>–</td>
</tr>
<tr>
<td>R&lt;sup&gt;2&lt;/sup&gt;</td>
<td>0.4073</td>
<td>–</td>
</tr>
<tr>
<td>Sample Size</td>
<td>1518</td>
<td>1518</td>
</tr>
</tbody>
</table>

Note: Figures in parentheses are robust ‘t’ statistics.

(a) IV estimator robust in the presence of correlated measurement errors.

(b) Variable not relevant.

(c) See Appendix Table A.1 for estimates pooled across identical and non-identical twins.

8 Voon and Miller (in press) argue that the differences across studies in the magnitude of the marriage effect is related to differences in the set of controls in the various estimating equations, in particular, the presence or absence of controls for the presence and number of children. However, this does not appear to be the reason for the insignificance of the marriage effect in the current analysis.

9 As the coefficient on the full-time employment dummy variable is large the algorithm proposed by Halvorsen and Palmquist (1980) whereby the percentage effect associated with the coefficient ‘b’ is calculated as 100.b=100.(exp(b)−1) is used.
workers in 1999 were 16 and for full-time workers 42, this earnings differential is not unreasonable.\footnote{11}

Since Griliches (1977), applied researchers have been acutely aware of the potential importance of errors of measurement in the schooling data when examining the returns to schooling. The usual response to such a situation has been to use an Instrumental Variables (IV) estimator, and this approach is followed here. The main interest in the IV estimates in the second column of results is the magnitude of the own-education variable, and this is 7.5 percent, or 20 percent (or three standard deviations) greater than the estimate obtained using OLS. A Hausman (1978) test of whether non-orthogonality of the error term and the self-reported schooling variable in the OLS specification constituted a problem returned a ‘t’ statistic of 5.12. This indicates that the notion of non-orthogonality needs to be taken seriously.

The final section of Table 3 presents estimates from fixed-effects models of the earnings of identical twins. In these estimates, the difference in the earnings of members of a set of twins is related to differences in the values of the explanatory variables. Obviously explanatory variables that have the same value for each twin will drop out of the estimating equation. For the identical twins, the estimating equation is as follows, where the subscript j refers to the family, and $\Delta$ indicates the difference in the values of the particular variable for the members of the set of twins in family j.

\[
\Delta \ln Y_j = \alpha_0 + \alpha_1 \Delta EDU_{ij} + \alpha_2 \Delta Z_j + \Delta e_j
\]  

\footnote{10 See Australian Bureau of Statistics (1999).}

As identical twins reared together have the same innate ability and family background, terms in ‘Ability’ and ‘Family’ drop out of the fixed-effects version of the model of earnings determination.

In the OLS estimates the return to schooling is 1.8 percent. This is, in principle, an estimate of the impact of schooling on earnings that is free of biases typically associated with the absence of measures of ability and family background. The recent concern over the role of unobserved ability differences between identical twins that are not removed in the within-twin model and which are correlated with twin differences in years of schooling indicates, however, that caution needs to be exercised in the strict interpretation of this estimate as being unbiased (Bound and Solon (1999) and Neumark (1999)). The 1.8 percent estimate of the return to schooling in Table 3 is slightly lower than that presented by Miller et al. (1995), of 2.5 percent.\footnote{12}

It is also noted that the constant term is significant in the fixed effects estimates. A constant term is included in the model to capture any effects associated with non-random selection of the twin to be the first twin for the within-twins difference operator. The first member of a twin pair encountered in the data set is designated the first twin. There are several possible explanations for this effect. First, the twins were recorded as “first
member” or “second member” according to the order in which their names appeared on the registration form. More of those listed first on this form were the first born twin, so this could be a birth-order effect. Second, it could reflect the initiative shown in being the first to register, and this could carry over to economic outcomes. It is also likely, however, given the small size of the effect for both identical and non-identical twins, the variability in the statistical significance of the effect for identical twins, and the insignificance for non-identical twins (Table 4), that this is just a type one error.

Errors of measurement have the potential to constitute a greater problem in the fixed-effects estimator than with the estimators applied to samples of individuals because the measurement error contributes relatively more to the within-family variance of schooling than to the overall variance of schooling. It is also possible that the standard within-twin IV estimators applied in the literature amplifies the bias that arises when differencing does not fully eliminate the ability differences between twins (Neumark (1999), Bound and Solon (1999), Black et al. (2000)). We return to this point in Section 5.13

When account is taken of measurement error through the use of Ashenfelter and Krueger’s (1994) IV estimators, the return to schooling obtained from the identical twins rises to between 4.4 and 5.4 percent. The 5.4 percent estimate is obtained using the difference between the twins’ reports on their co-twins’ levels of schooling as instruments. The 4.4 percent estimate is obtained using the IV estimator proposed by Ashenfelter and Krueger (1994) for the case where the two reports (on own and the co-twin’s educational attainment) contain a common measurement error. For this situation, a

<table>
<thead>
<tr>
<th>Variable</th>
<th>OLS</th>
<th>IV</th>
<th>OLS</th>
<th>IV</th>
<th>IV(a)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>9.321 (130.82)</td>
<td>9.030 (85.17)</td>
<td>0.020 (1.06)</td>
<td>0.019 (1.02)</td>
<td>0.000 (0.00)</td>
</tr>
<tr>
<td>Own Education</td>
<td>0.055 (14.16)</td>
<td>0.077 (11.06)</td>
<td>0.036 (5.16)</td>
<td>0.063 (4.55)</td>
<td>0.066 (5.63)</td>
</tr>
<tr>
<td>Age (e^{0.1AGE})</td>
<td>−5.939 (7.64)</td>
<td>−5.961 (7.75)</td>
<td>(b)</td>
<td>(b)</td>
<td>(b)</td>
</tr>
<tr>
<td>Female</td>
<td>−0.136 (4.97)</td>
<td>−0.151 (5.42)</td>
<td>−0.152 (3.80)</td>
<td>−0.168 (4.09)</td>
<td>−0.170 (4.13)</td>
</tr>
<tr>
<td>Married</td>
<td>0.098 (3.47)</td>
<td>0.106 (3.88)</td>
<td>0.119 (3.12)</td>
<td>0.122 (3.17)</td>
<td>0.117 (3.02)</td>
</tr>
<tr>
<td>Female × Married</td>
<td>−0.167 (4.14)</td>
<td>−0.160 (3.92)</td>
<td>−0.214 (3.88)</td>
<td>−0.215 (3.86)</td>
<td>−0.210 (3.75)</td>
</tr>
<tr>
<td>Employed Full-Time</td>
<td>0.807 (25.51)</td>
<td>0.800 (28.36)</td>
<td>0.780 (17.75)</td>
<td>0.775 (17.37)</td>
<td>0.781 (17.37)</td>
</tr>
<tr>
<td>R²</td>
<td>0.4546</td>
<td>–</td>
<td>0.4100</td>
<td>–</td>
<td>–</td>
</tr>
<tr>
<td>Sample Size</td>
<td>2062</td>
<td>2062</td>
<td>1031</td>
<td>1031</td>
<td>1031</td>
</tr>
</tbody>
</table>

Note: Figures in parentheses are robust ‘t’ statistics.
(a) IV estimator robust in the presence of correlated measurement errors.
(b) Variable not relevant.
(c) See Appendix Table A.1 for estimates pooled across identical and non-identical twins.

13 Analyses were also conducted that used birth weight as an endowment that might potentially inform on whether within-twin ability differences that are not removed in the fixed effects model could bias the results. While the results suggested that this was not the case, the analyses could only be conducted on the approximately one-third of the sample with valid birth weight data. See Miller et al. (2005) for details.
14 That is, (EDUC_{ij} − EDUC_{ci}) is used as an instrument for (EDUC_{ij} − EDUC_{ci}).
15 Neumark (1999, p. 146) notes that any amplification of the omitted variables bias in the fixed effects model is less with this IV estimator than for Ashenfelter and Krueger’s (1994) IV estimator that does not take into account the correlated measurement error.
consistent estimator is obtained by expressing the earnings difference as a function of the difference between the respondent’s own level of education and his or her report on the co-twin’s level of education, and instrumenting this using the difference between the co-twin’s report on the first twin’s level of education and the co-twin’s report on his or her own level of education. These estimates of the rate of return to education are only marginally lower than the 5.8 percent obtained in the conventional OLS study of the earnings of individuals.

The difference between the fixed-effects (by instrumental variables) estimate of the return to schooling of around 5 percent, and the IV estimate for the sample of individuals of 7.6 percent, provides a guide to the importance of the omission of measures of ability and family background when determining the return to education. It suggests that at most one-third of the usual estimate of the return to schooling is attributable to these forms of omitted variables bias.

Table 4 lists estimates of a number of models of the determinants of annual earnings for non-identical twins. The OLS and IV models for non-identical twins treated as a sample of individuals are presented to serve as a reference point. With fraternal twins who are reared together, terms in ‘Family’ will disappear from the fixed-effects version of the model. However, as fraternal twins are no more alike genetically than non-twin siblings, they will differ in genetic ability. Consequently, while the estimates obtained with the within-twin model will be free of bias associated with the absence of controls for common family background influences, they will be subject to biases associated with the absence of measures for ability.

The OLS estimate of the return to schooling from the fixed-effects model for fraternal twins is 3.6 percent. In comparison, the IV estimates are around 6.5 percent. These estimates contain the pure effect of education on earnings and the omitted variables bias effect associated with the ability variable. They are greater than the comparable estimates obtained for identical twins, by around 1.5 percentage points.

A feature of Table 3 is that the estimate obtained using the IV estimator that is robust to correlated errors of measurement is lower than the IV estimate that is inconsistent due to this feature of the data. This is also a feature of the studies by Ashenfelter and Krueger (1994) and Miller et al. (1995). Neumark (1999) provides a rationalisation for these results, as a bias that may arise from omitted ability in the within-twin models and which will vary according to the method of estimation (OLS, IV, IV allowing for correlated measurement errors). In the case of Table 4, the estimates obtained using the alternative IV estimators are broadly the same. A similar finding emerges from the study of non-identical twins by Miller et al. (1995).

The Table 4 results show that married male non-identical twins have earnings around 12 percent higher than their non-married counterparts, while married female non-identical twins have earnings around 10 percent lower than their non-married counterparts. These findings for non-identical twins are similar to those reported in the wider literature that uses a similar specification of the estimating equation (e.g. Voon and Miller (in press)).

However, the estimates of the links between marital status and earnings for male and

16 That is, \((\text{EDUC}_{ij} - \text{EDUC}_{-ij})\) is used as the explanatory variable, and \((\text{EDUC}_{ij}^{-1} - \text{EDUC}_{-ij}^{-1})\) is used as the instrument.
female non-identical twins contrast with those documented in Table 3 for identical twins.\textsuperscript{17} The difference in the marriage premia for identical and non-identical twins has been previously reported by Lee (2003), on the basis of study of twins covering all age groups. The origin of the difference is difficult to determine, as it needs to be associated with influences that are not removed by the within-twins estimators. It could perhaps be investigated through study of marriage patterns and the characteristics of spouses, in a way analogous to Meng and Gregory (2005), though this is beyond the scope of the current study.

The general patterns of the results reported in this study of the economic returns to education among a sample of young twins mirror those reported by Miller et al. (1995) for a sample of twins that covered all age groups. A similar finding is reported in Miller et al. (2004) from a study of older twins. The estimators proposed by Ashenfelter and Krueger (1994) appear to be robust to choice of age groups for study. Moreover, the mean occupational earnings data used by Miller et al. (1995) do not appear to be a major limitation of that study, a conclusion that is reinforced by Groshen’s (1991) argument that most of the earnings increments associated with extra years of education come about through job classification rather than intra-occupational channels.

4. Selection effects model

This section presents estimates from the selection effects model proposed by Ashenfelter and Krueger (1994). This model provide for more direct consideration of a number of the effects which are estimated indirectly using the within-twins, fixed-effects model. Consideration of this alternative model provides a means for assessment of the robustness of the findings obtained from the fixed-effects model.

In the selection effects model, the earnings of twin $i$ who is a member of family $j$ are modelled via Eq. (1). However, explicit account is taken of the unmeasured family effects ($\mu_j$), which can be modelled in general form as depending on the educational attainments of each twin member, the $Z_{ij}$ for each twin member, and on the $X_j$ variables that vary across families but not between twins. Hence, the model is given by the following equations:

$$\ln Y_i = \alpha_0 + \alpha_1 + \text{EDUC}^1_{ij} + \alpha_2 Z_{ij} + \beta X_j + \mu_j + \epsilon_{ij} \quad i = 1, 2; j = 1, \ldots, n.$$  \(3a\)

$$\mu_j = \gamma_1 \text{EDUC}^1_{ij} + \gamma_1 \text{EDUC}^2_{2j} + \gamma_2 Z_{ij} + \gamma_2 Z_{2j} + \delta X_j + \omega_j \quad j = 1, \ldots, n.$$  \(3b\)

Ashenfelter and Krueger (1994) set $\gamma_2 = 0$ (variables such as marital status and employment type are not linked to the family effects). This restriction is adopted here also to facilitate comparisons.\textsuperscript{17}

\textsuperscript{17} Krashinsky (2004) also reports the absence of marital status effects for both males and females in fixed effects models estimated on a sample of identical twins.
Substitution for the $\mu_j$ term in the earnings Eq. (3a) results in the following reduced form:

$$\ln Y_i = x_0 + (x_1 + \gamma)EDUC_i^j + x_2Z_{ij} + (\beta + \delta)X_j + \gamma EDUC_{-ij}^j + \epsilon_{ij}$$  

(4a)

$$\ln Y_{-ij} = x_0 + (x_1 + \gamma)EDUC_{-ij}^j + x_2Z_{-ij} + (\beta + \delta)X_j + \gamma EDUC_i^j + \epsilon_{-ij}$$  

(4b)

where, in the notation introduced above, $-i$ indicates the co-twin of respondent $i$. From this equation it is seen that the coefficient on the co-twin’s educational attainment ($\gamma$) provides an estimate of the impact of family effects. This estimate can be subtracted from the coefficient on the own-education variable ($\alpha_1 + \gamma$) to derive an estimate of the pure effect of schooling in the case of identical twins, or of the effect of schooling biased by the omission of a direct measure of ability in the case of fraternal twins.

Results from the selection effects model of Ashenfelter and Krueger (1994) are presented in Table 5. Here the first two columns refer to identical twins, and the second two to non-identical twins. In addition to an IV method of estimation, estimates were obtained using Generalised Least Squares (GLS). These estimates take account of the cross-equation restrictions apparent in Eqs. (4a) and (4b).

The estimate of the return to schooling is about one-half of a percentage point higher for non-identical twins under each method of estimation. This difference is due to the explicit control for ability and family background in the estimations for identical twins, but only for family background in the estimations for the non-identical twins who were reared together. As noted above, the estimate on the co-twin’s educational attainment is an estimate of the magnitude of selection effects. These are quite small in each set of results presented, but particularly so for the IV estimates which take account of errors of measurement.

A summary of the results from the selection effects model is that family background factors account for around one-half a percentage point of the total return to education of

<table>
<thead>
<tr>
<th>Variable</th>
<th>Identical twins</th>
<th>Non-identical twins</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>GLS</td>
<td>IV</td>
</tr>
<tr>
<td>Constant</td>
<td>9.243 (92.03)</td>
<td>9.177 (98.66)</td>
</tr>
<tr>
<td>Own Education</td>
<td>0.044 (9.37)</td>
<td>0.062 (4.47)</td>
</tr>
<tr>
<td>Co-Twin’s Education</td>
<td>0.027 (5.72)</td>
<td>0.011 (0.78)</td>
</tr>
<tr>
<td>Age ($e^{-0.1 Age}$)</td>
<td>−6.979 (6.71)</td>
<td>−6.736 (7.20)</td>
</tr>
<tr>
<td>Female</td>
<td>−0.179 (5.09)</td>
<td>−0.177 (5.38)</td>
</tr>
<tr>
<td>Married</td>
<td>0.061 (1.92)</td>
<td>0.065 (1.99)</td>
</tr>
<tr>
<td>Female × Married</td>
<td>−0.033 (0.71)</td>
<td>−0.026 (0.55)</td>
</tr>
<tr>
<td>Employed Full-Time</td>
<td>0.732 (21.46)</td>
<td>0.761 (21.60)</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.393</td>
<td>–</td>
</tr>
<tr>
<td>Sample Size</td>
<td>1518</td>
<td>1518</td>
</tr>
</tbody>
</table>

Note: Figures in parentheses are robust ‘t’ statistics.
between 6.5 and 7 percent. Genetic factors account for a further one-half of a percentage point.

5. Further empirical results: bounds on the return to schooling

Black et al. (2000) show that, asymptotically, the true value of the return to schooling in analyses such as Ashenfelter and Krueger (1994) and those reported above is bounded by the OLS and IV estimators employed in these studies. They draw attention to a likely negative correlation between the true value of the schooling variable and the errors of measurement, and propose an estimator that accommodates this. This alternative provides a tighter lower bound on the estimated return to schooling. The essence of the Black et al. (2000) estimator in the current context is to restrict the sample to observations where the twins agree on the schooling difference within the set of twins. OLS estimates are obtained from this restricted sample.

Black et al. (2000) applied their estimator to Ashenfelter and Krueger’s (1994) data. In this data set, of the 147 pairs of identical twins, 92 pairs (63 percent) agreed on the differences in the levels of schooling within the set of twins. Of these 92, 64 percent were cases where there were no differences in education levels. Following this method led to a marked tightening of the bounds on the returns to schooling, from a 9 – 18 percent range to a 13 – 18 percent range.

In the current analysis, the bounds on the return to schooling are already quite tight, and this return is bounded at a much lower level than in Ashenfelter and Krueger’s (1994) study. In the estimates for identical twins in Table 3, the OLS estimate is 0.018 and the IV estimate (in the presence of correlated measurement errors) 0.044—a gap of only 0.026. In the estimates for non-identical twins (Table 4), the OLS estimate is 0.036 and the IV estimate 0.066—a gap of only 0.03. As the bounds are already quite tight, the need to pursue alternatives that might narrow these further is less pressing than in analyses such as Ashenfelter and Krueger (1994).

The data set used in the current application of the Black et al. (2000) procedure has the same features as Ashenfelter and Krueger’s (1994) data. Among identical twins, 73 percent agree on the difference in levels of education, and in two-thirds of these cases there is no difference in the levels of education (compared to 52 percent of cases for the unrestricted sample). Among non-identical twins, 58 percent agree on the difference in levels of education, and in 56 percent of these cases there is no difference in the levels of education (compared to 43 percent of cases for the unrestricted sample). The tails of the distributions of the education variable in the restricted sample for both types of twins are also smaller than in the unrestricted sample.

The combination of the relatively low ‘t’ value on the coefficient on the schooling variable, the smaller sample size and the smaller dispersion of the schooling variable results in a statistically insignificant OLS estimate of the return to schooling for the restricted sample of identical twins. This estimate also does not differ significantly from the OLS estimate reported in Table 3.

The analyses were repeated for the sample of non-identical twins, where the issues alluded to above are less pressing. In this case the estimate of the return to schooling
obtained using the Black et al. (2000) procedure is 0.040 (‘t’ = 3.74). This is quite similar to the OLS estimate in the unrestricted sample presented in Table 4 (coefficient of 0.036, ‘t’ of 5.16).

Hence, in this case, given that the bounds on the estimated return to schooling obtained for the Ashenfelter and Krueger (1994) type of analysis are already quite narrow, there is less priority in attempting to tighten these. Black et al.’s (2000) procedure does not assist in this regard for the study of identical twins, and it would seem this is due to characteristics of the data. For non-identical twins, however, the procedure gives the expected results: the OLS estimate with the restricted sample is greater than with the unrestricted sample, but given the already tight bounds, the change is minor.

6. Conclusion

The economic return to schooling in Australia among a sample of young twins is between 5 and 7 percent when account is taken of genetic and family effects using either fixed-effects models or the selection effects model of Ashenfelter and Krueger. Estimates of this order of magnitude are similar to those obtained using conventional models, being at most one percentage point lower than the conventional estimates. Accordingly, it is concluded that application of a range of estimators to a sample of young Australian twins provides a consistent set of findings, and that these findings accord with the evidence reported for twins representing a broader range of ages in Ashenfelter and Krueger (1994) and Miller et al. (1995). Given the similarity of the findings in this and in related studies, it would appear that the models applied by Ashenfelter and Krueger (1994) are robust. Moreover, the comparison of the findings reported in this study with those reported by Miller et al. (1995) indicate that the use of mean occupational earnings in place of a measure of individual earnings when the latter are not available (as in Miller et al. (1995)) is a defensible research strategy when the focus is on the economic returns to schooling. The approach adopted in the earlier study has not impacted in any material way on the conclusions that may be drawn from the study of the returns to schooling in Australia, or on the place such study might stand in the growing twins research literature.

Finally, viewing the OLS and IV estimates as lower and upper bounds on the value of the return to schooling in the manner of Black et al. (2000), it is seen that in this study the bounds on the return to schooling are quite tight, and this return is bounded at a much lower level than in Ashenfelter and Krueger’s (1994) results.

Appendix A. Definitions of variables and comparison with 1991 census

The sample of twins analysed in this study were members of the young adult cohort of the Australian Twin Register. They constitute a volunteer twin panel born between 1964 – 1971. Nearly all were first registered with the panel between 1980 and 1982 by their parents. The data presented in this report are derived from responses to a telephone interview conducted by lay interviewers during the period 1996 – 2000. The individual response rate for these telephone interviews is over 80 percent.
Earnings: The annual earning data were collected in 12 categories, (i) $1,000 – $9,999; (ii) $10,000 – $19,999; (iii) $20,000 – $24,999; (iv) $25,000 – $29,999; (v) $30,000 – $34,999; (vi) $35,000 – $39,999; (vii) $40,000 – $49,999; (viii) $50,000 – $74,999; (ix) $75,000 – $99,999; (x) $100,000 – $149,999; (xi) $150,000 or more. Values of $7,500 and $225,000 are used for the bottom and top intervals, and mid-points for all other intervals.

Educational Attainment: All the education variables are coded to the following highest educational levels completed: (i) 7 or fewer years of schooling; (ii) 8 – 10 years of schooling; (iii) 8 – 10 years of schooling and apprenticeship or diploma; (iv) 11 – 12 years of schooling; (v) 11 – 12 years of schooling and apprenticeship or diploma; (vi) technical or teachers college; (vii) University first degree; (viii) University post-graduate training.

Full-time: Respondents were asked to list their employment type at the time of the interview. Five states were recognised: Student; Unemployed; Part-Time; Homemaker; Full-Time. The analysis is restricted to those who were employed on either a full- or part-time basis. The ‘Full-time’ variable takes a value of one if the individuals was employed on a full-time basis, otherwise it is assigned a value of zero.

Married: The survey collected information on the marital status at the time of the interview. The ‘Married’ variable takes a value of one for those who reported they were married, and is set equal to zero for all other respondents.

Table A.1 Estimates of Models of Log Annual Earnings: Young Individuals from the Australian Twins Survey and the 1991 Census

<table>
<thead>
<tr>
<th>Variable</th>
<th>Twins Survey</th>
<th>1991 Census</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>9.326 (163.92)</td>
<td>8.859 (311.56)</td>
</tr>
<tr>
<td>Own Education</td>
<td>0.058 (18.45)</td>
<td>0.064 (37.61)</td>
</tr>
<tr>
<td>Age ($e^{-0.1A^{0.5}}$)</td>
<td>– 6.379 (10.21)</td>
<td>– 2.104 (8.19)</td>
</tr>
<tr>
<td>Female</td>
<td>– 0.150 (7.02)</td>
<td>– 0.066 (5.25)</td>
</tr>
<tr>
<td>Married</td>
<td>0.083 (3.72)</td>
<td>0.098 (8.99)</td>
</tr>
<tr>
<td>Female × Married</td>
<td>– 0.103 (3.26)</td>
<td>– 0.280 (16.64)</td>
</tr>
<tr>
<td>Employed Full-Time</td>
<td>0.792 (31.61)</td>
<td>0.637 (59.64)</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.4336</td>
<td>0.3175</td>
</tr>
<tr>
<td>Sample Size</td>
<td>3,580</td>
<td>17,122</td>
</tr>
</tbody>
</table>

Note: Figures in parentheses are t’ statistics.

References


