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Genetic and environmental contributions to educational attainment in Australia

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Abstract

The genetic and environmental contributions to educational attainment in Australia are examined using a multiple regression model drawn from the medical research literature. Data from a large sample of Australian twins are analysed. The findings indicate that at least as much as 50 percent and perhaps as much as 65 percent of the variance in educational attainments can be attributed to genetic endowments. It is suggested that only around 25 percent of the variance in educational attainments may be due to environmental factors, though this contribution is shown to be around 40 percent when adjustments for measurement error and assortative mating are made. The high fraction of the observed variation in educational attainments due to genetic differences is consistent with results reported by Heath et al. (Heath, A.C., Berg, K., Eaves, L.J., Solaas, M.H., Corey, L.A., Sundet, J., Magnus, P., Nance, W.E., 1985. Education policy and the heritability of educational attainment. *Nature* 314(6013), 734–736.), Tambs et al. (Tambs, K., Sundet, J.M., Magnus, P., Berg, K., 1989. Genetic and environmental contributions to the covariance between occupational status, educational attainment and IQ: a study of twins. *Behavior Genetics* 19(2), 209–222.), Vogler and Fulker (Vogler, G.P., Fulker, D.W., 1983. Familial resemblance for educational attainment. *Behavior Genetics* 13(4), 341–354.) and Behrman and Taubman (Behrman, J., Taubman, P., 1989. Is schooling mostly in the genes? Nature-nurture decomposition using data on relatives. *Journal of Political Economy* 97(6), 1425–1446.), suggesting that the finding is robust.

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1. Introduction

Understanding the determinants of educational attainment may be the key to understanding much of the inequality and inter-generational transmission of inequality in society. The level of schooling in Australia, for example, has been argued to be an important determinant of the probability of being employed (Inglis & Stromback, 1986), of earnings (Chiswick & Miller, 1985), and of occupational attainment (Evans & Kelley, 1986). The level of education is also a key element in Broom, Jones, McDonnell, and Williams's (1980) three-generation model of the inheritance of inequality in Australia. While studies of schooling decisions in Australia have focused on measurable indices of family background, they have rarely considered the influence of genetic endowments (e.g., Miller & Volker, 1989; Williams, Harsel, Clancy, Miller, & Greenwood, 1987). As overseas research has shown that education outcomes are

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linked closely to genetic endowments (see, for example, Behrman, Hrubec, Taubman, & Wales, 1980; Behrman & Taubman, 1976, 1989; Heath et al., 1985), the policy relevance of education studies that do not take account of this factor is therefore restricted.

In this paper we use the model of DeFries and Fulker (1985) to examine schooling outcomes in Australia. This model was devised to provide direct estimates of genetic influences and of the proportion of variance due to shared environmental influences within a multiple regression framework.[1] The model also, in principle, permits the impact of measurable indices of family background (e.g., parents' educational attainments, father's occupational status) on schooling decisions to be estimated at the same time that the estimate of common environmental influence is obtained. It thus allows an assessment of the extent to which control for common environment is possible through inclusion of measurable indicators of family background in models of educational attainment.

Study of Australian twins using the model of DeFries and Fulker adds to the economics literature on the nature-nurture debate, providing evidence of robustness of findings to choice of method of estimation and sample. It provides an alternative linking of measured indicators of family background to the unobserved environmental factors to that provided in Behrman and Taubman (1989). We find that Australian studies using the conventional regression analysis of schooling tend to overstate the influence of family background. Our initial analysis reveals that heritability accounts for around 50 percent of the variance in educational attainments, and that shared family factors account for only 25 percent. It is also argued that to the extent there is assortative mating, this partition yields an understatement of the importance of genetic endowments: adjustment for assortative mating increases the contribution of genetic endowments to the variance in educational attainment to around 65 percent. When an adjustment is made for measurement error in the schooling data, the contribution of shared family factors rises to 50 percent (40 percent when assortative mating is also taken into account).

The structure of the paper is as follows. In Section 2 we outline a conventional model of educational attainment and provide a brief review of the model of DeFries and Fulker. In Section 3 we introduce the Australian Twin Registry data set upon which our analyses are based. Section 4 presents an empirical analysis of the data which addresses the basic issue of the roles of genetic endowments and environmentality in the determination of educational attainments. Section 5 extends the analysis to consider whether heritability has a differential influence across levels of educational attainments and by gender. In Section 6 we draw some conclusions and reflect on the broad implications of our findings for policy.

2. Models for the study of educational attainment

The starting point for many models of educational attainment is Becker's Woytinsky lecture (see Becker, 1975). In this lecture Becker set out a basic theory determining the optimal investment in human capital. The framework was based on a supply curve and a demand curve for human capital investment. Under the influence of diminishing returns to human capital production and rising cost of the time input into the production process, the demand curve is downward sloping. The supply curve represents the marginal cost of financing human capital investments and it will be upward sloping as individuals resort to progressively higher priced sources of funds to finance greater levels of human capital investments. Becker describes the position of the demand curve as being dependent on an individual's capacities, or genetic endowments, with the demand curve for individuals with greater genetic endowments being above that of other groups. Due to capital market imperfections, the position of the supply curve depends on family background, with the supply curve for individuals from more affluent backgrounds being below that of their less affluent counterparts. In other words, as argued by Behrman and Taubman (1989), at any given level of investment, cheaper funds are available to those from richer families.[2] The model determines equilibrium values of the returns to investment and the level of human capital as functions of genetic endowments and family background (Taubman, 1981). We may write $S=F(G,N)$, where S denotes the level of schooling, G genetic endowments and N is environment. A linear approximation to this is $S=\alpha_1 G+\alpha_2 N$.

Applied research has proposed various links between genetic endowments, environment and observed variables. Among the factors associated with environment are the education levels of parents, the occupational

[1] In the nature-nurture literature the term "environment" refers to both exogenous factors such as family wealth and the outcomes of behavioural choices such as choice of schools and neighbourhoods (see Behrman, Rosenzweig, & Taubman, 1994, (p. 1138)). This definition is wider than the exogenous factors that economists generally consider as part of the given "environment". To minimise confusion, the definition from the nature-nurture literature will be followed in this study.

[2] Becker (1975) outlines other factors that affect the location of the supply curve, including location, luck and contacts. These provide a variety of reasons why the equilibrium level of schooling will vary across families.

attainments of parents, the number of siblings, and household income.[3] The links between measures of socio-economic status and schooling levels are assumed to derive from the effects of the access to resources for financing investment in human capital in the presence of capital market imperfections and the environment in which individuals were raised (Becker, 1975 also notes that there may be effects which operate via the demand-side of the model). Consistent with Becker's investment model, educational attainment is expected to be positively related to parental income, greater among the more prestigious occupational groups, and negatively related to the number of siblings.[4] Some researchers have also proposed scholastic achievement as a measure of genetic endowments. Included here are Micklewright's (1989) maths and comprehension test scores. Less direct measures that have been advanced include the school-type variables employed by Micklewright (1989) and Miller and Volker (1989). A positive link is expected between the decision to continue at school and scholastic achievement, whether it is measured directly using measures of academic ability or indirectly through type of school attended. Through including the various sets of variables in regression models, control for the measurable aspects of genetic endowments and family environment can be effected. Naturally, any unmeasurable aspects such as motivation and the values placed on education within the home cannot be held constant in this approach. This is an important aspect, as Williams et al. (1987) (p. 11) conclude that it is "…differences in preferences for education that are associated with social status, rather than economic impediments, which underlie the always observed relationship between socio-economic status and educational participation."

The study of educational attainments using data on twins exploits two features of twins reared together that can minimise the adverse consequences of the omitted variables problem noted above. First, identical (monozygotic) twins reared together will have, by definition, the same genetic endowments and have a shared family environment. Differences in the educational attainments of such twins must therefore be accounted for by individual-specific environmental factors (e.g. childhood accidents and similar events specific to one twin only).[5] Non-identical (dizygotic) twins who are reared together will share the same family background, but will differ in genetic endowments: non-identical twins share, on average, one-half of their genes and are thus no more alike than other siblings. Differences in the schooling attainments of non-identical twins can therefore be attributed to either individual-specific environmental factors or differences in genetic endowments. In comparison, differences in the educational attainments of individuals can be attributed to any of (shared) family background factors, individual-specific factors or genetic differences. Comparisons of educational attainments across groups of monozygotic (MZ) twins, dizygotic (DZ) twins and individuals therefore offer an opportunity to assign weights to these sets of influences. This is the methodology that is used in the traditional study of variations in incomes using sets of twins (see, for example, Miller et al., 1995).[6] Behrman et al. (1980, 1994), Behrman and Taubman (1989) and Heath et al. (1985) have used kin correlations to determine the importance of genetic differences to variation in educational attainment.

In this study the model of DeFries and Fulker (1985) is used to analyse the variation in educational attainment of an Australian twin sample. Their model is best explained by first reviewing the basic variance components model used in twins research.

The principles of biometrical genetics permit theoretical twin correlations to be derived in a number of ways (see Neale & Cardon, 1992), with the most direct method being based on computation of the frequencies of all possible twin-pair genotypes that might arise in a random-mating population. These genotypes are held to account for variations across individuals in particular

[3] A more extensive list of possible influences is considered in Behrman and Taubman (1989, table 5). This list includes parental age, father's religion and various personal and job-related characteristics.

[4] Becker's investment model is used as the framework for this study. Behrman, Pollak, and Taubman (1982) and Behrman (1987) and Behrman (1987) propose a parental preference model which, under certain conditions, has similar implications for the distribution of educational resources among siblings as investment models. In the separable earnings-bequest version of the preference model, parental aversion to inequality among offspring affects the distribution of educational resources. They show that compensating (for smaller genetic endowments) and reinforcing strategies are each possible under alternative assumptions regarding parental aversion to inequality. In general, however, the distribution of educational resources depends on both parental preferences and the properties of the earnings function that relates a child's expected lifetime earnings to genetic endowments and human resource investments. The evidence reported by Behrman et al. (1982) and Behrman, Pollak, and Taubman (1995) favours the separable earnings-bequest model.

[5] Studies such as Loehlin and Nichols (1976) indicate that MZ differences in educational attainments are due to chance events. More recent work suggests that somatic mutation, somatic recombination, and differential methylation of DNA occurring early in embryogenesis or during development may be important influences on nonshared environmental factors (see, for example, Cote and Gyftodimou, 1991). See also Box 2 in Martin, Boomsma, and Machin (1997).

[6] For a more recent contribution on estimation of the return to schooling using MZ twins, see Behrman and Rosenzweig (1999), and for an overveiw, see Bound and Solon (1999a,b).

observable characteristics (phenotypes). Consequently, the expected covariances and variances constructed in this manner provide links between observable characteristics and genetic and shared environmental effects, the relative magnitude of which may be inferred through comparison of the covariance matrices of MZ and DZ twins.

With respect to a given observable characteristic such as schooling (S), the covariance between kin pairs may be written as:

$$\mathrm{COV}(S_i,S_{-i})=\mathrm{COV}(G_i,G_{-i})+\mathrm{COV}(E_i,E_{-i}) \tag{1}$$

where i denotes an individual, $-i$ his/her kin pair, G denotes genotype and E common environment.[7] Converting this into correlations by dividing through by the variance in schooling yields:

$$\mathrm{COV}(S_i,S_{-i})/\mathrm{VAR}(S_i)=\mathrm{COV}(G_i,G_{-i})/\mathrm{VAR}(S_i) \tag{2}$$
$$+\mathrm{COV}(E_i,E_{-i})/\mathrm{VAR}(S_i)$$

which may be rearranged as follows:

$$\mathrm{COV}(S_i,S_{-i})/\mathrm{VAR}(S_i)=\alpha_1\mathrm{VAR}(G_i)/\mathrm{VAR}(S_i) \tag{3}$$
$$+\alpha_2\mathrm{VAR}(E_i)/\mathrm{VAR}(S_i)$$

where α_1 is the correlation between the individuals' genotypes and α_2 is the correlation between the individuals' common environments.

In the case of identical twins reared together, α_1=1 and α_2=1. Thus,

$$\mathrm{COV}(S_i,S_{-i})/\mathrm{VAR}(S_i)=\mathrm{VAR}(G_i)/\mathrm{VAR}(S_i) \tag{4}$$
$$+\mathrm{VAR}(E_i)/\mathrm{VAR}(S_i)$$

or $r_{\mathrm{MZ}}=h^2+c^2$,

where r_{MZ} is the correlation coefficient between the twins on the observable characteristic, h^2 is the proportion of the phenotype variance due to genetic factors, and c^2 is the proportion of the phenotype variance due to environmental factors.[8] In the case of fraternal twins reared together, it is assumed here that α_2=1 so that the additive environmental influence is the same for DZ twins as for MZ twins. In addition, under standard assumptions (particularly random mating) the phenotypic correlation α_1=0.5. In this situation the DZ twin correlations are:[9]

$$r_{\mathrm{DZ}}=0.5h^2+c^2. \tag{5}$$

In the traditional approach to the derivation of h^2, the correlation coefficient for DZ twins is subtracted from that for MZ twins and multiplied by 2, viz.

$$h^2=2(r_{\mathrm{MZ}}-r_{\mathrm{DZ}})=2(h^2+c^2-0.5h^2-c^2) \tag{6}$$

and c^2 is obtained by subtracting the derived value of h^2 from r_{MZ}.

Estimates of the importance of heritability (h^2) and common environmental factors (c^2) in producing individual variations in educational attainments using the above or similar procedures have generally been computed to assess how equal educational opportunity is, or to assess the impact of policies that purportedly have equalised educational opportunity. Thus, Behrman and Taubman (1989) use this approach because "The share of the observed variation in schooling that is attributable to across-family variability in environment provides a measure of inequality in schooling opportunity". Tambs, Sundet, Magnus, and Berg (1989) use an estimate of the common environmental factors to assess the success of the greater egalitarianism introduced into the Norwegian educational system. Similarly, Heath et al. (1985) use the nature-nurture decomposition as part of an assessment of the extent to which family background can predict educational success when education opportunity is altered. For a discussion of the issues involved in assessing the types of questions that the various decompositions can and cannot be used to address, see Taubman (1978).

DeFries and Fulker (1985) present an alternative, regression-based method for deriving h^2 and c^2. The starting point of their analysis is the estimating equation:

$$S_{ji}=\beta_0+\beta_1S_{j-i}+\beta_2R_{ji}+v_{ji}\quad j=1,...n \tag{7}$$

where S_{ji} is an observable characteristic (hereafter assumed to be schooling) of a respondent who is a member of the jth twin pair, S_{j-i} is the schooling of the respondent's co-twin, R_{ji} is a coefficient of genetic relationship, which is defined using the fractions of gene frequencies derived in simple biometrical models, namely 1 for identical twins and 0.5 for non-identical twins, and v_{ji} is a stochastic disturbance term. As shown below, the definition of R employed here allows a simple interpretation to be attached to the estimates in the model of DeFries and Fulker (1985).

In the above equation, the OLS regression coefficient β_1 is interpreted in the usual manner as the net of zygosity covariance between the levels of schooling of twins divided by the net of zygosity variance in educational attainment. β_2 equals twice the difference between the means for MZ and DZ co-twins, *ceteris paribus*. DeFries and Fulker (1985) propose the t-statistic on this regression coefficient as a test of whether heritability matters in the explanation of the dependent variable. The

[7] This formulation assumes that the covariance between genotype and common environment is zero.

[8] The environmental effects incorporated into this model can be widened to include both common- and specific- environments. The common environment is that which is shared, while the specific component is essentially a residual. It will capture the non-shared environmental factors discussed in [5], and may be measured as $1-h^2-c^2$.

[9] Expected covariance matrices may also be constructed for twins reared apart and for other relatives. See, for example, Goldberger (1979).

model assumes that $E(v_{ji}S_{j-i})=0$. If $E(v_{ji}S_{j-i})\neq 0$ then the estimates obtained from the Defries and Fulker model will be inconsistent. An instrumental variables approach is considered to take account of this possibility.[10] A second regression equation considered by DeFries and Fulker (1985) is:[11]

$$S_{ji}=\beta_0+\beta_3S_{j-i}+\beta_4R_{ji}+\beta_5S_{j-i}R_{ji}+v_{ji} \tag{8}$$

β_5 is, by construction,[12] twice the difference between the MZ and DZ regression coefficients on the level of schooling of the co-twin. Accordingly, β_5, under the standard assumptions of an additive model, random mating, non-common environment of a DZ twin is not correlated with his/her co-twin's genes, provides a direct estimate of heritability, h^2, as given in Eq. (6).[13]

β_3 in this model is an estimate of the twin resemblance that is independent of genetic resemblance (as captured by the other model terms). β_3 therefore provides a direct estimate of common environmental influences, c^2.

The model of DeFries and Fulker (1985) outlined above, along with other standard variance components models, assumes independence of genetic and environmental effects. Studies may be designed to detect genotype-environmental covariance (covGE), and Eaves, Last, Martin, and Jinks (1977) outline and assess these. In almost all cases where covGE has been detected, it makes a quite minor contribution of the total variance (see Truett et al., 1994; Kendler et al., 1995 for recent studies based on large samples that have the power to detect covGE). The general finding from these studies suggests that inferences from models based on independence of genetic and environmental effects such as those used in this study are unlikely to be misleading.

DeFries and Fulker (1985) note (p. 472) that their regression model can be extended to include other independent variables, such as gender, age, ethnicity and environmental indices. The additional variables considered for inclusion in this analysis are the age, gender (GEN), and family background variables favoured by researchers in the educational attainment literature, namely parents' educational attainment (MumEd and DadEd), father's occupational status (DadOcc) and number of siblings (Sibs). The number of siblings variable may be simultaneously determined with the dependent variable in Eq. (9). To obtain a consistent estimate of β_{11} in Eq. (9) below would require the use of instrumental variables techniques, provided that there is available a variable that affects the quantity of children but not the level of schooling. The data set does not contain variables such as the price of child quality that could be used in this way, and this aspect of the specification needs to be borne in mind.

The inclusion of family background variables can be justified from the perspective of investment models of schooling outcomes. Age provides a partial control for cohort effects of the type reported by Heath et al. (1985). The gender variable will reflect factors that affect the (marginal) returns to education for males and females.[14] Hence, the specification of DeFries and Fulker's (1985) model employed in this study will be:

$$\begin{aligned}S_{ji}=&\beta_0+\beta_3S_{j-i}+\beta_4R_{ji}+\beta_5S_{j-i}R_{ji}+\beta_6\mathrm{GEN}_{ji}+\beta_7\mathrm{AGE}_j\\&+\beta_8\mathrm{MumEd}_j+\beta_9\mathrm{DadEd}_j+\beta_{10}\mathrm{DadOcc}_j+\beta_{11}\mathrm{Sibs}_j\\&+v_{ji}\end{aligned} \tag{9}$$

Comparison of the estimates associated with the variable constructed using information on the co-twin in Eqs. (8) and (9) will permit an assessment of the extent of the control for common environment achieved through inclusion in the estimating equation of measurable indicators of family background.[15] In other words, if the estimates of β_3 in Eqs. (8) and (9) do not differ appreciably,

[10] There is a lack of instruments in the data set, particularly when Eq. (9) is estimated. We use a twin's report on the co-twin's level of educational attainment as an instrument for the co-twin's educational attainment (Ashenfelter & Krueger, 1994, (p. 1163)). In addition, to reduce the effect of measurement error in the family background variables, we use the average of the twin's reports on the family background variables (see Ashenfelter & Krueger, 1994, (p. 1163)). Consideration was also given to using the average of the co-twin's own report and the respondent's report on the co-twin's level of educational attainment, though the results from this approach differ little from the OLS results. For commentary on the use of instrumental variables to correct for measurement error in the case of within-twins estimators, see Neumark (1999).

[11] See Behrman, Pollak, and Taubman (1989), table 3 for a similar approach.

[12] Recall that R_{ji} equals 1 for identical twins and 0.5 for non-identical twins.

[13] It is to be emphasised that the genetically determined difference in educational attainment can be influenced by social and educational policies (see, for example, Heath et al., 1985). Likewise, the environmental family effect will be sensitive to social conditions and educational policies. These issues have been themes in a number of studies investigating the genetic and environmental contributions to educational attainments (e.g., Vogler & Fulker, 1983; Heath et al., 1985; Tambs et al., 1989; Baker et al., 1996).

[14] The model terms in S_{j-1} and R_{ji} could be interacted with these additional regressors. Alternatively, separate analyses can be undertaken for males and females (see below) or various cohorts (see, for example, Baker, Treloar, Reynolds, Heath, & Martin, 1996; Heath et al., 1985).

[15] Behrman and Taubman (1989) use a two-step procedure to link the unobserved environmental factors to father's occupation and number of siblings. Their finding reported for father's occupation is similar to that reported in this paper while that for number of siblings is stronger than that contained in the present analysis.

then it would follow that measurable family background variables (father's educational attainment etc.) do not adequately capture the common environmental factors that actually influence educational outcomes. Such a situation would indicate that unmeasured factors which influence preferences for education are far more important (see the argument by Williams et al., 1987 noted earlier). Conversely, if the estimates do differ, then the change in the value of the estimate of β_3 indicates the contribution of measurable indicators of family background to the additive common environmental variance (c^2). This is an advantage of the approach adopted in this study.

The inclusion of direct measures of family background in the model of DeFries and Fulker (1985) has one consequence that needs to be noted. Where the direct measures of the environment are correlated with the genetic endowments that are identified by the co-twin's level of schooling, the genetic effect identified by the model will be distorted. Given the positive links between parent's genetic endowments and their level of socioeconomic status, the effect of the co-twin's genotype will tend to be minimised in this model. In other words, the results reported below will tend to provide a conservative indication of the genetic effect in educational attainment.

Comparison of the estimates of Eq. (9) with those derived from a conventional specification that does not include information on the co-twin will permit assessment of the bias associated with failure to control for genetic factors when examining educational attainments. For example, differences between the estimate of the impact of father's occupational status on educational attainment in the extended model of DeFries and Fulker and in the standard model estimated by researchers such as Micklewright (1989) or Miller and Volker (1989) would indicate that bias associated with failure to control for genetic endowments is present in the estimates derived from the more conventional models, as argued by Micklewright (1989), among others.

The focus of Eq. (9) is the variation in years of school. One dimension of schooling that cannot be analysed in this study is school quality. Schools are a state rather than local government function in Australia, and so the neighbourhood effects that might influence decisions are less important. Government funds are directed to both public and non-public schools, and government policy throughout this century has been to provide equal access to education for all individuals. Many statistics show few differences between private and public schools. For example, student/teacher ratios differ by only around two across the major types of schools, being lowest in the government schools (Freebairn, Porter, & Walsh, 1987 (p. 94)). Recent moves towards deregulation of the schools sector may result in larger differences across schools. These future changes, rather than the past, are argued by Baker et al. (1996) to be the appropriate focus of study that looks at extent to which parents' choice of schools based on quality differences affects the degree of genetic variability in educational attainments (p.101).[16]

DeFries and Fulker (1985) actually developed their model for the case where one twin has a deviant score on the variable of interest, thereby providing a natural index for assignment to the status of "twin" and "co-twin" in the context of Eqs. (8) and (9). The model has subsequently been generalised to random samples, and Cherny, Cardon, Fulker, and De Fries (1992) contains relevant details. Cherny et al. (1992) outline a number of ways that unbiased parameter estimates can be obtained with unselected samples. Thus, they note that each of the 2^N possible combinations of twin assignment (to the status of "twin" or "co-twin") could be considered and the average of estimates obtained used. Alternatively, the average of the estimates obtained from a smaller number of analyses based on random assignment could be used. Finally, each twin's score could be entered twice, once as "twin" and once as "co-twin", and Eq. (7) estimated on this double-entered data. Cherny et al. (1992) show that the parameter estimates obtained from the latter two, practical, methods are highly similar. The double entry method is used in this study. Following Cherny et al. (1992), all standard errors are adjusted for the correct degrees of freedom computed on the basis of the true sample size.[17]

3. Data

This study uses data from the Australian Twin Registry which were gathered in two surveys, in 1980–82 and 1988–89. This survey contains an exceptionally large sample of twins — around 3000 in all. The starting point for the data is a 12-page questionnaire mailed out to all 5967 twins aged over 18 years enrolled in the Australian National Health and Medical Research Council Twin Registry in 1980–82. Joining this registry and responding to the survey were both voluntary, but the twins were otherwise unselected. Replies were received from 3808 complete pairs (a 64 percent response rate). In 1988–89 this sample was followed up and 2943 twin pairs responded (a conditional response rate of 78 percent, and an unconditional rate of 49 percent) (see Appendix A for additional details on the data).

[16] For a study of school quality based on female twins, see Behrman, Rosenzweig, and Taubman (1996). They report that endowments, length of time at school and the quality of school time are complementary. Information on school quality is not available in the data set analysed in the current study.

[17] That is, all standard errors are computed using the number of twin pairs (n) as the sample size rather than the number of individuals which, given the double-entry system, is $2n$.

The Australian Twin Registry is a particularly rich data set, and contains information on a wide range of family background, demographic and labour market variables. It also contains large samples of both monozygotic and dizygotic twins. Zygosity determination for same-sex pairs was done on the basis of two self-report items in the 1980–82 survey (Jardine, Martin, & Henderson, 1984). If there were any inconsistencies with unequivocal zygosity assignment in the responses of the twins, they were contacted for further information and frequently supplied photographs which assisted in making the decision.

Analyses (Miller, Mulvey, & Martin, 1995) reveal that the average educational attainment of the twins recorded in the survey is a little more than one year higher than the national average recorded in the 1986 Australian Census of Population and Housing. On average, the twins are one year younger than the population of 20–64 year olds, and are more likely to be married than the total population. In part these differences are due to the attrition bias that is related to age, education, and marital status (see Baker et al., 1996). However, Baker et al.'s detailed analysis shows that the attrition bias does not lead to increased twin similarity in educational attainments, which is important to the present study. The samples used by Rouse (1999), Ashenfelter and Krueger (1994) and Behrman, Taubman, and Wales (1977) also have mean educational attainments in excess of the national average.

Only sets of twins where each member responded to the questions used in the study are included in the sample analysed. Hence, of the 2943 twin pairs, 170 were eliminated because they were older than 64 years[18] and 295 were eliminated because information on the various measures of family background was missing. This left 2478 complete pairs. Both males and females are included in the analysis. Mixed pairs of DZ twins are included in analyses pooled across males and females, but are omitted from all analyses done on separate samples of males and females.[19]

4. The estimates

We begin with a simple estimating equation representative of the type employed by a number of researchers in Australia. Hence, educational attainment is held to be determined by mother's level of education, father's level of education, father's occupational status, the number of siblings, gender and age. The first four variables largely capture socio-economic factors that affect access to education. Previous Australian research (e.g., Miller & Volker, 1989; Williams, Harsel, Clancy, Miller, & Greenwood, 1987) has established a strong, positive relationship between the socio-economic status of the family and school participation.[20] The empirical results relating to the influence of family size on educational attainment have been mixed. Thus, the analyses of Williams et al. (1987) and Rosier (1978) suggest that family size does not directly influence educational attainment while Miller and Volker (1989), (p. 55) argue that family size exerts "a strong, negative influence on school retention". A variable for gender is included in the model to capture any differences between males and females in household allocation of resources for education. Finally, as noted above, the age variable is intended to capture cohort effects. Our prior is that cohort effects may be modest for males but will, reflecting the change in attitudes towards female education in recent periods, be important for females. A linear term is used as this was found to be a statistically adequate description of the age influences in the data. The estimates are obtained using Ordinary Least Squares. An IV estimator is used later in this study.

Table 1 presents results for this type of equation. The first column lists results for a model estimated on data pooled across males and females while the second column augments this basic equation with an interaction term between age and gender. Columns (iii) and (iv) present results from analyses conducted on the separate samples of males and females. There are five salient features of the column (i) results.

First, the estimated effects associated with the level of education of the father and with the level of education of the mother are positive and virtually identical. These estimates indicate that the respondent's level of education rises by 0.16 of a year for each extra year of education possessed by one of the respondent's parents. For example, individuals whose father held a university degree would, on average, have one more year of education than an individual whose father left school at year 10.

[18] The age restriction yields a data set consistent with Miller et al. (1995). The absence of cohort effects in study of educational attaining using these data (see Baker et al., 1996) indicates that the age restriction should not affect the major findings of this study.

[19] This practice follows Miller, Mulvey, and Martin (1997), and it enables the gender/heritability interactions to be delineated.

[20] There has been some speculation over whether this relationship is attitude based or a reflection of differences in economic resources. Williams et al. (1987) suggest that the finding derives mainly from differences in preferences across social classes, with family wealth making only a modest net contribution to differences in educational attainment. Interpretation from the perspective of parental preference models is therefore warranted. See, for example, Behrman et al. (1982, 1995) and Behrman (1987).

Table 1
OLS estimates of educational attainment, Australian twins sample[a]

Variable	Total sample		Males[b]	Females[b]
Constant	10.709 (47.94)	9.754 (32.41)	9.011 (17.12)	10.221 (36.64)
Age	−0.041 (12.68)	−0.016 (2.37)	−0.001 (0.11)	−0.052 (13.37)
Fathers' education level	0.169 (11.33)	0.170 (11.42)	0.188 (5.70)	0.160 (8.06)
Mother's education level	0.156 (9.26)	0.154 (9.24)	0.166 (4.54)	0.156 (6.95)
Father's occupational status	0.013 (7.13)	0.013 (7.17)	0.011 (2.81)	0.013 (5.53)
Number of siblings	−0.095 (5.56)	−0.093 (5.50)	−0.055 (1.33)	−0.104 (4.81)
Female	−0.966 (14.35)	0.471 (1.77)	–[c]	–[c]
Female*age	–[c]	−0.038 (5.37)	–[c]	–[c]
Sample size	4956	4956	1184	2626
R^2	0.2556	0.2607	0.1516	0.2395

[a] Heteroscedasticity-consistent "*t*" statistics in parentheses.
[b] Mixed sex twin pairs are omitted from separate analyses of males and females.
[c] Variable not entered.

Second, the occupational status of the father has a strong, positive effect on educational outcomes. A 10-point rise in the status attainment index would lead to a 0.13-year rise in the level of schooling. Hence, comparing the children of a sales assistant (occupational status score of 25.3) with the children of a general manager (occupational status score of 76.0), we would expect the mean level of education of the latter to be 0.7 years greater than that of the former. Third, there is a strong, negative relationship between family size and the level of educational attainment. This is suggestive of a quantity-quality trade-off.[21] Fourth, females are shown to have, on average, one year less education than males. It is well known, however, that female school participation rates increased rapidly throughout the 1960s and 1970s and males and females now participate in education to a similar degree. This phenomenon is captured in the column (ii) results that we will discuss shortly. Finally, it is to be noted that the age variable is negative and highly significant: as expected, educational attainments are greater among the younger age groups.

Column (ii) augments the column (i) specification by adding an interaction term between the female and age variables. The two partial derivatives that are of interest in this table are $\partial EDUC/\partial Female=0.471-0.038Age$ and $\partial EDUC/\partial Age=-0.016-0.038Female$. The latter partial effect shows that the negative impact of age on educational attainment (i.e., the cohort effect) is more intense for females that for males. In other words, in terms of educational achievements, females have been catching up with their male counterparts. This is also evident in the partial effect of gender on educational attainment, which shows that there is a quite small gender impact among the younger age groups (e.g., about one-third of one year for 20 year olds) and a more substantial impact for older age groups (for example, one and one-half years for 50 year olds).

Columns (iii) and (iv) list results for males and females respectively.[22] The important findings from this disaggregated analysis are that the age variable (i.e., cohort effects) and the number of children variable do not significantly affect the educational attainment of males while these factors exercise significant influences on the educational attainments of females. As discussed previously, the finding with respect to the age variable is not surprising. However, the finding with respect to the family size variable is surprising. A quality-quantity trade off has been established in the literature. Yet according to these results, it is a finding that holds for females but not for males.

In summary, while containing an interesting finding with respect to the quality-quantity trade-off for males, these results are consistent with those reported in the literature. This model of educational attainment explains some 25 percent of the variation in educational attainment around its mean.

Estimates from the standard DeFries and Fulker (DFF) model are listed in column (i) of Table 2. Recall that the coefficient on the interaction term between the co-twin's

[21] Note, however, that the quality-quantity trade-off is not usually measured by including one demand variable (quantity) in the equation for the other demand variable (quality). Rather, this trade-off is quantified by there being opposite signs on variables in separate equations estimated for quantity and quality. The interpretation attached to the family size variable here is indicative of a general pattern only. For a detailed study of the quality/quantity trade-off using data on twins, see Rosenzweig and Wolpin (1980). On the use of schooling as a proxy for child quality, see Behrman (1987).

[22] The greater number of female respondents compared to that of males is typical in voluntary surveys of twins.

Table 2
OLS estimates of DeFries and Fulker model of educational attainment, Australian twins sample[a]

Variable	Total sample (i)	(ii)	Males[b] (iii)	(iv)	Females[b] (v)	(vi)
Constant	9.333 (13.94)	10.123 (15.15)	9.664 (5.43)	9.037 (5.27)	8.195 (8.14)	9.043 (9.14)
Co-twin's educational attainment (S_{j-i})	0.222 (4.03)	0.092 (1.75)	0.245 (1.81)	0.129 (0.99)	0.282 (3.18)	0.148 (1.78)
Coefficient of genetic relationship (R_{ji})	−5.864 (7.34)	−5.511 (7.23)	−5.150 (2.58)	−5.566 (2.93)	−4.769 (4.18)	−4.929 (4.57)
$S_{j-i}*R_{ji}$	0.485 (7.26)	0.474 (7.43)	0.406 (2.67)	0.448 (3.09)	0.420 (4.16)	0.437 (4.59)
Age	–[c]	−0.023 (5.64)	–[c]	0.002 (0.15)	–[c]	−0.025 (5.02)
Father's education level	–[c]	0.094 (4.83)	–[c]	0.098 (2.34)	–[c]	0.082 (3.22)
Mother's education level	–[c]	0.087 (3.97)	–[c]	0.090 (1.99)	–[c]	0.077 (2.60)
Number of siblings	–[c]	−0.056 (2.50)	–[c]	−0.032 (0.64)	–[c]	−0.053 (1.85)
Father's occupational status	–[c]	0.007 (3.12)	–[c]	0.006 (1.14)	–[c]	0.006 (2.21)
Female	–[c]	−0.632 (7.33)	–[c]	–[c]	–[c]	–[c]
Sample size	2478	2478	592	592	1313	1313
R^2	0.3557	0.4204	0.3331	0.3669	0.4015	0.4457

[a] Heteroscedasticity-consistent "t" statistics in parentheses.
[b] Mixed sex twin pairs are omitted from separate analyses of males and females.
[c] Variable not entered.

educational attainment and the coefficient of genetic relationship provides a measure of heritability (h^2) while the coefficient on the co-twin's level of education provides a measure of common environmentality (c^2). These estimates are, respectively, 0.49 and 0.22. In other words, almost one-half of the variance in educational attainments is linked to genetic endowments, while almost one-quarter is linked to common environmental factors. By way of comparison, Taubman (1976) uses the variance components method to present evidence on the links between educational attainment and genetic endowments and shared environmentality based on study of a sample of twins who served in the US Armed Forces during World War II. Taubman's "most plausible" values for the components of the variation in educational attainments due to genetic endowments and shared environment were 40 percent and 30 percent respectively. In Behrman and Taubman (1989) the genetic contribution to the variance in educational attainments is reported to be around 80 percent. Baker et al. (1996), on the basis of a psychometric model of educational attainment using Australian Twin Register data, report values for heritability of 0.5 to 0.6 and a value for shared environmental factors of 0.2.

How do these results change when information on observed characteristics of the family is included in the model?

Column (ii) contains relevant results. In this equation, estimated for the full sample, the basic DFF model is augmented with variables for father's educational attainment, mother's educational attainment, father's occupational status, and number of siblings and age. Each of these auxiliary regressors in the twins model is statistically significant, though the estimated impacts are only around one-half of those established in models estimated when the sample is treated as one of individuals. For example, each additional year of education of the father was estimated in Table 1 to lead to an extra 0.17 of a year of education for the respondent. Yet in the context of the DFF model, the impact of the level of education of the father is only 0.09. Similarly, each additional sibling was found to reduce the educational attainment of the respondent by 0.1 of a year in the Table 1 model, while in the augmented DFF model the reduction is estimated to be only 0.05. As outlined earlier, these changes indicate the presence of substantial bias in the estimates of the determinants of educational attainment derived from the standard model. This bias appears to be greater than that estimated by Micklewright (1989) through including measures of scholastic achievement in the estimating equation.

The estimate of the contribution of genetic variation (h^2) is not affected by the addition of the easily measured environmentality indices; the point estimate being 0.47 whereas the initial estimate derived was 0.49. Recall that this extended version of the model of DeFries and Fulker (1985) provides a conservative estimate of h^2.

The summary measure of common environmentality is affected by the addition of the specific measures of common environment. In the column (iv) results the estimate of common environmentality (c^2) is only 0.09, compared to the earlier estimate of 0.28. Similar findings hold when separate equations are estimated for males and females (columns (iii) to (vi)), where we observe that the genetic contribution is between 0.4 and 0.45 for each estimation, and the indirect estimate of common environmentality (c^2) falls by one-half when the measured indices of common environmentality are included in the estimation. This finding indicates that the variables for family background typically included in models of educational attainment estimated on random samples of individuals capture around one-half of the major influences in this regard.

It is noted that the results for females with respect to common environmentality (column (vi)) are much stronger than those for males (column (iv)), although the basic patterns are the same. This may be due to the smaller sample of males (592) than of females (1313), or be indicative of family effects being less influential with respect to the educational attainments of males. In this regard it is noted that Heath et al. (1985) report that the environmental impact of family background on educational attainment in Norway was quite strong for females, accounting for 41–50 percent of the variance, but relatively weak for males, accounting for only 8–10 percent of the variance.[23]

Hence, we conclude from these analyses that the major determinant of educational attainments is in fact genetic endowments, with at least 45 percent of the variance in educational attainments being attributed to this factor. Shared environmentality is less important, accounting for only about one-quarter of the variance in educational attainments. The importance of this finding for educational policy is canvassed in the final section of this paper. The standard measures of common environmentality used by many researchers (education levels of parents, father's occupational status etc.) provide a crude, though practical, approximation to these shared family effects. However, other unmeasured shared family factors, captured here through the inclusion in the estimating equation of the education level of the co-twin, are as important as the factors that can be easily measured. Importantly, there are considerable differences between the Table 1 results and those obtained using the more encompassing model of DeFries and Fulker (1985). An implication of this is that the traditional approach to estimation based on study of random samples of individuals may provide estimates that are misleading. For example, the estimate of the impact of the father's occupational status in Table 1 is a 100 percent over-estimate of the impact of this variable on educational outcomes recorded in Table 2. Accordingly, some revision of the conclusions of previous studies could be warranted.

While the evidence presented above shows that around one-half of the variance in educational attainments can be linked to genetic endowments, there is an important caveat which suggests that this is a lower bound. Thus, Plomin and Bergeman (1991) argue that the initial findings from research on the nature of nurture are sufficiently strong to challenge the assumption that the measures labelled in the current study as environment are in fact measures of environment. In particular, the biometric genetic model as fitted here assumes random mating with respect to the variable of interest. But we know this is far from the truth for educational attainment, there being a substantial correlation between husband's and wife's educational attainments (see Baker et al., 1996). Given that educational attainment is partly heritable, the effect of assortative mating will be to increase the genetic variance between families. Thus, what is formally estimated as shared family environment is completely confounded with extra additive genetic variance arising from assortative mating for educational attainment. However, if we have an independent estimate of the marital correlation for educational attainment (μ), and the random mating estimate of additive genetic variance and common environment (h^2_{Unadj} and c^2_{Unadj}), the adjusted effect of additive genetic factors is given by $c^2_{\mathrm{Adj}}=c^2_{\mathrm{Unadj}}-h^2_{\mathrm{Unadj}}A/(1-A)$, where $A=0.5(1-\sqrt{(1-4\mu h^2_{\mathrm{Unadj}})})$ (see Martin, 1978 (pp. 13–23) for further details). Application of Martin's (1978) adjustment factor in the current analysis indicates that genetic endowments account for around 65 percent of the variation in educational attainment.[24]

Table 3 lists instrumental variable estimates of the DeFries and Fulker (1985) model. The focus of this discussion will be on the coefficients for the co-twin's educational attainment ($S_{j\text{-}i}$) as the direct estimate of c^2, and the interaction term between schooling and the coefficient of genetic relationship ($S_{j-i}R_{ji}$) that provides the direct estimate of h^2. The IV estimates of h^2 do not differ appreciably from the OLS estimates: they are lower than, but well within one standard error of, the OLS estimates. However, the IV estimates of c^2 are appreciably larger than the OLS estimates in both the basic DFF model and the extended version of this model. Hence the IV estimates of c^2 are between 0.46 and 0.57, compared to the values of 0.22 and 0.09 obtained using OLS. Adjustment for assortative mating using the method outlined in Mar-

[23] These results hold for individuals born after 1940 and before 1961. For individuals born before 1940, family background effects were quite strong for both males and females. Heath et al. (1985) attribute the change to the more liberal social and educational policies introduced in Norway after the Second World War. Behrman et al. (1989) provide an examination of the impact of policies in the US that provided equal access to financing for education for WW II veterans and which encouraged the acquisition of greater levels of schooling. Their conclusion is similar to that of Heath et al. (1985).

[24] This level is slightly lower than the finding in Behrman and Taubman (1989). They report that around 80 percent of the variance in educational attainments is due to genetic endowments.

Table 3
IV estimates of DeFries and Fulker model of educational attainment, Australian twins sample[a]

Variable	OLS	IV	OLS	IV
Constant	9.333 (13.94)	5.121 (4.60)	10.123 (15.15)	6.346 (5.80)
Co-twin's educational attainment (S_{j-i})	0.222 (4.03)	0.573 (6.34)	0.092 (1.75)	0.455 (5.09)
Coefficient of genetic relationship (R_{ji})	−5.864 (7.34)	−5.503 (3.70)	−5.511 (7.23)	−5.050 (3.59)
$S_{j-i}*R_{ji}$	0.485 (7.26)	0.459 (3.79)	0.474 (7.43)	0.430 (3.75)
Age	–[b]	–[b]	−0.023 (5.64)	−0.009 (1.93)
Father's education level	–[b]	–[b]	0.094 (4.83)	0.035 (1.65)
Mother's education level	–[b]	–[b]	0.087 (3.97)	0.035 (1.47)
Number of siblings	–[b]	–[b]	−0.056 (2.50)	−0.024 (1.03)
Father's occupational status	–[b]	–[b]	0.007 (3.12)	0.003 (1.78)
Female	–[b]	–[b]	−0.632 (7.33)	−0.457 (4.67)
Sample size	2478	478	2478	2478
R^2	0.3557	0.2733	0.4204	0.3833

[a] Heteroscedasticity-consistent "*t*" statistics in parentheses.
[b] Variable not entered.

tin (1978) reduces the estimate of c^2 to between 0.3 and 0.4 and again raises the estimate of h^2 to around 0.65. But even with this adjustment for the extra additive genetic component due to assortative mating, the IV estimates ascribe a stronger role to common environmental factors. Ashenfelter and Krueger (1994) also find that the empirical findings in their study of wages are sensitive to the treatment of measurement errors in self-reported schooling data.

5. Differential heritability?

The analyses above have shown that the major factor in accounting for variance in educational attainments is genetic endowments. Cherny et al. (1992) argue that heritability may differ as a function of the phenotype and that such differences have important consequences for attempts to estimate heritability and for the policy implications derived from estimates. The first issue is easily interpreted in terms of model specification. Thus, Cherny et al. (1992) suggest that Eq. (8) may be written as:

$$S_{ji}=\beta_0+\beta_3S_{j-i}+\beta_4R_{ji}+\beta_5S_{j-i}R_{ji}+\beta_6S^2_{j-i}+\beta_7S^2_{j-i}R_{ji}+\gamma_{ji} \qquad (10)$$

where β_6 estimates the change in common environmentality as a function of S_{-i}.[25] and β_7 estimates the change in heritability as a function of S_{-i}. Higher order interaction terms may also be considered for inclusion in the estimating equation. Exclusion of the S^2_{j-i} and $S^2_{j-i}R_{ji}$ terms thus amounts to a misspecification of the estimating equation, and resulting parameter estimates may be biased. If β_6 is non-zero, then it implies that the impact of shared environment differs across educational attainments. A negative parameter, for example, would indicate that education policy would be more efficacious at lower levels of educational outcomes than at the tertiary level. Similar reasoning holds with respect to heritability.

Relevant estimates are presented in Table 4. Column (i) presents estimates obtained when the data are pooled across males and females, column (ii) lists results for males only while column (iii) lists results for females only.

The coefficient on $S^2_{j-i}R_{ji}$, which records differential heritability, is insignificant in each of the equations in Table 4. Hence, there is no evidence of differential heritability with respect to educational attainment in these data. In other words, while the evidence in this paper suggests that heritability is important (in that h^2 is sizeable), the degree of its importance does not vary across educational attainments for either males or females.

The coefficient on S^2_{j-i}, which records differential environmentality, is significant for males and for the total sample but is insignificant for females. The negative coefficient for males indicates that shared family effects are more important among males at the earlier educational attainments. Therefore, the totality of our evidence suggests that the family environment is important (and approximately equally so for males and females in that c^2 is of the same order of magnitude for these groups in the basic model of DeFries & Fulker, 1985), and that the degree of its importance varies across educational attainments, at least for males. The sensitivity of this effect to level of education, whereby shared family effects are more important among males at the earlier educational attainments, would be expected to accentuate the inter-generational transmission of inequality. In this

[25] This was measured in Eq. (8) as $\beta_3=\partial S_{ji}/\partial S_{j-i}$, net of zygosity influences (R_{ji}).

Table 4
OLS estimates of Cherny et al.'s model of educational attainment, Australian twins sample[a]

Variable	Total sample	Males[b]	Females[b]
Constant	5.530 (2.15)	−1.973 (0.32)	6.805 (1.72)
Co-twin's educational attainment (S_{j-i})	0.857 (2.05)	1.889 (1.97)	0.536 (0.81)
Coefficient of genetic relationship (R_{ji})	−1.803 (0.58)	3.816 (0.54)	−3.710 (0.83)
S_{j-i}*R_{ji}	−0.156 (0.30)	−1.077 (0.97)	0.219 (0.29)
Age	−0.021 (5.01)	0.003 (0.26)	−0.024 (4.70)
Father's education level	0.093 (4.76)	0.094 (2.24)	0.082 (3.20)
Mother's education level	0.085 (3.90)	0.090 (1.98)	0.074 (2.51)
Number of siblings	−0.052 (2.36)	−0.025 (0.51)	−0.050 (1.74)
Father's occupational status	0.007 (3.19)	0.006 (1.31)	0.006 (2.19)
Female	−0.631 (7.33)	–[c]	–[c]
S^2_{j-i}	−0.031 (1.84)	−0.068 (1.85)	−0.016 (0.60)
$S^2_{j-i}R_{ji}$	0.026 (1.23)	0.059 (1.38)	0.009 (0.30)
Sample size	2478	592	1313
R^2	0.4221	0.3720	0.4461

[a] Heteroscedasticity-consistent "*t*" statistics in parentheses.
[b] Mixed sex twin pairs are omitted from separate analyses of males and females.
[c] Variable not entered.

regard it is well recognised in the educational attainment literature that the important thresholds with respect to educational attainments occur at an early stage in the spectrum of levels of education. For example, James (1988), (p. 11) notes:

> It is also clear that aid alone will not bring about equal access. For example, until their earlier educational environments are equalised, disadvantaged groups will continue to have lower participation in higher education, particularly in the most selective schools.

6. Conclusion

An individual's educational attainment governs much of their lifetime prospects. Accordingly, understanding the determinants of educational attainment has much to recommend it. Application of the model of DeFries and Fulker (1985) shows that genetic endowments account for around 50 percent of the variance in educational attainments in Australia, and shared family factors between 25 and 50 percent, depending on the method of estimation. Adjustment for assortative mating using the correction provided by Martin (1978) results in a contribution of genetic endowments to the variance in educational attainments of around 65 percent, and a contribution of shared family factors of between 10 and 40 percent.

Most policy interest in Australia has to date centred on the impact that family background has on education outcomes, for much of government policy has attempted to affect the acquisition of human capital through altering access to financial resources for education (e.g. AUSTUDY[26]). The results from our estimation of an augmented DeFries and Fulker model shows that the family background measures traditionally included by many economists in models of educational attainment are crude approximations to the wider range of environmental effects incorporated in the models estimated in the behavioural genetics literature. Indeed, our analyses suggest that the partial effects associated with variables like the educational attainments of the parents, the occupational status of the father, the number of siblings etc in the augmented model of DeFries and Fulker are typically less than 50 percent of those obtained from models previously estimated in Australia. Thus, the bias in the traditional estimates is greater than the one-third to one-half suggested by Micklewright (1989). There is a possibility that the impact of common environmentality on educational outcomes may differ across educational attainments for males. In particular, it seems that the importance of shared family background in the determination of educational attainment is more intense at the early educational attainments than at the later educational attainments. This finding would be consistent with the

[26] AUSTUDY provides financial assistance to Australian students on a non-competitive basis. Eligibility is subject to means tests (on own income, parents' income, parents' assets). An interesting feature of the financial assistance scheme is that students can trade parts of their non-repayable grant for double the value in income-contingent loans.

education literature that stresses the importance of educational environments at the early schooling levels.

Our main empirical finding is the relatively low fraction of the variance in educational attainments in Australia that is due to shared environmental factors. It seems that only around 25 percent of the variance in educational attainments may be attributable to such factors, though this contribution is shown to be around 40 percent when adjustments for measurement errors and assortative mating are made. This provides a measure of inequality of schooling opportunity in the Becker tradition that is broadly similar to findings reported by Heath et al. (1985), Vogler and Fulker (1983), Tambs et al. (1989), Behrman and Taubman (1989) and Baker et al. (1996). Hence, this finding appears to be robust across data sets.

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Appendix A. The Australian twin registry data

The data used in this study begin with a mail survey undertaken in 1980-82 of all twins aged over 18 years enrolled in the Australian National Health and Medical Research Council Twin Registry at that time.

Joining this registry and responding to the survey were both voluntary. Replies were received from 3808 complete pairs. In 1988-89 this sample was followed up and 2934 twin pairs responded.

The survey gathered information on the extent of contact between twin pairs during childhood and during the eight years prior to the second survey, information on the respondent's family background (parents, siblings, marital status and children), socio-economic status (education, employment status, income, occupation), personal details (body size, smoking and drinking habits, general health), personality, feelings and attitudes.

Details on the construction of the key variables are set out below.

Educational Attainment: All the education variables (respondent's, father's, mother's) were coded in the survey to a seven-point scale: <7 years of schooling; 8–10 years of schooling, 11–12 years of schooling; Apprenticeship, Diploma, Certificate; Technical or Teachers' College; University — first degree; University — postgraduate degree. These categories have been recoded as 5, 9, 11.5, 11.5, 13, 15 and 17 years of education, respectively.

Father's Occupational Status: The survey data on occupation of employment of the respondent's father are used to derive an occupational status score using the information contained in Jones (1989). This score was scaled by Jones to range from 0 to 100. The scale has a positive skew, with high scores among administrators and professionals and low scores among labourers.

Coefficient of Genetic Relationship: This variable is set equal to one in the case of monozygotic twins and equal to one-half in the case of dizygotic twins.

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