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Genetics of Educational Attainment in Australian Twins: Sex Differences and Secular Changes

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The relative effects of genetic and environmental factors in producing individual differences in educational achievement are compared across women and men and over birth cohorts. In a large sample of Australian twin pairs, the heritability of self-reported educational attainment did not vary among women and men born before and after 1950. In a "psychometric" model of twin resemblance, based on separate self-reports in 1981 and 1989, genetic factors explained 57% of the stable variance in educational achievement, while environmental factors shared by twins accounted for 24% of the variance. Corrections for phenotypic assortative mating for educational level, however, suggested that estimated common-environmental effects could be entirely explained by the correlation between additive genetic values for mates. Taking this into account, heritability of "true" educational attainment in Australia may be as high as 82% with the remaining variation being due to individual environments or experiences. Unlike previous studies in Scandinavian countries, results in Australia suggest that factors influencing educational success are comparable between women and men and for individuals born at different points during this century.

KEY WORDS: Educational attainment; Australian twins; sex differences; secular changes.

INTRODUCTION

Comparing the relative effects of underlying genetic and environmental factors on educational attainment for women and men and over time can provide insight into the effects of social policy on groups of individuals. Attempts to provide equal educational opportunities to individuals within a population, if successful, should lead to more uniform environmental factors, and thus a greater relative genetic variation (i.e., heritability) (Scarr-Salapatek, 1971). As social policies change over time, or are applied differentially to different subgroups such as women and men or various social classes, different heritabilities should become apparent between these subgroups and/or over time.

Educational attainment is a phenotype that has been affected in Western society by several sociopolitical events in this century such as wars, changes in civil rights policies, changing reproductive practices associated *inter alia* with the introduction of the birth control pill, and the legalization of abortion. Changes in average levels of schooling have been apparent over the last 75 years, for both women and men, in most Western countries, including Australia (Castles, 1992a). Nonetheless, considerable individual differences within both sexes have continued to be apparent throughout this

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century. The causes of these individual differences, and how these causes may have changed over time for both women and men, are the focal points of this paper. To address both secular changes and sex differences, the relative impacts of genes and environment on educational attainment are studied in a large sample of female and male Australian twins born between 1893 and 1965.

There is already considerable evidence for the effects of genetic variation on educational attainment in men across various countries and birth cohorts. In an adoption study of Danish brothers (born between 1938 and 1947), heritable influences accounted for 68% of the variance, while shared family environmental factors explained another 27% of the variance in educational level reached between 18 and 26 years of age (Teasdale and Owen, 1984). In a joint analysis of parents and offspring, spouses, siblings, and MZ and DZ twins from several American twin and family studies, Vogler and Fulker (1983) have reported a heritability (h^2) of .56 for educational attainment in males (born primarily before 1945). Although there were also effects of environmental factors shared by siblings (explaining 19% of variation in twins and 21% of variation in nontwins), environmental transmission from parents to offspring was negligible. In a Norwegian sample of male twins, substantial heritable variation in educational achievement was also found for those born after 1935 (h^2 = .51). However, genetic effects were reported to be considerably lower (h^2 = .10) and shared twin environmental variation higher (c^2 = .62) in men born before 1935 (Tambs *et al.*, 1989). Thus, some secular changes in heritable influences in educational achievement may be apparent, at least in men.

There have been marked differences between educational attainments in women compared to men, although women's gains during this century (and even in the last decade) have reduced these mean differences considerably [see Castles (1992a) regarding these changes in Australia]. In spite of this, scant attention has been paid to the relative effects of genetic and environmental factors in women. One notable exception is the study of Norwegian twins, where Heath *et al.* (1985) examined both secular changes and sex differences in heritability and shared environmental factors. Although heritability for educational level was comparable for women and men born before 1940 (h^2 = .41), the relative impact of genetic factors increased dramatically for men (h^2 = .67–.74) but not for women (h^2 = .38–.45) born after 1940. These findings have been taken to imply that social policy affecting educational reforms after World War II have led to more (innate) ability-related success for men. Conversely, family environmental factors still play a major role in the educational success of Norwegian women, rather than genetically based individual differences. Heath *et al.* discussed the possibility that educational reforms in Norway may have led to more equal opportunities (i.e., less environmental variation) for men, but not for women born in the later part of this century.

The provocative findings in Norwegian twins led us to examine sex and cohort differences in genetic and environmental influences in the Australian educational system. Although there are many worldwide economic and political factors which may have influenced Norwegian and Australian societies in parallel, there are additional factors specific to Australia which may have influenced its educational system independently of that in Norway. For example, legislation between 1870 and 1900 was enacted to ensure that the government should provide compulsory, free, and secular education to its residents. Major changes in the Australian educational system since that time have primarily corresponded to increasing participation in private schooling, and the contribution of state funds to nonpublic schools, which has been legal in Australia since the late 1950s (see Anderson, 1993). Thus, the provision of equal access to education for all individuals in Australia since the turn of this century may suggest that heritable variation of educational attainments will be greater in older birth cohorts in Australia compared to Norway. Other factors which may have differentially affected women's and men's educational attainments throughout this century (e.g., sex-role attitudes) may lead to different effects of environmental and genetic variations in the two sexes.

METHOD

Subjects. The present study is based on a sample of 3808 pairs of Australian twins who participated in an initial mail survey in 1981, which included questions about their educational achievements (see Jardine *et al.*, 1984). From this total sample, 83.1% of the individuals (n = 6327) completed a similar follow-up survey in 1989 (including 2995 complete pairs, or 78.7% of the total sample). Excluding persons who had died or were too sick to participate (n = 217 individuals), or with whom contact could no longer be made by mail or phone (n = 270 individuals), return rates for those who received and were able to complete the 1989 survey were actually quite remarkable (88.8% personwise and 85.6% pairwise). The most severe attrition effects occurred among younger females (due to loss of contact, possibly as a function of changes in surname after marriage) and older males. Of the entire sample of 7616 individuals participating in the 1981 survey, there were significantly lower return rates for the 1989 survey for males (80.52%), opposite-sex twin individuals (79.88%), and older persons (born before 1939; 74.6%). Concurrent with this age-related bias were lower return rates for widowed (60.34%) and/or retired (51.75%) individuals. It is also noteworthy that lower return rates were present for unemployed persons (75.81%) and those individuals reporting lower educational attainments in 1981. The effects of these sampling biases in the longitudinal sample on the present analyses of twin similarity for educational attainment are discussed in greater detail below.

Twin pairs with educational data available from both the 1981 and the 1989 surveys were further divided into birth cohorts. The complete age distribution for this study is given by Martin *et al.* (1991). To achieve maximum power for detecting cohort differences, particularly for relative contributions of genetic and environmental variations, a median split was performed on birth year for the total sample. Thus, Cohort 1 consists of twins born between 1893 and 1950, and Cohort 2 consists of twins born between 1951 and 1965. This division also proves meaningful from a historical perspective, as the two groups represent, roughly, individuals born before and after the Second World War. As described previously, other major changes in the privatization of schooling in Australia also occurred shortly after 1950, which adds interest to the comparison of the two birth cohorts described here. Sample sizes (n pairs) for those participating in the follow-up 1989 questionnaire, separately for the five twin-pair types within each cohort are as follows: for Cohort 1, 520 MZ female, 216 MZ male, 299 DZ female, 94 DZ male, and 270 DZ opposite-sex; and for Cohort 2, 479 MZ female, 226 MZ male, 290 DZ female, 161 DZ male, and 388 DZ opposite-sex. The structural equation modeling analyses, which are of primary interest in this paper, are therefore based on a total of 2943 twin pairs, with educational attainment data available from both the 1981 and the 1989 surveys.

A subset of individuals in the original sample also participated in a pilot study of the 1981 questionnaire approximately 3 months prior to its mass distribution to the entire sample. These data provided us the opportunity to examine test–retest properties of the educational achievements scale (n = 93 individuals) in this subsample who reported on their educational achievement at both times of assessment. A second "repeat" sample was obtained for another subset of individuals responding to the 1989 survey, thereby allowing a separate evaluation of test–retest properties using the follow-up survey (n = 870 individuals who reported their educational attainments during both of their completions of the 1989 survey).

Measures. Both the 1981 and the 1989 surveys contained questions about the respondent's "Educational Achievement," each with seven response categories (1 = "less than 7 years schooling"; 2 = "8–10 years schooling"; 3 = "11–12 years schooling"; 4 = "apprenticeship, diploma, certificate, etc."; 5 = "technical or teachers college"; 6 = "university first degree"; 7 = "university postgraduate training"). These response categories form an ordinal scale, which is treated as such in most analyses presented here. Polychoric correlations or log-linear regression models of response frequencies were deemed appropriate based on the scale properties and distribution.

Additionally, each twin was asked to report on his/her cotwin's educational achievement (using the same 7-point scale) in the 1989 survey, which is used in this paper to evaluate further the measurement properties of the self-report scale itself.

Zygosity determination for same-sex pairs was done on the basis of two self-report items in the 1981 survey (Jardine *et al.*, 1984). Such methods have been shown to give at least 95% agreement with diagnosis based on extensive blood-typing (Cederlof *et al.*, 1961; Kasriel and Eaves, 1976; Magnus *et al.*, 1983; Martin and Martin, 1975; Nichols and Bilbro, 1966; Ooki *et al.*, 1990). 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.

RESULTS

Preliminary Analyses. To facilitate examination of the distributions of educational level in 1981 and 1989, the 7-point educational attainment scale was recoded at each time point to form three categories of achievement, based upon whether or not the individual had (1) completed at least 11–12 years of school (i.e., most or all of secondary school), (2) completed at least a university first degree, and (3) completed some postgraduate education. The proportions of respondents falling in each of these three categories are presented in Table I, separately by sex, zygosity, and birth cohort. (Note that these three categories are not mutually exclusive. For example, all persons receiving at least a university first degree in the second category are also included in the first category of completing at least 11–12 years of school. Proportions in each mutually exclusive category could be easily obtained from the cumulative percentages presented in Table I.)

A cursory examination of Table I draws attention to several noteworthy points. First, there are clear cohort differences within all zygosity groups, whereby postwar individuals are more likely to complete both secondary and tertiary education. In comparing results for 1981 and 1989 responses, however, it is also apparent that considerable censorship occurred for the younger cohort in the 1981 survey. That is, many individuals had not completed their education by the time of the initial survey. Although there is little change in the proportions of Cohort 2 who completed secondary school, there are considerably more individuals, both women and men, who completed university first degrees and postgraduate training between the two surveys.

Second, there are also sex differences in these proportions, although these vary somewhat by cohort and category. In particular, the difference between achievements of women and men is far less apparent in the younger cohort, at both secondary and tertiary levels, in both the 1981 and the 1989 surveys. In comparing the two cohorts, the greatest differences occur between the older and the younger women, with postwar women being far more likely to complete both secondary education and university degrees. These findings are most dramatic in the 1989 survey, where less censorship of the younger cohort is apparent. In contrast, educational achievements of men, especially at the university level, do not vary much across birth cohorts.

Statistical analysis of these trends was performed using log-linear regression models; predicting each categorical response from Cohort (older, younger), Zygosity (MZ, DZ same-sex, DZ opposite-sex), Sex (male, female), and their respective interactions. These analyses were performed only on the first two categories, due to the relatively small proportions observed in the third (postgraduate) category. Results are summarized in Table II. Sex and cohort main effects are confirmed for completion of at least 11–12 years of school and for completion of university first degree, in both the 1981 and the 1989 responses. Overall, men and postwar individuals obtained higher educational levels than women and prewar individuals, respectively. However, the significant Sex × Cohort interactions for the second category (university first degree or higher) confirm that the sex difference is significantly less in the younger cohort. It is noteworthy that this attenuated sex difference in Cohort 2 is seen both in the 1981 responses and in the 1989 responses, where considerably less censorship for educational attainment occurred.

It should also be emphasized that no significant Zygosity main effects were observed for either educational achievement category in either the 1981 or 1989 surveys. This is important, as the biometrical analyses described below rest upon the assumption that level of response does not vary across zygosity group. However, some attention must also be drawn to the significant Zygosity X Sex and Zygosity X Sex X Cohort interactions obtained for the second achievement category ($\geq$11-12 years of school) in both 1981 and 1989 responses. A look at Table II shows the largest sex difference occurring for MZ twin individuals in the younger birth cohort, with an absolute difference of 30% between women and men, compared to only about a 1–16% difference between males and females in DZ twins in both cohorts.

The unusual elevation of secondary school achievement in the older MZ males led us to consider further the possible biases which may have resulted from sample truncation of less educated individuals. Namely, if the probability a given twin

Table I. Distribution (%) of Educational Levels by Twin Type and Cohort in Longitudinal Sample[a]

	Cohort 1 (1893–1950)			Cohort 2 (1951–1965)		
Group	≥ 11–12 yr	University	Postgrad.	≥ 11–12 yr	University	Postgrad.
			Education reported in 1981			
MZF	48.2	6.3	2.3	82.7	11.2	2.7
DZF	53.9	6.3	2.6	84.4	9.2	1.5
DZOF	59.7	3.8	.6	89.3	12.6	2.3
MZM	78.3	24.3	10.9	91.4	21.0	4.1
DZM	68.2	26.5	11.4	94.2	21.5	2.6
DZOM	66.3	17.5	6.4	90.4	20.0	4.2
			Education reported in 1989			
MZF	44.7	7.6	3.6	79.4	19.6	6.8
DZF	48.2	7.9	3.7	80.9	17.1	5.0
DZOF	52.4	5.4	1.9	87.9	21.4	7.5
MZM	74.3	28.3	13.6	91.4	34.7	9.9
DZM	66.8	30.8	17.1	93.9	36.3	11.6
DZOM	64.7	19.9	8.4	91.4	32.8	12.4

[a] Opposite-sex twins are separated into females (DZOF) and males (DZOM).

volunteers her/his participation in this study is correlated with educational level, the biometrical analyses in this paper could yield biased estimates of heritability and common environmentality (see Martin and Wilson, 1982; Neale *et al.*, 1989). We explored this possibility by examining educational levels in a separate sample of 555 "singleton" respondents (n = 342 females; n = 213 males) for the 1981 questionnaire (i.e., where only one twin participated) and comparing to the individuals where both twins participated ("concordant-cooperative" respondents). As Neale and Eaves (1993) have suggested, if the criterion variable under study (here, educational level) is causally related to liability to volunteer for study, one would expect mean (educational) differences between singleton and concordant-cooperative respondents. Furthermore, these mean differences will be exaggerated as the correlation between the two twins increases. Thus, truncation of less educated individuals in this study would be evidenced by lower educational levels in the 1981 singletons, particularly in the MZ individuals.

We found that distributions of self-reported educational attainment were remarkably similar (and not significantly different) for singleton and concordant cooperative females, for both MZ and DZ twins. In the women, therefore, there appeared to be little or no effect of truncation. This was also the case for males on the whole (ignoring zygosity). However, there was some suggestion of truncation of lower educational levels in MZ males, where distributions of singleton and concordant-cooperative men were significantly different (χ^2 = 17.38, df = 6, p = .008), with a lower proportion (.67) of MZ singleton males reporting completion of at least 11–12 years of school as compared to MZ concordant-cooperative males (.83). There were no significant differences between concordant-cooperative and singleton DZ males. The implications (particularly for our genetic analyses) of these differential effects of truncation in women and men are highlighted under Discussion.

Table II. Log-Linear Analysis of Educational Categories

		1981 report		1989 report	
Source	df	≤ 11–12 yr	University	≤ 11–12 yr	University
		χ^2	χ^2	χ^2	χ^2
Cohort	1	405.5*	14.4*	464.2*	105.0*
Zygosity	2	.3	3.9	1.8	5.0
Sex	1	80.8*	183.0*	125.7*	232.2*
C × Z	2	5.5	7.7*	5.9	7.3*
C × S	1	.2	24.2*	.1	24.7*
Z × S	2	23.9*	1.2	16.0*	2.3
C × Z × S	2	7.4*	.8	4.5	.4

* $p < .05$.

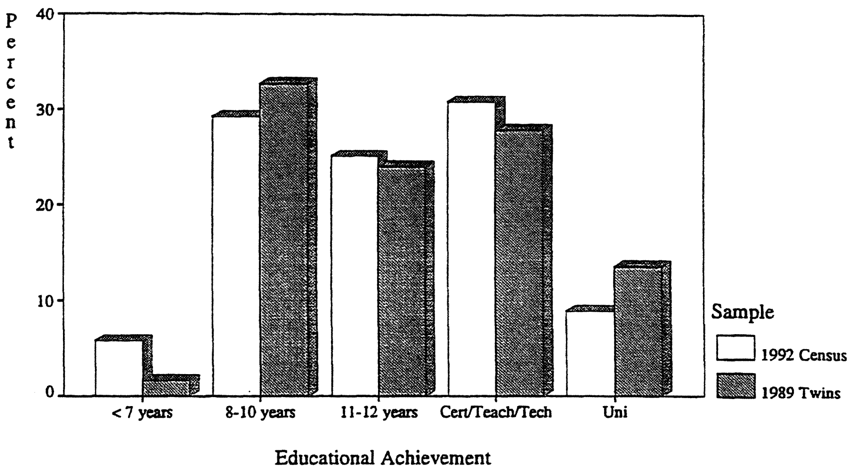

Fig. 1. Educational distributions for Australian census and twin samples: females. n = 3917 for twin sample; n = 4694.7K for census sample.

Reliability and Stability of Self-Reported Educational Achievement. Measurement properties of the educational achievement scale used in the 1981 and 1989 surveys were examined in two different ways. First, PRELIS 2.1 (Joreskog and Sorbom, 1993) was used to compute test–retest (polychoric) correlations in the 1981 and 1989 surveys. For the pilot sample of 93 individuals, these were r = .85 ± .12 (n = 58) for Cohort 1 and r = .89 ± .10 (n = 35) for Cohort 2, which did not vary appreciably between the cohorts. For the repeat sample of 870 individuals for the 1989 survey, test–retest correlations were also quite high and remarkably similar across cohorts: r = .94 ± .01 (n = 498) for Cohort 1 and r = .94 ± .02 (n = 372) for Cohort 2. Second, for pairs where each twin reported on educational level for self and cotwin in 1989, polychoric correlations between self-report and cotwins' report were computed within birth cohort. Again, these correlations were quite high and remarkably similar across cohort (r = .87 ± .01 for Cohort 1, n = 1238; r = .88 ± .01 for Cohort 2, n = 1331). Thus, both self-report/twin-report and short-term test–retest correlations indicate considerable reliability of self-reported educational achievements as measured in this study.

Comparing the test–retest and multiple rater indicators of reliability with the longitudinal (polychoric) correlations between 1981 and 1989 individual responses provides some additional insights into some important differences between birth cohorts in these data. Some cohort differences in the longitudinal correlations between 1981 and 1989 are apparent, with the older cohort showing 1981–1989 correlations about as high as the test-retest and self-report/twin-report correlations (r = .88 ± .01; n = 3041), while the 1981–1989 correlations are somewhat attenuated in the younger group (r = .77 ± .01; n = 3229). We attribute this largely to the greater censorship of educational achievements in the 1981 survey of the younger individuals. The attenuated longitudinal correlations in the younger cohort are quite consistent with the findings of greater changes between the two surveys for level of education in the younger cohort as discussed above (and see Table III). Nonetheless, the reliability of the scale itself is quite satisfactory within both cohorts for both 1981 and 1989 responses.

Sampling Bias Analyses. Several preliminary analyses were undertaken to investigate possible sampling biases in these data. First, the distribution of self-reported educational level for the sample of 1989 respondents (born in 1920 or later) was contrasted with census data (for a sample of men and women with a comparable age range) from the Australian Bureau of Statistics (ABS) (Castles, 1992).[6] As shown in Fig. 1, the proportions of women whose highest achievements were for sec-

[6] For purposes of comparison to the census data, twin survey responses 4 and 5 in the educational achievements scale were combined into one "Certificate" category, and responses 6 and 7 were combined into one "University" category (see Figs. 1 and 2).

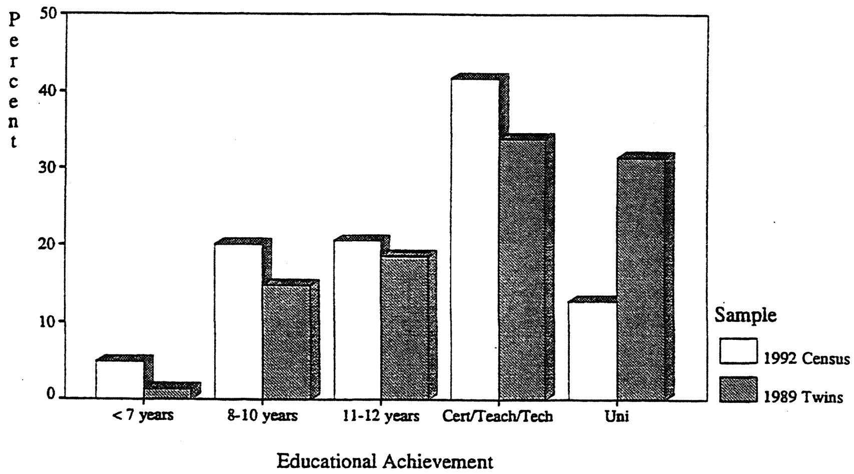

Fig. 2. Educational distributions for Australian census and twin samples: males. $n = 2114$ for twin sample; $n = 4725.5\text{K}$ for census sample.

ondary schooling (8–10 and 11–12 years) are remarkably similar between the present twin and the ABS census samples. There are also comparable proportions of females in both samples receiving some kind of postsecondary certification (i.e., combining the two highest categories in Fig. 1): 39.8% of the ABS women and 41.5% of the female twins. However, there is a greater tendency for women's postsecondary education to bear a university degree (or higher) in the twin sample than in the census sample. There are correspondingly fewer female twins than census individuals receiving teaching, technical, or other (nonuniversity) certification. A similar pattern is also found for men (see Fig. 2), whereby a notable difference between the twin and the census samples is seen in the distribution of postsecondary levels of achievement. However, unlike in women, the men in the twin sample are somewhat more likely in the first place to have achieved some kind of postsecondary education compared to the census sample: 54.5% of the ABS males and 65.3% of the male twins fell into the one of the two highest categories in Fig. 2.

Thus, there was a slight upward bias in educational achievement in the present study, in that more individuals (both women and men) in the present sample held university degrees than in the census sample (see Figs. 1 and 2). This tendency is somewhat more severe in the men in our study. The noted differences between ABS and twin samples described here could be attributed, in part, to different age distributions in the two samples (in spite of the comparable age range in the two groups). Specifically, the twin sample has an overrepresentation of younger individuals compared to the ABS sample, especially in males. Therefore the sampling biases with respect to education in this study quite possibly stem more directly from *age biases* rather than level of schooling per se.

In a further exploration of possible sampling biases, the effects of attrition in the follow-up survey in 1989 were examined by comparing education levels (reported in 1981) for twins who participated in both 1981 and 1989 (the "longitudinal sample") to those twins who participated in 1981 only (the "attrition sample"). Figure 3 shows that the longitudinal sample reports somewhat higher levels of education in 1981 (note the slightly higher proportions of individuals receiving 11–12 years of schooling, teacher or technical certificates, and university first degrees) compared to the attrition sample (who have a greater proportion of individuals receiving 8–10 years of schooling or less). These distribution differences are statistically significant ($\chi^2 = 48.90$, df $= 6$, $p < .001$), as are the mean education levels (3.59 and 3.33 for the longitudinal and attrition samples, respectively; $t = 5.88$, df $= 7614$, $p < .001$).

Although the elevated educational level of the longitudinal respondents may limit generalizations of the present findings to the general Australian population, of equal or greater importance to the biometrical analyses presented here is the extent to which sampling biases may have led to increased

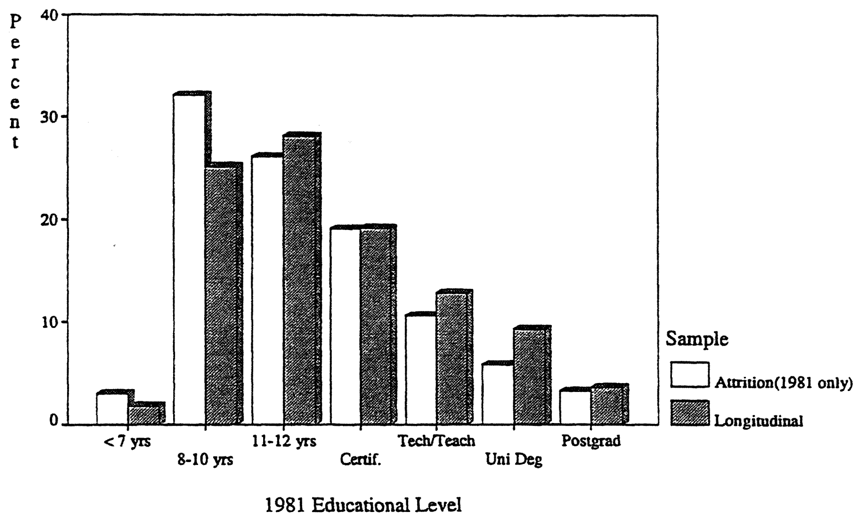

Fig. 3. Distribution of educational achievement for longitudinal and attrition samples. $n = 1289$ for attrition sample; $n = 6327$ for longitudinal sample.

twin *similarity*, for educational level in particular here. This issue was addressed in a series of hierarchical multiple regression analyses within each of the five twin types. Within each zygosity group, the following hierarchical regression equation was used to predict self-reported educational level of Twin-a in 1981 (ED81a) in three steps:

$$\begin{aligned} \text{ED81a} = {} & \text{ED81b} + \text{COHORT} + \text{SAMPLE} && \text{(Step 1)} \\ & + \text{ED81b} * \text{COHORT} + \text{ED81b} && \text{(Step 2)} \\ & \quad * \text{SAMPLE} + \text{COHORT} * \text{SAMPLE} && \\ & + \text{ED81b} * \text{COHORT} * \text{SAMPLE} && \text{(Step 3)} \end{aligned}$$

where ED81b is the self-reported 1981 educational level of Twin-b, COHORT reflects the birth cohort (0 = 1893–1950; 1 = 1951–1965), SAMPLE reflects longitudinal response status (0 = one or both twins did not respond in 1989; 1 = both twins responded in 1989), and the remaining product terms represent the interactions among the three variables entered in Step 1. While the three main-effect terms in Step 1 are interesting in that they provide, respectively, twin resemblance in educational level, mean cohort differences, and mean differences between longitudinal and attrition samples, the interaction terms are of most importance in these analyses. In particular, differential twin resemblance for longitudinal and attrition subsamples would be evidenced by a significant ED81b* SAMPLE interaction or by a significant ED81b*COHORT*SAMPLE interaction in Step 3 if sample differences in twin resemblance vary by cohort. Although there was significant twin resemblance across all five twin types, none of the interactions in Steps 2 and 3 were even marginally significant for any group. Thus, the apparent sampling biases resulting in a slightly more educated sample of longitudinal respondents did not have any corresponding effects on twin resemblance for educational attainment. Given this, further structural analysis is confined to the longitudinal sample.

It is also noteworthy that the regression analyses did not reveal any significant ED81b*COHORT interactions for any of the five zygosity groups, indicating a lack of change in twin resemblance across birth cohorts. This finding is next explored further in more rigorous biometrical model-fitting analyses.

Structural Modeling Analyses. The relative effects of genetic and environmental variation, and how these effects may vary over sexes and birth cohorts, were explored using a weighted least squares model-fitting procedure described by Neale and Cardon (1992). PRELIS 2.1 (Joreskog and Sorbom, 1993) was used to estimate 4×4 matrices of polychoric correlations among self-reported 1981 and 1989 educational achievement for each cotwin. Correlation matrices were computed separately for each of the five twin types within the two birth cohorts and are presented in Appendix A. Asymptotic variance–covariance matrices of the esti-

Table III. Polychoric Correlations Between Cotwins' Educational Achievements

	Cohort 1 (1893–1950)			Cohort 2 (1951–1965)		
	(*n* pairs)	*r* (1981)	*r* (1989)	(*n* pairs)	*r* (1981)	*r* (1989)
MZF	(520)	.70	.77	(479)	.76	.75
MZM	(216)	.70	.70	(226)	.71	.74
DZF	(299)	.51	.55	(290)	.46	.49
DZM	(94)	.42	.53	(161)	.45	.47
DZO	(270)	.33	.41	(388)	.41	.41

mated polychoric correlations within each group were also computed. The 10 matrices of polychoric correlations, along with their corresponding matrices of asymptotic variances and covariances, were then used to fit a series of biometrical models in LISREL VII (Joreskog and Sorbom, 1989).

Table III shows the polychoric correlations between cotwins (extracted from the full correlation matrices contained in Appendix A), separately by twin type, sex, birth cohort, and survey time point. It is most notable that (a) twin similarity (within sex and zygosity) appears to be quite similar in 1981 and 1989, (b) correlations for male and female same-sex pairs do not differ appreciably for either MZ or DZ twins, and (c) there are no remarkable differences in twin similarity across cohorts for any of the five twin types. Thus, in contrast to the Norwegian twin data, a cursory glance at the Australian twin correlations does not suggest either appreciable cohort or sex differences in the relative effects of heredity and environment in educational achievement.

These preliminary observations were further confirmed in model-fitting analyses. We began by fitting a full model of genetic and environmental effects, allowing estimates to vary across sex and cohort. A "psychometric factors" model (described by McArdle and Goldsmith, 1991; Neale and Cardon, 1992) was used to capture the shared covariation between the 1981 and the 1989 reports of education as well as the unique variance at each time of measurement. As shown in Fig. 4, this model specifies a common factor (P_{ED}) representing a measured phenotype of educational achievement, derived from 1981 and 1989 survey responses. Parameter b estimates the square root of reliability of

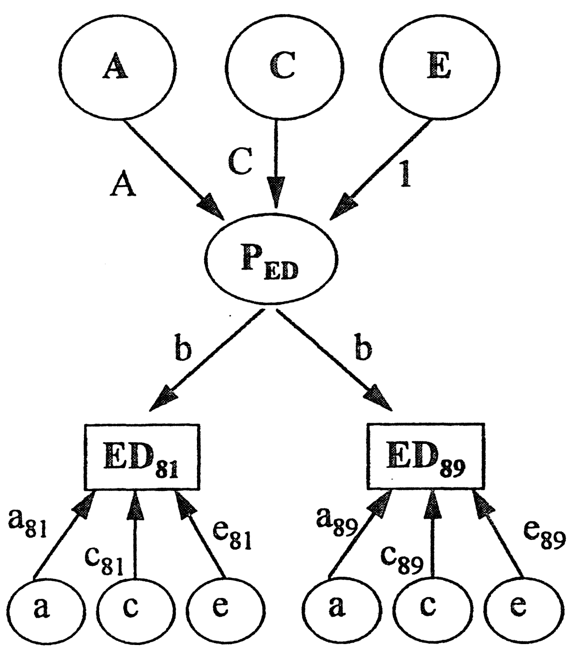

Fig. 4. Psychometric factor model of educational achievement in twins.

the measured phenotype. Variation in the latent ED phenotype is influenced by additive genetic factors (A), environmental factors common to cotwins (C), and environmental factors specific to each individual twin (E). (For purposes of model identification, the path from E to P_{ED} must be fixed at unity, although the relative contributions of this factor to P_{ED} can be examined in standardized parameter estimates in LISREL.) Residual variation in 1981 and 1989 measures of educational achievement left unexplained by the covariation between the two survey responses is also decomposed into sources of additive genetic (a), common environmental (c), and specific environmental (e) influences. To the extent that lack of covariation between 1981 and 1989 represents measurement error, rather than real change in educational achievement, most residual variation should be explained by "e" at each time point, rather than by "a" or "c" factors.

In the older cohort, where there is negligible censorship for educational achievement measures in either survey, the latent phenotype in Fig. 4 may be thought of as the "true score" of educational achievement—in other words, the phenotype without measurement error. Because our preliminary analyses suggested a high reliability (within survey time points) as well as a high stability between survey time points, any genetic or common environmental effects are expected to appear for the latent phenotype only (A and C), with residual variation at each time point being a function of only specific environmental effects (e_{81} and e_{89}). In the younger cohort, on the other hand, where considerably more censorship for educational achievement is apparent, some of the residual variation at each time point will be partly a function of real change in educational level between 1981 and 1989 and, thus, may appear in a_{81}, a_{89}, c_{81}, and c_{89}. These expectations were confirmed in the analyses described below.

A full model allowing different parameter estimates across birth cohorts and sexes fit the data from the 10 groups very well ($\chi^2 = 30.51$, df = 24, $p > .10$). Constraining estimates to be equal for females and males within Cohort 1 did not significantly affect the goodness of fit compared to the full model (χ^2 difference = 4.51, df = 9, $p > .10$). However, a similar equality constraint in Cohort 2 did significantly worsen the model fit (χ^2 difference = 17.08, df = 9, $p < .05$). Constraints of equality across birth cohorts also provided a significant worsening of fit, both for males (χ^2 difference = 45.54, df = 9, $p < .05$) and for females (χ^2 difference = 30.69, df = 9, $p < .05$).

It is apparent from the goodness-of-fit statistics that some sex differences (particularly in the younger cohort), as well as cohort differences, must exist for the psychometric model in Fig. 4. Examination of the parameter estimates, along with the fitted correlations within each group, however, revealed that the primary source of both the sex and the cohort differences lay in the (phenotypic) longitudinal correlation between 1981 and 1989 reports of educational achievement. As noted earlier, the younger cohort demonstrated less stability between the two occasions of measurement than the older cohort. Fitted correlations in the full psychometric factor model revealed a longitudinal correlation of .87 for males and .88 for females in the older cohort, which are quite comparable to the observed polychoric correlation of .88 (sexes combined) discussed above. Although the longitudinal correlation is lower for both sexes in the younger cohort, it was most attenuated for younger males ($r = .71$) compared to younger females ($r = .80$).

Subsequently, a more restrictive model was fit to the 10 groups, whereby the genetic and environmental effects in the psychometric common factor were constrained to be equal across sex and cohort. In this model the phenotypic correlation between 1981 and 1989 educational levels was allowed to vary across both sex and cohort by freeing the b parameter (see Fig. 2) across these four groups. Specific genetic and environmental effects (i.e., a, c, e for 1981 and 1989) were also allowed to vary across sex and cohort. This model also fit the data very well ($\chi^2 = 35.55$, df = 30, $p = .58$) and did not differ significantly from the full model with separate effects in A, C, and E across men, women, and birth cohorts (χ^2 difference = 5.92, df = 6, $p > .10$). Thus, genetic and environmental contributions to educational level, at least as measured in the common factor for 1981 and 1989 reports, do not appear to be affected by gender or secular changes, unlike previous findings in the Norwegian twins.

Table IV. WLS Standardized Parameter Estimates for "Psychometric Model" of Educational Achievements in Two Cohorts[a]

	Cohort 1 (1893–1950)		Cohort 2 (1951–1965)	
Parameter	Women	Men	Women	Men
b	.94	.94	.90	.84
a^2_{81}	.00	.01	.10	.14
c^2_{81}	.00	.00	.01	.00
e^2_{81}	.12	.10	.09	.14
a^2_{89}	.05	.00	.01	.18
c^2_{89}	.00	.03	.07	.01
e^2_{89}	.07	.10	.11	.10
A^2		.57		
C^2		.24		
E^2		[.19]		
Derived estimates				
h^2_{81}	.52	.53	.57	.56
c^2_{81}	.20	.20	.19	.16
h^2_{89}	.57	.51	.48	.59
c^2_{89}	.20	.23	.26	.17

[a] Parameters A^2 and C^2 constrained to be equal across sex and cohort. Parameter E^2 fixed at 1.0 in the unstandardized solution.

Parameter estimates for this more restrictive model are presented in Table IV. As shown, additive genetic effects account for 57% of the reliable variation in educational achievement, environmental effects shared by twins account for 24% of variance, and nonshared environmental effects account for the remaining 19% of variance. Stability between the two measurement occasions is quite good and comparable for women and men in Cohort 1, and lower for Cohort 2, particularly for men. As predicted, occasion-specific variance was explained almost entirely by nonshared environmental factors in the older women and men (i.e., note that parameters a^2_{81}, c^2_{81}, a^2_{89}, and c^2_{89} are all negligible and nonsignificant in Cohort 1). In the younger cohort, on the other hand, 1981-specific variance is explained by both genetic and (nonshared) environmental factors, for both women and men. Thus, some of the "measurement error" for the common psychometric factor apparently may be due to censorship at the first survey restricting the expression of genetic potential or (genetically influenced) changes which have occurred between occasions. Some sex differences do appear for 1989-specific variance, which shows effects of common environment (but not genes) for women, and vice versa for men.

Further exploration of sex and cohort differences in the relative contributions of genes and environment can be made by examining total heritabilities and environmentalities for observed measures of educational achievement in 1981 and 1989. These were derived from parameter estimates in the more restrictive model just described and are also presented in Table IV. Note that heritability estimates are quite comparable for 1981 and 1989 survey responses and that there are not appreciable differences between men and women or between young and old cohorts. Looking at parameter estimates for 1989, when censorship of educational attainment is minimized, it is clear that the relative contributions for genetic and environmental factors are quite similar for women and men in both age groups. Although the heritabilities in Cohort 2 are at least in the same order for men and women as in the Norwegian data (i.e., where men born after World War II showed a higher heritability for educational level than women), univariate biometrical analyses of 1989 data confirmed that these relative components of genetic and environmental variance do not significantly differ ($\chi^2 = .57$, df = 2, $p > .50$). Parameter estimates from the univariate analyses were as follows: $h^2 = .45$ for Cohort 1 females, .59 for Cohort 1 males, .51 for Cohort 2 females, and .67 for Cohort 2 males; and $c^2 = .32$ for Cohort 1 females, .12 for Cohort 2 males, .24 for Cohort 2 females, and .08 for Cohort 2 males. Pooling across sexes gives $h^2 = .52$ and $c^2 = .23$ in Cohort 1 and $h^2 = .59$ and $c^2 = .16$ in Cohort 2, which are very much in agreement with model-fitting results presented in Table IV.

In an effort to understand further the psychometric factor model results in Table IV, we reanalyzed the twin correlations for both 1981 and 1989 educational attainment, using the same psychometric model, but limiting the sample to only those aged 25 and older in 1981 (i.e., a noncensored sample). This sample thus included all of Cohort 1 and over one-third of Cohort 2. The most significant aspect of this analysis was that it produced the exact same estimates of relative genetic and environmental influence on the latent "education" factor as was obtained in the full analysis of the entire two cohorts presented in Table IV. In the analysis of the noncensored sample, the specific factor variance was entirely explained by "*e*" (nonshared environment) for both 1981 and 1989—a pattern of results identical to those for Cohort 1 in Table IV. This reanalysis confirmed the main finding of the present paper—that a composite measure of educational attainment, based on 1981 and 1989 reports, is explained largely by heritable factors (57%), with common environmental effects accounting for another sizable portion of the variance (24%).

It is also interesting that in Cohort 1, the distribution of Time2–Time1 differences in educational attainment are such that there are roughly equivalent proportions of individuals reporting increases (15% of females and 19% of males) as those reporting decreases (16% of females and 16% of males). Thus, most of the discrepancies reported by Cohort 1 probably stem from measurement error, rather than real change per se. In Cohort 2, however, there are considerably greater proportions of individuals who report increased educational attainment between 1981 and 1989 (24% of females and 33% of males) than who report decreases (12% of females and 10% of males). The preponderance of increases relative to decreases again supports the idea that real changes have more than likely occurred in this younger cohort. Based on the psychometric factor model results, where occasion-specific genetic variance occurs in this younger cohort, this would suggest twin concordance for such increases, particularly for MZ pairs. In Cohort 1, where occasion specific variance is completely explained by "*e*" (nonshared environment), the temporal changes in educational attainment reported are evidently random and not consistent between cotwins.

DISCUSSION

These results suggest that genetic factors play a large role in producing individual variation in educational achievements in Australia. Structural-modeling analyses estimate that well over half of the reliable variance in self-reported educational level could be explained by genetic variations. However, given that strong assortative mating has been repeatedly found for educational attainment in spouses (e.g., Watkins and Meredith, 1981; Plomin *et al.*, 1977), the model-fitting analyses (which assume random mating) have quite possibly underestimated the relative genetic variance (Eaves *et al.*, 1978). That is, if spouse correlations for educational level arise through a phenotypic assortment process, then genetic factors important to educational level become correlated across spouses, consequently leading to increased parent–offspring, sibling, and DZ twin resemblance. Correspondingly, the common environmental effects in our analyses may be biased upward as a function of phenotypic assortative mating. [See Reynolds *et al.* (1996) for a discussion of phenotypic and alternative models of assortment.]

Using a correction procedure described by Martin (1978), estimates of common environmental variation from classical twin models may be adjusted to remove the confounding effects of phenotypic assortment as follows:

$$c^2_{\mathrm{Adj}} = c^2_{\mathrm{R}} - h^2_{\mathrm{R}}A/(1 - A)$$

where h^2_{R} and c^2_{R} are relative effects of additive genetic and common environmental factors estimated from the twin model under the random mating assumption, and A is the correlation between additive genetic values of mates. One is required to estimate A as a function of the observed mate correlation (μ) and h^2_{R}:

$$A = 1/2\ [1 - \mathrm{sqrt}(1 - 4\ \mu h^2_{\mathrm{R}})]$$

An estimate of the spouse correlation for educational achievement was available in the present study, since twins were also asked to report on their spouse's educational achievement as part of the 1989 survey. This correlation (r = .426, n = 4830 couples) is comparable to estimates of spouse similarity from other studies using self-report measures for both husband and wife (e.g., Plomin *et al.*, 1977; Watkins and Meredith, 1981). Substituting the marital correlation estimated from the present study for μ, and h^2_{R} = .57 from above, A is estimated to be .50, and the entire c^2_{R} in Table V appears to be explained by phenotypic assortative mating. We conclude, then, that environmental factors shared by twins probably have little or no effect on their educational achievements. Moreover, if one then combines c^2_{R} and h^2_{R} as a more accurate estimate of relative genetic effects, the heritability of (reliably measured) educational attainment appears to be over .80. Remaining variation must be entirely explained by individual environmental factors which vary for siblings in the same family.

The finding that educational attainment in Australia is heritable is not surprising. Other studies in Western countries have confirmed this. What is new is that the relative contributions of genetic and environmental factors appear comparable for women and men in Australia. This is at odds with the one other study which has examined this issue, namely, the study of Norwegian twins (Heath *et al.*, 1985). This discrepancy in findings could stem from a number of factors, including cultural differences between Norway and Australia, sampling differences, and statistical power. It should be noted, however, that American and Danish studies of men born before 1945 have all shown considerably higher heritability estimates than the older male cohort in the Norwegian study. Although comparable results are not available for American or Danish women, these findings may suggest the presence of something specific to the Norwegian sample or culture accounting for the lower heritabilities for individuals born in the earlier part of this century.

Although it is possible that sex and cohort effects may be too small to be detected given the current sample sizes, examination of the parameter estimates within each of these groups do not show any striking pattern of differences as found in the Norwegian twin sample. Sex or cohort differences might also have been missed because of our split of cohorts at 1950 or because of limitations in the scale of measurement for educational achievement. Even larger samples would be required for a more detailed examination of cohorts spanning fewer birth years. Of greater concern are the possible biases which may have resulted from differential truncation between the sexes for educational levels in this volunteer sample of twins. As truncation (of less educated individuals) appears to have occurred in this study for men but not for women, this may have led to biased estimates of genetic and/or environmental variation in males. More complex analyses would be required to determine the degree of bias, if any, in variance components estimates in these data. We note, however, that for Australian women in this study (for whom there was little or no indication of truncation effects for educational level), heritable variation was considerably higher than that found in the Norwegian women, for both younger and older cohorts. Moreover, the considerable comparability between the present twin sample and representative (census) samples from the general Australian population, coupled with the surprising lack of evidence of truncation effects (particularly for females), gives us reasonable confidence in these findings and even allow generalizability of our findings to the Australian population, particularly women.

As a crude indicator of social equality in Australia, then, heritability estimates suggest that the success of women and men are both equally attributed to genetically based individual abilities and that these have not changed notably during this century. It remains to be seen how more recent changes in the Australian educational system, such as increased privatization and government deregulation of public schools (see Anderson, 1993), may affect the relative effects of environmental and genetic factors in producing individual differences in educational attainment. Some have even speculated that the trend toward reduced phenotypic variation in number of years of schooling will extend to a point of zero variance, such that "years of schooling" per se will no longer be a social class indicator (Western, 1993). Although not of primary concern in the birth cohorts studied in the present paper, it may become necessary in the future to distinguish more carefully among educations obtained at various types of schools, which cater, for example, to different aspects of pupils such as academic ability, religion, ethnicity, handicaps, social status, or geographical (neighborhood) location (Anderson, 1993). The extent to which parents' choice of schools based on these characteristics may covary with intellectual ability is likely to determine the degree of genetic variability in educational attainments to be maintained (see Tambs *et al.*, 1989).

APPENDIX A

Table AI. Matrices of Polychoric Correlations Among Reports of Educational Attainment (ED) for Twin-a and Twin-b in 1981 and 1989, Used in Structural Equation Modeling, Separately for Cohort 1 (Below Diagonal) and Cohort 2 (Above Diagonal) by Five Twin Types

	ED_{81a}	ED_{81b}	ED_{89a}	ED_{89b}
MZF	1.0	.81677	.76357	.69630
	.84016	1.0	.63247	.74835
	.69971	.70419	1.0	.80221
	.72286	.77016	.89628	1.0
MZM	1.0	.71691	.70871	.59535
	.85679	1.0	.53864	.74325
	.69844	.70559	1.0	.74791
	.63615	.69466	.86541	1.0
DZF	1.0	.79308	.46211	.41119
	.88970	1.0	.39296	.49396
	.50709	.49643	1.0	.78210
	.54856	.54468	.85682	1.0
DZM	1.0	.74584	.44470	.38134
	.87602	1.0	.41064	.46890
	.42295	.45434	1.0	.62499
	.50145	.52762	.90106	1.0
DZO	1.0	.79232	.40187	.31660
	.85693	1.0	.36189	.40222
	.35540	.43608	1.0	.69150
	.39585	.45368	.88970	1.0

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NOTE ADDED IN PROOF

A parallel analysis of these data from an economics perspective may be found in Miller *et al.* (1995).

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