Research Article

GENETIC AND ENVIRONMENTAL CONTRIBUTIONS TO VARIANCE IN AGE AT FIRST SEXUAL INTERCOURSE

M.P. Dunne,[1] N.G. Martin,[1] D.J. Statham,[1] W.S. Slutske,[2] S.H. Dinwiddie,[2] K.K. Bucholz,[2] P.A.F. Madden,[2] and A.C. Heath[2]

[1]*Epidemiology Unit, Queensland Institute of Medical Research, Brisbane, Queensland, Australia, and* [2]*Department of Psychiatry, Washington University School of Medicine*

Abstract—*Little is known about the relative importance of genetic and environmental factors as determinants of age at first sexual intercourse. In this study, subjects were 5,080 individuals from the Australian Twin Registry (3,310 females, 1,770 males; age range: 27–70 years, median: 40 years) who completed a semistructured interview by telephone in 1992–1993. Self-reported age at first intercourse correlated higher for identical (monozygotic) twins than for nonidentical (dizygotic) twins. Structural equation model fitting found that the genetic contribution to variance was considerably greater among twins aged 40 years or less (72% for males and 49% for females) than for those aged from 41 to 70 years (0% for males and 32% for females). Among the older cohort, there was evidence that somewhat different aspects of the shared social environment influenced age at onset in males and females. In a more laissez-faire social climate in recent decades, it is likely that biological and psychological characteristics that are partly under genetic control significantly influence the age at which a person commences sexual activity.*

The age at which people have their first sexual intercourse is probably determined by a mix of social, cultural, psychological, and biological influences. To a greater or lesser extent, factors such as social class, educational background, religiosity, personality, self-image, peer influences, ethnicity, gender, physical maturation, and pubertal hormone levels correlate with age at first sexual interest and activity (Bingham, Miller, & Adams, 1990; Dunne, Edwards, Lucke, Donald, & Raphael, 1994; Johnson, Wadsworth, Wellings, & Field, 1994; Seto, Lalumiere, & Quinsey, 1995; Udry, 1988). However, little is known about the relative importance of these factors as determinants of the age of onset of sexual behavior.

It is possible that the salience of particular factors varies at different points in time. As the age of onset of sexual intercourse in Western societies has steadily declined over the past 50 years (ACSF Investigators, 1992; Johnson et al., 1994), the relative importance of environmental and biological (including genetic) factors may have changed.

One example of an interaction between social and biological determinants of early sexual intercourse appears in work by Udry, Halpern, and their co-workers (Halpern, Udry, Campbell, & Suchindran, 1993; Udry, 1988). Although an early rise of pubertal androgen levels, particularly testosterone, correlates with early onset of sexual ideation and masturbation in males (Udry, Billy, Morris, Grof, & Raj, 1985), the hormone surge might not lead directly to sexual intercourse. Pubertal physical development may act as a social signal to potential partners, and whether this signal is acted upon may largely be determined by social factors that regulate partners' willingness and availability (Halpern et al., 1993). It may be that as social controls on adolescent and premarital sexual activity have diminished in recent decades (Johnson et al., 1994; Sundet, Magnus, Kvalem, Samuelsen, & Bakketeig, 1992), genetic factors that underlie the rate of physical maturation (Olsen, 1992; Treloar & Martin, 1990) explain more of the variance in age at first sexual involvement.

Research with identical and nonidentical twins provides an opportunity to test the idea that temporal changes in environmental context moderate the expression of genetic predispositions. However, few twin studies have examined the age of onset of sexual behaviors. In the only directly relevant data published to date, Martin, Eaves, and Eysenck (1977) reported that genetic determination of early sexual intercourse was possibly greater among older compared with younger twins. However, with only 246 twin pairs in the study, there was insufficient power to examine generational differences adequately.

The aim of this study was to examine the relative contribution of genetic and environmental sources of variance in age at first intercourse. The data were derived from a large telephone survey of Australian twins aged between 27 and 70 years. Many opposite-sex twin pairs were included. We examined the goodness of fit of genetic structural equation models with a focus on heterogeneity between sexes and two age cohorts. This approach enabled study of sex differences in heritability and environmental covariation among men and women who became sexually active at different times over the past 50 years.

METHOD

Sample

Subjects were twins from the Australian Twin Registry, a volunteer register that began recruitment in the late 1970s. The twins were members of an original "1981 cohort" and had participated in surveys that concentrated on alcohol dependence and general health (Heath, Bucholz, et al., 1994). Three major alcohol-related surveys have been conducted to date: a mailed questionnaire in 1980–1981 (n = 8,183 individual returns, response rate = 67%; Martin & Jardine, 1986); a 1988–1989 mailed questionnaire to intact pairs in the original cohort (n = 6,329 individual returns, response rate = 83%; Heath, Cloninger, & Martin, 1994); and a telephone interview in 1992–1993 (n = 5,995 individuals, response rate = 89%). A follow-up telephone survey of 590 individuals who participated in the 1992–1993 survey was completed in 1993–1994, which enabled the assessment of item test-retest reliability. All subjects in the follow-up study were members of twin pairs in which at least one twin reported a positive history of alcohol dependence.

Data on age of onset of sexual activity were obtained in the 1992–

Address correspondence to Michael P. Dunne, Epidemiology Unit, Queensland Institute of Medical Research, Post Office, Royal Brisbane Hospital, Herston, Queensland, 4029 Australia; e-mail: michaeld@qimr.edu.au.

1993 telephone interviews and the follow-up interviews. There was a strong positive skew in the age distribution of the sample (range: 27–90 years, median = 41 years). In order to examine age-cohort differences, we split the sample into two groups on the basis of median present age: a younger cohort aged from 27 through 40 years (n = 2,918, mean age = 34.3 years, sd = 3.34, born from 1952 through 1965) and an older group aged from 41 through 70 years (n = 2,824, mean age = 51.3 years, sd = 8.37, born from 1922 through 1952). A minority of subjects aged from 71 to 90 years (n = 253, 4.2%) was excluded to reduce generational heterogeneity and potentially less reliable reporting by the older people (Dunne et al., 1997), while losing a minimal number of subjects. Even with this adjustment, use of the median cut point produced cohorts in which the younger group had a span of 14 years, but the older group spanned 30 years.

Reported age at first intercourse was analyzed for 2,540 intact twin pairs, comprising 3,310 females and 1,770 males. Twins aged less than 71 years who refused to answer the question about sexual intercourse (n = 101, 1.8%) were excluded, and individuals whose co-twins did not complete the interview (n = 561) were excluded from the genetic analysis, although responses from singles were used to check for sampling bias.

Measure

Age at sexual initiation was obtained in the 1992–1993 telephone interview with the question: "How old were you when you first had sexual intercourse?" The same question was asked in the follow-up telephone survey.

Data Analysis

Twin data analysis rests on several basic assumptions. For a given characteristic, if there is no correlation between identical (monozygotic, MZ) twins or between nonidentical (dizygotic, DZ) twins, then it is assumed that individual specific (i.e., random) environmental factors influence the trait. If the correlations between MZ twins and between DZ twins are equal, but greater than zero (i.e., $r_{MZ} = r_{DZ} > 0$), then there is evidence of resemblance due to shared environmental factors. Further, if MZ correlations are greater than DZ correlations (i.e., $r_{MZ} > r_{DZ} > 0$), it is assumed that genetic factors are involved in addition to influences of shared family environment. Error of measurement is subsumed within the nonshared environment component (Neale & Cardon, 1992). The variance components are estimated by minimizing the discrepancy between observed and expected covariances in a measure (such as age at first intercourse) among family members who differ in genetic relatedness (Waller & Shaver, 1994).

One of the benefits of this large twin panel sample is that, in addition to same-sex MZ and DZ pairs, there are many opposite-sex DZ pairs. These opposite-sex pairs enable the examination of sex limitation in genetic and environmental influences. For example, if correlations between opposite-sex DZ pairs are lower than those for same-sex DZ pairs, it is possible to conclude that either genetic or environmental influences on the trait in males and females are different. Sex limitation reflects sex differences in biological, psychological, or social sources of variation, a plausible hypothesis for many phenotypes.

The analysis was completed in several steps. First, sampling biases, the distribution of age at first intercourse, and the test-retest reliability of the measure were described. For genetic analyses, age at initiation was converted into a 10-point ordinal scale, with each category containing approximately one decile. This strategy had the effect of minimizing skew while overcoming a data-censoring problem by including virgins in the highest age-of-onset category. PRELIS 2.12 (Jöreskog & Sörbom, 1994) was used to compute twin pair polychoric correlations. Correlation and asymptotic covariance matrices from PRELIS 2.12 were then imported into Mx (Neale, 1994) to test the goodness of fit of genetic structural equation models separately for younger and older age cohorts (see Neale & Cardon, 1992, for further details about fitting genetic and environmental models to twin data).

Genetic Model

A structural equation model for the analysis of twin data that included allowance for sex limitation (Neale, 1994) was used. This model allows specification of separate parameter estimates for males and females. Differences between men and women in the influence of genetic or shared environmental factors were tested by freeing either the genetic or environmental components of the opposite-sex DZ twin correlations and examining the relative fit of the constrained and free models.

Regression of current age on age at first intercourse was also specified to estimate the independent contribution of age to total variation. If this is not specified, estimates of the influence of common environment will be inflated (Neale & Cardon, 1992).

RESULTS

Analysis of Possible Sampling Biases

Sampling bias would be indicated if there were a significant difference in age at first intercourse between twins from intact pairs and singletons whose co-twins did not participate (Neale & Cardon, 1992). In this study, the proportion of virgins among people in intact pairs (2.1%; 107/5,080) was not significantly different from the proportion in incomplete pairs (2.3%; 13/561). Among the full sample of nonvirgins, the mean ages at first intercourse in intact pairs (20.05 years, sd = 3.98) and singletons (20.19 years, sd = 4.23) were not significantly different, $t(654.4) = 0.53$, $p = .596$. The variance tended to be greater for singletons than for twins in intact pairs, $F(547, 4959) = 1.17$, $p = .011$, although this effect size was small and unlikely to introduce any major biases into inferences drawn from this study.

Enrollment bias exists if there are significant differences between zygosity groups in the means and variances of the trait being studied (Heath, Neale, Hewitt, Eaves, & Fulker, 1989). One-way analysis of variance using log-transformed age at first intercourse revealed no significant differences in means between females in MZ, same-sex DZ, and opposite-sex DZ pairs, $F(2, 3222) = 1.31, p = .271$. Among males, there was a small but significant association with zygosity group, $F(2, 1732) = 3.12$, $p = .044$. The means for males in MZ, same-sex DZ, and opposite-sex DZ pairs were 20.3, 19.8, and 19.7 years. The differences between means were no greater than 0.14 of a standard deviation, so it is unlikely that this small sampling bias would invalidate analyses of twin pair covariance in age at first sexual intercourse.

Distribution of Age at First Sexual Intercourse

The predictable trend toward earlier sexual intercourse among younger people was present in our sample, with a significant positive correlation between present age and log-transformed age at first intercourse, $r = .33$, $p < .0001$. Excluding virgins, mean ages of first intercourse for the younger twin cohort (aged 27–40 years) and older twin cohort (aged 41–70 years) were 18.94 years ($sd = 3.24$) and 21.10 years ($sd = 4.33$), respectively; these means were significantly different, $t(4695.7) = 20.12$, $p = .0001$. There are no comparable data on age at first intercourse among a random sample of adults in the general Australian population. A recent national study of adolescents (Dunne, Donald, et al., 1994) indicated that in the early 1990s, the median age at first intercourse reached 17 years. In the twin sample, there was a small main effect of gender, with mean age of first intercourse being slightly lower for males (19.9 years, range: 9–49) than for females (20.1 years, range: 6–52), which was statistically significant with log-transformed data, $t(2942.4) = 2.20$, $p = .03$. The interaction between gender and age cohort was not significant.

Reliability of Self-Reported Age at First Sexual Intercourse

A total of 570 subjects reported age at first intercourse in both the initial and the follow-up telephone interviews. The test-retest Pearson correlation was .88 ($p < .0001$), indicating high reliability in self-reported age over an average follow-up period of 15 months (range: 2.3–24.7 months). Older subjects and those with a history of sexual abuse were slightly less likely than other subjects to give consistent answers in the two interviews (Dunne et al., 1997).

Twin Pair Polychoric Correlations

Table 1 shows the twin pair correlations for the younger and older age cohorts. Several trends are apparent. Among subjects ages 40 or less, the MZ correlations were substantially higher than the DZ correlations. In contrast. $r_{MZ} - r_{DZ}$ differences in the older cohort were much smaller, with a very small difference in the size of correlations among MZ and DZ males. It is also clear that same-sex DZ correlations were higher than opposite-sex DZ correlations, particularly in the older cohort.

Genetic and Environmental Contributions to Variance

Younger cohort

Table 2 summarizes the hypotheses that were tested and the goodness-of-fit statistics for each structural equation model. The upper panel shows nine different models that examined the fit of male and female twin data. The lower panel specifies the critical hypotheses and the outcomes of likelihood ratio tests (i.e., the significance of the simple difference in goodness-of-fit statistics, χ^2 and *df*, of the models in each comparison). The logic of the hypothesis-testing process is to test the relative fit of two models and if the likelihood ratio χ^2 is not significant, accept the model with the fewer parameter estimates (i.e., the higher number of degrees of freedom). If the likelihood ratio is significant, the model with the better fit (lower χ^2 value) is accepted.

The fit of the model for the younger cohort of twins was improved when age regression was specified. Model 3, which set the age regression to be equal in males and females, was preferred over Model 2 because it included one less parameter and the likelihood ratio test was not significant (see the tests of critical hypotheses in Table 2).

The hypothesis of no shared environmental effects among the younger twins could not be rejected, although there is a trend for the comparison between Models 3 and 4 to be significant ($p = .059$). Subsequently, we found that there were no shared environmental effects for men, although shared environment remained important for women. It was very clear that genetic effects could not be dropped from the model. Testing of sex limitation models indicated that similar genetic factors influence men and women, although the magnitude of the effect was greater in men than in women.

Proportions of variance attributable to genetic factors, shared environment, nonshared environment, and age were obtained by squaring the standardized parameter estimates from Model 6. That model was chosen because it had the lowest value for Akaike's information criterion (AIC; Akaike, 1987) and the fit was as good as any other. Genetic factors accounted for 49% of the variance among females and 72% among males. For both males and females, the regression of age on the dependent variable accounted for less than 1% of the variance (0.49%). The remainder of the variance in young females' age at first intercourse was attributed to shared environment (24%) and nonshared environment or measurement error (26%). Among young males, the remainder of the nongenetic variance (27%) was attributed to nonshared environment plus error.

Table 1. *Twin pair polychoric correlations for age at first sexual intercourse*

	Type of twin pair				
Statistic	MZ_{Female}	DZ_{Female}	MZ_{Male}	DZ_{Male}	$DZ_{Opposite\ sex}$
			Younger cohort (27–40 years)		
n	411	240	185	141	320
r	.74	.49	.73	.35	.29
se	.01	.01	.01	.04	.02
			Older cohort (41–70 years)		
n	456	264	197	78	248
r	.65	.50	.50	.53	.24
se	.01	.03	.02	.05	.03

Note. MZ = monozygotic; DZ = dizygotic.

Table 2. *Results for the younger cohort (aged 27–40 years)*

A. Genetic model fitting results

	Model				Goodness of fit			
Number	Genetic effects	Shared environment	Nonshared environment	Age regression	*df*	χ^2	*p*	AIC[a]
1.	M ≠ F	M ≠ F	M ≠ F	No	11	16.51	.123	−5.5
2.	M ≠ F	M ≠ F	M ≠ F	M ≠ F	9	7.91	.544	−10.1
3.	M ≠ F	M ≠ F	M ≠ F	M = F	10	9.30	.504	−10.7
4.	M ≠ F	M = F = 0	M ≠ F	M = F	12	14.95	.244	−9.2
5.	M ≠ F	M ≠ F, F = 0	M ≠ F	M = F	11	14.34	.215	−7.7
6.	M ≠ F	M ≠ F, M = 0	M ≠ F	M = F	11	9.30	.594	−12.7
7.	M = F = 0	M ≠ F, M = 0	M ≠ F	M = F	13	404.60	.000	378.6
8.	M ≠ F[b]	M ≠ F, M = 0	M ≠ F	M = F	10	9.28	.506	−10.7
9.	M = F	M ≠ F, M = 0	M ≠ F	M = F	12	13.22	.354	−10.8

B. Likelihood ratio tests of critical hypotheses

Hypothesis	Model comparison	Likelihood ratio test *df*	χ^2	*p*	Outcome
No age regression	2 vs. 1	2	8.60	.014	Reject
No gender difference in age regression	3 vs. 2	1	1.39	.238	Accept
No shared environmental effects for men or women	4 vs. 3	2	5.65	.059	Accept
No shared environmental effects in women	5 vs. 3	1	5.04	.025	Reject
No shared environmental effects in men	6 vs. 3	1	0	1.000	Accept
No genetic effects in men and women	7 vs. 6	2	395.3	.000	Reject
Same genetic factors influence men and women	8 vs. 6	1	.02	.888	Accept
No gender difference in magnitude of genetic effects	9 vs. 6	1	3.92	.048	Reject

Note. M = males; F = females. M ≠ F means that the model allows the genetic (or shared environmental) effects to differ for males and females.
[a]AIC is Akaike's information criterion, which is a measure of goodness of fit derived by subtracting 2**df* from χ^2 (Akaike, 1987).
[b]The model allowed for sex-specific genetic effects not shared by siblings of unlike sex.

Older cohort

Results of model fitting with data from the older cohort are shown in Table 3. Again, best fit was achieved when age regression was specified in the model, and Model 3 was preferred because it was more parsimonious than Model 2.

Among the older cohort, it was clear that shared environment could not be dropped from the model. There was evidence that somewhat different factors in the shared environment influenced men and women and that the shared environmental effects were significantly greater in men than in women. It was not possible to drop genetic effects from the general model for older twins, although a model that specified no genetic effects for men fitted best to the data.

The estimates of genetic and environmental influences were derived from Model 9. Genetic factors accounted for 32% of the variance among women and 0% among men. In this cohort, a relatively greater proportion of the variance among males (42%) than females (25%) was attributed to shared environmental factors, and nonshared environment accounted for 48% and 34% of the variance among males and females, respectively. It is notable that in the older cohort, age independently accounted for approximately 9% of the variance in age at first intercourse, compared with 0.49% among younger twins. This difference is probably due to the wider age span among the older cohort (30 years) than the younger cohort (14 years).

DISCUSSION

We believe these are the first data on age at first sexual intercourse to be reported for a large sample of male and female twins. Our results are clearly consistent with the hypothesis that genetic factors contribute significantly to variation in age at first sexual intercourse. The data are also novel in showing significant generational change both in the overall importance of genetic and environmental sources of variance in the onset of intercourse and in the relative influence of these factors in males and females.

The younger twins were born from 1952 through 1965 and, on average, would have commenced sexual activity in the early 1970s to mid 1980s. Covariation among the younger twins in age at onset was strongly associated with genetic similarity, with the magnitude of genetic effects being significantly greater among males (72%) than females (49%). Among the young females, shared environment (i.e., common family and cultural influences) exerted roughly the same influence as nonshared environment and measurement error (approximately 25%), whereas all of the nongenetic variance among young males (27%) was attributed to nonshared environment and measurement error.

The great majority of the older twins would have commenced sexual activity prior to the 1970s, and it is notable that the pattern of genetic and environmental variation for this cohort is quite different

Table 3. *Results for the older cohort (aged 41–70 years)*

A. Genetic model fitting results

	Model				Goodness of fit			
Number	Genetic effects	Shared environment	Nonshared environment	Age regression	*df*	χ^2	*p*	AIC[a]
1.	M ≠ F	M ≠ F	M ≠ F	No	11	183.69	.000	161.7
2.	M ≠ F	M ≠ F	M ≠ F	M ≠ F	9	14.60	.103	–3.4
3.	M ≠ F	M ≠ F	M ≠ F	M = F	10	14.84	.138	–5.2
4.	M ≠ F	M = F = 0	M ≠ F	M = F	12	23.14	.027	–0.9
5.	M ≠ F	M = F	M ≠ F	M = F	11	21.41	.029	–0.6
6.	M ≠ F	M ≠ F[b]	M ≠ F	M = F	9	10.35	.323	–7.7
7.	M = F = 0	M ≠ F[b]	M ≠ F	M = F	11	16.88	.111	–5.1
8.	M ≠ F, F = 0	M ≠ F[b]	M ≠ F	M = F	10	16.88	.077	–3.1
9.	M ≠ F, M = 0	M ≠ F[b]	M ≠ F	M = F	10	10.35	.411	–9.7

B. Likelihood ratio tests of critical hypotheses

		Likelihood ratio test			
Hypothesis	Model comparison	*df*	χ^2	*p*	Outcome
No age regression	2 vs. 1	2	169.09	.000	Reject
No gender difference in age regression	3 vs. 2	1	0.24	.624	Accept
No shared environmental effects for men or women	4 vs. 3	2	8.30	.016	Reject
No gender differences in magnitude of shared environmental effects	5 vs. 3	1	6.57	.010	Reject
Same shared environment factors influence men and women	6 vs. 3	1	4.49	.034	Reject
No genetic effects in men and women	7 vs. 6	2	6.53	.038	Reject
No genetic effects in women	8 vs. 6	1	6.53	.011	Reject
No genetic effects in men	9 vs. 6	1	0	1.000	Accept

Note. M = males; F = females. M ≠ F means that the model allows the genetic (or shared environmental) effects to differ for males and females.
[a]AIC is Akaike's information criterion, which is a measure of goodness of fit derived by subtracting 2**df* from χ^2 (Akaike, 1987).
[b]The model allowed for sex-specific shared environmental effects not shared by siblings of unlike sex.

from that revealed among the younger cohort. Among older males, both nonshared environment and shared environment accounted for large proportions of variance, but there was no evidence of heritability. Among older females, in contrast, the relative importance of shared environment (25% of variance) was similar to that found among younger women (24%). Variance attributed to nonshared environment was somewhat greater for the older (34%) than the younger (26%) females, but genetic variance was lower for older women (32%) than for younger women. Model fitting with the older cohort also indicated a sex difference in the type of within-family and cultural influences on early sexual intercourse.

Of course, the observation that genetic contribution to variance depends on the environmental contribution is not new. Similar age-cohort differences in heritability estimates have been noted for characteristics as diverse as educational attainment (Heath et al., 1985), alcohol consumption patterns (Kaprio, Rose, Romanov, & Koskenvuo, 1991), and risk of tonsillectomy (Martin, Kehren, Battistutta, & Mathews, 1991). The greater heritability of early sexual intercourse among younger women and men may be related to a general phenomenon of personality; that is, characteristics such as sociability and impulsivity that correlate with early sexual activity (Rawlings, Boldero, & Wiseman, 1995; Seto et al., 1995) may be more heritable among adolescents than adults (Dworkin, Burke, Maher, & Gottesman, 1976; McCartney, Harris, & Bernieri, 1990). In this study, significantly more of the younger than the older twins recalled that first intercourse occurred during adolescence. Heritability may be higher during periods of physical and psychological development than in adulthood (Dworkin et al., 1976). Perhaps the strongest influence, though, is that first intercourse for younger twins occurred during a time when there were fewer social controls on adolescent sexual behavior (Johnson et al., 1994; Sundet et al., 1992), and heritability may be increased when environments change from being suppressive and become more expressive (Dworkin et al., 1976).

This effect may be most noticeable among males. The large decline across cohorts in shared environmental effects on early sexual intercourse for men but not women may have arisen because the social and familial variables that influence onset of sexual behavior among men have changed more rapidly than those shared environmental factors that influence the early sexual behavior of women. One caveat is that the differences between age cohorts in the parameter estimates for men appear surprisingly large in this study. It is possible that these estimates are imprecise because of small numbers of subjects. For example, there were less than 200 pairs in each of the four same-sex male groups (see Table 1), including only 78 pairs of same-sex DZ males in the older cohort. It will be interesting to see in future twin studies whether a similar pattern is observed for the onset of sexual behavior among men and women.

Twin research with multivariate theoretical models should attempt to identify which specific psychological and physical predispositions can explain the patterns in twin similarity. Apart from heritable aspects of pubertal hormonal activity, physical development, and personality, genetic similarity in physical attractiveness and other morphological features may be important, given some evidence that adolescents and young adults who have been ''objectively'' rated as physically attractive engage in more sexual activity than those who have not (Curran, Neff, & Lippold, 1973). It is also likely that biological and psychological factors interact, with behavior in adolescence (including sexual interest and activity) being moderated by genetic covariation in hormones, temperament, and personality (Buchanan, Eccles, & Becker, 1992).

There are many plausible lines of further inquiry into the sociobiology of early sexual involvement. Comprehensive studies must go beyond an exclusive focus on social and demographic variables because in total these may account for a relatively small proportion of the variance and be declining in significance.

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