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Adolescent psychological problems, partnership transitions and adult mental health: an investigation of selection effects

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ABSTRACT

Background. Marital status is a strong correlate of psychiatric morbidity in adulthood, but debate continues on how far this association reflects causal influences or selection effects based on prior psychological characteristics.

Method. Prospective data from the National Child Development Study were used to examine effects of adolescent emotional and behavioural problems on transitions into and out of first partnerships, and their implications for psychiatric morbidity at age 33.

Results. Emotional and behavioural problems in adolescence showed systematic links with early partnership transitions (age at partnership formation, type of first partnership, and risks of first partnership breakdown). More detailed tests suggested that these effects only accounted for a modest proportion of the associations between partnership status and psychiatric morbidity at age 33.

Conclusions. In a non-referred community sample selection effects associated with adolescent emotional and behavioural problems appear to play only a modest role in links between partnership status and adult mental health.

INTRODUCTION

Marital status is among the best-established correlates of psychiatric morbidity in adult life (Bebbington, 1988). Marital breakdown is associated with increased risks of psychological distress (Bruce & Kim, 1992; Aseltine & Kessler, 1993), while marriage, especially perhaps for men, is associated with more positive psychological well-being. Findings for the never married vary, but often fall between the extremes of the stably married and those whose relationships have broken down (see e.g. OPCS, 1995).

Both social causation and social selection models have been advanced to account for these effects (Pearlin & Johnson, 1977). From a social causation perspective, variations in psychological well-being are assumed to represent outcomes of differing partnership statuses, with their varying satisfactions, supports and stressors. From a selection perspective, the prior characteristics of individuals are assumed to play at least some part in influencing the marital status categories that they occupy. Although the possibility of selection effects has been acknowledged for many years, direct tests of marital selection on the basis of psychological characteristics have remained quite limited. Early cross-sectional studies (see e.g. Gove *et al.* 1983) concluded that some selection effects might be involved, but that they were of limited importance at most. Longitudinal designs, which track individuals across changes in marital status, provide more direct tests. With appropriate longitudinal data two related issues can be examined: first, how far prior characteristics influence marital transitions, and second, how far associations between marital status and subsequent mental health are attributable to, or

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at least attenuated by, posited selection factors. Rather in the manner of tests for mediator variables (Baron & Kenny, 1986), a full test of selection effects depends not simply on demonstrating effects on marital transitions, but also on evidence that bivariate associations between marital status and psychiatric morbidity are reduced by selection factors.

To date, most studies have focused on the first of these steps. Because selection accounts assume that more psychologically healthy individuals form partnerships earlier than their peers, the timing of marital transitions has attracted particular attention here. Several studies have now identified effects of prior disorder or psychological distress on marital timing. In the National Comorbidity Survey (Forthofer *et al.* 1996), for example, retrospective reports of DSM-III-R depression, conduct disorder, drug abuse and alcohol problems were all associated with a reduced likelihood of 'on-time' marriage (between ages 19 and 24), and social phobias and dysthymia with lower rates of marriage from age 25 onwards. Using a prospective design, Horowitz and colleagues found that problem drinkers were less likely to marry in the late teens and early twenties (Horowitz & White, 1991), and that for women, prior depression was also associated with a reduced likelihood of marriage in the late twenties and early thirties (Horowitz *et al.* 1996). Mastekaasa (1992) found strong effects of prior psychological well-being on transitions into marriage, most marked for women in the 20–25 age-range, and for men aged 26–39. Hope *et al.* (1999), using the data-set we report on here, examined the impact of psychological distress at age 23 on marital transitions in the next 10 years. Those who remained single throughout this period had significantly higher initial distress scores; although selection effects were evident in both sexes, they were especially marked for men.

An important qualification to this picture relates to very early partnership formation. Here, pre-existing adolescent psychiatric disorders – especially antisocial behaviours and conduct problems – have been associated not with a reduced likelihood of partnership formation, but instead with very early moves into relationships (Bardone *et al.* 1996; Forthofer *et al.* 1996; Stattin & Magnusson, 1996). Prior problem behaviours also seem elevated among those who cohabit. Yamaguchi & Kandel (1985), for example, found links between cohabitation and prior illicit drug use, while Booth & Johnson (1988) found associations with a wider range of deviant characteristics including personality problems, alcohol or drug use, irresponsibility with money and trouble with the police. Early problem behaviour has also been associated with increased risks of marital disruption (Earls, 1994). Retrospective reports from the Epidemiological Catchment Area Studies (Robins *et al.* 1991) identified strong links between childhood conduct problems and later marital breakdown in both men and women. Kiernan & Mueller (1998), using the data-set we report on here, found that emotional/behavioural problems at age 16 were important predictors of subsequent divorce; and dysphoric mood in adolescence has been associated with problems in establishing close interpersonal relationships in early adulthood, and with increased risks of divorce among women (Kandel & Davies, 1986).

Although complex, these findings suggest that prior psychological characteristics are indeed consequential for transitions into and out of intimate relationships. To assess their implications for later mental health, two further steps are needed. First, it is necessary to control for effects of other childhood factors known to be associated both with patterns of partnership formation and with later mental health. Social adversity in childhood, educational attainments, and exposure to parental divorce might all be important confounds here (see e.g. Rodgers & Pryor, 1998). Secondly, more explicit tests of selection effects are needed. Few studies have attempted these to date. In a demographic analysis, however, Lillard *et al.* (1995) examined the role of selection processes on the well-established link between pre-marital cohabitation and partnership breakdown (Bennett *et al.* 1988; DeMaris & Rao, 1992). Uncorrected for selection factors, pre-marital cohabitation had strong implications for marital dissolution. Corrected for selectivity – modelled in this instance as unobserved heterogeneity – these effects were completely eliminated. Focusing on mental health outcomes, Horowitz *et al.* (1996) reported much more modest effects: associations between marital status and depression were attenuated when pre-marital measures of depression were taken into account, but remained significant. So far as we are aware, few other tests of this kind have been reported in the psychiatric literature to date.

The present study was designed to provide further evidence on marital selection, using data from the National Child Development Study (NCDS). NCDS has traced a national sample of individuals from birth to age 33. It includes measures of emotional and behavioural problems in adolescence, before the great majority of cohorts members began their first co-residential partnership; histories of partnership formation and dissolution from the teens to the early thirties; and self-reports of depressed mood at age 33. (We use the term 'partnership' as a generic label to refer to all types of co-residential partnership: legal marriage; marriage preceded by cohabitation; and cohabitation alone.) Overall, these measures provide for tests of a wide range of potential selection influences. We focused on two central issues: (*i*) how far adolescent emotional and behavioural problems influenced transitions into and out of first partnerships, controlling for effects of potential confounds; and (*ii*) how far these influences attenuated associations between marital status and depressed mood at age 33.

METHOD

Sample

The sample was drawn from the National Child Development Study (NCDS), a prospective study of all children in Britain born in one week in March 1958. NCDS began as a perinatal mortality survey (Butler & Bonham, 1963) and has subsequently involved follow-ups at ages 7, 11, 16, 23 and 33 years (Fogelman, 1983; Ferri, 1993).

Measures

Sociodemographic data

Measures of childhood social class (categorized as non-manual and manual, including no male head) and region of residence were derived from parental reports at age 11.

Adolescent emotional and behaviour problems

Teacher behaviour ratings (Rutter, 1967; Elander & Rutter, 1996) were completed when cohort members were aged 16. Factor analyses identified two main factors: emotional difficulties (worries, solitary, miserable, fearful, fussy) and antisocial behaviours (destructive, fights, not much liked by other children, irritable, disobedient, lies, steals, resentful/aggressive, bullies). Scales summing these items were dichotomized as closely as possible to the top 20th percentile (emotional factor, 17·7%; antisocial factor, 18·3%) to identify adolescents with definite, though not necessarily clinically significant, difficulties.

Education

Childhood ability and behaviour was assessed at age 11. Children completed a standardized general ability test (Douglas, 1964) and teachers rated their behaviour on the Bristol Social Adjustment Guides (BSAG, Stott, 1966).

Age at completion of full-time education was obtained from cohort member interviews at age 23 and categorized as: (*a*) up to age 16; (*b*) age 16 to 18; (*c*) over 18.

Parental divorce

Retrospective reports of parental divorce were collected from interviews at age 33. Parental divorce was coded as a dichotomy, scored positively for all subjects (8·3%) reporting that their parents had divorced before they were aged 16.

Partnership variables

A partnership history was obtained from all cohort members at age 33. Partnerships were defined as any co-residential relationships, including both legal marriage and non-marital cohabitation, lasting for 1 month or more. The histories provided details of age at first partnership, type of first partnership (marriage, cohabitation leading to marriage and cohabitation only), and a summary measure of partnership history/current status, coded as: (*i*) never married/cohabited; (*ii*) one partner – still together; (*iii*) partnership breakdown(s), cohort member currently alone; (*iv*) partnership breakdown(s), cohort member currently in second or subsequent partnership.

Parenthood

Parenting histories were also collected at age 33. A simple dichotomous measure was created reflecting whether cohort members had children within a given partnership.

Adult depressed mood: the Malaise Inventory

A 24-item self-completion questionnaire (Rutter *et al*. 1970), was used to assess depressed mood at ages 23 and 33. Mean total Malaise scores provide an index of low mood, while scores of $\geqslant 6$ show good sensitivity and specificity by comparison with interview measures of clinically significant depression (Rodgers *et al*. 1999). In the present sample 20·7% of women and 7·9% of men scored above this cut-point at age 23, and 15·2% and 8·2% respectively at age 33. Individual differences in self-rated mood showed considerable stability, with total score correlations of 0·51 for men and 0·56 for women between ages 23 and 33. Men and women who reported consulting a GP or specialist for emotional problems over the previous 10 years had significantly elevated Malaise scores at age 33 (men, 36·4% *v*. 10·3%; women, 61·8% *v*. 21·6%), suggesting that many of the difficulties tapped by high Malaise scores were relatively chronic.

Exclusions

Several exclusions were made to ensure that the findings were not unduly influenced by extreme groups that may experience particular problems in relationship formation. These were: (*a*) children scoring in the bottom 1·0% on the general ability test at age 11 (likely to have shown learning disabilities); (*b*) subjects assessed as showing severe physical or mental health problems on medical assessments at age 16 (0·5%); (*c*) subjects who reported 'hearing or seeing things' in adulthood, and so possibly suffering from psychotic illnesses (1·0%); and (*d*) any additional subjects registered as disabled at age 33 (1·0%). Two partnership categories were too small for adequate analysis, and were also excluded, namely, widows/widowers (0·6%) and those reporting same sex partnerships (0·7%).

Attrition

Preliminary analyses of response bias at the age 33 contact in NCDS (Shepherd, 1993) indicated that sample loss was limited, and primarily located in groups with low socioeconomic status and poor school achievements in childhood, and those from ethnic minority backgrounds. The interviewed sample at age 33 was broadly comparable in basic demographic and economic characteristics with other national surveys carried out at the time.

Our study drew on measures from four study waves, and a range of informants. This inevitably compounded problems of subject attrition with those of missing data at individual waves. Complete data on all the measures required for the analyses were available on 2780 men and 2993 women, 67·3% of the age 33 sample with valid adolescent behaviour measures. The final sample slightly under-represented cohort members whose first relationship involved cohabitation only, and over-represented those who married directly, compared with all respondents at age 33. Losses among those whose first relationships had broken down, and particularly among those remaining alone at age 33, were more marked. Adolescent antisocial behaviour, and to a lesser extent emotional problems, were also associated with subsequent attrition.

We used weights to take account of potential biases due to selective attrition. Non-response weights were derived in two stages, the first adjusting for losses between ages 11 and 33, the second for loss of cases with partnership history data at age 33, but where other covariates were missing. Both adjustments were derived using logistic regression models to predict response, with the weights being the inverse of the probability of response obtained from the predicted values of the models. The first stage adjustment identified significant effects of sex, region, childhood social class, general ability and total behaviour problems at age 11 on non-response at age 33. The second stage adjustment included effects of region and ability, along with effects associated with type of first partnership, first partnership breakdown and having had a second partnership. The final weight was derived as the product of these two weights, rescaled to the achieved sample size, and was used in all analyses.

Analysis strategy

The analyses fell into three main parts. First, we examined the general pattern of partnership formation in the NCDS cohort, and associations between partnership status and depressed mood at age 33. Next we explored three routes for effects of adolescent emotional and behavioural problems on partnership transitions: (*i*) on the timing of first partnerships; (*ii*) on the types of first partnerships that cohort members entered; and (*iii*) on the risks that those partnerships would break down before age 33. Thirdly, bringing these various findings together, we identified subgroups where selection effects on partnership transitions might have influenced later mental health, and tested these effects directly, using both continuous Malaise scores and categorical high/low cut-offs as dependent measures.

All models were estimated using STATA for Windows Version 5.0 (StataCorp, 1997), with robust variance estimates to give correct inferences for weighted models. Logistic regression models were used in tests of categorical variables, and poisson regression models with over-dispersion for the continuous Malaise scores. Cox proportional hazards models were used to examine effects on rates of partnership transitions, and predictors of the three-category measure of partnership type (marriage, cohabitation leading to marriage, and cohabitation only) were tested in multinomial logit models, with marriage as the baseline category. Model results are given in terms of odds ratios (OR), rate ratios (RR) or hazard ratios as appropriate, with 95% confidence intervals. The RR estimates (poisson regression models) estimate ratios of means when used with categorical predictors, and can be interpreted as testing mean differences. Hazard ratios (proportional hazards models) are estimates of the risk of the event of interest (such as the start and ending of first partnerships) occurring at any given time. This risk is generally assumed to be constant over the observation period unless time-varying effects are included in the model. The tables show the significance level of a Wald or *t* test of each model parameter being equal to zero. Selection effects are reported in terms of the percentage reduction in RR or OR for marital status when adolescent emotional/behavioural problems and other confounds are included in the models.

RESULTS

Background

First partnership formation and dissolution by age 33

Table 1 gives an overview of cohort members' partnership histories by age 33. The great majority had established partnerships by this stage; direct entry into marriage was the dominant pattern, but a fifth of sample members had lived with their partner before marrying, and just over one in eight first partnerships were non-marital cohabitations. Women entered partnerships earlier than men, and direct entry to marriage occurred earlier than other types of partnership formation. Among women, median ages at marriage, cohabitation leading to marriage, and cohabitation were 21·2 years, 21·8 years and 22·0 years respectively; for men, comparable figures were 23·1, 23·6 and 24·3 years.

A fifth of marriages, whether 'direct' or preceded by cohabitation, had broken down by age 33; breakdown rates for cohabitations were markedly higher. Around 60% of those who separated or divorced had entered second or subsequent partnerships by their early thirties.

Partnership status and psychological morbidity at age 33

Table 2 shows mean Malaise scores, and proportions with high scores (reflecting clinically significant problem levels) in each of the main marital status categories at age 33. Men and women in stable first partnerships had the lowest levels of difficulty on both measures. With the exception of women who had never cohabited, all other groups had significantly increased mean Malaise scores; in addition, those who were alone at age 33 following a partnership breakdown had increased risks of clinically significant depression. As the lower part of Table 2 shows, Malaise scores also varied by first partnership type for those still in first partnerships at age 33. Problem levels were lowest among those who married directly; compared with this group, men who cohabited prior to marriage had significantly increased risks of high Malaise scores, and both cohabiting

Table 1. *Partnership histories by age 33*

Partnership history	Type of first partnership: Never cohabited %	Married %	Cohabited then married %	Cohabited %	Total %
First partnership					
Men ($N = 2780$)	11·6	50·7	23·6	14·2	100·0
Women ($N = 2993$)	5·8	61·1	20·4	12·8	100·0
Breakdown of first partnership					
Men	—	18·6	17·5	64·7	25·7
Women	—	23·8	20·6	75·8	30·1
Second/subsequent partnership of those experiencing breakdown					
Men ($N = 602$)	—	58·6	61·9	63·2	61·1
Women ($N = 795$)	—	57·7	52·4	59·4	57·5

Table 2. *Partnership history and Malaise scores at age 33*

	Mean Malaise scores and proportions with high Malaise scores at age 33							
	Men ($N = 2780$)				Women ($N = 2993$)			
(A) Partnership history	Mean score	RR	High score %	OR	Mean score	RR	High score %	OR
Never cohabited	2·1	1·2*	9·6	1·4	2·2	0·9	12·7	1·0
1 partner, still together	1·7	1·0†	6·9	1·0†	2·5	1·0†	13·1	1·0†
Breakdown, currently alone	2·5	1·5**	15·4	2·4**	4·0	1·6**	25·9	2·3**
Breakdown, 2nd+partner	2·1	1·2**	8·4	1·2	2·9	1·2**	16·5	1·3
	Men ($N = 1861$)				Women ($N = 2020$)			
(B) Partnership type	Mean score	RR	High score %	OR	Mean score	RR	High score %	OR
Directly married	1·5	1·0†	5·4	1·0†	2·4	1·0†	12·2	1·0†
Cohabited prior to marriage	1·9	1·2**	9·8	1·9*	2·7	1·1*	15·2	1·3
Cohabited only	2·2	1·4**	8·5	1·6	2·5	1·1	14·7	1·2

* $P < 0·05$; ** $P < 0·01$.
† Omitted category.

groups had higher mean Malaise scores. Among women, those who cohabited prior to marriage had higher mean scores than their directly married counterparts.

Effects of adolescent behaviour and emotional problems on partnership transitions

Age at first partnership

To assess the impact of adolescent emotional and behaviour problems on partnership transitions, we began by examining age at first co-residential partnership. NCDS offered an extended observation period here, beginning with relationships formed in the teens, and ending with those established in the early thirties. Because previous studies suggested that effects might vary by age, we divided the observation period into three phases, reflecting 'early', 'modal' and 'later' ages at partnership formation, based on interquartile ranges in age at first partnership for men and women separately. We then used Cox proportional hazards models to estimate effects of adolescent problems (controlling for other childhood background confounds) on age at entry into partnerships within each phase.

Table 3 shows the results. In the earliest phase of partnership formation (covering the mid to late teens for women, and the teens and very early twenties for men), the confounding factors

Table 3. *Proportional hazards model for age at first partnership formation*

Variable	Hazard ratio	95% CI
Men		
First quartile: age ⩽ 21 yr 7 mo ($N = 2780$)		
Childhood social class: manual *v.* non-manual	1·37**	1·14, 1·64
Parental divorce	1·70**	1·32, 2·19
Age left full-time education		
16 to 18 *v.* up to 16	0·51**	0·40, 0·65
Over 18 *v.* up to 16	0·52**	0·41, 0·66
Antisocial behaviour age 16	1·43**	1·18, 1·73
Interquartile range: ages > 21 yr 7 mo to 27 yr 6 mo ($N = 2089$)		
Parental divorce	1·32**	1·07, 1·62
Age left full-time education		
16 to 18 *v.* up to 16	0·80**	0·70, 0·92
Over 18 *v.* up to 16	0·75**	0·66, 0·85
Emotional problems age 16	0·66**	0·55, 0·79
Last quartile: age > 27 yr 6 mo ($N = 706$)		
Emotional problems age 16	0·69**	0·52, 0·91
Women		
First quartile: age ⩽ 19 yr 6 mo ($N = 2993$)		
Parental divorce	1·74**	1·39, 2·18
Age left full-time education		
16 to 18 *v.* up to 16	0·46**	0·37, 0·56
Over 18 *v.* up to 16	0·20**	0·15, 0·26
Antisocial behaviour age 16	1·87**	1·55, 2·26
Interquartile range: ages > 19 yr 6 mo to 24 yr 4 mo ($N = 2292$)		
Age left full-time education		
16 to 18 *v.* up to 16	0·78**	0·68, 0·88
Over 18 *v.* up to 16	0·58**	0·51, 0·65
Parental divorce	1·22*	1·00, 1·50
Last quartile: age > 24 yr 4 mo ($N = 756$)		
No variables significant		

* $P < 0·05$; ** $P < 0·01$.

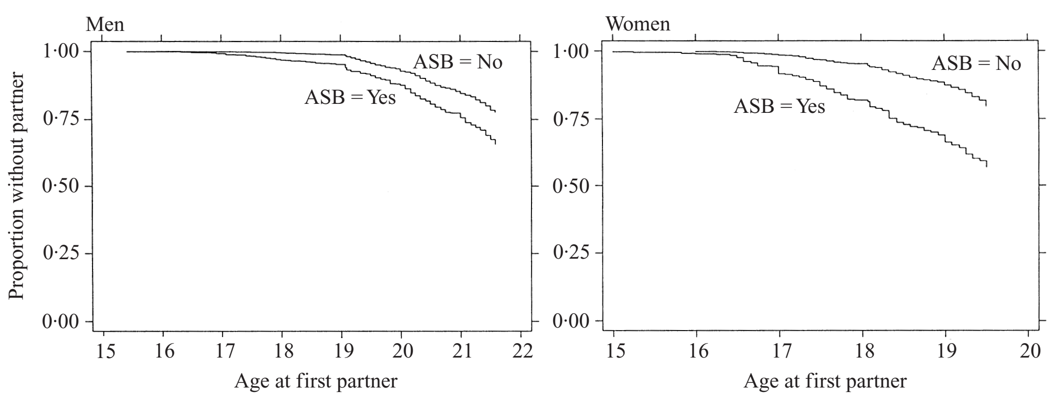

FIG. 1. Kaplan–Meier survival estimates for age at first partnership in first quartile by adolescent antisocial behaviour (ASB) for men and women.

of childhood social class, education, and parental divorce showed a variety of significant effects. Taking these into account, adolescent emotional problems had no influence on rates of partnership formation, but antisocial behaviour did. For both men and women, antisocial behaviour at age 16 was strongly associated with increased rates of partnership formation later in the teens. Fig. 1 illustrates these effects with Kaplan–Meier estimates of the survival curves. To test how far effects were constant across this age-range, we added an interaction term that allowed the

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Table 4. *Multinomial logit model for type of first partnership: pre-marital cohabitation and cohabitation only compared with direct entry into marriage*

Variable	Odds ratio	95% CI
Men ($N = 2463$)		
Cohabitation → Marriage		
Parental divorce before age 16	1·90**	1·31, 2·75
Age left full-time education		
16 to 18 *v.* up to 16	1·29	0·99, 1·68
Over 18 *v.* up to 16	1·90**	1·50, 2·42
Age started partnership (years)	1·05**	1·02, 1·08
Antisocial behaviour age 16	1·47**	1·12, 1·94
Cohabitation		
Parental divorce before age 16	2·69**	1·77, 4·07
Age left full-time education		
16 to 18 *v.* up to 16	1·64**	1·19, 2·26
Over 18 *v.* up to 16	2·49**	1·84, 3·37
Age started partnership (years)	1·12**	1·08, 1·17
Antisocial behaviour age 16	2·48**	1·81, 3·39
Women ($N = 2815$)		
Cohabitation → Marriage		
Childhood social class: manual *v.* non-manual	0·69**	0·56, 0·84
Parental divorce before age 16	2·33**	1·71, 3·17
Age left full-time education		
16 to 18 *v.* up to 16	0·83	0·65, 1·07
Over 18 *v.* up to 16	1·50**	1·16, 1·93
Age started partnership (years)	1·06**	1·03, 1·09
Antisocial behaviour age 16	1·75**	1·31, 2·33
Cohabitation		
Childhood social class: manual *v.* non-manual	0·69**	0·53, 0·89
Parental divorce before age 16	1·73**	1·16, 2·58
Age left full-time education		
16 to 18 *v.* up to 16	1·07	0·79, 1·45
Over 18 *v.* up to 16	1·52*	1·10, 2·12
Age started partnership (years)	1·10**	1·06, 1·15
Antisocial behaviour age 16	2·62**	1·90, 3·63

* $P < 0{\cdot}05$; ** $P < 0{\cdot}01$.

hazard ratio to vary between the first and second halves of the observation period. Effects for men were markedly increased up to age 19 (hazard ratio = 3·16, CI = 2·03–4·92), but became non-significant thereafter. For women, effects were significantly elevated both before and after age 18 (up to 18 years, hazard ratio = 2·89, CI = 2·13–3·93; ages 18 to 19·50, hazard ratio = 1·49, CI = 1·18–1·88).

In the twenties and early thirties, findings differed for men and women. Among women, age at completion of full-time education and parental divorce were the only factors to show continuing effects on rates of first partnership formation: neither antisocial nor emotional difficulties in adolescence had any impact beyond the late teens. For men, both parental divorce and age at completion of education showed significant effects during the central, interquartile partnership formation phase. In addition, effects of adolescent emotional problems emerged during this central time-period, and persisted – this time as the only significant predictors in the models – for relationships formed from the late twenties onwards. In both instances, young men who had shown high levels of emotional difficulties in their teens were less likely to establish partnerships than their peers.

Influences on first partnership type

Next, we examined how far adolescent emotional and behavioural problems affected the types of first partnerships that cohort members entered: marriage, cohabitation leading to marriage, and cohabitation alone. As we had already found variations in the ages at which these different types of relationship were formed, age at first partnership was included as a covariate throughout.

Table 4 shows the results, using direct entry into marriage as the baseline category. Parental divorce, more extended full-time education and later age at partnership formation were associated with increased rates of both pre-marital and non-marital cohabitation in both sexes. For women only, manual social class background was associated with a reduced likelihood of cohabitation. Taking these background confounds into account, adolescent emotional problems had no significant effect on partnership type, but antisocial behaviour did. In all analyses, teenage antisocial difficulties were associated with an increased likelihood of pre-marital and non-marital cohabitation, with odds ratios ranging from 1·47 for pre-marital cohabitation among men to 2·62 for non-marital cohabitation in women.

Influences on first partnership breakdowns

Finally, we explored how far adolescent behavioural and emotional characteristics were associated with risks of breakdown in first partnerships. Because cohort members had established relationships at a wide range of ages, their exposure to risks of breakdown also varied. To take this into account, initial analyses used length of first partnership as the dependent variable. In a univariate test antisocial problems in adolescence emerged as a significant predictor for both men and women (men: $\chi^2(1) = 23{\cdot}17$, $P < 0{\cdot}001$, women: $\chi^2(1) = 9{\cdot}85$, $P < 0{\cdot}002$), while adolescent emotional problems were only associated with higher rates of breakdown among men ($\chi^2(1) = 8{\cdot}13$, $P < 0{\cdot}004$). Next, we added a series of potential confounds, beginning with childhood social class, parental divorce, and length of full-time education. In addition to these childhood factors, earlier stages of the partnership formation process also seemed likely to influence risk of breakdown. It was already clear, for example, that cohabitations were at greater risk of breakdown than legal marriages in this cohort, and much previous research suggested that age at partnership formation would prove a strong predictor of relationship stability. Finally, we anticipated that the presence of children would constitute a bar to relationship breakdown.

Table 5 shows significant results from proportional hazards models testing this full range of factors. Taking other significant influences into account, antisocial behaviour in adolescence continued to show direct effects on rates of partnership breakdown for men, but not for women. More detailed tests suggested that partnership type constituted the most important mediator here. As noted earlier, antisocial problems in adolescence were strong predictors of cohabitation rather than legal marriage, and cohabitation in turn was associated with markedly increased risks of breakdown. Among the women, these prior associations essentially served to eliminate any direct effects of antisocial behaviour on rate of first partnership breakdown; for men, direct effects remained, though these too were considerably reduced by the inclusion of the indicator of partnership type. The interaction between partnership type and antisocial problems in adolescence for men suggested that there was a ceiling on the risk of partnership breakdown among men with antisocial problems whose first partnership was a cohabitation: their risk was less than the addition of the separate main effects (hazard ratio = 5·48).

Selection effects on mental health at age 33

These findings suggested that adolescent emotional and behavioural problems influenced first partnership transitions in a variety of ways. The final series of analyses was designed to assess how far these effects attenuated associations between marital status and later mental health. In some instances, of course, this was difficult to determine, as cohort members had made further partnership transitions before the mental health assessments at age 33. The most unambiguous evidence for selection effects thus involved individuals whose partnership status at 33 still reflected their first partnership transition (e.g. those still in their first relationship at age 33, or who had remained alone throughout), and where we had detected some effects of adolescent emotional/behavioural problems on these first transitions. Three particular subgroups met these criteria: (*i*) men who remained alone to age 33; (*ii*) men and women still in first cohabitations, or in marriages preceded by cohabitation; and (*iii*) cohort members who had experienced the breakdown of a first partnership by age 33. In the first group, we anticipated that single men's increased vulnerability to low mood in the early 30s would reflect selection effects associated with emotional problems in the teens; in the second and third groups, we anticipated selection effects associated with adolescent antisocial behaviour.

Table 5. *Proportional hazards model for breakdown of first partnership*

Variable	Hazard ratio	95% CI
Men ($N = 2463$)		
Age left full-time education		
16 to 18 *v.* up to 16	1·27*	1·01, 1·60
Over 18 *v.* up to 16	0·86	0·68, 1·09
Age started partnership (years)	0·89**	0·86, 0·91
Had children in partnership	0·21**	0·17, 0·25
Type of partnership – cohabitation *v.* marriage	6·36**	5·06, 8·00
Antisocial behaviour age 16	1·79**	1·39, 2·32
Interaction (antisocial and partner type): antisocial and cohabited *v.* not antisocial and married	0·48**	0·30, 0·77
Women ($N = 2815$)		
Age started partnership (years)	0·88**	0·85, 0·91
Had children in partnership	0·23**	0·19, 0·27
Type of partnership – cohabitation *v.* marriage	4·66**	3·82, 5·69

* $P < 0·05$; ** $P < 0·01$.

Table 6. *Selection effects*: *adolescent/early adult emotional problems*, *single status and mean Malaise scores age* 33 (*men*)

	Model 1a	Model 1b
	RR (95% CI)	RR (95% CI)
Age 16–33 ($N = 2178$)		
Not cohabited *v.* one stable partnership	1·22* (1·03, 1·45)	1·18 (0·99, 1·41)
Emotional problems age 16		1·28** (1·07, 1·53)
Ages 23–33 ($N = 1250$)		
Not cohabited *v.* one stable partnership	1·27* (1·07, 1·50)	1·22* (1·02, 1·45)
High Malaise at 23		1·19** (1·17, 1·21)

Model 1a, unadjusted for prior emotional problems; Model 1b, adjusted for prior emotional problems.
Models for ages 16–33 compare men who had not cohabited by age 33 with all those in stable first partnerships, and use high emotional score at age 16 to predict Malaise at 33. Models for age 23–33 compare men who had not cohabited by age 33 with those in stable first partnerships beginning after age 23, and use high Malaise scores at 23 to predict Malaise at 33.
* $P < 0·05$; ** $P < 0·01$.

Men: single status at age 33

To assess how far adolescent emotional problems accounted for single men's higher mean Malaise scores at age 33 we contrasted results from two poisson regression models. The first estimated effects of single *versus* partnered status, using those still in first partnerships as the comparison group (Table 6, Model 1a); the second re-assessed those effects with indicators of adolescent emotional difficulties included in the model (Model 1b). As a check on the pattern of the findings, a parallel series of analyses was undertaken focusing on men who began their first partnerships after age 23, and using the age 23 Malaise scores as the indicator of 'prior' emotional difficulties. These basic models were then repeated to include other significant covariates of age 33 Malaise (parental divorce and length of education – not shown).

Each set of analyses suggested similar conclusions. As expected, earlier emotional difficulties were strongly associated with age 33 Malaise scores, but the inclusion of these indicators had only a minor impact on the estimates for single status. Adolescent emotional problems were associated with a 3·3% reduction in the RR for single status, and the more proximal 23-year measures with somewhat larger effects (3·9%). Including other significant covariates in the analyses reduced these effects slightly (to 2·9% and 3·2% respectively). Although

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prior emotional difficulties were associated with a significant reduction in men's likelihood of establishing partnerships in their twenties and early thirties, this nonetheless appeared to play only a minor role in single men's reported rates of psychological distress at age 33.

Type of first partnership

Next we used similar approaches (with the directly married again forming the comparison group), to assess how far the increased vulnerability of men and women who had cohabited before marriage, or continued to live in non-marital cohabitations, was mediated by selection effects deriving from adolescent antisocial behaviours. As outlined earlier, men who cohabited prior to marriage had significantly elevated risks of both categorical high Malaise scores and mean levels of low mood at age 33 by contrast with their directly married counterparts. Although antisocial behaviour in the teens was significantly associated with pre-marital cohabitation, the inclusion of this term in the models had a negligible effect on the estimates for partnership type (0·9% reduction in OR for high Malaise, 0·7% reduction in RR for mean Malaise). Cohabitation alone was associated with increased mean Malaise scores among the men. Here, adolescent antisocial behaviour was associated with a 5·5% reduction in the RR for cohabitation (4·1% with other childhood covariates included). Among the women, the only significant effects of partnership type were in terms of increased mean Malaise scores for those who had cohabited prior to marriage. Including adolescent antisocial behaviour in the models reduced these effects by $< 1·0\%$. (Full details available from the authors.)

Table 7. *Selection effects for partnership breakdown*

Variable	Malaise at 33 OR	95% CI
Men		
Model 1a ($N = 2463$)		
Any breakdown	1·69*	1·21, 2·36
Model 1b		
Any breakdown	1·59*	1·13, 2·23
Antisocial behaviour age 16	1·75*	1·20, 2·54
Model 2a ($N = 2083$)		
Breakdown and alone at age 33	2·44*	1·58, 3·78
Model 2b		
Breakdown and alone at age 33	2·27*	1·46, 3·52
Antisocial behaviour age 16	1·87*	1·23, 2·83
Model 3a ($N = 2463$)		
Breakdown before age 24	3·12*	1·98, 4·90
Model 3b		
Breakdown before age 24	2·83*	1·79, 4·47
Antisocial behaviour age 16	2·21*	1·54, 3·18
Women		
Model 1a ($N = 2815$)		
Any breakdown	1·73*	1·37, 2·18
Model 1b		
Any breakdown	1·67*	1·33, 2·11
Antisocial behaviour age 16	1·64*	1·23, 2·21
Model 2a ($N = 2342$)		
Breakdown and alone at age 33	2·32*	1·72, 3·13
Model 2b		
Breakdown and alone at age 33	2·20*	1·62, 2·97
Antisocial behaviour age 16	1·81*	1·31, 2·50
Model 3a ($N = 2815$)		
Breakdown before age 24	2·24*	1·70, 2·95
Model 3b		
Breakdown before age 24	2·05*	1·55, 2·71
Antisocial behaviour age 16	2·12*	1·63, 2·74

* $P < 0·05$; ** $P < 0·01$.

First partnership breakdown

Finally, we estimated a series of models to examine the impact of selection processes on links between first partnership breakdown and later mental health. Three rather different approaches were possible here. First, we examined a general model comparing outcomes for all cohort members who had experienced a relationship breakdown (irrespective of their subsequent relationship history) with those who remained in stable first partnerships at age 33 (Table 7, Models 1a and 1b). Secondly, we focused on the smaller subgroup still living alone after a breakdown at age 33 (Models 2a and 2b). Thirdly, we explored effects on very early breakdowns, using Malaise scores at age 23 as the dependent variable (Models 3a and 3b).

Table 7 shows the results, reporting reductions in ORs for clinically significant high Malaise scores; poisson regression models for mean total scores showed a similar pattern of findings. In all models selection effects associated with antisocial behaviour in adolescence were present, but small in magnitude. Effects were largest for early breakdown, where adolescent antisocial problems attenuated the effect of breakdown on high Malaise scores at age 23 by 9·2% in men and 8·3% in women. For any breakdown, and breakdowns where cohort members were still alone at 33, effects were smaller, showing attenuations of 6·1% and 7·3% respectively in men, and 5·4% and 3·3% in women. As in all previous analyses, the findings suggested that small but persistent selection effects contributed to the links between partnership experience and adult mental health.

DISCUSSION

This study set out to explore how far selection processes based on adolescent psychological characteristics might contribute to the well-documented associations between marital status and adult mental health (Pearlin & Johnson, 1977; Gove *et al*. 1983; Bebbington, 1988). The NCDS data set had a number of advantages for a study of this kind: measures of behaviour problems in the teens, before the vast majority of partnerships were formed; measures of mental health in adulthood; and detailed partnership histories over an extended observation period, spanning 17 years. The cohort included substantial minorities of young people who cohabited before marriage, whose first relationships were non-marital cohabitations, and whose first partnerships had ended before age 33, making it possible to examine effects of adolescent characteristics on a variety of partnership indicators. In addition, the data set included measures of a number of other childhood factors known to influence partnership formation and to be associated with mental health, so that effects could be assessed net of these other influences.

As with the majority of large-scale longitudinal studies, the main disadvantage of the data set lay in the effects of attrition and missing data. These resulted in a sizeable reduction in the absolute size of the samples available for analysis, and in some clearly identifiable patterns of selective loss. We used weights to correct so far as possible for any bias this might have introduced. As we were unable to weight back to the full original birth cohort, we regard the results as indicative rather than definitive; our estimates of selection effects may be somewhat conservative as a result. A further potential limitation stemmed from the fact that while the partnership histories provided detailed information on the specific ages at which partnerships began and ended, the mental health assessments were made at fixed points for all cohort members. We return to the implications of this issue later. Finally, it is important to note that the study was based on a single birth cohort. The rapid secular changes in patterns of partnership formation and dissolution over recent years (Hess, 1995) may mean that some of the associations we found are specific to this particular cohort.

Preliminary analyses confirmed the links between marital status and psychiatric morbidity found in many previous studies. The most marked effects were associated with partnership breakdowns: men and women who had experienced the breakdown of a relationship and were alone at the time of the interview had the highest rates of mental health difficulties of any in the cohort, and those who had formed new partnerships after a breakdown continued to show some increased risk of low mood. For cohort members who had remained alone into their early 30s, effects varied by gender: single men showed increased vulnerability to low mood by comparison with men in stable first partnerships, but single women showed no increase in risk. The analyses also highlighted important distinctions between partnership types, with cohabitation associated with higher levels of psychological distress than marriage. These varying associations emphasize the need to specify partnership histories in some detail if we are to elucidate links with mental health problems in adult life.

We assessed the possibility of selection effects in two stages, first exploring associations between adolescent problems and partnership transitions, then testing how far these associations might mediate links between partnership status and later mental health. Most previous selection studies have concentrated on testing for associations; as our findings underlined, including tests of mediation affected the conclusions in some important ways. In addition, throughout the analyses we took account of other childhood factors – social class, education, and exposure to parental divorce – known from previous studies to be associated both with patterns of partnership formation and with adolescent and adult mental health. So far as we are aware, our study is also the first to control for potential confounds of this kind.

We explored three routes whereby prior behavioural characteristics might influence partnership transitions: through effects on the age at which young people began first partnerships; on the types of partnerships they established; and on the risks that those partnerships would break down during the period of the study. Each series of analyses highlighted significant effects of adolescent emotional or behavioural problems. In terms of age at partnership formation, most previous studies had assumed that psychological vulnerability would function to delay partnership entry, though support for this model had been mixed. Our findings suggested some reasons for these inconsistencies, showing that processes of this kind do occur, but in relatively focused ways. In the NCDS cohort, they were important for men but not for women; in relation to prior internalizing but not externalizing difficulties; and in the central and later parts of the observation period, from the mid-twenties onwards, but not in the teens. Not surprisingly, perhaps, influences on partnership formation showed both age and gender specificity, arguing that models for selection effects also need to be framed to take such variations into account.

By examining each phase of partnership formation separately, our analyses also highlighted a second series of effects, rather different from those posited in many selection models. In the earliest years of partnership formation, and especially in the teens, we found a marked tendency for antisocial behaviour to increase the rate of partnership formation among both men and women. Similar 'precocious' patterns of early adult role transitions have been noted in studies of clinically defined groups (see e.g. Bardone *et al*. 1996; Forthofer *et al*. 1996; Stattin & Magnusson, 1996), where conduct disordered young people have been found to begin both sexual and cohabiting relationships earlier than their peers. In addition, and again confirming previous reports, we found that antisocial tendencies were associated with increased rates of both pre-marital and non-marital cohabitation.

As yet, the basis for these patterns is unclear: they may reflect behavioural impulsivity, a lack of perceived alternatives to marriage for some vulnerable groups of young people, or, in more extreme cases, a desire to escape from unsatisfactory home conditions. Whatever their exact basis, their prime interest here lay in the possibility that they might form the first steps in a pathway to increased risks of partnership breakdown. The findings bore this out. Among women, significant bivariate links between partnership breakdown and adolescent antisocial behaviour were largely mediated by partnership type. For men, a more complex series of influences emerged, including both indirect links via age at partnership and partnership type, and additional direct effects of earlier antisocial behaviour.

Taken together, these 'transition' analyses highlighted a range of effects of prior psychological characteristics on patterns of partnership formation and dissolution. As a result, we anticipated that a substantial proportion of the psychological vulnerability associated with differing partnership statuses at age 33 might be traceable to selection processes. In practice, the findings suggested a persistent but far from dominant role. Although adolescent characteristics showed significant links with later mental health, little of this impact seemed mediated via partnership transitions. Across a range of tests, the magnitude of selection effects on later psychological morbidity was modest – never rising above 10% – and was further reduced when other childhood and adolescent factors were included in the models.

Methodological factors may have played some part here. Our measures of adolescent characteristics relied on teacher ratings of emotional and behavioural problems in the teens. Although ratings of this kind are known to be quite robust predictors of later mental health, they may nonetheless be less than ideal indicators of the characteristics of prime interest here. Other measures – especially perhaps self-reports of adolescent emotional and behavioural difficulties – might have highlighted somewhat stronger selection effects. In addition, as outlined above, the NCDS data-set covered an extended observation period, and involved different measures of mental health problems in adolescence and early adult life. Both of these factors may have attenuated the effects that emerged. It was clear, for example, that effects were strongest when we were able to constrain the observation period (as in the analysis of partnership breakdown before age 23), or when the same measure was used at each assessment (as in the analyses involving Malaise scores at ages 23 and 33). Estimates from our other analyses reflect average values, pooled across relationships that could range from under one year to over 16 years' duration. It seems quite plausible, however, that the relative importance of selection and causation influences may vary at different stages of a relationship. Selection effects may be most salient in the early years, but decrease with time, as the effects of shared experience take prominence. In a somewhat related way Booth & Johnson (1998) explored how far the quality of marital relationships reflects characteristics of the individual partners, or emergent properties of the relationship itself. They concluded that relationship quality was largely an emergent feature; in the present context, a model of this kind would suggest that although patterns of partnership formation and dissolution may be considerably influenced by individual characteristics, the longer-term implications of partnership status for mental health may reflect additional processes.

We conclude with two caveats. First, our analyses focused on a population sample, with adolescent measures designed to indicate serious but not necessarily clinically significant problems in the teens. Previous research based on diagnostic groupings has documented strong links between adolescent psychopathology and negative outcomes in the domains of both relationships and mental health in adult life (see e.g. Robins *et al*. 1991). Selection effects may thus be more marked in clinical samples. Secondly, we followed the epidemiological literature in focusing on links between partnership status and adult mental health. But partnership status categories alone are almost certainly too crude and too restricted to reflect the full range of processes involved. A better understanding of both selection and causation processes will almost certainly require more sensitive measures, including indicators of the quality of relationships, and of related supports and stressors. More detailed measures of this kind might also highlight a different pattern of effects. At this stage, however, our findings suggest that although individual characteristics do indeed influence both the timing and the type of early adult relationships, it is those relationships themselves (or the lack or loss of them) that hold important implications for adult mental health.

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