Children's age at parental divorce and depression in early and mid-adulthood

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Running head: Age at parental divorce and later depression

Abstract

This study aimed to assess whether children's age at their parents' divorce is associated with depression in early and mid-adulthood, as indicated by medication purchase. A sibling comparison method was used to control for unobserved factors shared between siblings. The data were extracted from the Norwegian Population Register and Norwegian Prescription Database and included about 181,000 siblings aged 20–44 who had experienced parental divorce and 636,000 who had not. Controlling for age in 2004, sex, and birth order, children who were aged 15–19 when their parents divorced were 12 per cent less likely to purchase antidepressants as adults in 2004–08 than those experiencing the divorce aged 0–4. The corresponding reduction for those aged 20+ at the time of divorce was 19 per cent. However, the association between age at parental divorce and antidepressant purchases was only evident among women and those whose mothers had low education.

Keywords: children; age at divorce; depression; prescription; register data; sibling analysis; antidepressants

Introduction

The implications of parental divorce for children's well-being have attracted much scholarly interest, not least because of large increases in divorce rates in many Western countries over the past half century (Amato 2000). In addition to examining general effects, some previous studies have sought to investigate whether effects vary by the child's age at the time of the divorce (Furstenberg and Kiernan 2001; Sigle-Rushton et al. 2014). An important motive for this is that the existence of such differences would suggest that children's age at parental divorce should be taken into account in setting priorities for the provision of support to families experiencing divorce.

Several previous studies on how parental divorce generally affects the well-being of offspring have focused on mental health, including depression (Amato 2000). Most scholars have analysed implications for mental health in childhood, although some have considered the mental health of young adults in their 20s (Chase-Lansdale et al. 1995) or 30s (Cherlin et al. 1998; Amato and Sobolowski 2001; Furstenberg and Kiernan 2001; Gilman et al. 2003, Sigle-Rushton et al. 2005). Only a few of these studies of the effects of parental divorce on offspring's mental health have considered variations in associations by children's *age* at parental divorce. Those that *have* examined this have failed to find such variations (Chase-Lansdale et al. 1995; Amato and Sobolowski 2001; Ermisch and Francesconi 2001; Størkersen et al. 2006; Uphold-Carrier and Ute 2012; Gähler and Palmtag 2015), but most of them included only a few hundred children and so may have lacked statistical power to identify effects that might nevertheless be important.

A major problem in this research area is that children who experience parental divorce are likely to come from families with various characteristics—many unobserved—that themselves cause poorer child outcomes. Similarly, the *age* at parental divorce may be linked to factors of importance for the outcomes. Ideally, we would like to control as well as possible for such selective influences. In this study, sibling fixed effects models were estimated to account for unobserved factors shared by siblings, such as genetic similarities and the stable components of the parents' attitudes, behaviours, and socio-economic resources. This approach has been used in some studies of effects of children's age at parental divorce on subsequent educational attainment (e.g., Björklund and Sundström 2006; Sigle-Rushton et al. 2014), but in only one previous study of the effects of children's age at parental divorce on later mental health (Ermisch and Francesconi 2001). Their relatively small investigation found no significant relationship between children's age at divorce and a summary indicator of distress when they were around age 20.

The present study is a full-sibling analysis of the association between children's age at their parents' divorce and their chance of suffering from depression in early and mid-adulthood (ages 20–44 in 2004), as measured by purchases of antidepressants registered in the Norwegian Prescription Database between 2004 and 2008. As discussed further in the 'Limitations' subsection, the focus on medication purchases excludes many cases, especially of less severe depression, that might be included in studies using surveys where depressive symptoms are self-reported (Fournier et al. 2010; Hämäläinen et al. 2009). However, survey data rarely allow the estimation of sibling models, and register data on medication have the advantage of constituting an objective indicator not influenced by recall or reporting bias.

In addition to exploring how children's age may modify the effect of divorce on their later mental health, earlier investigations have considered a number of other variations in the effect of divorce. (The word 'effect' is used here and elsewhere, for simplicity. Realistically, we can hardly ever be sure that a truly causal effect has been estimated.) For example, there has been an interest in whether divorce affects boys and girls differently, and whether the effects vary by the parents' education (see references cited in 'Background' section). Additionally, some attention has been given to whether the implications of divorce depend on whether the mother or father forms a new relationship (Chase-Lansdale et al. 1995; Gilman et al. 2003; Shafer et al. 2017; Sweeney 2010), and a few authors have considered interactions with number of siblings (Sun and Li 2009) or the parents' divorce probability predicted from socio-demographic characteristics (Amato and Anthony 2014). However, all these studies have dealt with variations in the effects of divorce in general, rather than in the effects of the child's *age* at divorce. In the present analysis, three factors potentially modifying the effects of the child's age at divorce on later depression are considered: the child's sex, maternal education, and paternal education. Furthermore, in some models, variables that may mediate the effect of age at parental divorce on depression are included. These potential mediators are the child's educational attainment, their own marital status, and the number of children they themselves had by 2004. Such variables have been included in a few earlier studies of the effects of divorce on mental health, but these studies have considered divorces in general without regard to the children's age (Amato and Sobolewski 2001; Gilman et al. 2003).

The Norwegian setting

Divorce is now common in Norway, as in most other high-income countries (Prioux 2006). A couple who experience the marital-duration-specific divorce rates observed in Norway in 2016 throughout their marriage have a 38 per cent chance of ever divorcing (Statistics Norway 2018a). To the extent that Norwegian children are influenced by the age at which they experience parental divorce (or disruption of their parents' consensual union), there is no obvious reason why such effects should not also exist in other settings. In fact, we might expect even stronger effects elsewhere, as the rather generous social welfare policies in Norway (Baran et al. 2014) may serve to dampen some of the potentially adverse implications of parental disruption.

The rising divorce rates have occurred in tandem with a strong increase in informal cohabitation (Noack 2010). For example, at the end of the period studied in this paper, 38 per cent of all unions among women and men aged 25–49 were consensual (calculations from tables in Statistics Norway 2018b). It has also become increasingly common to have children outside marriage: already in 1986, two years after the youngest children included in this study were born, 28 per cent of the births in Norway were to single or, more commonly, cohabiting women (calculated from tables in Statistics Norway 2018c). Many single or cohabiting mothers later marry their partner and are therefore included in this analysis, but there are also many who do not. This is discussed further in the 'Data' section and 'Limitations' subsection. The chance of a break-up is much higher in informal unions than in marriages, even if the couple have a child together (Jensen and Clausen 2003).

Background

A number of pathways may contribute to the relatively adverse outcomes observed among children who have experienced parental divorce. The marital discord typically underlying the divorce (which, of course, may also characterize some marriages that remain intact) may be harmful to children (Amato and Sobolewski 2001). Parental conflict may make children sad, angry, or frightened, and they may blame themselves for the situation (Pryor and Rodgers 2001). Parents in a troubled marriage may also have less time to care for and supervise their children (Amato 2000). Additionally, although the divorce itself (or, rather, the actual separation that typically precedes it) may reduce or change the nature of the original conflict between the parents, it may give rise to new conflicts about various practical arrangements, including child custody and how to raise children. Such post-separation conflicts may also be harmful to children (Gilman et al. 2003; Kalmijn 2016) and cause further distress to parents, which might have additional adverse effects (although an earlier Norwegian study found that parents' mental health after divorce did not mediate the effect of divorce on children's well-being; Størkersen et al. 2006). Furthermore, children may receive less parental attention simply because, at any one point in time, they live with only one of their parents and, if the parent–child relationship is weakened, this may impede children's development of social skills (Cavanagh and Huston 2008). Divorce may also involve children in various other transitions, such as moving to another area and changing school, which may increase distress (Amato 2000). Additional family changes may also be experienced, and although a step-parent may be very supportive, having a step-parent and perhaps stepsiblings may be a source of stress as well (Sweeney 2010; Shafer et al. 2017). Finally, fewer economic resources may be available to children after a divorce, despite welfare support, maintenance payments from the noncustodial parent, and possibly increased work activity on the part of the custodial parent (Andress et al. 2001). However, the unfavourable situation after a parental divorce should be considered relative to the most reasonable counterfactual, and some studies have indeed indicated that children fare better after divorce than if their parents remain in a poor relationship (Booth and Amato 2001; Strohschein 2005; Musick and Meyer 2010).

These effects of divorce and the underlying marital discord may contribute to weakening children's educational progression (Steele et al. 2009; Potter 2010; Ermisch and Pronzato 2011) and also have implications for later life through other mechanisms. For example, parental divorce may increase children's chances of having a child early or outside marriage (Reneflot 2009) or their own chance of divorcing (Dronkers and Härkönen 2008). All these may in turn have implications for their later mental health (Bulloch et al. 2009; Kessler and Bromet 2013; Kravdal et al. 2015). However, results from previous empirical studies of such causal pathways have been inconclusive. Some have found that parental divorce affects children's mental health partly through their education (Gilman et al. 2003); others have concluded that neither children's education nor their own marital status appear to mediate the effects of parental divorce on adult psychological well-being (Amato and Sobolewski 2001).

Age at parental divorce

Given these types of effects of parental divorce, it would also be reasonable to expect that the age at which it is experienced matters. One argument is that, if the divorce happens at an early age, children will spend a larger part of their childhoods with the disadvantages resulting from the divorce and parental discord (although the intensity of these disadvantages may change over time). A longer period with disadvantage in childhood may contribute negatively to well-being in adulthood, including mental health. This is referred to from now on as the '*duration of childhood disadvantage'* argument. A similar argument is that, the younger the age of a child at parental divorce, the greater the likelihood of further family changes, such as repartnering of a parent, which are potentially stressful (although also potentially beneficial).

Additionally, it is possible that the intensity and character of the disadvantages depend on how old the children are when the divorce takes place. This is referred to as the argument about *'age at divorce affecting disadvantage intensity and character'*. To be more specific, the life changes that occur around the time of divorce may have particularly adverse immediate effects on children if the divorce happens very early in life, or such life changes at an early age may lead to particularly strong disadvantages through the remaining childhood. Also, because of the stronger intensity, and perhaps different character, of the childhood disadvantages that are caused by early parental divorce, there may be a larger chance of implications for adult well-being. For example, it has been pointed out that a parental divorce early in life is particularly likely to disrupt normal development, with possibly long-lasting consequences for social behaviour and otherwise (Cavanagh and Huston 2008; Ryan and Claessens 2013). This may be considered as 'version 1' of the argument.

However, the opposite is also possible: perhaps reduced parental attention, emotional strains, or other changes around the time of divorce have the most harmful implications in the short term, or for the remaining childhood, if the divorce happens when the children are older and understand more. In support of that idea, Diener at al. (2008) found that the immediate (at least) consequences of reduced attention from or attachment to the parents were strongest in mid-childhood. Furthermore, because older children typically have to spend more time on schoolwork than the younger children, lack of support and emotional distress at that age because of a recent parental disruption may be particularly likely to weaken school performances, with possibly serious implications for later life. This may be considered as 'version 2' of the argument. Clearly, if parental divorce at a relatively older age leads to particularly large disadvantages over the remaining childhood, or leads to types of disadvantages that are particularly likely to leave a mark on later life, and if this is not counterbalanced by the other mechanism mentioned ('duration of childhood disadvantage' argument), it would mean that we would see worse adult outcomes, the higher the children's age at the parental divorce.

As noted, earlier studies have not reported significant associations between age at divorce and later depression among the children of divorce, but there is evidence from sibling models and similar analyses about effects of age at parental divorce on educational achievements (Steele et al. 2009; Sigle-Rushton et al. 2014). Furthermore, relationships between age at divorce and own family formation (and some other outcomes) have been found in more traditional regression analysis (Furstenberg and Kiernan 2001) and, as mentioned, education and own family formation (or dissolution) may in turn have implications for later risks of depression.

Interactive effects

Effects of divorce may vary by the child's sex. Possibly boys tend to be more attached to their fathers and benefit from having a male role model (Diener et al. 2008), and it is the father who is usually the parent ceasing to be co-resident after a divorce. In Norway, 82 per cent of children in disrupted families in 2004 lived with their mother and 8 per cent with their father, while custody was shared for 10 per cent (Lyngstad et al. 2014). Additionally, it has been argued that there may be differences between girls' and boys' reactions to stress and their coping strategies (Dedovic et al. 2009; Seiffge-Krenke 2011), although it is far from apparent what the total contribution from this would be, in particular with respect to the stress resulting from parental divorce. It is clearly also hard to predict how the importance of *age* at parental divorce will vary between the sexes. For example, if the disadvantages resulting from divorce were generally most pronounced for boys, we would expect a particularly sharp negative association with age for them, based on the 'duration of childhood disadvantage' argument. Conversely, if parental divorce at a relatively high age led to particularly large disadvantages during the remaining childhood, or types of disadvantage that are particularly likely to have longer-term influences (version 2 of the argument about 'age at divorce affecting disadvantage intensity and character'), and if this were the case especially for boys, an opposite pattern would be seen.

Amato (2001) concluded in a review paper that the evidence suggests moderately greater effects of divorce on boys' well-being than on girls', at least in some domains, while two more recent studies showed larger effects on mental health for girls (Størkersen et al. 2006) or effects exclusively for girls (Huurre et al. 2006). Cavanagh and Huston (2008) found in a study of social development up to age eleven that experiencing parental divorce at a young age was particularly bad for boys. Sex differences in the importance of children's age at divorce for mental health in the longer run have not been examined.

With respect to parental education, some studies have indicated that the effect of divorce on children is weaker if the mother, who is usually the main custodial parent, has a high level of education and is therefore more economically (and otherwise) resourceful. By contrast, for fathers—often the parent who leaves the child's household—high education makes divorce more disadvantageous (Mandemakers and Kalmijn 2014; for a study of children's educational achievements, see Jonsson and Gähler 1997). However, this may be offset if better educated fathers tend to be more involved with their children after divorce, as suggested in a recent Norwegian study (Lyngstad et al. 2015).

In principle, parents' education may condition the effect of divorce on children's well-being partly by affecting repartnership. However, a precise picture of educational gradients in repartnership has not emerged so far, and it is unclear how parental repartnering affects children's well-being (Sweeney 2010).

Turning again to the children's *age* at divorce, the modifying effect of parental education would depend on whether the disadvantages resulting from divorce are generally influenced by education (so that the implications of having divorce disadvantages over a longer part of the childhood depend on education; the 'duration of childhood disadvantage' argument)and whether education affects the link between age at divorce and the intensity and character of these disadvantages (the other main argument). It is, of course, difficult to make specific predictions about the modifying effects from these general ideas, and the existing literature provides little help. Earlier studies of how age at parental divorce affects mental health have not considered interactions with parental education.

Data

The core data source for this analysis was the Norwegian Population Register, which includes everyone who has lived in Norway at any time after 1964. For each individual, the following information from the 2008 and older versions of the register was extracted: personal identification number (PIN), year of birth, years of death, emigration, or immigration (if applicable), marital status at the beginning of every year since 1970, PIN of spouse (if married, from 1975), and PINs of parents (almost complete for those born after about 1953). All information needed about the reproductive biographies of mothers, that is, the years when their children were born and the PINs of their co-parents, could be derived from these variables.

The educational level attained as of each year since 1980 was added from the Educational Database operated by Statistics Norway, and purchases of medicine were added from the Norwegian Prescription Database (NorPD) (Furu et al. 2010). NorPD was started in 2004 and covers all purchases of prescription medicine by Norwegian residents, except individuals living in institutions, of whom there are very few in the age groups considered in this study. These purchases are reported to NorPD by pharmacies. Only prescription data up to 2008 were included in the data file available for this analysis. Permission to establish and use this data file was obtained from the data owners, the Regional Committees for Medical and Health Research Ethics, and the Norwegian Data Protection Authority.

In accordance with the definitions of Kuo et al. (2011), medicines with Anatomical Therapeutic Chemical (ATC) codes N06AA, N06AB, N06AF, N06AG, and N06AX (except N06AX01 and N06AX02) were considered antidepressants. The only way to purchase such medicines in Norway is to first obtain a prescription from a doctor.

In a first step, 1,950,042 individuals born in 1960–84 (and so aged 20–44 years old in 2004) were selected. Then, this sample was further restricted to the 1,193,610 who lived in Norway at the beginning of each of the years 2004–09, whose father was identified, and whose mother was born in 1935 or later. The reason for the exclusion of individuals with mothers from older cohorts was that there was incomplete information about the birth histories for these women. Subsequently, the 68,353 individuals (5.7 per cent) whose mother was not reported as married at 1 January the year after the birth (and thus presumably to the child's father), or as married to the child's father some time afterwards, were excluded. This was because the data allowed identification of the timing of only marital dissolutions, not the breakdown of cohabiting relationships. Some of these excluded individuals had parents who never married (they may have lived in a consensual union and, in some cases, it is obvious that a disruption must have taken place, but its timing is unknown). The other excluded individuals were those whose parents divorced before 1975 (when there was no information about spousal identities) and who were themselves either born before 1970 (when there was no information on marital status) or had parents who were not yet married the year after birth. The relative sizes of these excluded groups are unknown.

The remaining 1,125,257 were grouped into three categories: (i) those who did not experience separation, divorce, or death of a parent before 2004 (743,849); (ii) those who experienced parental separation or divorce before 2004, and whose parents were still alive in 2004 (242,887); and (iii) all others (138,521). Note that 'separation' refers to formal separation, not the time when the partners actually moved apart. After further exclusion of individuals whose mother had experienced an earlier disruption, as judged from their mother either having had a child with another man before their birth or having divorced or separated, the first two groups included 726,322 and 230,116 individuals, respectively.

Finally, those without at least one full sibling were excluded (although they would automatically drop out of the estimation of the fixed effect models). Thus, the analysis was based on sets of two or more full siblings, none or all of whom had experienced parental separation or divorce. These two groups included 636,294 and 181,241 individuals, respectively (see Table 1). Only the sets of siblings that included at least one who had purchased antidepressant medication and at least one who had not purchased antidepressants contribute to the estimates. There were 148,275 such siblings among those who had not experienced separation or divorce and 54,137 among those who had experienced this event.

(Table 1 about here)

‘Divorce or separation’ is referred to from now on, for simplicity, as ‘divorce’. The time of this ‘divorce’ was set to the first time either of these events was recorded. The most common pattern was for separation to be recorded first and then divorce (92 per cent of the divorces between 1975 and 2003 were preceded by a registered formal separation). The individuals under study are sometimes referred to as ‘children’, even though they were adults when the medication measurements were made.

Methods

The statistical model

In the first part of the analysis, both linear probability models and logistic models were estimated, since both absolute and relative differences in the probability (odds) of purchasing antidepressants are of interest. However, logistic models were not estimated when the importance of interactions and mediators was assessed, as they are not well suited to this (Ai and Norton 2003). A technique that can be used to analyse the importance of mediating variables has been developed for logistic models (Karlson et al. 2012), but cannot be used when the models include fixed effects, such as those here.

To be more specific, the following logistic model for the chance, p_{ij}, that child i of mother j purchased antidepressants at least once during the years 2004–08 was estimated as:

$$\log (p_{ij}/(1-p_{ij})) = b_0 D_j + b_1 A_{ij}^{(5)} D_j + b_2 A_{ij}^{(10)} D_j + b_3 A_{ij}^{(15)} D_j + b_4 A_{ij}^{(20)} D_j + b_5 C_{ij} + \mathbf{b}_6 \mathbf{X}_{ij} + m_j \qquad (1)$$

where $A_{ij}^{(5)}$, $A_{ij}^{(10)}$, $A_{ij}^{(15)}$, and $A_{ij}^{(20)}$ are dummies for whether the child was aged 5–9, 10–14, 15–19, or 20+ years old at the time of divorce (age 0–4 being the reference category). D_j is a divorce indicator ('1' if experienced divorce; otherwise '0'), C_{ij} is the child's birth cohort (in two-year categories), and $\mathbf{X}_{ij}$ is a vector of child characteristics that can vary between siblings (further specified later in this subsection). The 'b's are the corresponding coefficients, and m_j is a mother-level fixed effect. Note that b_0 cannot be estimated, as D_j is either '0' for all siblings or '1' for all siblings, so the term is subsumed into m_j (and included in the equation for pedagogical purposes). The estimation was done using the 'Logistic' procedure in SAS. In the linear probability models, the outcome was a dummy variable for purchase of antidepressants and a child-specific error term was added. These models were estimated with the 'GLM' procedure in SAS.

Using less formal language, outcomes in a fixed period (2004–08) were measured for adult siblings who had experienced parental divorce but at different ages. It is clearly sensible to control for age in 2004, because among a group of siblings, those who were oldest at the time of divorce would also be oldest in 2004, and several studies have shown an increase in depression up to ages in the 40s or 50s (Jorm 2000). In principle, year of birth might also have an impact on the use of antidepressants, above and beyond the effect of age. When birth cohort (which is 2004 minus age in 2004) is included in the model, a combination of the age effect and the possible cohort effect is captured.

However, the effects of the child's age at parental divorce and the child's birth cohort are impossible to separate in a fixed effects model estimated only for children who have experienced divorce, unless we are willing to make some bold assumptions about the functional forms. This is a because of a linear dependence: a child's birth cohort plus a child's age at parental divorce equals the calendar year of divorce, which is the same for all siblings and whose effect can be seen as part of the fixed effect. Thus, with no essential change to the model, an arbitrary linear trend may be added both to the effect of the child's birth cohort and the effect of the child's age at divorce (see discussion of a similar situation by Keiding and Andersen 2016). In this study, the problem was handled by also including the children who did *not* experience parental divorce. They contributed to the estimation of the cohort effect, but only indirectly through that to the effect of age at divorce.

The models also include birth order and sex (in $\mathbf{X}_{ij}$). Females have a higher prevalence of depression than males (Kessler and Bromet 2013), and results from an analysis of suicide suggest that higher birth order might increase the chance of depression (Rostila et al. 2014), while an analysis of mental health among young children has pointed in the opposite direction (Grinde and Tambs 2016). Furthermore, among siblings, birth order is clearly linked to age at parents' divorce, but not caused by it. Due to the lack of information about month of birth, siblings born in the same calendar year, most of whom were probably twins, were assigned the same birth order (and only contributed to the estimation of effects of birth order, birth year, and age at parental divorce if they had younger or older siblings, with whom they were then compared).

Mediators and interactions

To carry out a simple analysis of mediating factors, children's education and indicators of their own family formation behaviour (marital status, parenthood, and interactions between these two factors) were added to the model.

The last step was to consider interactions with parental divorce. In one model, a dummy indicator, G_{ij}, for whether the child is a girl, was interacted with divorce, age at divorce, and birth cohort (an important control variable), but not birth order (which turned out to be much less important to control for). Thus, the following terms were added to the linear probability version of equation (1):

$$b_7 D_j G_{ij} + b_8 A_{ij}^{(5)} D_j G_{ij} + b_9 A_{ij}^{(10)} D_j G_{ij} + b_{10} A_{ij}^{(15)} D_j G_{ij} + b_{11} A_{ij}^{(20)} D_j G_{ij} + b_{12} C_{ij} G_{ij}.$$

In a second model, the following interactions with whether the mother had low education (E_j = '1' if low; otherwise '0') were added to the linear probability version of equation (1):

$$b_7 A_{ij}^{(5)} D_j E_j + b_8 A_{ij}^{(10)} D_j E_j + b_9 A_{ij}^{(15)} D_j E_j + b_{10} A_{ij}^{(20)} D_j E_j + b_{11} C_{ij} E_j.$$

Note that the analogue of $D_j G_{ij}$ could not be included, as maternal education, unlike offspring sex, does not vary between siblings. For the same reason, the main effect of Ej could, of course, not be included.

In a third model, father's education was included as an interacting variable instead of mother's education. Finally, a model with all these interactions was estimated, as well as models including a three-way interaction (see in 'Results' section for details).

Results

Distributions over the outcome variable (purchase of antidepressants) and the independent variables are shown in Table 1 for those with at least one sibling (left hand side) and for the subgroup where at least one sibling purchased antidepressants and at least one did not (right hand side). In the former group, 11.7 per cent had purchased antidepressants, while this proportion, of course, was higher (41.5 per cent) in the latter group, which was constructed by conditioning on such purchases.

Estimates from the sibling fixed effects logistic model show that the chance (odds) of purchasing antidepressants between 2004 and 2008, among women and men aged 20–44 in 2004, was 19 per cent lower for those who had experienced parental divorce after age 20 than for those who had experienced it before they were five years old (Table 2, Panel A, left hand column). The reduction was 5, 6, or 12 per cent for those who had experienced parental divorce at ages 5–9, 10–14, or 15–19, respectively. However, only the differences between the lowest and the two highest age groups (15–19 and 20+) were significant at the 5 per cent level or higher.

(Table 2 about here)

Being born in an earlier cohort (higher age in 2004) and being female raised the chance of antidepressant usage (i.e., ‘positive effects’ of these variables) as expected, and the highest usage of such medicines among those of the highest birth order accords with the pattern reported in some, but not all earlier studies (see ‘Methods’ section). However, evidence about a relationship between birth order and mental health based on sibling models has been lacking. Experimentation with alternative specifications showed that the inclusion of sex had no impact on the estimated effects of the child’s age at parental divorce, while inclusion of birth order made them slightly weaker.

Linear probability models gave similar results (Table 2, Panel B, left hand column). For example, the coefficient for age 20+ was −0.0169 and significant at the 5 per cent level. This accords quite well with a 19 per cent reduction in the odds when the overall probability of the outcome is around 12 per cent. The coefficient for age 15–19 was −0.0091 and not significant (p=0.12). In other words, the chance of purchasing antidepressants was reduced by about 1–2 percentage points for those who were older teenagers or aged 20+ when their parents divorced.

In comparison, when the sibling fixed effects were excluded, older age at parental divorce was associated with a much larger reduction in purchases of antidepressants (Table 2, Panels A and B, right hand sides). Adding mother’s education to these ‘naïve’ models reduced the effect, as we would expect, but not much. For example, the coefficients for the two highest age groups in the logistic models increased by 0.02 (not shown). If several other observed family- or community-level factors shared between siblings had also been added, the effects could, of course, have come even closer to those obtained with the fixed effects approach. (Using the 'naïve' models, individuals without siblings can also be included and contribute to the estimates. The results remained almost unchanged when this was done.)

When the child's education was added to the linear probability model including sibling fixed effects, the coefficients for the two oldest age groups were reduced by about one-third and were no longer significant (Table 3, second column of data). This reflects the fact that an older age at parental divorce has a positive effect on educational attainment, which in turn lowers the chance of later depression. However, inclusion of the child's own marital status and parenthood status, and the interaction between these two factors, *amplified* the effects of age at divorce (Table 3, third column of data). The reason is that an older age at parental divorce increases the chance of never marrying (not shown), which is positively related to depression. It also increases the chance of becoming divorced oneself (not shown), which has a similar impact on depression. Effects of widowhood on depression are less interesting, as very few of these relatively young people were widowed, but the large positive estimate is highly reasonable. Parenthood appeared to reduce the chance of later depression only among this small group of widowed individuals, while it had an opposite effect for the divorced. When both education and family indicators were added, effects of age at parental divorce were quite similar to those appearing without any of these variables included (Table 3, fourth column of data).

(Table 3 about here)

Sensitivity tests

Some sensitivity tests were carried out (not shown). These showed, for example, that the effects of children's age at parental divorce were almost identical if one-year instead of two-year categories were used for birth cohort (an important control variable), but were slightly weaker if five-year categories were used. Also, the inclusion of an indicator for twins was unimportant, and if twins were left out of the analysis, the results changed little.

Furthermore, when the 1985–89 cohorts (aged 15–19 in 2004) were added, using otherwise the same restrictions used when constructing the sample, the effects of children's age at parental divorce were slightly weaker. Adding the 1950–59 cohorts as well expanded the sample by only about 5 per cent and the key effects were essentially unchanged. (Recall that the analysis had to be restricted to children of mothers born no earlier than 1935, and such individuals could not have been born before 1950.)

Finally, three additional family variables were included in the analysis of mediators: (i) whether the child was married at the beginning or end of the year when their own first child was born; and whether they had (ii) been married or (iii) had a first child before age 24 (women) or 26 (men). This had almost no impact on the results.

Interactions

There was a strong interaction with offspring's sex, regardless of whether interactions with parents' education were also included: an older (rather than younger) age at parental divorce had a significantly more negative effect on antidepressant purchase for females than for males (Table 4, Model 1). In fact, the main effects of age at parental divorce in the models where there were only interactions with sex—and which can be interpreted as the effect of age at parental divorce among males—were not significant. There was only a weak indication of relatively high chance of antidepressant purchase among boys who were 10–14 years old when their parents divorced. Antidepressant purchases decreased monotonically with increasing age at parental divorce among females (seen by adding main and interaction effects).

Furthermore, there was a significant negative interaction between sex and the 'general' effect of divorce (Model 1). In other words, we know at least that the effect on antidepressant purchase of divorce at age 0–4 compared with not having experienced divorce (which cannot be estimated) is less positive or more negative for females than for males. This difference between the sexes is largest among the children who are oldest at the time of divorce, given the larger decline in the chance of purchasing antidepressants with age at divorce among females.

(Table 4 about here)

There were also quite clear interactive patterns with respect to parental education. More specifically, the effect of the child's age at parental divorce on depression was significantly more negative among children whose mother had relatively low education than for those whose mother had a higher level of education (Table 4, Model 2). The effects for those with better educated mothers (i.e., the main effects of the child's age at divorce in the model with no other interactions) were not significant. In contrast, there were no significant interactions with the *father's* education (Model 3). The same patterns appeared when *all* three interactions were included in the model (Model 4).

According to the point estimates of a three-way interaction between maternal and paternal education and age at parental divorce, having a mother with low education mattered less for the effect of age at divorce when the father also had low education (Appendix Table A1). However, this interaction was far from significant. A similar analysis with a three-way interaction involving offspring's sex and parental education showed that the interaction between offspring's sex and age at divorce did not depend on the mother's or father's education (not shown in tables).

Inclusion of offspring's own education weakened the interaction between offspring sex and age at parental divorce (as judged from the point estimates in models where the effect of offspring education was allowed to vary with offspring sex; see Appendix Table A2, Model 2). In other words, the more negative relationship between age at divorce and depression for females than for males is partly a result of a stronger relationship between age at divorce and own education for females, or a more negative effect of own education on depression. Similarly, the more negative relationship between age at divorce and depression among those whose mother had low education is partly linked to offspring's own education. However, none of these differences in point estimates were larger than 0.009 and on average they were only around 0.005. The addition of family formation variables left the interaction effects almost unchanged (Appendix Table A2, Model 3).

Summary and conclusion

The key message from this analysis is that a child's age at parental divorce seems to have long-term implications for mental health, as indicated by purchases of antidepressants in early and mid-adulthood. Those who were already aged 20+ when their parents divorced had a 19 per cent lower chance of purchasing antidepressants than those aged under five at the divorce. In terms of absolute differences, the chance was 1.7 percentage points lower. This difference was about one-third of the difference between individuals with primary education and those with upper secondary education, or between individuals with upper secondary education and those with a master's degree (Table 3). Effects of this size can only be revealed in investigations including many observations (as opposed to the smaller studies carried out earlier by other researchers).

This variation across children's age at parental divorce may reflect the youngest children having experienced disadvantages, such as post-divorce conflict between the parents, less parental attention, and fewer economic resources, over a longer part of their childhoods ('duration of childhood disadvantage' argument). It is also possible that parental divorce at a low age leads to particularly large disadvantages during the remaining childhood, or leads to types of disadvantages with strong implications for later life (version 1 of the argument about 'age at divorce affecting disadvantage intensity and character').

The estimates are consistent with an idea that children's age at parental divorce influences their chance of depression in early and mid-adulthood partly because it affects their educational achievements, which in turn have consequences for depression, and partly through more direct routes. This would fit with some of the (somewhat mixed) findings in earlier studies of the association between children's age at parental divorce and educational outcomes (Sigle-Rushton et al., 2014; Steele et al. 2009). However, the causal pathways involving education may also be more complex: some children may develop depression at a young age—as a result of divorce or for other reasons—which may last at least until the time of measurement, or there may be a recurrence at that time. This early depression may, along with the parental divorce, have implications for educational achievements, which in turn have implications for later depression. The data do not allow exploration of such pathways.

While we might expect that the child's own family formation behaviour accounts for some of the association, evidence of this did not appear. On the contrary, the effect of children's age at parental divorce would have been stronger if those experiencing the divorce at an older age had not also had a lower chance of marrying and remaining married. This result might appear surprising in light of existing literature on intergenerational transmission of divorce and how parental divorce affects own family formation, but these studies have not considered the importance of the child's age at divorce (Dronkers and Härkönen, 2008; Reneflot, 2009). In any case, the relationship between age at parental divorce and depression changed very little when both education and the family formation indicators were included.

The results suggest that parents, or others with responsibility for supporting children who have experienced parental divorce, should perhaps give special priority to helping children who have had this experience at a young age. However, to give more specific advice about this, we would need knowledge about *why* parental divorce at a young age has particularly adverse implications for the children's mental health, which this study has not provided.

One should not conclude from the estimates that delaying a divorce until the children are older would generally be beneficial. A difference in depression between, for example, children who were 10–14 vs. 15–19 years old at the time of their parents' divorce—estimated in the sibling models controlling for cohort, sex, and birth order—does not tell us what the outcome (measured in 2004–08) would have been for the former if the parents had delayed the divorce for five years. The main reason for this is that a divorce (or separation) typically happens after a period when the parents have had poor relationship quality, which likely affects the children negatively. A five-year delay may prolong this exposure to parental discord, and the children would then not be in the same situation as the siblings aged 15–19 with whom they are being compared. In other words, if the sibling models suggest an advantage from a delayed divorce, the actual benefit is probably smaller than this—although we cannot know by how much—and there could even be a disadvantage. Such a conclusion about delayed divorce is only justified in the hypothetical case where it is only the divorce itself that has an impact on child well-being, or if the parental relationship was already poor from the time they had their first child, so that all children experienced a low-quality parental relationship through their entire life.

Interactive effects

There appears to be considerable variation in the effect of a child's age at parental divorce on later depression across the sexes. Women generally use more antidepressants than men, but there is a smaller difference after divorce according to this analysis, and usage declines more with increasing age at parental divorce. As explained earlier, it is far from evident theoretically what kind of pattern one should have expected, and results from previous studies of sex differences in the association between children's divorce experiences and their later depression—none of which have considered the age at the divorce—have been mixed. One possible explanation of the differential age pattern is that if the effects of divorce are generally more harmful for girls, then living a larger part of childhood in a one-parent family has larger consequences for them ('duration of childhood disadvantage' argument). However, that may not accord well with the generally weaker effects of divorce seen for girls. Alternatively, children's sex may have a modifying impact (only) through the other main argument (about 'age at divorce affecting disadvantage intensity and character'; version 2): Reduced contact with the father—which often happens after parental disruption—may have particularly adverse consequences for boys if this happens at a relatively older age.

There is also an interaction with the parents' education: the advantages associated with older age at parental divorce are particularly pronounced among those whose mothers have relatively low education, while there is no such modifying effect of father's education. One possible explanation could be that maternal education generally weakens the disadvantages resulting from divorce ('duration of childhood disadvantage' argument). Perhaps it also dampens the specific problems caused by divorces that occur at a young age (version 1 of the other main argument). In principle, maternal education may also *aggravate* the disadvantages produced in particular by divorces at higher ages, but this seems less plausible. In any case, the differences between the modifying effects of maternal and paternal education accord with an earlier study of parental divorce and child well-being (Mandemakers and Kalmijn 2014; which did not consider the child's age), and make sense in the light of mothers' stronger involvement with children after divorce. (Unfortunately, the available data did not include information about children's living arrangements.)

Limitations

There are four main weaknesses of this study, in addition to the lack of data on the quality of the parental relationship already alluded to. One is that the proportion who purchase medication may not adequately reflect the prevalence of depression. Rather, purchases reflect a combination of being depressed, going to the doctor, and getting a medication-based treatment instead of another kind of treatment or no treatment. According to Swedish and Finnish studies, only about one-quarter of those classified as depressed based on interviews used antidepressants, although the proportion was larger among the subgroup with the most severe depression (Henriksson et al. 2006; Hämäläinen et al. 2009). In fact, there is probably underuse of all kinds of treatment (so studies based on healthcare usage would have similar limitations). Some authors have concluded that only half of depressed people receive any kind of acceptable treatment (Hämäläinen et al. 2009) and even among the severely depressed only two-thirds do (Shim et al. 2011). On a more positive note, 12 per cent of the relatively young adults included in the present study purchased antidepressants during a five-year period, so given the lifetime prevalences of about 15–20 per cent according to clinical interviews in surveys from Norway and other countries (Mykletun et al. 2009), a relatively high proportion of those with depression may actually have been included.

In any case, our main concern should be whether the relationship between the child's age at parental divorce and the actual prevalence of depression is different from that suggested by the estimates. In theory, this could be the case if those who experience parental divorce at an early age are more or less likely than their older siblings to go to the doctor if they are depressed, or to get medical rather than other types of treatment. However, this seems implausible and there is no evidence to support such an idea.

A related concern is that the available data provide very little information about the severity of the depression. One step that can be taken to get an impression of the effects on severe vs. less severe types is to consider the number of years in which antidepressants were purchased. In supplementary analysis, two outcome variables were considered (see Appendix Table A3, Panel A). The first was whether antidepressants had been purchased in only one of the years 2005–07, but not in either 2004 or 2008. Such a pattern would mean that purchases had definitely not been made in two consecutive years, which may signal a short episode of depression, perhaps as a result of acute stress. The second outcome was whether antidepressants had been purchased in at least two of the years from 2004 to 2008. Age at parental divorce had a significant effect on the second outcome, while an effect was only weakly indicated for the first outcome ($p=0.11$ for the highest age group). This points towards the possibility that age at divorce is important primarily for the development of more severe episodes of depression. On the other hand, according to models estimated for individuals who had made at least one purchase, there was no effect of age at divorce on the number of years in which purchases were made (not shown in tables).

A second limitation of the study is that antidepressants are sometimes used for conditions other than depression, for example anxiety. According to some studies, this may be the case for as many as one-third of antidepressant prescriptions (Gardarsdottir et al. 2007; Milea et al. 2010). Supplementary analysis showed that there was no significant relationship between the child's age at parental divorce and purchases of anxiolytics (drugs to relieve anxiety) (Appendix Table A3, Panel B). If the association between age at divorce and the use of antidepressants for anxiety is similar to that between age at divorce and the use of anxiolytics, then age at divorce will be more strongly associated with the use of antidepressants for depression than with the total use of these medicines (i.e., the association estimated in this study).

Third, it was only possible to study disruption of formal marriages and, as mentioned, an increasing proportion of Norwegian children have been born outside marriage, largely to cohabiting parents, whose relationships tend to be less stable. However, less than 6 per cent of children in the selected birth cohorts (who satisfied the requirements about residence and maternal birth cohort) were excluded because their parents never married. Also, there are no obvious reasons why the effect of child's age when a parental cohabiting union is dissolved should be different from the effect of age when a parental divorce takes place. Ideally, future Norwegian studies of this subject should include children in consensual unions.

Finally, one should keep in mind that the focus is on the timing of the formal separation or (if a separation had not been recorded earlier) the divorce. The actual disruption may have taken place several months, or even years, before that. That said, the timing of the actual disruption may not be the most crucial factor either, as there has usually been an earlier period with marital discord, which may have affected the child's well-being as much as or more than the disruption itself.

Conclusion

Although the study has certain limitations, it also has clear strengths. The latter include the use of a large data set with objective measurements of treatment for depression. Additionally, the study is based on a method taking shared family characteristics into account, which has only been used in one earlier study of how age at parental divorce may affect children's chances of later depression. Controlling for these characteristics, as well as some child-specific characteristics, a negative association between age at parental divorce and later depression was found. This suggests that the youngest children of divorce may need special help to cope with the new situation. According to models including interactions, this is especially the case among girls and among children of mothers with low levels of education. However, while those who experience parental divorce at a relatively older age fare better than those who experience it at a younger age, one cannot conclude that it would be helpful to delay the dissolution. A delay could be an advantage, but could also be a disadvantage, depending on the implications of exposing the child to a prolonged period under the same roof as two parents with a poor relationship. This is an issue that needs further exploration, using data with more detailed information on relationship quality than available here.

Notes and acknowledgements

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Table 1 Number of individuals by whether they have experienced parental divorce, age at parental divorce, sex, birth order, birth cohort, and whether they purchased antidepressants in 2004–08: Norwegian women and men aged 20–44 in 2004 with at least one sibling in this age group

			Subgroup where at least one sibling purchased antidepressants and at least one did not	
	Have experienced parental divorce	Have not experienced parental divorce	Have experienced parental divorce	Have not experienced parental divorce
Age at parental divorce				
0–4	14,461	–	4,368	–
5–9	32,217	–	9,518	–
10–14	39,698	–	11,898	–
15–19	43,182	–	12,655	–
20+	51,683	–	15,698	–
Not experienced divorce	–	636,294	–	148,275
Sex				
Men	92,060	325,272	25,969	70,391
Women	89,181	311,022	28,168	77,884
Birth order				
1	73,310	239,114	20,016	50,041
2	74,489	246,936	20,559	52,458
3	25,188	107,266	9,431	29,811
4+	8,254	42,978	4,131	15,965
Birth cohort				
1960–61	4,182	21,803	1,703	6,671
1962–63	6,320	32,295	2,684	9,749
1964–65	10,404	44,348	3,846	12,829
1966–67	14,128	54,693	4,975	15,118
1968–69	18,703	63,386	6,122	16,579
1970–71	19,722	66,398	6,200	16,335
1972–73	20,338	67,066	6,017	15,327
1974–75	19,218	63,643	5,537	13,618
1976–77	17,671	57,355	4,632	11,803
1978–79	17,271	56,080	4,471	10,739
1980–81	15,755	50,973	3,810	9,184
1982–83	12,214	40,223	2,907	7,191
1984	5,315	18,031	1,233	3,132
Purchases of antidepressants				
No	153,125	569,124	30,490	87,958
Yes	28,116	67,170	23,647	60,317
Total	181,241	636,294	54,137	148,275

Source: Authors' calculations based on Norwegian Population Register and NorPD.

Table 2 Effects of age at parental divorce, sex, birth order, and birth cohort on the probability of purchasing antidepressants at least once during 2004–08: Norwegian women and men aged 20–44 in 2004

Panel A: logistic models (with 95 per cent confidence intervals)

	Model including sibling fixed effects	Model not including sibling fixed effects[a]
Age at parental divorce		
0–4 (ref)	1	1
5–9	0.95 (0.88–1.03)	0.88 (0.84–0.93)***
10–14	0.94 (0.85–1.03)	0.82 (0.78–0.86)***
15–19	0.88 (0.79–0.98)*	0.76 (0.72–0.80)***
20+	0.81 (0.72–0.92)***	0.68 (0.64–0.71)***
Not experienced divorce	–	0.49 (0.47–0.51)***
Sex		
Men (ref)	1	1
Women	1.71 (1.68–1.74)***	1.68 (1.65–1.70)***
Birth order		
1 (ref)	1	1
2	1.06 (1.04–1.09)***	1.08 (1.06–1.09)***
3	1.10 (1.06–1.15)***	1.14 (1.12–1.17)***
4+	1.12 (1.05–1.20)***	1.23 (1.20–1.27)***
Birth cohort		
1960–61	1.60 (1.42–1.80)***	1.79 (1.72–1.87)***
1962–63	1.59 (1.44–1.77)***	1.75 (1.68–1.82)***
1964–65	1.53 (1.40–1.68)***	1.62 (1.57–1.68)***
1966–67	1.46 (1.35–1.59)***	1.53 (1.48–1.58)***
1968–69	1.33 (1.23–1.43)***	1.41 (1.37–1.46)***
1970–71	1.30 (1.22–1.39)***	1.36 (1.31–1.40)***
1972–73	1.21 (1.14–1.29)***	1.25 (1.21–1.29)***
1974–75	1.18 (1.12–1.25)***	1.20 (1.16–1.24)***
1976–77	1.14 (1.09–1.20)***	1.15 (1.11–1.19)***
1978–79	1.08 (1.03–1.14)***	1.09 (1.06–1.13)***
1980–81 (ref)	1	1
1982–83	0.91 (0.87–0.96)***	0.92 (0.89–0.96)***
1984	0.82 (0.76–0.88)***	0.84 (0.80–0.89)***

Panel B: linear probability models (with standard errors)

	Model including sibling fixed effects	Model not including sibling fixed effects[a]
Age at parental divorce		
0–4 (ref)	0	0
5–9	−0.0042 (0.0044)	−0.0164 (0.0032)***
10–14	−0.0028 (0.0052)	−0.0245 (0.0031)***
15–19	−0.0091 (0.0059)	−0.0336 (0.0031)***
20+	−0.0169 (0.0067)*	−0.0459 (0.0030)***
Not experienced divorce	–	−0.0813 (0.0027)***
Sex		
Men (ref)	0	0
Women	0.0508 (0.0009)***	0.0522 (0.0007)***
Birth order		
1 (ref)	0	0
2	0.0064 (0.0011)***	0.0075 (0.0008)***
3	0.0099 (0.0021)***	0.0138 (0.0011)***
4+	0.0112 (0.0033)***	0.0216 (0.0015)***
Birth cohort		
1960–61	0.0466 (0.0057)***	0.0598 (0.0024)***
1962–63	0.0465 (0.0051)***	0.0572 (0.0020)***
1964–65	0.0422 (0.0045)***	0.0486 (0.0018)***
1966–67	0.0371 (0.0040)***	0.0423 (0.0017)***
1968–69	0.0273 (0.0035)***	0.0335 (0.0017)***
1970–71	0.0253 (0.0031)***	0.0293 (0.0016)***
1972–73	0.0184 (0.0028)***	0.0209 (0.0016)***
1974–75	0.0150 (0.0025)***	0.0168 (0.0017)***
1976–77	0.0113 (0.0023)***	0.0124 (0.0017)***
1978–79	0.0072 (0.0021)***	0.0081 (0.0017)***
1980–81 (ref)	0	0
1982–83	−0.0069 (0.0023)**	−0.0072 (0.0019)***
1984	−0.0153 (0.0031)***	−0.0148 (0.0024)***

[a]Estimated from the same data as the model including sibling fixed effects (sibling model), that is, data only including individuals who have at least one sibling in the chosen age group (although individuals without a sibling would automatically have dropped out when estimating the sibling model). Almost the same estimates appeared when those without a sibling were also included: the effects of age at divorce in the logistic model were then 0.88, 0.82, 0.76, 0.68, and 0.50, while those in the linear probability model were −0.0163, −0.0232, −0.0325, −0.0435, and −0.0784. As explained in the ‘Methods’ section, those who have not experienced divorce are also included when estimating sibling models. They contribute indirectly to the estimation of the effects of age at parental divorce, but a coefficient for ‘not experienced divorce’ cannot be estimated.

* $p<0.05$; ** $p<0.01$; *** $p<0.001$. Ref is the reference category.

Source: As for Table 1.

Table 3 Effects of age at parental divorce, child's education, marital status, and parenthood on purchases of antidepressants 2004–08, according to linear probability models where the outcome is whether antidepressants was purchased at least once: Norwegian women and men aged 20–44 in 2004[a]

	As Table 2, Panel B, fixed effects model	Control for for child's education	Control for child's marital status and parenthood	Control for child's education, marital status and parenthood
Age at parental divorce				
0–4 (ref)	0	0	0	0
5–9	−0.0042 (0.0044)	−0.0031 (0.0044)	−0.0056 (0.0044)	−0.0044 (0.0044)
10–14	−0.0028 (0.0052)	−0.0011 (0.0052)	−0.0058 (0.0053)	−0.0038 (0.0052)
15–19	−0.0091 (0.0059)	−0.0060 (0.0060)	−0.0130 (0.0060)*	−0.0098 (0.0059)
20+	−0.0169 (0.0067)*	−0.0108 (0.0068)	−0.0216 (0.0067)**	−0.0153 (0.0066)*
Child's education				
Primary	–	0	–	0
Lower secondary	–	−0.0359 (0.0022)***	–	−0.0352 (0.0022)***
Upper secondary	–	−0.0610 (0.0013)***	–	−0.0598 (0.0013)***
Some tertiary	–	−0.0881 (0.0015)***	–	−0.0862 (0.0015)***
Master's degree	–	−0.1131 (0.0023)***	–	−0.1104 (0.0023)***
Child's family formation				
Never married	–	–	0.0232 (0.0024)***	0.0202 (0.0024)***
Married (ref)	–	–	0	0
Widowed	–	–	0.1580 (0.0394)***	0.1507 (0.0393)***
Divorced	–	–	0.0394 (0.0059)***	0.0338 (0.0059)***
Childless (ref)	–	–	0	0
Parent	–	–	0.0013 (0.0025)	−0.0026 (0.0025)
Never married × parent	–	–	0.0021 (0.0027)	0.0019 (0.0027)
Widowed × parent	–	–	−0.0881 (0.0418)*	−0.0841 (0.0418)*
Divorced × parent	–	–	0.0281 (0.0063)***	0.0284 (0.0062)***

[a]The models also include birth cohort, sex, and birth order.

* p<0.05; ** p<0.01; ***p<0.001. Ref is the reference category. Standard errors in parentheses.

Source: Authors' calculations based on Norwegian Population Register, Statistics Norway Educational Database, and NorPD.

Table 4 Effects of age at parental divorce and interactions between age at divorce and other factors on purchases of antidepressants 2004–08, according to linear probability models where the outcome is whether antidepressants were purchased at least once: Norwegian women and men aged 20–44 in 2004[a]

	Model 1[b]	Model 2[c]	Model 3[d]	Model 4[e]
Age at parental divorce				
0–4 (ref)	0	0	0	0
5–9	0.0006 (0.0060)	−0.0033 (0.0070)	−0.0070 (0.0064)	−0.0002 (0.0088)
10–14	0.0115 (0.0066)	0.0110 (0.0085)	−0.0056 (0.0077)	0.0210 (0.0102)*
15–19	0.0005 (0.0071)	0.0085 (0.0096)	−0.0110 (0.0087)	0.0143 (0.0114)
20+	−0.0042 (0.0077)	−0.0001 (0.0108)	−0.0191 (0.0098)	0.0084 (0.0126)
Additional effect of age at parental divorce if <u>child is girl</u>				
0–4 (ref)	0	–	–	0
5–9	−0.0102 (0.0084)	–	–	−0.0102 (0.0084)
10–14	−0.0295 (0.0081)***	–	–	−0.0295 (0.0081)***
15–19	−0.0204 (0.0079)*	–	–	−0.0204 (0.0080)*
20+	−0.0263 (0.0079)***	–	–	−0.0263 (0.0079)***
Additional general effect of parental divorce if <u>child is girl</u>	−0.0354 (0.0071)***	–	–	−0.0354 (0.0071)***
Additional effect of age at parental divorce if <u>mother had primary or lower secondary education</u>				
0–4 (ref)	–	0	–	0
5–9	–	−0.0013 (0.0089)	–	−0.0035 (0.0092)
10–14	–	−0.0222 (0.0105)*	–	−0.0257 (0.0111)*
15–19	–	−0.0281 (0.0122)*	–	−0.0316 (0.0126)*
20+	–	−0.0264 (0.0138)	–	−0.0300 (0.0143)
Additional effect of age at parental divorce if <u>father had primary or lower secondary education</u>				
0–4 (ref)	–	–	0	0
5–9	–	–	0.0051 (0.0087)	0.0060 (0.0090)
10–14	–	–	0.0052 (0.0105)	0.0118 (0.0109)
15–19	–	–	0.0034 (0.0119)	0.0112 (0.0123)
20+	–	–	0.0043 (0.0134)	0.0118 (0.0138)

[a]All four models also include birth cohort, sex, and birth order.
[b]Model 1 also includes an interaction between birth cohort and sex.
[c]Model 2 also includes an interaction between birth cohort and whether mother had primary or lower secondary education.
[d]Model 3 also includes an interaction between birth cohort and whether father had primary or lower secondary education.
[e]Model 4 also includes the interactions between birth cohort and sex, between birth cohort and whether mother had primary or lower secondary education, and between birth cohort and whether father had primary or lower secondary education.

* p<0.05; ** p<0.01; ***p<0.001. Ref is the reference category. Standard errors in parentheses.

Source: As for Table 3.

APPENDIX TABLES

Table A1 Effects of age at parental divorce and two- or three-way interactions involving age at divorce and other factors on purchases of antidepressants 2004–08, according to linear probability models where the outcome is whether antidepressants were purchased at least once: Norwegian women and men aged 20–44 in 2004[a]

Age at parental divorce	
0–4 (ref)	0
5–9	−0.0010 (0.0090)
10–14	0.0224 (0.0114)
15–19	0.0173 (0.0128)
20+	0.0102 (0.0142)
Additional effect of age at parental divorce if child is girl	
0–4 (ref)	0
5–9	−0.0102 (0.0084)
10–14	−0.0295 (0.0081)***
15–19	−0.0204 (0.0080)*
20+	−0.0263 (0.0079)**
Additional general effect of parental divorce if child is girl	−0.0355 (0.0071)***
Additional effect of age at parental divorce if mother had primary or lower secondary education	
0–4 (ref)	0
5–9	−0.0020 (0.0127)
10–14	−0.0284 (0.0154)
15–19	−0.0379 (0.0174)*
20+	−0.0333 (0.0196)
Additional effect of age at parental divorce if father had primary or lower secondary education	
0–4 (ref)	0
5–9	0.0080 (0.0144)
10–14	0.0083 (0.0175)
15–19	0.0031 (0.0199)
20+	0.0074 (0.0225)
Additional effect of age at parental divorce if both parents had primary or lower secondary education	
0–4 (ref)	0
5–9	−0.0033 (0.0185)
10–14	0.0056 (0.0223)
15–19	0.0133 (0.0253)
20+	0.0068 (0.0286)

[a]The model also includes birth cohort, sex, birth order, interaction between birth cohort and sex, interaction between birth cohort and whether mother had primary or lower secondary education, interaction between birth cohort and whether father had primary or lower secondary education, and interaction between birth cohort and whether both parents had primary or lower secondary education.

* p<0.05; ** p<0.01; *** p<0.001. Ref is the reference category. Standard errors in parentheses.

Source: As for Table 3.

Table A2 Effects of age at divorce and interactions between age at divorce and other factors on purchases of antidepressants 2004–08, controlling for child's own education and family formation behaviour, according to linear probability models where the outcome is whether antidepressants were purchased at least once: Norwegian women and men aged 20–44 in 2004[a]

	Model 1 (as Model 4, Table 4)	Model 2: controls for child's own education[b]	Model 3: controls for child's own education and family formation behaviour[c]
Age at parental divorce			
0–4 (ref)	0	0	0
5–9	−0.0002 (0.0088)	−0.0023 (0.0088)	−0.0037 (0.0087)
10–14	0.0210 (0.0102)*	0.0170 (0.0102)	0.0143 (0.0102)
15–19	0.0143 (0.0114)	0.0098 (0.0113)	0.0060 (0.0113)
20+	0.0084 (0.0126)	0.0053 (0.0126)	0.0004 (0.0125)
Additional effect of age at parental divorce if <u>child is girl</u>			
0–4 (ref)	0	0	0
5–9	−0.0102 (0.0084)	−0.0070 (0.0083)	−0.0068 (0.0083)
10–14	−0.0295 (0.0081)***	−0.0251 (0.0080)**	−0.0251 (0.0080)**
15–19	−0.0204 (0.0080)*	−0.0143 (0.0080)	−0.0138 (0.0080)
20+	−0.0263 (0.0079)***	−0.0178 (0.0078)*	−0.0174 (0.0078)*
Additional general effect of parental divorce if <u>child is girl</u>	−0.0354 (0.0071)***	−0.0267 (0.0070)***	−0.0245 (0.0070)***
Additional effect of age at parental divorce if <u>mother had primary or lower secondary education</u>			
0–4 (ref)	0	0	0
5–9	−0.0035 (0.0092)	−0.0015 (0.0092)	−0.0015 (0.0092)
10–14	−0.0257 (0.0111)*	−0.0214 (0.0111)	−0.0214 (0.0111)
15–19	−0.0316 (0.0126)*	−0.0270 (0.0126)*	−0.0273 (0.0126)*
20+	−0.0300 (0.0143)*	−0.0252 (0.0142)	−0.0254 (0.0142)
Additional effect of age at parental divorce if <u>father had primary or lower secondary education</u>			
0–4 (ref)	0	0	0
5–9	0.0060 (0.0090)	0.0068 (0.0090)	0.0069 (0.0090)
10–14	0.0118 (0.0109)	0.0134 (0.0108)	0.0134 (0.0108)
15–19	0.0112 (0.0123)	0.0139 (0.0122)	0.0141 (0.0122)
20+	0.0118 (0.0138)	0.0154 (0.0138)	0.0158 (0.0138)

[a]All models also include birth cohort, sex, birth order, interaction between birth cohort and sex, interaction between birth cohort and whether mother had primary or lower secondary education, and interaction between birth cohort and whether father had primary or lower secondary education.

[b]Model 2 also includes the child's education (four dummy variables), interaction between child's education and sex, interaction between child's education and whether mother had primary or lower secondary education, and interaction between child's education and whether father had primary or lower secondary education.
[c]Model 3 includes the same additional variables as Model 2, plus the child's family formation behaviour (seven dummy variables), interaction between child's family formation behaviour and sex, interaction between child's family formation behaviour and whether mother had primary or lower secondary education, and interaction between child's family formation behaviour and whether father had primary or lower secondary education.

*p<0.05; ** p<0.01; *** p<0.001. Ref is the reference category. Standard errors in parentheses.

Source: As for Table 3.

Table A3 Effects of age at parental divorce on the probability of purchasing medication during 2004–08: Norwegian women and men aged 20–44 in 2004[a]

Panel A: Frequency of purchases

	Purchased antidepressants in only one year, 2005–07, but no purchases in 2004 or 2008	Purchased antidepressants in two or more years 2004–08
Age at parental divorce		
0–4 (ref)	0	0
5–9	−0.0025 (0.0022)	−0.0043 (0.0034)
10–14	−0.0025 (0.0026)	−0.0021 (0.0041)
15–19	−0.0037 (0.0029)	−0.0071 (0.0046)
20+	−0.0054 (0.0033)	−0.0113 (0.0052)*

Panel B: Antidepressants vs. anxiolytics

	Purchased antidepressants in at least one year (as Table 2)	Purchased anxiolytics in at least one year
Age at parental divorce		
0–4 (ref)	0	0
5–9	−0.0042 (0.0044)	0.0001 (0.0040)
10–14	−0.0028 (0.0052)	0.0025 (0.0047)
15–19	−0.0091 (0.0059)	0.0042 (0.0053)
20+	−0.0169 (0.0067)*	−0.0012 (0.0061)

[a]The models also include birth cohort, sex, and birth order.

*p<0.05. Ref is the reference category. Standard errors in parentheses.

Source: As for Table 1.