Teenage Daughters as a Cause of Divorce*

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Abstract

Evidence from the U.S. that couples with daughters are more likely to divorce than couples with sons has not been found for other Western countries. Using 1995-2015 Dutch marriage registry data, we show that daughters are associated with higher divorce risks, but only when they are 13 to 18 years old. There are no detectable gender differences before or after those ages. These age-specific findings are at odds with son-preference and selection explanations for differences in divorce risks. Instead, the findings point to explanations which involve family relationship dynamics associated with teenage sons and daughters. We find supporting evidence of relationship explanations in supplemental analyses of Dutch survey data. We also find that teenage daughters are associated with higher divorce in the U.S. in analyses of the Current Population Survey Marriage and Fertility Supplements.

JEL classification: J12, J13, J16

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Research by sociologists (Spanier and Glick 1981; Morgan et al. 1988) and economists (Ananat and Michaels 2008; Bedard and Deschênes 2005; Dahl and Moretti 2008; Mammen 2008) has found that U.S. couples with daughters are more likely to divorce than couples with sons. The results seem to complement evidence from developing countries that preferences for sons alter family behaviour (e.g., Arnold 1997; Ben-Porath and Welch 1976; Das Gupta et al. 2003). However, associations between children's genders and divorce do not appear in all U.S. data (Diekmann and Schmidheiny 2004; Morgan and Pollard 2002; Reichman et al. 2004), and methodological weaknesses affect several studies that do find associations. In addition, studies by Andersson and Woldemicael (2001, Sweden), Diekmann and Schmidheiny (2004, 16 European countries and Canada), Flouri and Malmberg (2010, U.K.), and Leigh (2009, Australia) have not detected associations in other Western countries. Thus, there is a genuine question regarding whether divorce risks of parents in developed countries are affected by gender of their children.

If associations are present, the next logical question is why they exist. Lundberg (2005) differentiates between two general categories of economic explanations. One category involves parents' preferences, such as an overarching preference for sons or a preference among fathers to spend time with sons. The other category involves the constraints parents face, such as higher time or money costs of raising girls, more stressful parent-child interactions with girls, or worse developmental consequences of divorce for boys. Sociologists have pointed to greater father involvement with sons, the reinforcement of gendered specialisation within households, and the effects of daughters on parents' gender-role attitudes (Katzev et al. 1994; Morgan et al. 1988; Raley and Bianchi 2006). Besides the causal mechanisms, selection into live birth might also explain associations between children's gender and divorce if pre-existing relationship conflict and stress differentially affect whether girls or boys will be miscarried (Hamoudi and Nobles 2014). Although the studies of children's gender and divorce have considered these explanations, few have had the necessary data or statistical power to distinguish among them.

We examine how children's gender affects divorce risks in the Netherlands, using administrative data that cover every marriage and registered partnership that existed in that country from 1995 to 2015.[1] The data include more than 2 million marriages, allowing us to estimate effects precisely and consider how effects vary with the children's ages, parities, and conditional on the parents' backgrounds. The administrative data are highly accurate with the exact dates of weddings, births, and divorces. This contrasts with nearly all the previous studies which have relied on retrospective self-reports that are subject to recall errors and other misreporting (Mitchell 2010). It also contrasts with most of the economic studies which could identify parenting relationships or children's parities only approximately (e.g., Ananat and Michaels 2008; Bedard and Deschênes 2005; Dahl and Moretti 2008; Mammen 2008).

Unlike previous studies that have used reports from a single point in time, our data take the form of prospective longitudinal observations of marriage and registered partnership spells. This allows us to estimate event-history models that account for the duration of the marriage or registered partnership, changing divorce risks with children's ages, and right-censoring in spells. Morgan et al. (1988) and Morgan and Pollard (2002) had previously estimated event-history models but with point-in-time, retrospective data.

As with many of the U.S. studies, we find that divorce risks are higher among Dutch couples with girls than those with boys. The gender differences are larger for first-born children than for subsequent children but are evident regardless of parity. The novel finding in our analysis is that the gap in divorce risks only appears when the daughters are 13 to 18 years old; there are no detectable gender differences before or after those ages. This pattern of

[1] For brevity, we use "divorce" to refer to the final, legal dissolution of either a marriage or registered partnership.

results is inconsistent with selection into live birth, general preferences for sons, and other causal mechanisms which operate with foreseeable differences in preferences or costs. Instead, the results are consistent with unexpected changes of constraints or family processes, possibly involving relationship strains with teenage daughters.

Additional analyses buttress this explanation. We find that the gender gap depends on the father's sibship. The disparity in divorce risks is substantial for fathers who have no sisters, but is absent for fathers with sisters. This suggests that men who experienced more mixed gender relationships when growing up may be better prepared for similar relationships in their own households. In addition, we find that the gender gap is wider among earlier cohorts of parents and among parents with dissimilar immigration backgrounds; for both of these groups, differences in gender-role attitudes between parents and daughters or between mothers and fathers may contribute to relationship strains. We analyse the Longitudinal Internet Studies for the Social Sciences (LISS) panel, a large socio-economic survey of Dutch households, and uncover direct evidence of relationship strains between fathers and teenage daughters. Fathers of teenage daughters report worse relationships with their families and more parenting disagreements with their partners than fathers of teenage sons, and teenage daughters report worse relationships with their fathers.

Finally, we re-examine the 1980, 1985, 1990 and 1995 Current Population Survey Marriage and Fertility Supplements that have been analysed in U.S. studies. We corroborate evidence of higher divorce risks for parents of teenage daughters, though the estimates are less precise than the Dutch results, owing to the smaller samples of the U.S. surveys.

Theory

We use Weiss' (1997) rational-choice model of marriage and divorce to consider differences in how sons or daughters might affect marital stability (see also Becker et al. 1977). Weiss theorised that couples initially choose to wed and subsequently choose period-by-period whether to remain married by comparing the expected value of being married to the value of being single or in another relationship. The benefits within marriage include the enjoyment of time that a married couple can spend together; the extra consumer goods that they might be able to purchase or produce because of household efficiencies, economies of scale, and specialisation (Becker 1985); and special goods, such as the number and quality of children, that their marriage might produce. Weiss showed that the net benefits of marriage depend on characteristics of each spouse, the marriage-specific goods (e.g., children), and a match-specific quality. In each period, the couple knows the current values of these variables but does not know all the future values. Divorce occurs when one or more variables suffers a sufficiently negative unexpected realisation, such as a drop in match-specific quality, a loss of earning power, or a change in outside opportunities.

As mentioned, Lundberg (2005) proposed preference- and constraint-based explanations for the possible effects of children's gender on divorce. In the context of Weiss' model, children are a type of marriage-specific capital, and an overarching preference for boys over girls would raise the value of this capital more when a boy is born. An overarching son preference might also lead parents to invest more in boys' development, which would increase the value of the couple's marriage-specific capital. Either effect would increase the benefits of marriage and reduce the chances of divorce. Similar effects occur if fathers more strongly prefer spending time with sons than daughters (Lundberg et al. 2007; Mammen 2011), as this would raise their valuations of marriage-specific capital and their incentives to invest in it. Lundberg (2005) and Raley and Bianchi (2006) summarise evidence that fathers' preferences for sharing activities with sons are stronger when the boys are school-aged. This would increase the value of marriage-specific capital and strengthen marriages at those ages, but it would also strengthen marriages with sons at earlier ages because the benefits would be anticipated and fathers would need to remain in the marriage to realise them.

Lundberg (2005) discussed another set of explanations that involve differences in the costs and constraints of raising sons and daughters. Daughters might be more expensive to raise or require more time than boys. Baker and Milligan (2016) found that parents spend more time in teaching activities for pre-school girls than boys. Durante et al. (2015) conducted experiments in which people allocated more resources to daughters than sons, and Moffitt and Ribar (2016) found that disadvantaged families were more protective of daughters' food needs. Other things held equal, higher costs of daughters would reduce the amount of marriage-specific capital that a couple could produce and weaken the incentives to remain married. Another possibility is that parent-child interactions are more stressful with girls than boys. VanderValk et al. (2007) found that adolescent girls' emotional problems strained their parents' marriages but boys' problems did not. A different constraint involves the possibility that boys are more susceptible to developmental problems if parents divorce, which would lower the value of parents' alternatives to marriage. Lundberg (2005) and Raley and Bianchi (2006) summarised evidence of increased vulnerability, but a newer review by Amato (2010) suggests that the results might not apply to recent cohorts of children.

As with the preference explanations, age-related differences in costs and constraints could lead to age-specific divorce patterns. Higher relative costs of very young girls or greater vulnerability among young boys could raise the divorce risks for families with girls at those ages. Similar reasoning applies for families with older children, although disparities among older children may have further-reaching consequences. Within the rational-choice framework, foreseeable differences in costs or vulnerabilities enter the valuation of marriage at the point when the gender is revealed, and so the differences at older ages would affect divorce risks not only during that period but also earlier in the child's life. An isolated pattern of elevated divorce risks which is not preceded by a gradual build-up would have to be attributed to unexpected shocks to the net valuation of marriage or to myopic behaviour.

Sociologists have offered other explanations that involve family processes. One possibility is that fathers' preferences for spending time with sons increase their overall household involvement, which might improve a couple's communication, increase the time a couple shares, and strengthen the solidarity of their marriage (Katzev et al. 1994; Morgan et al. 1988). These processes could lead to gender differences in divorce risks that vary with children's ages if fathers are more involved at some ages but not others. Another possibility is that the presence of sons or daughters changes parents' preferences and attitudes. The presence of boys may reinforce traditional gender-role or family attitudes (Baxter et al. 2016; Katzev et al. 1994; Morgan et al. 1988), which might directly strengthen marriages or increase gender specialisation and couples' interdependence.

While preferences, constraints, and family processes explain why a child's gender might affect divorce, it is also possible that characteristics related to divorce affect child gender. Hamoudi and Nobles (2014) described how girls in utero have survival advantages under conditions of stress relative to boys. They found that mothers who reported high levels of relationship conflict prior to their children's births were more likely to give birth to girls. Thus, the association between children's gender and divorce could arise from selection.

Previous research

Descriptive studies using the 1975 (Spanier and Glick 1981) and 1980 (Morgan et al. 1988) Marriage and Fertility Supplements of the U.S. Current Population Survey (CPS-MFS) found that married couples with daughters had divorce rates that were modestly higher than, yet statistically different from, couples with sons. However, the results were not robust. Spanier and Glick found that the differences by children's genders mainly appeared among mothers who had not graduated high school, and an analysis of later CPS-MFS data by Morgan and Pollard (2002) failed to detect differences in divorce for couples who married after 1975.

Using larger samples from the 1960-2000 U.S. Decennial Censuses, Dahl and Moretti (2008) found that first-born girls aged 12 years and younger were 0.5 per cent more likely to live without a father than similarly aged first-born boys. And using data from the 1960-1980 Censuses and 1980, 1985, 1990, and 1995 CPS-MFS, they found that first-born girls were 1 per cent more likely than boys to be living with parents who divorced. These analyses had several weaknesses. The Census data only identified children who lived in each household at the time of the enumeration and may not have indicated all of a couple's children. Dahl and Moretti did not restrict their analyses to children born within a given relationship, so the higher divorce risks could have occurred among blended families, and some divorces may have occurred before the children were born. Mammen (2008) also found differences in girls' and boys' living arrangements in the 1988-2006 Current Population Surveys.

Ananat and Michaels (2008) and Bedard and Deschênes (2005) examined data from the 1980 U.S. Census on white ever-married mothers whose oldest children were minors who co-resided with them. The studies further restricted their analyses to mothers whose oldest children were born during the mothers' first marriages. Both studies found that first-born girls were associated with slightly higher risks of divorce than first-born boys. Although the data restrictions addressed some issues from Dahl's and Moretti's study, the Census data did not distinguish between biological, adopted and step-children, and first-born children could only be identified indirectly. Also, the Census data relied on self-reported marriage and childbearing histories, which are subject to misreporting (Mitchell 2010).

Corroborating evidence of the effects of children's gender on behaviour of American parents has appeared in analyses of other marriage outcomes. Dahl and Moretti (2008) and Lundberg and Rose (2003) reported that unwed parents were more likely to marry if their children were boys than girls. And Katzev et al. (1994) found that married mothers of sons perceived higher chances that their marriages would break up than mothers of daughters.

Research from other countries, however, has not uncovered differences in divorce between parents of sons and daughters. Diekmann and Schmidheiny (2004) analysed data from 16 European countries, Canada, and the U.S. but found no associations between divorce and children's gender in any of the countries. Flouri and Malmberg (2010) failed to find differences in divorce in a U.K. survey of parents of infants, and Kalmijn (1997) found no differences in perceived marriage risks in a Dutch survey. Although modest sample sizes may have accounted for the null findings in these studies, Leigh (2009) also failed to uncover differences in divorce risks when he pooled data from five Australian censuses, and Andersson and Woldemicael (2001) could only find associations for high-parity births in Swedish registry data.

Our review indicates several weaknesses and gaps in the existing studies. Many studies suffer from data problems, including the inability to identify all of a couple's children and biological relationships. These problems lead to other limitations, such as restricting the analyses to parents with co-resident children and parents whose oldest children are minors. Few studies consider the timing of divorce relative to the birth of children (age effects), account for the duration of marriages, or adjust for right censoring in observed marriage spells. Our analysis overcomes all these issues; in addition, it examines children's gender and divorce in a developed country outside the U.S.

Marriage and divorce in the Netherlands

As with other Western countries, marriage and divorce in the Netherlands have been subject to several changes over the last 50 years. In 1971 the country introduced a no-fault divorce law, which replaced the law which only granted divorce on grounds of adultery, cruelty or other issues (Boele-Woelki et al. 2003). In 1998, registered partnerships were introduced as an alternative civil arrangement for couples who want to live together.

Dutch marriages and registered partnerships offer similar legal benefits and protections. Couples who want to enter either arrangement must first register their intention with a municipal authority at least two weeks before the wedding or partnership occurs. The lone exception is that registered partners can convert their arrangement into a marriage. The notice requirements and effective waiting periods for marriages are more stringent than those of the U.S.

The marriage rate in the Netherlands has fallen over the last half century from 9.5 different-sex marriages per 1,000 inhabitants in 1970 to 3.8 per 1,000 inhabitants in 2015 (CBS 2017a). New different-sex partnership registrations rose from fewer than 2,000 in 2001 to about 13,000 in 2015. The marriage patterns are similar to trends in the U.S. and other developed countries. Dutch attitudes on marriage and its alternatives have become less traditional over time (Treas et al. 2014). These are consistent with the "deinstitutionalisation" of marriage (Cherlin 2004) in the U.S. and Europe.

Divorce in the Netherlands requires a formal legal proceeding, but dissolution of a registered partnership does not. Between 2001 and 2009, couples could also take advantage of a "flash divorce" procedure under which they could convert marriages into registered partnerships and then quickly dissolve the partnerships. Divorces and partnership dissolutions in the Netherlands take effect once they have been recorded in the municipal register.

An analysis of Dutch couples' divorce motivations (de Graaf and Kalmijn 2006a) found that most reported relationship issues, such as growing apart and partners not providing enough attention. Overall, the personal determinants of divorce in the Netherlands are similar to those in other countries (see the review in de Graaf and Kalmijn 2006b). The number of divorces was relatively constant at around 35,000 per year, or about 10 per 1,000 married couples, over the period that we study.

Data

We construct our analysis dataset primarily from Dutch municipal register data

(*Gemeentelijke Basis Administratie*). The data are maintained by Statistics Netherlands and cover all Dutch residents between 1995 and 2015. They describe each person's date of birth, gender, immigration background, marital history, living arrangements, place of residence, and family relationships. Personal identifiers allow us to link records of people who are married or registered partners and create couple-specific measures. They also allow us to link parents and children.

The underlying records cover all the marriages and registered partnerships that existed in the Netherlands from 1995 to 2015. This includes 3,609,495 marriages that were on-going on 1 January 1995 and 1,839,504 marriages or registered partnerships that began on or after 1 January 1995.[2] For each marriage and registered partnership, we observe the date of the wedding or registration and the type of union. For marriages and registered partnerships that end within the analysis period, we also observe the termination dates and causes for termination (divorce, death, or a change of civil status).

The municipal registers enable us to match children to their legal parents, starting with the children born in 1966.[3] Each child is assigned paternal and maternal identifiers, provided that the parental records are observed in the data. Most children can be linked to both parents; only seven per cent cannot, usually because the father's identifier is missing. Parental records may be missing if the parent died before 1995, the parent resides in another country, or the child has no legal mother or father. Leveraging the parent-child identifiers, we construct measures of the gender- and age-composition of children of each couple.

For our analyses, we impose several restrictions on the data. We drop marriages and registered partnerships of couples who have older children with prior partners, couples who re-married, couples who adopted children, couples whose first-born children were twins, and

[2] We also observe marriages which were terminated before 1 January 1995 among the people who were alive and living in the Netherlands on or after that date. We drop these marriage spells from our analyses because we do not have information on other explanatory variables before 1995.
[3] Parent identifiers are available for all children born after 1965 but only for some children born earlier.

same-sex couples. We further restrict ourselves to marriages that began on or after 1 October 1971 and were therefore initiated under the no-fault divorce regime. Lastly, we drop marriages and partnerships if either spouse was born before 1935 or after 1985. The results are not sensitive to the choice of cutoff years.

Is the first-born child's gender random?

To see whether the gender of the child is random, we compare the average characteristics of couples with first-born boys and girls. We restrict the comparison to children born in or after 1995 (1.1 million children) because we do not have complete data on family characteristics before then. Table 1 lists the gender-specific averages, differences, and p-values from two-sided t-tests of the differences of characteristics.

[Table 1 about here]

The ratio of male to female children within this sample is 1.05, which is identical to the population-wide sex-ratio at birth (CBS 2017b). There are no statistically significant differences between the ages, immigration backgrounds, education, employment, or earnings of parents with first-born sons or daughters. In addition, the share of first-born children who were born out of wedlock or registered partnership is identical across the two genders, as is the share of couples who entered registered partnerships. Thus, the observable characteristics provide no evidence that the genders of Dutch couples' first-born children are selective.

The first-born's gender is, however, associated with subsequent family outcomes. Parents of first-born girls had slightly fewer children than parents of boys, and of particular relevance to our study, the divorce rate for parents with first-born girls was 0.16 percentage points higher (1.04 per cent in relative terms) than the rate for parents with first-born boys.

Multivariate event-history models of divorce and results

We investigate the association between divorce and child gender further in a series of multivariate event-history analyses. For these analyses, we use the sample of marriage and registered partnership spells which includes couples with children born prior to 1995. This allows us to track divorce outcomes in families with adolescent and adult children. The sample of couples with children born in or after 1995 is analyzed separately in the robustness section, yielding the same results.

We estimate complementary log-log discrete-time hazard models of marriage durations. The models allow us to jointly model several duration-dependent processes, including the duration of the marriage and the ages of the couple's children (effectively durations following the children's births). They also allow us to control for other fixed and time-varying observed characteristics of the couples. For these models, we split the marital spells into a series of yearly records which contain the characteristics of the couple and their children on the day of the wedding anniversary and an indicator for the event of divorce or dissolution in the following 12 months.

The spell records begin on the wedding or registration date if the couple married in or after 1995 or are left-truncated in 1995 if the couple married earlier. For the left-truncated spells, we observe and condition on the elapsed duration of the marriage or partnership, the couple's childbearing history, and the age progression of their children through 1995. We also observe the couple's time-invariant background characteristics. The sequence of records continues until a divorce or partnership dissolution occurs, or until a right-censoring event occurs. Spells are right-censored (a) on 31 December 2015 if the relationship was on-going then, (b) when a spouse dies, (c) when one or both spouses emigrate from the Netherlands, (d) when the spouses reach 40 years of marriage, or (e) when the youngest child reaches age 27. We censor at these last two events because our other selection criteria lead to few spells extending beyond these points. Our results are robust to the choice of censoring cutoffs.

Unconditional results. As an initial analysis, we estimate a descriptive hazard model of marriage durations for the couples with children in which the age of the first-born child (equivalently, the time since the first child's birth) is the lone duration measure. We specify the duration pattern non-parametrically by including dummy variables for the child's age, and we allow for gender-specific effects by interacting the age dummies with an indicator for the child being a daughter. The reference group contains couples whose child is less than one year old. For this initial analysis, we ignore the risks faced by the parents in the years preceding the birth, and the risks faced by childless couples.

Exponentiated estimates of the age-specific coefficients are graphed in Figure 1a. Appendix Table A1 (column 1) lists the estimated coefficients. The estimates, which can be interpreted as approximate odds ratios, indicate that the unconditional hazard probabilities of divorce for couples with first-born sons and daughters follow indistinguishable trajectories through the children's 12th year, rising sharply until the children reach age 7 and falling thereafter. At age 13, however, the trajectories diverge, with higher hazard probabilities attained by families with daughters. The disparity peaks when the children are 15 years old but remains until they reach age 19. After age 19, the hazards are again indistinguishable. The dots indicate the ages at which the excess divorce risks for families with daughters are statistically significant. In Figure 1b, we plot the exponentiated coefficients and 95 per cent confidence intervals for the age-specific interactions. These confirm the statistical significance of excess divorce risks for 13- to 18-year-old daughters. At the peak of the effect the odds ratio is 1.088, which means that families with first-born girls are facing 8.8% higher hazard risks of divorce compared to families with first-born sons of the same age. Overall, the descriptive results point to a gender difference in divorce risks but one that only appears during a child's teenage years—not early in the child's life nor after the child's 19th birthday.

[Figures 1a and 1b about here]

The substantive importance of hazard probabilities is difficult to ascertain because they are calculated conditionally upon the duration of the marriage. To get a better sense of the magnitudes, we use the hazard estimates to calculate cumulative probabilities that a marriage fails at different ages of the first-born child and list these probabilities and confidence intervals in Table 2.

[Table 2 about here]

As with the hazard trajectories, couples' cumulative divorce risks are identical until their first child reaches age 13, when the risks faced by couples with first-born daughters start to accelerate. The difference between the cumulative risks reaches 0.47 percentage points at age 19, which renders the couples with 19-year-old first-born girls 1.87 per cent more likely to be divorced than couples with first-born boys of the same age.

Conditional results. Estimated duration dependence patterns can be sensitive to omitted variables. For that reason, we estimate multivariate models that add controls for marriage duration, parental immigration, parental education, parental age at marriage, being in a registered partnership, the first child being born pre-maritally, and cohort and year effects. We also expand the analysis data set to include childless couples and periods before couples become parents. To make the results comparable to the previous unconditional estimates for couples with children, we continue to use the first-born child's year of birth as the reference category and add a time-varying dummy indicator for the couple being childless. The gender- and age-specific estimates of the divorce risks associated with first-born children from this model are graphed in Figure 2a, and the full set of coefficient estimates are reported in Appendix Table A1 (column 2).

[Figures 2a and 2b about here]

The magnitude and general shape of the age-dependent pattern change considerably in the multivariate specification. Accounting for the duration of marriage and other covariates leads to divorce risks that increase until the first-born reaches age 18 or 19 and that fall afterwards. This pattern is consistent with parents trying to delay divorce until the children are adults ('staying together for the sake of the kids'). It is also consistent with increasing marital stress during children's adolescence and diminished stress as young adult children leave and parents experience the 'empty nest' (Heaton 1990; Hiedemann et al. 1998).

Although the shape of the age pattern changes, the age-specific gender differences remain. The conditional age-specific excess divorce risks for first-born daughters, which are plotted in Figure 2b, closely follow the pattern in Figure 1b. As with the unconditional results, divorce risks from daughters and sons are indistinguishable through age 12 and after age 18, but the risks from daughters are higher from ages 13 to 18. The differences retain both the original magnitudes and statistical significance.

The coefficient estimates for the other characteristics of couples in Appendix Table A1 reveal that Dutch couples' divorce risks rise through the first four years of marriage and fall thereafter. Childless couples are estimated to have the same divorce risks as parents whose first-born children are 12 years old. For couples with children, divorce risks are higher if the first child was born pre-maritally. Registered partnerships are more likely to dissolve than formal marriages, which is consistent with the easier dissolution procedures for partnerships. The divorce risks rose in 2001-2007 when flash divorces were available. Couples in which both spouses immigrated to the Netherlands have lower divorce risks than couples in which both spouses are natives, while couples whose parents immigrated and couples with mixed immigration backgrounds (e.g., a first-generation immigrant married to a native) have higher divorce risks. Divorce is negatively associated with the spouses' education levels.

Higher-parity children. A potential confounding factor in our analysis of divorce is the effect of the first-born's gender on subsequent childbearing. One implication of son preferences is that couples whose first child is a daughter will have incentives for another child. These types of fertility effects have been found in developing countries (Arnold 1997; Das Gupta et al. 2003), and Dahl and Moretti (2008) found evidence of them in the U.S., while Angrist and Evans (1998) found evidence that American couples had preferences for variety in genders. As additional children generally reduce the risks of divorce (Amato 2010), part of the effect of the first child's gender might operate through family size.

Our descriptive analyses in Table 1 did reveal a small association between the first-born's gender and the number of children: families with first-born boys had marginally more children. This finding is inconsistent with childbearing decisions being influenced by preference for sons. To address the conditional association between fertility and the first-born's gender, we extend our multivariate model by adding counts of the number of higher-parity children of each age from 0 to 26 years, a count of higher-parity children aged 27 or older, and counts of the number of higher-parity daughters at each of these ages. Besides controlling for family size, these measures also allow us to examine the association between divorce risks and the gender and age-composition of higher-parity children. We use count variables instead of dummies to account for twins and closely-spaced siblings. The rest of the model specification is kept unchanged.

Estimates of the gender-specific age effects of first-born children on divorce risks from this extended specification are plotted in Figure 3a, while the full set of coefficient estimates are reported in Appendix Table A1 (column 3). Adding controls for higher-parity children increases the amplitude of the first child's age effects from ages 1 to 18 and leads to a hump in the age trajectory from ages 2 to 9. However, as with the estimates from the previous specifications, the extended model continues to indicate that there are excess divorce risks from daughters who are 13 to 18 years old.

[Figures 3a, 3b and 3c about here]

The age-specific estimates of the divorce risks associated with higher-parity children are graphed in Figure 3b. Higher-order children are associated with a similar age-specific pattern of the divorce risks as the first-borns, with risks rising through age 19 and falling thereafter. The principal difference between the patterns is that the magnitudes of the associations for higher-parity children are smaller.[4] The gender pattern in Figure 3b is also similar to the pattern found among the first-borns—between the ages of 13 and 18, families with higher-order daughters face higher divorce risks. Figure 3c shows the excess divorce risks for higher-parity daughters. The statistical significance is retained by most of the dummies within this age range, although the overall magnitude of the effects are lower than those for first-borns.

Discussion of the results

The finding of a gender effect which only appears when children are teenagers seems to rule out several causal mechanisms discussed in the theory section. The absence of differences in infancy or early childhood goes against the hypothesis regarding overarching cultural or social preferences for sons. The lack of gender differences in marriage rates for parents of children who were born pre-maritally and the higher rates of subsequent childbearing for parents of sons are also inconsistent with overarching son preferences.

More generally, our divorce findings run counter to hypotheses that operate with forward-looking behaviour and foreseeable differences in parents' marriage valuations. If the utilities of marriage with teenage daughters and sons are foreseeably different, then such differences should be incorporated into rational parents' decision making from the moment the child's gender is revealed. Accordingly, families with daughters would be more likely to divorce throughout the entire childhood, not only in the teenage period.

The teenage effect is also inconsistent with non-random sex selection in utero. If

[4] The coefficients for the age patterns of first-born and higher-parity children have different scalings in Figures 3a and 3b. Our assessment of magnitudes come from models that are respecified to have comparable scalings.

daughters are more likely to be born to mothers who experience stress and conflict during the pregnancy, then the divorce risks should be high in the period closely following the birth. The lack of such an effect, together with the lack of evidence of selection on parental observable characteristics in Table 1, leads us to reject the sex-selective hypothesis.

Based on this evidence, we conclude that the effect is more likely to occur because of unexpected changes in the constraints or family processes faced by the couples with teenage boys and girls. The teenage years are a period of tremendous and sometimes tumultuous change, not only in terms of the child's physical, emotional and behavioural development, but also in his or her relationship with parents, independence, and the consequences of actions. Some of these changes might be more expected or more successfully navigated by the parents than others. In principle, the mechanism might lie in any of the age-specific constraints, including differential money costs, time requirements, parental involvement, developmental vulnerabilities to divorce, and parent-child relationship strains. However, the final explanation—parent-child relationship strains—seems especially potent. Teenage girls are likely to be subject to stricter regimes, less autonomy and more parental supervision than teenage boys, and these are likely to be met with child's opposition. The parental beliefs about a daughter's lifestyle choices may go unnoticed when the child is younger and dependent, but in the teenage period they may translate into worse relationships between the parents and the child, and also between the parents themselves - particularly if the father's and the mother's opinions are not aligned. The lower enjoyment of the shared family time reduces the utility from marriage, and may nudge the parents towards divorce.

Subsample analyses

We investigate the parent-child conflict hypothesis further by estimating versions of our hazard model for various subpopulations which are likely to differ in terms of the gender norms and experiences of fathers and mothers. We analyse whether the parents from groups with more traditional norms or less mixed-gender experience are subject to stronger teenage effects. For our subgroup analyses, we estimate hazard specifications with controls that are identical in almost every respect to the models underlying Figures 2a and 2b. However, to facilitate comparisons across the groups, we replace the 27 age-specific interactions of the daughter dummy with a simpler set of three age-group interactions corresponding to the daughter being a child (aged 0-12), teenager (aged 13-18), or young adult (aged 19-26).

As a benchmark, we estimate the simplified model for the full sample of couples and report the exponentiated coefficients for the three daughter-age interactions in the first row of Table 3 (complete results are provided in Supplemental Appendix Table S1). In line with the results in Figure 2b, the only statistically significant coefficient is the one for teenage daughters, who are associated with divorce hazard risks that are 5.2 per cent higher than those of first-born teenage sons.

[Table 3 about here]

Our first subsample analysis differentiates couples by their immigration backgrounds. We split couples into four groups: spouses who were both born in the Netherlands, spouses who were both first-generation immigrants, couples with a native husband and immigrant wife, and couples with an immigrant husband and native wife. Our supposition is that differences in gender-role attitudes will be larger between immigrant parents, who may hold attitudes from their country of origin, and their children raised in the Netherlands. Another supposition is that these differences may be compounded for immigrants married to native spouses. The coefficients of daughter-age interaction terms are shown in the second through fifth rows of Table 3. Consistent with our suppositions, the teenage differential in divorce risks is weakest among couples who were born in the Netherlands, stronger among immigrant couples, and strongest among couples with mixed immigration backgrounds, especially those with immigrant fathers. For immigrant couples, the elevated divorce risks persist even in the adult period, which may reflect gender-role strains continuing past the teenage years.

We next consider how the relationship between child gender and divorce differs with wives' and husbands' education levels. Spanier and Glick (1981) and Dahl and Morretti (2008) found that child gender had stronger associations with family structure among less-educated parents. Gender-role attitudes tend to be more egalitarian among people with more schooling (see Treas et al. 2014). More educated parents might also be better informed about parent-child relationships or have better capability to navigate difficulties. Education is reported in four categories, including a category for missing educational records.[5] The results show that parents with lower education levels face higher relative divorce risks with teenage daughters.

Gender-role attitudes have also become more egalitarian over time, and we might expect that earlier cohorts of parents would have more relationship strains with daughters than later cohorts. The next four rows of Table 3 report estimates for models that consider fathers who were born in 1935-49, 1950-59, 1960-69, and 1970-85. Consistent with the trends in attitudes, the divorce risk differential for daughters was highest for the earliest cohort of fathers and lowest for the last two cohorts.

In the last set of models, we investigate whether the teenage effect varies with the gender composition of the parents' siblings. Parents' early exposure to mixed-gender relationships may influence their gender norms, with men becoming more egalitarian if they have sisters and women becoming less egalitarian if they have brothers. More generally, growing up with an opposite-sex sibling may provide more insight and sensitivity regarding mixed-gender relationships. For these analyses, we link the spousal records to records for their mothers and quantify the number and genders of each spouse's maternal siblings. The

[5] The administrative records of educational attainment come from municipal authorities. All municipalities provide records for people born after 1986, but municipal participation is incomplete for earlier cohorts, with the number of participating areas falling for successively earlier cohorts. The availability of education information does not appear to be associated with other personal characteristics except for birth cohort.

sample of couples with sibling records is smaller than the original sample. We drop spouses who were born before 1966 because maternal identifiers are incomplete before then. We also exclude first-generation immigrants because we do not have maternal records unless their mothers also emigrated to the Netherlands. We split the husbands and wives into groups based on whether they have at least one opposite-gender sibling born less than ten years apart from themselves. We apply this spacing restriction because near-aged siblings interact more. Results from these models are presented in the last segment of Table 3.

For fathers, having a sister has a pronounced effect on the size of the teenage coefficient. Teenage daughters increase the relative risks of divorce by 7.9 per cent among fathers who grew up without near-age sisters, but they have no association with divorce for fathers with near-age sisters. In contrast, mothers with near-age brothers face the same gender gaps as other mothers. Looser spacing restrictions do not change the conclusions of the analysis, though they reduce the differences between the fathers' sibship groups somewhat. To dispel concerns about the endogeneity of sibship size and gender composition, we also estimate models for subsamples of spouses with twin siblings, conditioning on the gender of the twin. Even within this restricted subsample, we find that fathers with twin brothers have a larger gender gap in divorce risks than fathers with twin sisters.

The sibship analyses point to fundamental differences in the behaviour of fathers with and without sisters. It appears that fathers with sisters are better equipped to raise daughters than fathers who lack such exposure. Accordingly, the fathers with sisters might better anticipate the problems that may arise during daughter's teenage years, or be better at resolving problems with daughters or partners.

LISS panel

The subsample analyses provide indirect evidence that the higher divorce risks stem from parent-child, and possibly father-daughter, relationship strains. However, they cannot fully rule out other explanations such as different costs and parental involvement for teenage sons and daughters. To examine potential mechanisms more directly, we turn to the LISS, a panel survey, which followed a representative sample of Dutch households, totaling 11,500 individuals, for nine years. The LISS asked household members many questions regarding the quality of their relationships, attitudes, children's behaviour, time-use, and expenditures.

Following the results of our principal analysis, we focus on married couples whose first-born child is younger than 19 years of age and is the biological child of both spouses. The full set of sample restrictions and the number of observations lost in every step of the sample selection procedure are described in Appendix B. The resulting sample has approximately 6,500 parent-year observations.

We estimate multivariate models of the outcomes reported by the mothers and fathers. For responses involving relationships, attitudes, and children's behaviour, we estimate ordered logit models, and for reports of expenditures and time-use, we estimate OLS regressions. Our principal explanatory variables are dummy variables for first-born children aged 0-12 and 13-19 and interactions with an indicator for a first-born daughter. The models also control for the parent's age (cubic), education and immigrant background; the numbers of higher-parity boys and girls; and wave fixed effects. We report the estimated coefficients of the first-born daughter/age-group interactions separately for fathers and mothers in Table 4. Complete results are reported in Supplemental Appendix Table S2.

[Table 4 about here]

The LISS asked several questions about conflict and relationship strains. Both fathers and mothers of teenage daughters report more parenting disagreements with their partners than fathers and mothers of teenage sons, even as mothers and fathers of younger girls report fewer disagreements. Fathers of teenage daughters also report worse family relationships. Mothers of teenage daughters report more arguments with partners over household expenditures, less life satisfaction, and more positive attitudes towards divorce than mothers of teenage sons. These results indicate that marital and parenting relationships are more strained in households with teenage daughters.

Parents report different types of behavioural problems for their teenage sons and daughters. Mothers are more likely to report that teenage daughters are quarrelsome, feel sad or depressed, or feel worthless or inferior than teenage boys. Mothers and fathers are each less likely to report that teenage daughters have trouble concentrating. Additionally, fathers are less likely to report that their teenage daughters are clingy or disobedient in school.

The LISS provides no evidence of higher time of money costs of teenage daughters. Parents of teenage daughters are no more likely than parents of teenage sons to report children's care as a burden, though mothers of younger daughters report greater burdens. Parents also do not report differences in expenditures for children if they have teenage daughters (though the measure only describes expenditures for children under 15 years) nor significant differences in the time spent with children.

The LISS asked children who were age 16 or older about relationships with their parents. We estimate ordered logit models of the responses with two sample specifications: one for first-born children aged 16-18 and one for all children aged 16-18. The models include a dummy variable for daughters and controls for the other parent, household, and temporal characteristics from our previous specifications. Estimates of the daughter coefficients are reported in Table 5. Teenage daughters report worse relationships than sons with their fathers, but they do not report worse relationships with their mothers. These responses bolster the evidence that the relationships between fathers and daughters are particularly important for understanding the gender divorce gap. Complete results are reported in Supplemental Appendix Table S3.

[Table 5 about here]

Robustness checks

We subject our baseline finding of excess divorce risks faced by families with teenage daughters to many robustness checks. These include model specifications in which we: estimate logit specification of the model; drop marriages of couples with children born before 1995; extend the sample to include families with children that are not shared by the same parents; replace the divorce date by the date of separation (extracted from the personal cohabitation records); and add employment and earnings controls among the set of covariates (available as of 1999). Neither the magnitude nor the significance of the teenage daughter effect is affected by the changes to the baseline specification. Full sets of results are listed in Supplemental Appendix Table S4.

We also investigate whether similar patterns exist among parents who are living out of wedlock. Cohabitation is a household arrangement that is steadily growing in popularity, and it is therefore important to see whether the cohabiting couples are subject to the same effects as the married couples. The analysis of cohabiting couples is complicated by the lack of precise information on the starting date and ending dates of the relationships. The date of separation can be approximated by the date when one of the parents is observed to move out of the shared residence (provided that he or she does not move back in afterwards). A similar strategy can be adopted to approximate the date of initiation; however, this is further complicated by left-censoring of the cohabitation data (residential histories were not collected prior to 1995). For this reason, we replicate the strategy used in the unconditional analysis and use the age of the first-born child to approximate the duration of the relationship. The specification of the hazard model for cohabiting couples includes dummies for the three age groups of the first-born, gender-age group interaction terms, cubic functions of parental age at birth, and year and cohort fixed effects. Results are shown in Figure 4.

[Figure 4 about here]

The hazard rates show that cohabiting couples with first-born daughters are also subject to excess divorce risks in the teenage period. The coefficient is higher than the one corresponding to the married sample, but the smaller sample of cohabiting couples renders the estimate less precise. Nevertheless, it is clear that the finding of excess risks of separation applies to couples with children regardless of their type of union.

What about the U.S.?

To test the external validity of our findings, we replicate the event-history analysis using U.S. data from CPS Marriage and Fertility Supplements extracted from IPUMS-CPS (Flood et al. 2015) and NBER CPS archives. The marriage and fertility supplements were asked of all women living in the CPS households who were older than 15 years of age at the time of interview. Following Dahl and Moretti (2008) we use the 1980, 1985, 1990 and 1995 surveys. Women in these surveys were asked to list their full marriage and fertility histories (up to three marriages and five children), whereas other and subsequent waves only asked about the youngest child and current marital status. This data limitation makes our event-history analysis intractable for more recent CPS surveys. Women in the four selected surveys were also asked about the dates of wedding, separation, and divorce for each marriage. The fertility information includes the genders and the years of birth of the first four children and the most recent child.

The sample selection criteria are amended to account for the unavailability of fathers' characteristics. We analyse durations of first marriages of women whose first child was born between the date of the first wedding and (if observed) the date of the first divorce. The sample also contains first marriages of women who did not have children. Women with children born before the date of the first marriage are dropped from the sample because it is less certain that the first husband is the biological father. We also drop marriages of women who were born before 1935, and women who married either before the age 16 or after the age

45. The results are not sensitive to the choice of the cut-off ages. Each marital spell is assigned a failure time if the marriage ends in divorce prior to the year of the survey. Otherwise, the spells are treated as right-censored. Censoring is applied either in the year of the survey, or in the year when the husband dies, the spouses reach 30 years of marriage, or their youngest child reaches age 27. The final sample contains 127,236 marital spells.

We estimate a specification of discrete-time hazard model which is very similar to the specification used in the subsample analysis. The covariates include dummies for first-born's age-groups; interactions of age-groups with a daughter dummy; controls for marriage duration, non-parenthood, maternal age at marriage; and cohort and year fixed effects. The observations are weighted by the CPS fertility supplement weights. Estimates for the daughter/age-group interactions are graphed in Figure 5.

[Figure 5 about here]

The age-dependent pattern from the CPS-MFS closely resembles the pattern from the Dutch data. U.S. couples with teenage daughters are more likely to divorce than couples with teenage sons. The effect is three times as large as the effect for Dutch married couples. For daughters aged 0-12 we are unable to find a statistically significant increase in divorce odds. The large standard errors however make it impossible to rule out existence of an effect of the magnitude found by Dahl and Moretti (2008). Their analysis of the U.S. census data yielded an estimate of a 1.3 per cent increase in divorce odds for couples with first-born daughters aged 0-12. Interestingly, the analysis of the CPS data by the same authors yielded a daughter effect more than six times larger than the baseline result. The authors attributed this difference to the small sample size of the CPS, and focused on the baseline effect found in the census. However, our analysis indicates that this interpretation is incorrect—the difference between the two coefficients is driven by the fact that the CPS sample was not restricted to mothers with children aged 0-12. In contrast to the census analysis, the CPS sample included mothers with older children as well, and the resulting estimate of excess divorce risk for daughters therefore incorporated the teenage effect shown above. This rendered the CPS coefficient much larger than the census coefficient for families with children younger than 13 years of age. Further evidence supporting this interpretation can be found in Bedard and Deschênes (2005). Their analysis of the 1980 U.S. Census did not restrict first-born children to be younger than 12 years of age and yielded a coefficient four times larger than the coefficient found by Dahl and Moretti.

Our analysis shows that the effect for young daughters is dwarfed by the effect which emerges in the teenage years. The teenage effect is large in both relative and absolute terms, since families with older children face higher risks of divorce. It is therefore likely to contribute to a much greater disparity between the total divorce rates of families with sons and families with daughters. Our findings from the CPS-MFS are robust to multiple changes of the sample and model specification, including choice of the censoring thresholds, exclusion of marriages of childless mothers, extension to second and third marriages, and inclusion of covariates controlling for presence of higher-order children.[6]

Conclusion

In this paper, we show that parents with teenage daughters are more likely to divorce than parents with teenage sons. The effect peaks at the age of 15, when Dutch married couples with first-born daughters face an 8.8 per cent higher hazard probability of divorce compared to couples with first-born sons. The daughter-specific divorce risks remain elevated throughout the teenage years, so that by the age of 19, the cumulative divorce rate for couples with first-born daughters is 0.47 percentage points (1. 87 per cent) higher than the cumulative rate for couples with first-born sons. Similar effects are found for children of higher parities.

[6] Estimates corresponding to the robustness checks are available upon request.

We find no evidence of gender-specific divorce risks among families with children aged 0-12, or children older than 18. This finding is of particular importance for pinning down the causal mechanism responsible for the disparity observed in the teenage period. It allows us to rule out mechanisms which assume the existence of time-invariant preferences for boys, or of other foreseeable differences in the parental valuations of marriages with sons and daughters. Within a rational-choice framework, such mechanisms would render the daughter to be a risk factor straight from her birth, leading to higher divorce rates among families with young daughters. Using similar reasoning we can also rule out stress-induced selection into live birth, because this mechanism would also raise divorce rates among families with very young daughters.

The isolation of the gender effect in the later stages of a child's life implies the existence of a dynamic mechanism which influences marital stability through unexpected changes in the constraints faced by the parents of teenage children. Although analysts have suggested expenditures, time requirements, and parental involvement as possible constraints, we see stressful parent-child relations as a more likely mechanism. Teenage daughters are likely to be subject to stricter regimes and more parental supervision than teenage sons, and this involvement may instil conflict between the parents and their child. The conflict can be further exacerbated if the gender-role attitudes and expectations maintained by the parents are not aligned with those of their daughters, or with each other.

Our subsample analyses support this hypothesis, showing that families in which parents and children are more likely to hold conflicting beliefs about the gender roles are subject to larger teenage effects than more homogenous families. The effects are particularly strong among couples with mixed immigration backgrounds, and among couples in which the father grew up without near-age sisters. Conversely, the effect is attenuated for native couples, and it disappears completely if the father's sibship includes one or more sisters. The conflict hypothesis is also supported by our analysis of household survey data, which shows that families with teenage daughters report significantly lower levels of satisfaction with their inter-personal relationships. The responses of parents with teenage sons and daughters show no significant difference in both realized and desired levels of expenditures on children, and we also find no differences in the levels of parental time-use.

We show that our findings are robust to various alteration of the sample, and estimation methods, and we show that the same effect of elevated divorce risks for families with teenage daughters holds also for cohabiting couples in the Netherlands, and also for married couples in the United States. In both cases, the effects are substantially stronger than the effect for married couples in the Netherlands. This result contrasts with some of the previous U.S. studies, which find marginal divorce disparities among families with younger daughters, and points to the importance of including families with teenage children in the analyses of gender-specific divorce risks.

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Table 1: Average characteristics of couples with first-born sons and daughters

	Sons	Daughters	Diff.	P-val.
Father's birth year	1971.59	1971.60	-0.01	0.398
Mother's birth year	1974.11	1974.12	-0.01	0.389
Year of wedding	2001.14	2001.15	-0.01	0.635
Father's age at wedding	30.04	30.04	0.00	0.702
Mother's age at wedding	27.53	27.53	0.00	0.704
Mother's age at birth of the first-born	31.91	31.91	0.00	0.859
Father's age at birth of the first-born	29.39	29.39	0.00	0.871
Marriage duration at birth of the first born (conditional on being married at birth)	2.78	2.78	0.00	0.424
Father's immigration background				
Native	77.31%	77.23%	0.09%	0.284
1st Generation Immigrant	16.08%	16.14%	-0.06%	0.415
2nd Generation Immigrant	6.61%	6.64%	-0.03%	0.547
Mother's immigration background				
Native	75.55%	75.51%	0.04%	0.587
1st Generation Immigrant	17.43%	17.44%	-0.01%	0.870
2nd Generation Immigrant	7.01%	7.05%	-0.03%	0.518
Father's completed education				
Less than High School	3.17%	3.20%	-0.03%	0.397
High-school	29.81%	29.77%	0.04%	0.622
University	23.66%	23.65%	0.01%	0.880
Missing records	48.35%	48.38%	-0.05%	0.776
Mother's completed education				
Less than High School	3.52%	3.54%	-0.02%	0.438
High-school	31.40%	31.49%	-0.09%	0.315
University	26.01%	25.92%	0.09%	0.278
Missing records	39.07%	39.04%	0.03%	0.784
Labor supply 1yr prior to birth of the first-born				
Father employed	85.25%	85.20%	0.05%	0.587
Mother employed	84.57%	84.63%	-0.06%	0.478
Father's annual earnings, in 1000€	30.38	30.35	0.03	0.644
Mother's annual earnings, in 1000€	22.19	22.21	-0.02	0.531
Share of first-borns who were born prior to marriage / registered partnership	14.13%	14.10%	0.03%	0.597
Share of first-borns whose parents engaged in registered partnerships	3.12%	3.07%	0.04%	0.206
Number of siblings	**1.07**	**1.06**	**0.007**	**0.000**
Birth spacing between the first two children	2.80	2.79	0.004	0.213
Share of first-borns whose parents divorced (marriage or registered partnership)	**15.33**	**15.49**	**-0.16**	**0.023**
Number of observations		1,067,067		

Note: Authors' estimates of average characteristics of couples with first-born sons and daughters born in 1995-2015. The sample excludes same-sex couples, and couples whose first-born children are either adopted, twins, or step-children. The sex ratio is 1.052. Labour supply statistics are restricted to years 2000-2015 due to limited availability of the employment records. P-values correspond to means comparison t-tests with unequal variance.

Table 2: Cumulative unconditional divorce rates following the birth of the first child

	Couples with sons		Couples with daughters		
Age	% divorced	(95% CI)	% divorced	(95% CI)	Difference
0	0.18	(0.23, 0.25)	0.18	(0.23, 0.25)	0.00
1	0.78	(1.00, 1.05)	0.79	(0.98, 1.03)	0.01
2	1.71	(2.13, 2.20)	1.73	(2.11, 2.18)	0.01
3	2.89	(3.47, 3.57)	2.90	(3.45, 3.55)	0.01
4	4.18	(4.91, 5.02)	4.20	(4.91, 5.02)	0.02
5	5.60	(6.45, 6.58)	5.61	(6.44, 6.57)	0.01
6	7.11	(8.05, 8.19)	7.11	(8.02, 8.16)	0.00
7	8.65	(9.66, 9.81)	8.67	(9.65, 9.81)	0.03
8	10.18	(11.24, 11.41)	10.21	(11.24, 11.41)	0.03
9	11.70	(12.81, 12.98)	11.75	(12.81, 12.99)	0.05
10	13.22	(14.35, 14.54)	13.25	(14.33, 14.52)	0.03
11	14.64	(15.79, 15.98)	14.70	(15.81, 16.01)	0.06
12	16.01	(17.17, 17.37)	16.09	(17.20, 17.41)	0.08
13	17.39	(18.56, 18.77)	17.49	(18.62, 18.84)	0.11
14	18.72	(19.91, 20.12)	18.89	(20.03, 20.25)	0.17
15	20.04	(21.23, 21.46)	20.30	(21.44, 21.67)	0.25
16	21.32	(22.52, 22.75)	21.66	(22.81, 23.05)	0.34
17	22.60	(23.81, 24.04)	23.00	(24.16, 24.40)	0.40
18	23.87	(25.09, 25.33)	24.30	(25.47, 25.72)	0.43
19	25.09	(26.32, 26.57)	25.56	(26.74, 27.00)	0.47
20	26.25	(27.50, 27.76)	26.71	(27.91, 28.18)	0.47
21	27.30	(28.58, 28.85)	27.77	(29.00, 29.28)	0.47
22	28.27	(29.60, 29.87)	28.74	(30.02, 30.30)	0.48
23	29.12	(30.54, 30.81)	29.59	(30.96, 31.24)	0.47
24	29.90	(31.41, 31.69)	30.39	(31.86, 32.15)	0.49
25	30.65	(32.26, 32.55)	31.11	(32.70, 32.99)	0.47
26	31.34	(33.07, 33.36)	31.82	(33.51, 33.81)	0.47

Note: Authors' estimates of cumulative divorce rates using linked marriage, divorce, and other registry data for marriages of different-sex couples with children who married after 1971, and did not have children with other partners prior to their current spouse.

Table 3: Excess hazard probabilities of first-born daughters in subsample analyses

Model specification	Age 0-12	Age 13-18	Age 19-26	Spells
Baseline, full sample	1.005	1.052***	0.995	2,732,223
	(0.004)	(0.007)	(0.007)	
Immigration background and homogeneity				
Both spouses native	1.007	1.043***	0.985*	2,180,235
(incl. 2nd gen. immigrant)	(0.005)	(0.007)	(0.008)	
Both spouses immigrants	0.991	1.077***	1.059**	315,582
	(0.014)	(0.024)	(0.025)	
Father native, mother	1.014	1.115***	1.047	129,055
immigrant	(0.019)	(0.036)	(0.041)	
Mother native, father	0.989	1.125***	1.033	97,351
immigrant	(0.018)	(0.036)	(0.039)	
Education, husband				
Less than High School	0.997	1.069**	1.053*	78,183
	(0.018)	(0.029)	(0.032)	
High School	0.999	1.045***	0.987	585,197
	(0.008)	(0.013)	(0.015)	
University	1.014	1.029	1.014	418,115
	(0.013)	(0.022)	(0.025)	
Missing records	1.007	1.057***	0.989	1,640,728
	(0.006)	(0.009)	(0.010)	
Education, wife				
Less than High School	0.995	1.066***	0.996	102,362
	(0.17)	(0.024)	(0.024)	
High School	1.005	1.044***	0.991	650,852
	(0.007)	(0.011)	(0.013)	
University	0.987	1.012	0.953	421,438
	(0.012)	(0.021)	(0.028)	
Missing records	1.012*	1.064***	1.004	1,547,571
	(0.007)	(0.010)	(0.011)	
Birth cohort, husband				
Cohorts 1935-1949	0.984	1.166***	1.017	256,568
	(0.055)	(0.038)	(0.021)	
Cohorts 1950-1959	1.005	1.057***	0.992	809,118
	(0.013)	(0.011)	(0.010)	
Cohorts 1960-1969	1.003	1.041***	0.993	895,247
	(0.006)	(0.010)	(0.014)	
Cohorts 1970-1985	1.007	1.053**	1.005	761,290
	(0.007)	(0.024)	(0.069)	
Sibship, husband				
No sisters	1.011	1.079***	0.948	365,798
	(0.010)	(0.025)	(0.057)	
At least one sister	1.012	1.005	0.965	520,846
	(0.007)	(0.020)	(0.047)	
Sibship, wife				
No brothers	0.996	1.039**	0.974	407,470
	(0.009)	(0.020)	(0.042)	

At least one brother	1.008	1.038**	0.953	630,304
	(0.008)	(0.016)	(0.031)	
Sibship, husband - twin sample				
Twin brother	1.007	1.134**	1.067	27,318
	(0.044)	(0.071)	(0.098)	
Twin sister	1.019	1.030	1.043	13,977
	(0.062)	(0.093)	(0.127)	
Sibship, wife - twin sample				
Twin brother	1.065	1.047	0.808	13,750
	(0.064)	(0.090)	(0.108)	
Twin sister	1.012	1.029	1.084	29,361
	(0.043)	(0.065)	(0.094)	

Note: Authors' estimates of exponentiated coefficients corresponding to first-born daughters in the three age-groups from the simplified specification of cloglog model of marriage durations. The model uses linked marriage, divorce, and other registry data for different-sex couples who married after year 1971, did not have children with other partners prior to the marriage

Table 4: Regression analysis of parental responses in the LISS panel, excess coefficients for first-born daughters aged 0-12 and 13-18

	Father		Mother	
Subjective questions	Age 0-12	Age 13-18	Age 0-12	Age 13-18
How satisfied are you with your current relationship?	0.032	0.030	-0.185**	-0.071
[Did] you and your partner (have) any differences of opinion regarding money expenditure over the past year?	-0.113	0.148	0.189**	0.258**
[Did] you and your partner (have) any differences of opinion regarding raising the children over the past year?	-0.152	0.298**	-0.108	0.181*
A woman is more suited to rearing young children than a man	-0.170*	-0.371***	0.014	-0.189*
Divorce is generally best solution if a married couple cannot solve their marital problems	-0.102	0.085	0.052	0.323***
Married people are generally happier than unmarried people	-0.100	-0.155	-0.079	-0.073
How would you generally describe the relationship with your family?	-0.216	-0.406**	0.032	-0.111
Caring for my child is not such a burden	0.116	0.100	0.380***	0.048
How satisfied are you with the life you lead at the moment?	0.012	-0.060	-0.146*	-0.212**
Behaviour of first-born child over last three months				
Is too quarrelsome	-0.358**	-0.074	-0.432***	0.420**
Has trouble concentrating cannot keep his/her attention focused on something for long	-0.797***	-0.870***	-1.023***	-0.662***
Has trouble relating to other children	-0.228	-0.254	-0.529***	-0.118
Is easily confused	-0.359**	-0.975***	-0.353**	-0.650***
Feels worthless or inferior	0.154	0.017	-0.267*	0.521**
Is not liked by other children	-0.619***	0.072	-0.316*	-0.154
Is headstrong sullen or irritable	-0.417***	0.134	-0.371***	0.232
Is unhappy sad or depressed	-0.295	0.053	-0.207	0.504**
Clings to adults	-0.070	-0.518*	0.142	0.026
Is too dependent on others	-0.227	-0.308	-0.082	0.127
Is disobedient in school	-1.244***	-1.032***	-1.093***	-0.388
Has trouble relating to teachers	-0.363	-0.272	-0.454	-0.117
Expenditures and time-use				
How much time did you spend in the last seven days on activities with own child	0.320	1.375	-2.013**	-0.679
Log total expenditure per month for children living at home, children 0-15	0.235	-0.017	0.038	0.100

Note: Authors' estimates of coefficients of the daughter-age interactions from ordered logit and regression models of responses from different-sex couples with a first-born biological child younger than 19 at the time of the survey. LISS data 2008-2015. The models include controls for parent's age, education and immigration background, and number and gender of children of higher parities. Significance based on robust standard errors.
*** = 0.01 significance level, ** = 0.05 significance level, * = 0.1 significance level.

Table 5: Regression analysis of children's responses in the LISS panel, excess coefficients for daughters aged 16-18

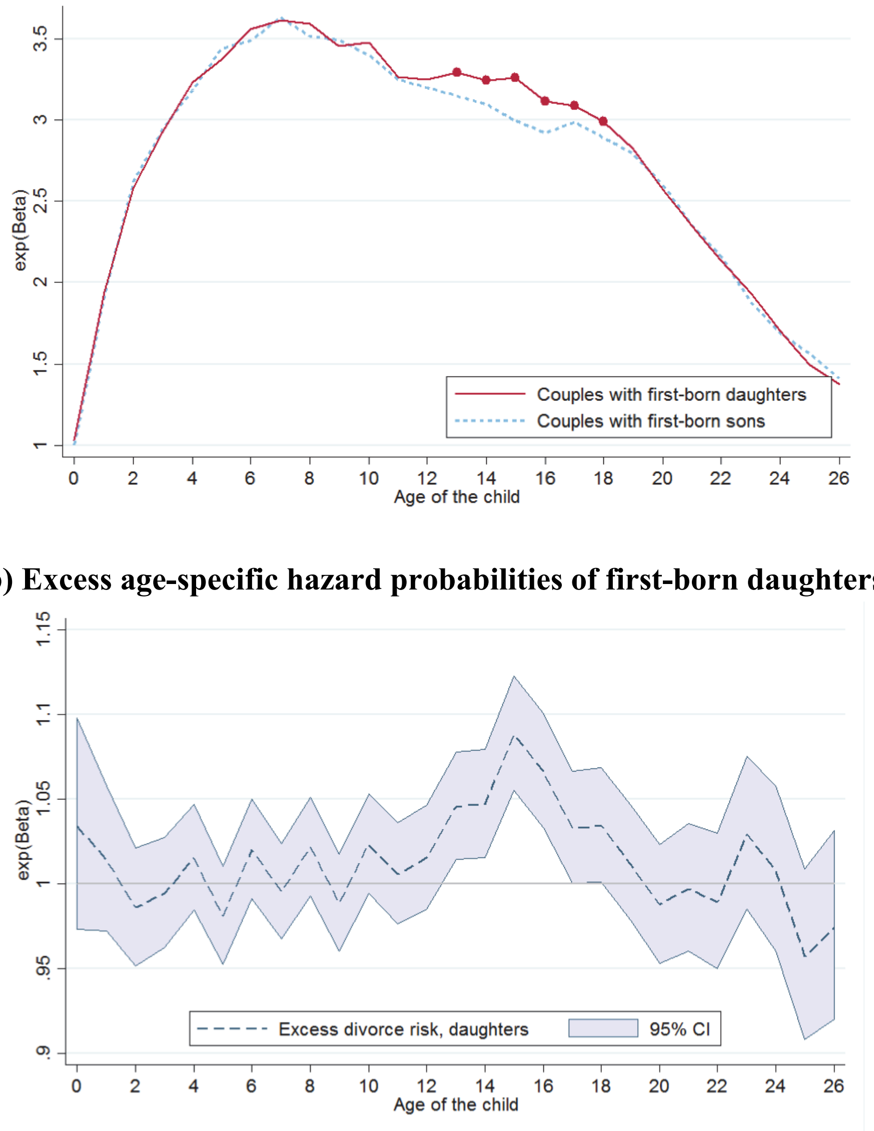
Question	Teenage daughters aged 16-18, first-born only	Teenage daughters aged 16-18, all parities
How would you describe your overall relationship with your father	-0.277*	-0.265**
How would you describe your overall relationship with your mother	-0.034	-0.043

Note: Authors' estimates of coefficients of the daughter-age interactions from ordered logit models using the responses of 16-18 year old children born to different-sex couples whose first-born is a biological child. LISS data 2008-2015. The models include controls for parent's age, education and immigration background, and number and gender of children of higher parities. Models use ordered logit specification with robust standard errors.
*** = 0.01 significance level, ** = 0.05 significance level, * = 0.1 significance level.

Figure 1: Unconditional hazard probabilities of divorce

(a) By age and gender of first-born children

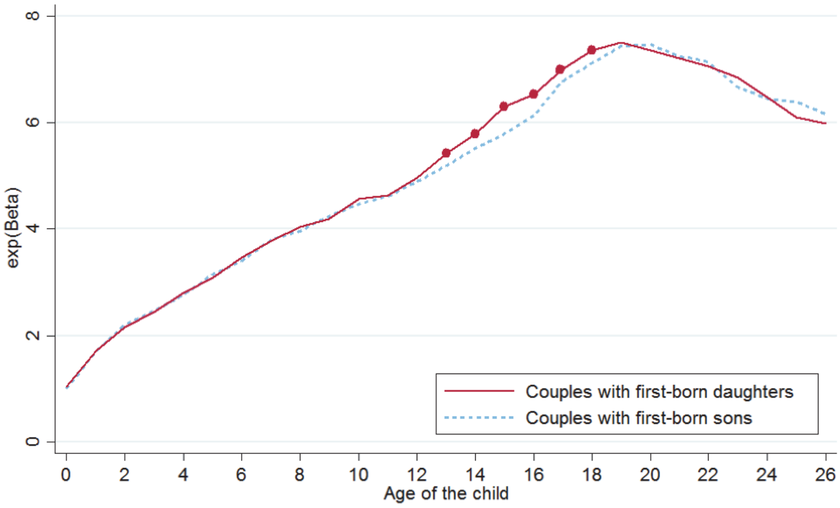
(b) Excess age-specific hazard probabilities of first-born daughters



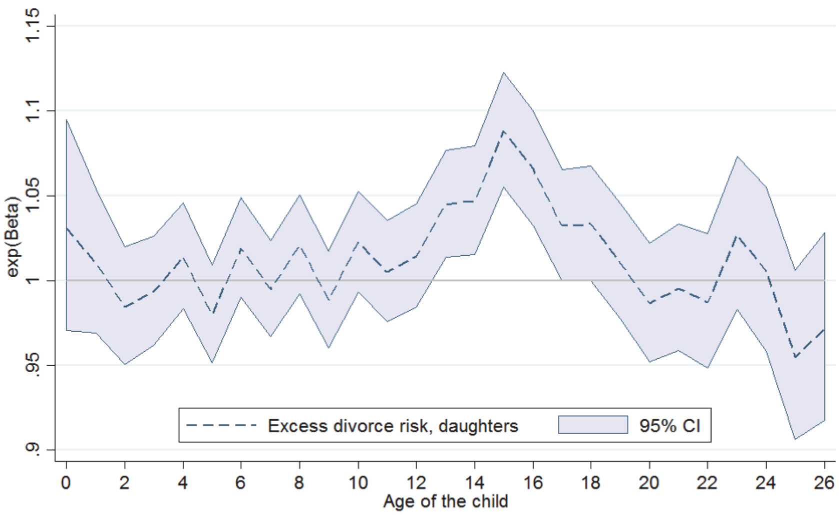
Note: Authors' estimates of exponentiated coefficients of dummy indicators of ages of the first-born daughters and sons from the cloglog model of marriage durations with no other controls. The model uses linked marriage, divorce, and other registry data for different-sex couples who married after year 1971, did not have children with other partners prior to the marriage, and remained married up to the point of having children.

Figure 2: Conditional hazard probabilities of divorce

. (a) By age and gender of first-born children



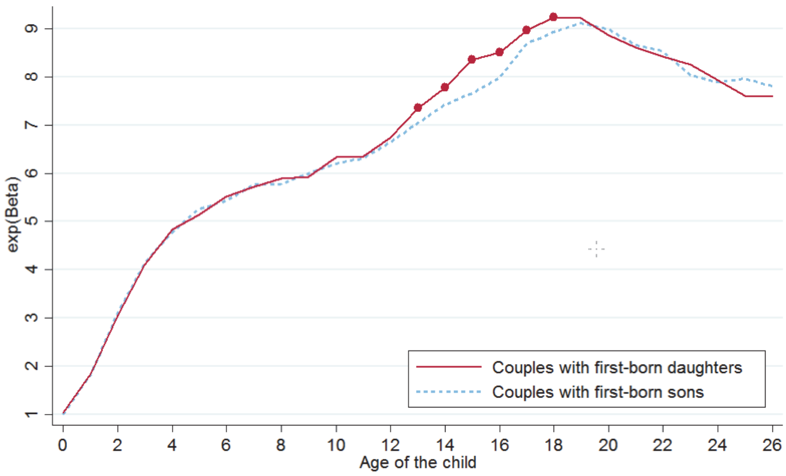
(b) Excess age-specific hazard probabilities of first-born daughters



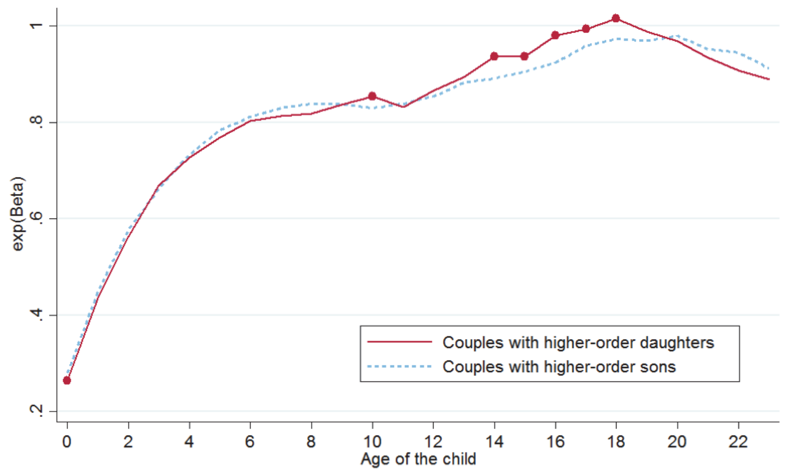
Note: Authors' estimates of exponentiated coefficients of dummy indicators of ages of the first-born daughters and sons from the principal specification of cloglog model of marriage durations. The model uses linked marriage, divorce, and other registry data for different-sex couples who married after year 1971 and did not have children with other partners prior to the marriage.

Figure 3: Conditional hazard probabilities of divorce

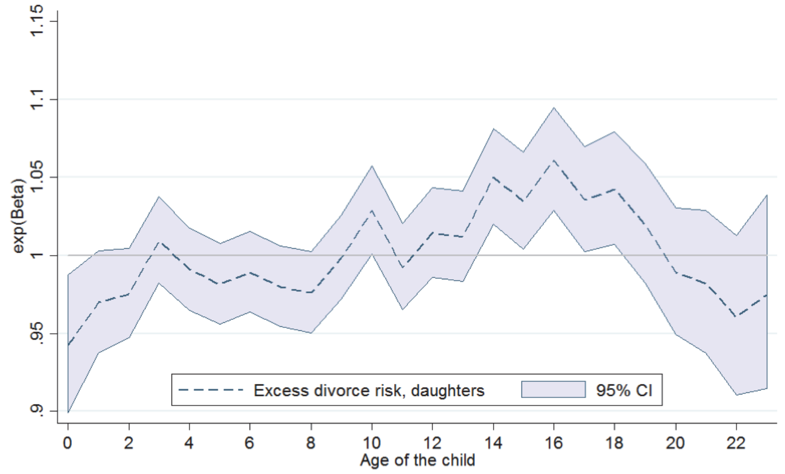
(a) By age and gender of first-born children



(b) By age and gender of higher-order children

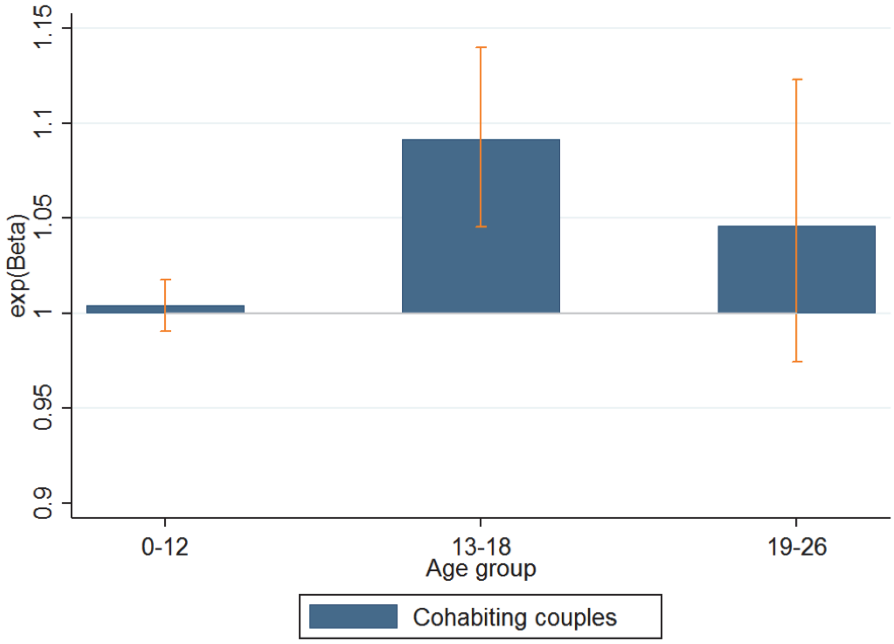


(c) Excess age-specific hazard probabilities of higher-order daughters



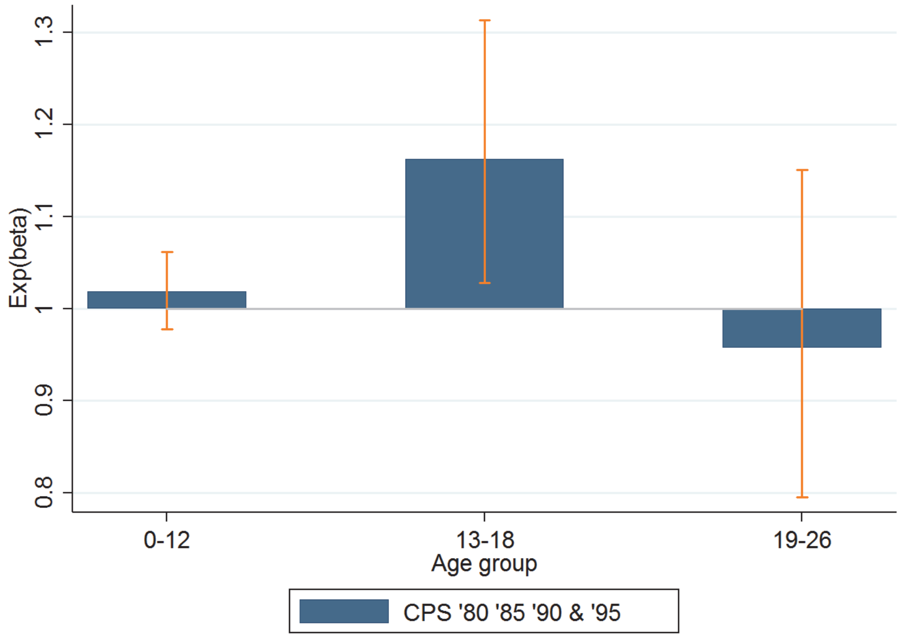
Note: Authors' estimates of exponentiated coefficients from cloglog hazard model of marriage durations which includes covariates for higher-parity children. The model uses linked marriage, divorce, and other registry data for different-sex couples who married after year 1971, and did not have children with other partners prior to the marriage.

Figure 4: Excess divorce risks of cohabiting couples with first-born daughters, by child's age group



Note: Authors' estimates of exponentiated coefficients of dummy indicators of age groups of higher parity daughters from the simplified specifications of cloglog model of cohabitation duration. The model uses linked data for cohabiting different-sex couples with children who lived in the Netherlands between 1995 and 2015. A separation is recorded when one of the spouses moves out of the shared residential address, and does not move back in within more than two years. The data for years 2014 & 2015 is not used since it cannot be determined whether the recorded separations were definitive. 95% confidence intervals indicated by shaded regions.

Figure 5: Excess divorce risks of CPS families with first-born daughters, by child's age group



Note: Authors' estimates of exponentiated coefficients of dummy indicators of age groups of higher parity daughters from a specification of cloglog model of marriage durations which adjusts for limitation of the CPS data. The model uses retrospective marital histories of American women aged 15-65, collected in CPS-MFS waves 1980, 1985, 1990 and 1995. The sample consists of first marital spells of women who were born in or after 1935, who married between the age of 16 and 45, and who did not give birth to any children prior to the date of the first marriage.

Appendix Table A1. Regression results, baseline specifications

	1		2		3	
VARIABLES	exp(Beta)	St.e.	exp(Beta)	St.e.	exp(Beta)	St.e.
First-born's age dummies						
age 1	1.915***	0.051	1.697***	0.045	1.828***	0.048
age 2	2.620***	0.066	2.197***	0.055	3.104***	0.078
age 3	2.942***	0.072	2.461***	0.061	4.152***	0.103
age 4	3.185***	0.077	2.763***	0.068	4.776***	0.118
age 5	3.444***	0.083	3.144***	0.077	5.262***	0.130
age 6	3.491***	0.084	3.403***	0.083	5.433***	0.134
age 7	3.630***	0.087	3.797***	0.092	5.771***	0.142
age 8	3.518***	0.084	3.962***	0.097	5.777***	0.143
age 9	3.495***	0.084	4.245***	0.104	6.003***	0.149
age 10	3.403***	0.082	4.468***	0.110	6.196***	0.155
age 11	3.252***	0.078	4.617***	0.115	6.316***	0.159
age 12	3.202***	0.077	4.897***	0.122	6.655***	0.169
age 13	3.151***	0.076	5.195***	0.130	7.042***	0.180
age 14	3.098***	0.075	5.521***	0.139	7.424***	0.190
age 15	2.997***	0.073	5.794***	0.147	7.672***	0.198
age 16	2.923***	0.072	6.133***	0.157	7.981***	0.208
age 17	2.987***	0.073	6.783***	0.174	8.688***	0.227
age 18	2.894***	0.071	7.119***	0.184	8.939***	0.236
age 19	2.798***	0.069	7.437***	0.194	9.131***	0.244
age 20	2.608***	0.065	7.466***	0.197	8.996***	0.245
age 21	2.360***	0.060	7.249***	0.195	8.664***	0.241
age 22	2.167***	0.056	7.147***	0.196	8.539***	0.244
age 23	1.884***	0.050	6.669***	0.189	8.045***	0.239
age 24	1.692***	0.047	6.452***	0.190	7.906***	0.245
age 25	1.568***	0.045	6.389***	0.194	7.974***	0.259
age 26	1.411***	0.042	6.151***	0.195	7.799***	0.273
*First-born's age * daughter dummies*						
age 0	1.034	0.032	1.031	0.032	1.031	0.032
age 1	1.014	0.022	1.011	0.022	1.011	0.022
age 2	0.986	0.018	0.984	0.018	0.982	0.018
age 3	0.994	0.017	0.994	0.017	0.991	0.016
age 4	1.015	0.016	1.014	0.016	1.013	0.016
age 5	0.981	0.015	0.980	0.015	0.978	0.015
age 6	1.020	0.015	1.019	0.015	1.016	0.015
age 7	0.995	0.014	0.995	0.014	0.992	0.014
age 8	1.022	0.015	1.021	0.015	1.019	0.015
age 9	0.989	0.015	0.989	0.015	0.987	0.015
age 10	1.023	0.015	1.023	0.015	1.021	0.015
age 11	1.005	0.015	1.005	0.015	1.004	0.015
age 12	1.015	0.016	1.014	0.016	1.014	0.016

age 13	1.046***	0.016	1.045***	0.016	1.045***	0.016
age 14	1.047***	0.016	1.047***	0.016	1.047***	0.016
age 15	1.088***	0.017	1.088***	0.017	1.088***	0.017
age 16	1.067***	0.017	1.066***	0.017	1.067***	0.017
age 17	1.034**	0.017	1.033**	0.017	1.034**	0.017
age 18	1.034**	0.017	1.033**	0.017	1.034**	0.017
age 19	1.011	0.017	1.010	0.017	1.011	0.017
age 20	0.987	0.018	0.986	0.018	0.986	0.018
age 21	0.997	0.019	0.995	0.019	0.995	0.019
age 22	0.989	0.020	0.987	0.020	0.987	0.020
age 23	1.030	0.023	1.027	0.023	1.028	0.023
age 24	1.008	0.025	1.006	0.025	1.006	0.025
age 25	0.958	0.026	0.956*	0.026	0.956*	0.026
age 26	0.977	0.029	0.974	0.028	0.974	0.028
No Children dummy			6.099***	0.134	6.216***	0.136
Marriage duration dummies						
1 year			3.882***	0.067	3.896***	0.068
2 years			5.864***	0.107	5.927***	0.108
3 years			7.157***	0.143	7.289***	0.145
4 years			8.018***	0.179	8.233***	0.183
5 years			7.741***	0.194	7.971***	0.200
6 years			7.593***	0.213	7.819***	0.219
7 years			7.109***	0.222	7.326***	0.228
8 years			6.642***	0.229	6.842***	0.236
9 years			6.235***	0.236	6.413***	0.243
10 years			5.879***	0.243	6.027***	0.249
11 years			5.517***	0.247	5.629***	0.252
12 years			5.099***	0.246	5.171***	0.250
13 years			4.822***	0.250	4.855***	0.252
14 years			4.585***	0.254	4.582***	0.254
15 years			4.378***	0.259	4.340***	0.256
16 years			4.163***	0.261	4.093***	0.257
17 years			3.880***	0.257	3.781***	0.251
18 years			3.721***	0.261	3.595***	0.252
19 years			3.513***	0.259	3.363***	0.248
20 years			3.322***	0.257	3.149***	0.244
21 years			3.112***	0.252	2.922***	0.237
22 years			3.011***	0.255	2.803***	0.238
23 years			2.861***	0.253	2.642***	0.234
24 years			2.766***	0.255	2.539***	0.234
25 years			2.525***	0.242	2.309***	0.222
26 years			2.400***	0.239	2.190***	0.218
27 years			2.262***	0.234	2.062***	0.214
28 years			2.134***	0.229	1.943***	0.209
29 years			2.099***	0.234	1.908***	0.213

30 years	1.901***	0.220	1.729***	0.200
31 years	1.883***	0.226	1.717***	0.206
32 years	1.868***	0.233	1.711***	0.214
33 years	1.700***	0.222	1.567***	0.205
34 years	1.554***	0.213	1.447***	0.198
35 years	1.409**	0.204	1.327*	0.192
36 years	1.097	0.172	1.046	0.164
37 years	0.983	0.167	0.950	0.161
38 years	0.827	0.155	0.809	0.151
39 years	0.837	0.168	0.827	0.166
40 years	0.953	0.201	0.950	0.201
Registered Partnership	1.479***	0.019	1.465***	0.019
Child born prior to marriage	1.455***	0.009	1.461***	0.009
Spousal immigration background				
Husband native, Wife 1st gen.	1.368***	0.010	1.317***	0.009
Husband native, Wife 2nd gen.	1.389***	0.009	1.366***	0.009
Husband 1st gen. Wife native	1.921***	0.014	1.898***	0.014
Husband 1st gen. Wife 1st gen.	0.679***	0.004	0.689***	0.004
Husband 1st gen. Wife 2nd gen.	1.386***	0.016	1.344***	0.015
Husband 2nd gen. Wife native	1.376***	0.009	1.357***	0.009
Husband 2nd gen. Wife 1st gen.	1.216***	0.017	1.156***	0.016
Husband 2nd gen. Wife 2nd gen.	1.501***	0.022	1.464***	0.021
Age at wedding				
Husband, linear	0.888***	0.010	0.892***	0.010
Husband, quadratic	1.003***	0.000	1.003***	0.000
Husband, cubic	1.000***	0.000	1.000***	0.000
Wife, linear	1.042***	0.011	1.029***	0.011
Wife, quadratic	0.998***	0.000	0.998***	0.000
Wife, cubic	1.000***	0.000	1.000***	0.000
Education levels				
Husband, High School	0.877***	0.007	0.869***	0.007
Husband, University	0.626***	0.006	0.643***	0.006
Husband, Missing	0.836***	0.007	0.828***	0.007
Wife, High School	0.999	0.007	0.978***	0.007
Wife, University	0.744***	0.006	0.756***	0.006
Wife, Missing	0.601***	0.004	0.590***	0.004
Calendar year				
1996	0.970***	0.010	0.969***	0.010
1997	0.952***	0.012	0.950***	0.012
1998	0.976*	0.015	0.973*	0.015
1999	1.022	0.018	1.020	0.018
2000	1.090***	0.023	1.088***	0.023
2001	1.141***	0.028	1.138***	0.028
2002	1.111***	0.031	1.107***	0.031
2003	1.072**	0.034	1.067**	0.034

2004	1.081**	0.038	1.074**	0.038
2005	1.067*	0.041	1.057	0.041
2006	1.023	0.043	1.011	0.043
2007	1.009	0.047	0.995	0.046
2008	0.933	0.046	0.920*	0.046
2009	0.881**	0.047	0.869***	0.047
2010	0.892**	0.051	0.880**	0.050
2011	0.873**	0.053	0.860**	0.052
2012	0.902	0.058	0.888*	0.057
2013	0.929	0.064	0.912	0.062
2014	0.930	0.067	0.911	0.066
2015	0.907	0.069	0.886	0.067
Number of higher-order children by age				
age 0			0.280***	0.005
age 1			0.448***	0.006
age 2			0.577***	0.006
age 3			0.663***	0.007
age 4			0.734***	0.007
age 5			0.784***	0.008
age 6			0.812***	0.008
age 7			0.830***	0.008
age 8			0.839***	0.009
age 9			0.838***	0.009
age 10			0.830***	0.009
age 11			0.839***	0.009
age 12			0.854***	0.009
age 13			0.883***	0.010
age 14			0.892***	0.010
age 15			0.905***	0.011
age 16			0.924***	0.011
age 17			0.959***	0.012
age 18			0.974*	0.013
age 19			0.969**	0.015
age 20			0.979	0.016
age 21			0.951***	0.018
age 22			0.943***	0.020
age 23			0.914***	0.024
age 24 and older			0.923**	0.029
Number of higher-order daughters by age				
age 0			0.943**	0.023
age 1			0.970*	0.017
age 2			0.975*	0.015
age 3			1.009	0.014
age 4			0.991	0.013
age 5			0.982	0.013

age 6					0.989	0.013
age 7					0.980	0.013
age 8					0.976*	0.013
age 9					0.999	0.014
age 10					1.029**	0.014
age 11					0.992	0.014
age 12					1.015	0.015
age 13					1.012	0.015
age 14					1.050***	0.016
age 15					1.035**	0.016
age 16					1.061***	0.017
age 17					1.036**	0.017
age 18					1.041**	0.018
age 19					1.020	0.020
age 20					0.989	0.021
age 21					0.981	0.023
age 22					0.959	0.026
age 23					0.970	0.032
age 24 and older					0.935*	0.037
Constant	0.004***	0.000	0.004***	0.001	0.005***	0.002
Observations	30,180,829		36,541,693		36,541,693	
Marriage spells	2,174,182		2,722,223		2,722,223	
Cohort FE	NO		YES		YES	
ln likelihood	-1,964,835		-2,427,247		-2,412,257	

Note: Authors' estimates of exponentiated coefficients from cloglog hazard models of marriage durations. The models use linked marriage, divorce, and other registry data for different-sex couples who married after year 1971, and did not have children with other partners prior to the marriage.
*** = 0.01 significance level, ** = 0.05 significance level, * = 0.1 significance level.

Appendix Table A2. Regression results, CPS-MFS sample

	Age & gender only		Duration controls		More controls	
VARIABLES	exp(Beta)	St.e.	exp(Beta)	St.e.	exp(Beta)	St.e.
First-born daughter aged 0-12	1.026	0.022	1.028	0.022	1.021	0.022
First-born daughter aged 13-18	1.147**	0.073	1.149**	0.073	1.150**	0.073
First-born daughter aged 19-26	0.937	0.091	0.938	0.091	0.941	0.091
First-born's age dummies						
1 year	1.250***	0.057	1.162***	0.053	1.133***	0.052
2 years	1.282***	0.059	1.194***	0.057	1.143***	0.054
3 years	1.264***	0.059	1.164***	0.057	1.086*	0.053
4 years	1.260***	0.060	1.248***	0.064	1.146***	0.059
5 years	1.093*	0.055	1.139**	0.062	1.024	0.056
6 years	1.078	0.056	1.242***	0.070	1.100*	0.063
7 years	0.957	0.052	1.194***	0.072	1.042	0.063
8 years	0.998	0.055	1.348***	0.085	1.161**	0.073
9 years	0.872**	0.052	1.297***	0.088	1.103	0.075
10 years	0.815***	0.052	1.317***	0.096	1.108	0.081
11 years	0.786***	0.052	1.404***	0.109	1.166**	0.091
12 years	0.750***	0.052	1.538***	0.128	1.275***	0.106
13 years	0.588***	0.05	1.295***	0.127	1.067	0.105
14 years	0.671***	0.057	1.520***	0.153	1.246**	0.125
15 years	0.478***	0.045	1.165	0.132	0.954	0.108
16 years	0.621***	0.055	1.691***	0.188	1.395***	0.155
17 years	0.522***	0.050	1.546***	0.187	1.281**	0.154
18 years	0.625***	0.063	1.975***	0.259	1.648***	0.215
19 years	0.587***	0.065	1.890***	0.275	1.582***	0.227
20 years	0.573***	0.076	1.895***	0.321	1.589***	0.266
21 years	0.476***	0.064	1.578***	0.278	1.328	0.229
22 years	0.436***	0.064	1.452*	0.284	1.218	0.234
23 years	0.424***	0.063	1.579**	0.324	1.332	0.268
24 years	0.350***	0.062	1.351	0.322	1.134	0.265
25 years	0.371***	0.075	1.731*	0.488	1.447	0.402
26 years	0.295***	0.073	1.923*	0.649	1.594	0.530
No children dummy			1.436***	0.053	1.627***	0.061
Marriage duration dummies						
1 year			1.631***	0.064	1.654***	0.065
2 years			2.057***	0.08	2.098***	0.083
3 years			1.951***	0.079	1.998***	0.084
4 years			2.302***	0.093	2.370***	0.103
5 years			2.033***	0.086	2.096***	0.097
6 years			2.089***	0.091	2.162***	0.106
7 years			1.831***	0.084	1.891***	0.099
8 years			1.680***	0.081	1.731***	0.099
9 years			1.589***	0.081	1.626***	0.099
10 years			1.421***	0.077	1.444***	0.095
11 years			1.310***	0.076	1.321***	0.094
12 years			1.258***	0.078	1.257***	0.096
13 years			1.014	0.071	1.001	0.085

14 years			0.881*	0.066	0.862	0.078
15 years			0.903	0.069	0.872	0.082
16 years			0.882	0.075	0.843*	0.087
17 years			0.761***	0.070	0.710***	0.079
18 years			0.690***	0.068	0.634***	0.075
19 years			0.604***	0.067	0.543***	0.071
20 years			0.603***	0.071	0.534***	0.073
21 years			0.579***	0.077	0.503***	0.077
22 years			0.554***	0.077	0.470***	0.074
23 years			0.644***	0.103	0.536***	0.095
24 years			0.513***	0.091	0.417***	0.082
25 years			0.567***	0.104	0.449***	0.090
26 years			0.494***	0.113	0.382***	0.092
27 years			0.405***	0.097	0.304***	0.077
28 years			0.170***	0.066	0.126***	0.050
29 years			0.169***	0.093	0.123***	0.068
30 years			0.202***	0.112	0.135***	0.076
Mother's age at wedding - linear term					0.423***	0.042
Mother's age at wedding - quad. term					1.023***	0.004
Mother's age at wedding - cubic term					1.000***	0.000
Birth cohort 1940-1944					1.085**	0.036
Birth cohort 1945-1949					1.079*	0.049
Birth cohort 1950-1954					1.111*	0.068
Birth cohort 1955-1959					1.075	0.084
Birth cohort 1960-1964					1.001	0.098
Birth cohort 1965-1969					0.876	0.106
Birth cohort 1970+					0.806	0.135
Calendar years 1955-1959					1.011	0.163
Calendar years 1960-1964					1.087	0.173
Calendar years 1965-1969					1.413**	0.229
Calendar years 1970-1974					1.865***	0.314
Calendar years 1975-1979					1.807***	0.318
Calendar years 1980-1984					1.883***	0.350
Calendar years 1985-1989					2.049***	0.403
Calendar years 1990+					3.379***	0.707
Constant	0.015***	0.001	0.008***	0.000	79.164***	63.618
Observations	962,065		1,565,855		1,565,855	
Marriage spells	82,167		133,022		133,022	
ln likelihood	-109,143,653		-194,254,953		-182,229,045	

Note: Authors' estimates of exponentiated coefficients from cloglog hazard models of marriage durations. The model uses retrospective marital histories of American women aged 15-65, collected in CPS-MFS waves 1980, 1985, 1990 and 1995. The sample consists of first marital spells of women who were born in or after 1935, who married before the age of 16 and 45, and who did not give birth to any children prior to the date of the first marriage. *** = 0.01 significance level, ** = 0.05 significance level, * = 0.1 significance level.

Appendix B. LISS dataset characteristics and sample selection

The Longitudinal Internet Studies for the Social Sciences dataset consists of 4500 households comprising 11,500 individuals who are followed over 8 years (2008-2015). Each household member older than 15 years of age is surveyed individually. Children up to the age 15 do not participate actively, their presence in the household (and a basic set of characteristics) is reported by the parents.

Individual household members may opt out from the survey. On average, 15% of the household members opt out. Their presence and characteristics are reported by an appointed household member. The subjective responses of members who opted out are not solicited. In the key demographic of 16-19 year olds (that is, teenagers who are surveyed), the opt-out has been slightly higher, averaging 20%. Conditional on participating, the response rates of household members are good, averaging 75-80%.

Our sample is restricted to couples with children who are married and whose first-born is younger than 19 years of age, is alive, is neither adopted nor a step-child, and lives in the same household as the parents. This sample consists of 6,603 person-year records of participating parents, and 632 person-year records of participating first-born teenagers (aged 16-18). The extended sample of all participating teenagers (regardless of birth parity) contains 1,178 person-year records. The loss of parental observations due to the sample restrictions is documented in Table B1.

Table B1. Person-year records of coupled individuals who live with at least some of their children in the same household, LISS panel, years 2008-2015

Sample	Number of person-year records
Coupled adults in LISS households who live with their children	24,467
Out of whom:	
Adults who participate in the survey	19,084
Adults who filled out the family module (necessary to identify biological children)	14,507
Out of whom:	
Adults who share biological first-born with their current partner	12,886
Adults whose biological first-born is alive and at most 18 years old	8,079
Adults whose biological first-born is at most 18 years old and lives in the same household	8,029
Out of whom:	
Adults who are married	6,603
Adults who are married and their first-born is a teenager	2,521

The numbers of observations corresponding to individual regression models listed in Supplemental Appendix Table S2 may differ from the numbers of person-year records corresponding to the selected sample. This is partially due to individual non-response to specific questions, and partially due to changes to the structure of the LISS questionnaire across waves. Several questions have been asked only in a subset of waves (columns 7-22), which lowers the numbers of observations. Furthermore, the questions regarding first child's behaviour were asked to a random subset of families, and the questions regarding expenditures were asked only to the adult household member who is usually responsible for shopping.