Full title: Daughters and Divorce*

Short title: Daughters and Divorce

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Abstract: Are couples with daughters more likely to divorce than couples with sons? Using Dutch registry and U.S. survey data, we show that couples with daughters face higher risks of divorce, but only when daughters are 13 to 18 years old. These age-specific results run counter to explanations involving overarching, time-invariant preferences for sons and sex-selection into live birth. We propose another explanation that involves relationship strains in families with teenage daughters. In subsample analyses, we find larger child-gender differences in divorce risks for parents whose attitudes towards gender-roles are likely to differ from those of their daughters and partners. We also find survey evidence of relationship strains in families with teenage daughters.

Keywords: Marriage, divorce, gender, son preference, Netherlands, registry data

JEL Codes: J12, J13, J16

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Research by sociologists (Spanier and Glick, 1981; Morgan *et al*., 1988) and economists (*e.g.*, Bedard and Deschênes, 2005; Dahl and Moretti, 2008) has found that, in the U.S., couples with daughters face modestly higher risks of divorce than couples with sons. Their findings seem to complement evidence from developing and non-Western countries that preferences for sons alter family behaviour (Ben-Porath and Welch, 1976; Das Gupta *et al.*, 2003), with Dahl and Moretti (2008, p. 1112) concluding "that the age-old favouring of boys is not confined to the past."

However, associations between children's genders and divorce do not appear in all studies of U.S. data (Morgan and Pollard, 2002; Diekmann and Schmidheiny, 2004), and associations have not been detected in other Western countries (*e.g.*, Andersson and Woldemicael, 2001; Leigh, 2009). Thus, there is a question as to whether the associations found in some U.S. studies are idiosyncratic or whether they reflect a fundamental characteristic of modern families.

There is also a question of why associations between children's genders and divorce might exist. Dahl and Moretti (2008) developed evidence of overarching preferences for sons. Morgan *et al*. (1988), Katzev *et al*. (1994), and Lundberg *et al*. (2007) found evidence of greater father involvement with sons. More broadly, Lundberg (2005) has differentiated between explanations based on preferences (including parents' preferences for sons and fathers' preferences for spending time with sons), and explanations based on constraints (including higher time or money costs of raising girls or worse developmental consequences of divorce for boys). Besides the causal mechanisms, sex-selection into live birth has been also proposed to explain the associations (Hamoudi and Nobles, 2014).

In this article, we examine how children's gender affects the risk of divorce among couples in the Netherlands, using administrative data that cover the near universe of marriages and registered partnerships that began in the country between October 1971 and

December 2015. The data include nearly 3 million marriages and partnerships, allowing us to estimate effects precisely and consider how effects vary with children's ages, parities, and conditional on parents' backgrounds. The data are highly accurate with the exact dates of weddings, births, and divorces. This contrasts with most previous studies which have relied on retrospective self-reports that are subject to recall errors and other misreporting (Mitchell, 2010), and with several studies which could not fully identify parenting relationships or children's parities (*e.g.*, Bedard and Deschênes, 2005; Dahl and Moretti, 2008). Unlike most previous economic studies, we estimate event-history models that: account for the duration of the marriage or registered partnership; allow the risks of divorce to change with children's ages; and accommodate right-censoring in marriage spells.

We find that having a daughter increases the risks of divorce among Dutch couples—the first robust finding of this association for a European country. An even more novel and intriguing result is that the increased risks of divorce only appear when daughters are teenagers (aged 13-18)—there is no detectable gender difference at earlier or later ages. We observe this pattern among both firstborn and subsequent children. We also find the same age pattern in analyses of the 1980, 1985, 1990 and 1995 U.S. Current Population Survey Marriage and Fertility Supplements (CPS-MFS).

From ages 13 to 18, the divorce risks for couples with firstborn daughters are about five percent higher, on average, than the risks for couples with firstborn sons. However, because there are no differences at younger or older ages, the cumulative effect of the child's gender on parental divorce is modest. For instance, the cumulative divorce rates of couples with firstborn sons and daughters aged 19 only differ by 1.8 percent (0.36 percentage points).

Our results prompt us to revisit the explanations for gender-related divorce disparities. The 'teenage daughter effect' is at odds with many possible explanations, including i) overarching, time-invariant preferences for sons, ii) sex-selection into live birth, and iii)

rational forward-looking behaviours based on age-specific differences in preferences for or costs of raising sons and daughters. This is because each of these explanations predicts that gender disparities would also appear earlier in the child's life. Instead, the precisely-estimated pattern of zero effects during childhood and non-zero effects during the teenage years leads us to consider explanations based on unexpected changes in parents' valuation of marriage during children's adolescence. We give special consideration to family conflicts that might arise from differences in family members' gender-role attitudes. These differences can become a more salient source of conflict as girls mature, and conflict in this dimension of family life may spill over to the parents' marital relationship.

Additional analyses of Dutch administrative data support this explanation. We find that the excess divorce risks associated with having teenage daughters are higher for couples whose gender-role attitudes are likely to differ from those of their daughters (for example, for less-educated couples and immigrant couples), and that the risks are exacerbated further for couples in which the two spouses are likely to hold different gender-role attitudes (for example, for parents with different immigration backgrounds or different levels of education). In a further analysis, we examine the gender composition of the parents' siblings and find that the teenage daughter effect only appears among fathers who grew up without sisters. In contrast, we find no differences with regard to the gender composition of the mother's siblings, which suggests that the father's experiences are critical.

Survey data also support this explanation. We analyse the Longitudinal Internet Studies for the Social Sciences (LISS) panel, a large survey of Dutch households, and uncover direct evidence of relationship strains between fathers and teenage daughters, more disagreements over parenting among couples with teenage daughters, and more favourable attitudes towards divorce among mothers of teenage daughters.

1. Related literature

1.1 *Empirical findings*

A host of U.S. studies, including Spanier and Glick (1981), Morgan *et al*. (1988), Bedard and Deschênes (2005), Ananat and Michaels (2008), Dahl and Moretti (2008), and Mammen (2008), have found that parents of daughters face higher risks of divorce or single parenthood than parents of sons. However, analyses of U.S. data by Morgan and Pollard (2002), Diekmann and Schmidheiny (2004), and Reichman *et al*. (2004) and analyses of data from other Western countries by Andersson and Woldemicael (2001), Diekmann and Schmidheiny (2004), Leigh (2009), Flouri and Malmberg (2010) did not find significant differences.

To gain more insight into why the studies reach different conclusions, we have collected multiple characteristics from these studies, including their data sources, outcomes of interest, and sample restrictions. Table 1 lists these characteristics together with the baseline divorce rate observed in the respective samples and estimates of the relative and absolute effects of having a daughter on the probability of divorce (and/or separation). For the sake of comparability, we restrict our attention to results corresponding to firstborn children.[1]

First, we note that when significant differences have been detected, they have tended to be modest. Estimates of the differences from U.S. Census data by Bedard and Deschênes (2005), Ananat and Michaels (2008), and Dahl and Moretti (2008) range from 0.5 percentage-points (2.2 percent) to 0.8 percentage-points (4 percent). A few larger but less-precise differences have been estimated with smaller data sets, including a 1.0 percentage-point difference in Dahl and Moretti's (2008) analysis of the CPS-MFS. The prevalence of

[1] In Spanier and Glick (1981), Morgan *et al*. (1988), Andersson and Woldemicael (2001), and Flouri and Malmberg (2010), the firstborn results correspond to families with one child only. Morgan and Pollard (2002) assume that the child-gender associations are the same for firstborn children and children of higher parities. We exclude the study by Reichman *et al*. (2004) because it focuses on a selective sample of mostly unwed families.

modest-sized estimates suggests that big samples are needed to detect differences reliably. Among the studies with insignificant findings, only Andersson and Woldemicael (2001) used a large sample.

Second, the magnitudes of the reported differences vary with the data restrictions the studies put on children's ages and marriage durations. Studies that include families with teenage children tend to find larger and more significant differences than studies that only consider families with younger children. The U.S. Census analysis by Dahl and Moretti (2008) focused on households in which the oldest child was age 12 or younger and estimated relatively small firstborn daughter differences. Ananat and Michaels (2008) and Bedard and Deschênes (2005) considered households with children up to 16 or 17 years of age and found larger differences. Supplementary analyses by Dahl and Moretti (2008) showed that the absolute differences in divorce risks were small and only marginally significant near the time of the child's birth but increased as the child aged.

One explanation for these patterns is mechanical: samples of parents with older children contain more divorces and so we might expect to find larger absolute differences in their divorce rates, assuming that the relative effect of child's gender on divorce is time-invariant. However, the relative effect does not appear to be time-invariant. Studies that focus on young children find the relative effect to be either small or missing altogether; studies that also consider families with teenage children tend to find larger relative effects; and studies that consider families with adult children report relative effects that fall back in magnitude. This leads us to consider the possibility that the association between children's genders and parents' divorce risks is time-varying, and possibly age-specific.

1.2 *Possible channels*

A few theoretical channels predict an association between child gender and parental divorce. We frame our discussion of these channels using Weiss' (1997) rational-choice

model of marriage (see also Becker *et al.*, 1977). Weiss theorised that couples choose whether to marry and subsequently whether to remain married by comparing the expected value of being married to the value of being single or in another relationship. The expected value of being married depends on the characteristics of each spouse, match-specific quality, and marriage-specific capital (which includes the number and quality of their children). In each period, the couples know the current values of these variables but do not know all the future values. Divorce occurs when one or more variables is subject to a sufficient change, such as a drop in match-specific quality or a change in outside opportunities.

We can examine Lundberg's (2005) preference- and constraint-based explanations through the lens of Weiss' model. Consider an overarching preference for boys over girls. Such a preference would raise the value of the marriage-specific capital for couples with sons, thereby lowering their incentives to divorce. It could also lead couples to invest more in boys' development, further increasing the value of this capital. Similar effects would occur if fathers prefer spending time with sons more than spending time with daughters (Lundberg *et al.*, 2007; Mammen, 2011). These preferences may be time-invariant, but they also may be stronger at different ages. Lundberg (2005) and Raley and Bianchi (2006) summarise evidence that fathers' preferences for sharing activities with sons are stronger when the sons are school-aged. This would strengthen the marriages with sons at the respective ages. However, if parents are rational and forward-looking, it would also strengthen marriages with sons at earlier ages because the fathers would need to stay in the marriages to realise the benefits.

The constraint-based explanations posit that daughters are more costly to raise than sons. Baker and Milligan (2016) found that parents spend more time in teaching activities for pre-school girls than pre-school boys. Durante *et al.* (2015) conducted experiments in which people allocated more resources to daughters, and Moffitt and Ribar (2018) found that

disadvantaged families were more protective of daughters' food needs. Higher costs of daughters would reduce the amount of marriage-specific capital that a couple could produce, weakening the incentives to remain married. Another possibility is that boys are more susceptible to developmental problems if parents divorce, which would lower the value of parents' alternatives to marriage (see the reviews by Lundberg, 2005, and Raley and Bianchi, 2006, and findings by Bertrand and Pan, 2013). It is also possible that parenting interactions are more strained with girls than boys and that these strains lower the match-specific quality. Morgan *et al*. (1988) reported that fathers had more distant relationships with their adolescent daughters than their adolescent sons, and VanderValk *et al*. (2007) found that adolescent girls' emotional problems strained their parents' marriages but boys' problems did not. Alternatively, adolescent girls may perform more housework or require less supervision than same-age boys (Kalenkoski *et al.*, 2011), which might directly reduce the costs of divorce or indirectly reduce them by allowing mothers to work more in the labour market.

As with the preference explanations, age-related differences in the costs of sons and daughters could lead to age-specific divorce patterns. However, higher costs at later ages would also affect divorce risks at earlier ages, assuming that the costs are anticipated. To entertain the possibility that the gender differences in divorce risks appear only at later ages, we may have to abandon the assumption of rational forward-looking parents. If we assume that parents are unable to anticipate the age-specific effects of sons and daughters, we can invoke several of the explanations listed above. For example, older boys may require more-than-expected amounts of supervision by the parents, which strengthens the marriages at later ages (this is in line with the "compensation hypothesis" of Dahl and Moretti, 2008). Parents may also not anticipate having strained interactions with their adolescent daughters, or they may not consider how these interactions could affect their marriages. Another possibility is that the presence of older sons reinforces traditional gender-role or family attitudes (Morgan

et al., 1988), thereby making the marriages more stable; whereas the presence of older daughters might contribute to more egalitarian attitudes (Warner 1991; Washington 2008; Oswald and Powdthavee 2010; Shafer and Malhotra 2011; Borrell-Porta *et al.* 2019) which may have the opposite effect.[2]

The explanations listed so far assumed that the child's gender causally affects divorce. However, it is also possible that the association is not causal. 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. This means that the gender-related divorce disparities may result from sex-selection into live birth.

2. Context and Data

2.1 *Marriage and divorce in the Netherlands*

The Netherlands has operated under a no-fault divorce law since 1971 (Boele-Woelki *et al.*, 2003). In 1998, the country introduced registered partnerships 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 their municipality at least two weeks before the wedding or partnership occurs.

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, 2017). New different-sex partnership registrations rose from fewer than 2,000 in 2001 to about 13,000 in 2015. The marriage trends are similar to those in the U.S. and other Western countries.

[2] There is some contradictory evidence—Katzev *et al*. (1994) and Perales *et al*. (2018) have found that parents of daughters have more traditional attitudes than parents of sons.

Divorce in the Netherlands requires a formal legal proceeding, but dissolution of a registered partnership does not. Between 2001 and 2009, married couples could also take advantage of a "flash divorce" procedure under which they could bypass the legal proceeding by converting their marriages into registered partnerships and then quickly dissolving the partnerships (Kabátek 2018). Divorces and partnership dissolutions in the Netherlands take effect once they have been recorded in the municipal register. Overall, the determinants of divorce in the Netherlands are similar to those in other countries (see De Graaf and Kalmijn 2006). Throughout the rest of the article, we will use the term "divorce" to refer to both divorces and partnership dissolutions.

2.2 *Dutch Registry Data*

The primary source of data for our analysis is the Dutch municipal register. The register data are maintained by Statistics Netherlands and cover all Dutch residents between 1995 and 2015. The data 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 (or were) married or registered partners and create couple-specific measures.

The registers also allow us to match children to their legal parents, starting with children born in 1966. Each child is assigned paternal and maternal identifiers, provided that the parental records are present in the data. Most children can be linked to both parents; only 7% cannot, usually because the father's identifier is missing. Parental records may be missing if the parent died before 1995, the parent lives in another country, or the child has no legal mother or father. Using the parent-child identifiers, we construct measures of the gender- and age-composition of each couple's children.

Our empirical analyses focus on first different-sex marriages and registered partnerships that began between 1 October 1971 (the date when the Netherlands adopted no-

fault divorce) and 1 January 2016. We drop marriages and registered partnerships of couples who had children with prior partners, who adopted children, whose firstborn children were twins, or in which either spouse was previously married. We also drop marriages and partnerships if either spouse was born before 1935 or after 1985. 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 ended before 2016, we also observe the termination dates and whether the relationships ended because of divorce, death, or a change of civil status. Our resulting dataset includes 2,950,986 first marriages and registered partnerships. These represent nearly all the first marriages and registered partnerships that began within our analysis window among people in our selected birth cohort.[3] Our results are not sensitive to changes in the cut-off years for the analysis window.

We use these data to create marriage and partnership spells, and we analyse the annual hazard risks of divorce or dissolution. Each spell begins on the day of the wedding or registration and continues until a divorce or censoring event occurs. Spells are censored on 31 December 2015, when a spouse dies or emigrates, when the marriage duration reaches 40 years, or when the firstborn child reaches age 27. Our results are not sensitive to alternative censoring cut-offs.

3. Main Results

3.1 *Characteristics of Parents*

We begin our analysis by examining whether there are differences in the average observable characteristics of couples with firstborn sons and daughters. This helps us to determine whether associations between the firstborn child's gender and divorce are causal.

[3] We are unable to observe marriages of people who died or left the Netherlands before 1995. Using a competing-risk model of mortality and emigration decisions, we estimate that this leads to a loss of approximately 4% of marriage records. In the robustness section, we show that our results are unlikely to be affected by this selectivity.

The observable characteristics include each parent's birth year, age and calendar year at wedding/registration, age and marriage/partnership duration at first childbirth, immigration background, education, and—for children born after 1999—the parents' employment status and earnings in the year before the birth. We also observe the registered partnership status, whether the child was born before the start of marriage/partnership, and whether the relationship ended in divorce. Table 2 lists averages conditional on the child's gender, boy-girl differences in the averages, and p-values for the corresponding pairwise t-tests.

There are no large differences and few statistically significant differences between the average characteristics of couples with firstborn sons and daughters. Marriage durations at childbirth were 0.1 months (3 days) longer for parents of firstborn daughters. Parents of firstborn daughters are 0.11% more likely to both be natives and 0.15% less likely to both be first-generation immigrants. Mothers of firstborn daughters are slightly less likely to have primary or tertiary educations and more likely to have a secondary education. These results provide little evidence of sex-selection into live birth on the basis of observable characteristics.

The nearly identical shares of firstborn sons and daughters who were born before the start of the marriage or partnership suggest that the incidence of so-called shotgun weddings is not affected by the firstborn's gender. An auxiliary test (not shown) reveals that there were no gender differences in legitimisation rates among firstborn children who were born out of the wedlock. The findings of no significant associations differ from those of Dahl and Moretti (2008) and imply that our results should not be affected by selective entry into marriage.

Gender-specific differences do appear, however, in other outcomes that follow the births. First, parents of firstborn daughters have slightly *fewer* children than parents of firstborn sons. Although this association is statistically significant, its direction is inconsistent with an overarching preference for sons. Interestingly, the same association has been recently

found in the U.S. (Blau *et al*., 2020). Second, parents of firstborn daughters are more likely to divorce, with their divorce rate being 0.27 percentage points (1.3% percent) higher than the rate for parents of firstborn sons.

3.2 *Divorce Hazards for Parents of Firstborn Sons and Daughters*

To explore the divorce association further, we turn to event-history analysis. Specifically, we estimate complementary log-log (cloglog) discrete-time hazard models of divorce. The key advantage of cloglog models is that they allow us to estimate the extent to which the divorce association varies with children's ages. They also allow us to jointly model several duration-dependent processes (including the duration of the marriage/partnership and the couple's realised fertility), and control for other observable characteristics. The functional form of the cloglog model is

$$\Pr[y_{it} = 1 \mid \mathbf{x}_{it}] = 1 - \exp\left(-\exp\left(\mathbf{x}_{it}'\boldsymbol{\beta}\right)\right), \tag{1}$$

where the hazard probability of a divorce y_{it} for a couple i observed in year t is defined as a function of covariates $\mathbf{x}_{it}$ that are specific to the given couple and year. The influence of the covariates on the hazard probability is captured by a vector of parameters $\boldsymbol{\beta}$, which is estimated by maximum likelihood (the log-likelihood function is included in Appendix Section 1). The model works with discrete time, and so we split the original dataset of 2,950,986 first marriage/partnership spells into a dataset of marriage-year and partnership-year observations. The resulting dataset contains 57,267,483 observations.

Our principal model is a multivariate hazard model of divorce that controls for the presence of children, age of the firstborn child, gender of the firstborn child, duration of marriage, and a set of other observable characteristics. We specify the age pattern non-parametrically by including dummy variables for the child's calendar age (k, measured in years), and we interact the age dummies with an indicator for the child being a daughter.

Similar to the age pattern, divorce risks at different marriage durations are approximated by a set of yearly dummies. Other controls include each partner's immigration background, education, and age at marriage; being in a registered partnership; the first child being born pre-maritally; and cohort and year effects.[4] Note that the model does not control for the genders and ages of higher-parity children because of concerns regarding the endogeneity of the corresponding fertility outcomes (though we examine such births later).

The corresponding specification of $\mathbf{x}_{it}'\boldsymbol{\beta}$ is

$$\begin{aligned}\mathbf{x}_{it}'\boldsymbol{\beta} = \beta_0 + \sum_{k=0}^{26}\left[\mathbf{1}\left(FB\,age_{it}=k\right)\cdot\left(\beta_{1k}+\beta_{2k}\cdot FB\,daughter_i\right)\right] + \beta_3\cdot Childless_{it} \\ + \sum_{j=1}^{40}\beta_{4j}\cdot\mathbf{1}\left(Duration_{it}=j\right) + \beta_5\cdot RegPar_{it} + \sum_{y=1972}^{2015}\beta_{6j}\cdot\mathbf{1}\left(Year_{it}=y\right) + \mathbf{z}_i'\boldsymbol{\beta}_7,\end{aligned} \tag{2}$$

where the set of time-invariant characteristics is represented by the vector $\mathbf{z}_i$.
We set the coefficient β_{10} to zero, meaning that the reference group for the age coefficients (β_{1k}) and the childless dummy coefficient (β_3) consists of couples whose firstborn sons are less than one year old (k=0). The reference groups for the interacted daughter-age coefficients (β_{2k}) consist of couples whose firstborn sons are in the respective age categories (k). Throughout the article, the coefficient estimates are presented in the exponentiated form and can be interpreted as approximate odds ratios. Standard errors corresponding to these estimates are robust to heteroskedasticity. We derive them by bootstrapping, using 1000 draws from the underlying parameter distribution.

Estimates of the age-specific coefficients are graphed in Figure 1 (detailed results can be found in Appendix Table A1, column 2). The first panel shows how the hazard probabilities of divorce vary with the ages of firstborn sons and daughters. Conditional on marital duration and other observables, the hazard probabilities increase until the firstborn

[4] To show that the child-gender patterns are not affected by the inclusion of other control variables, we estimate an auxiliary model which excludes the controls for marriage duration and other observables. Results corresponding to this model can be found in column 1 of Appendix Table A1

reaches age 18 or 19 and fall afterwards. This pattern is consistent with parents trying to delay divorce until the children are adults. It is also consistent with increasing marital stress during children's adolescence and diminished stress when young adult children leave the household and parents become "empty nesters" (Heaton, 1990; Hiedemann *et al*., 1998).

With respect to the gender differences, we see that couples with firstborn sons and daughters follow nearly identical trajectories through the children's 12th year. At age 13, however, the trajectories diverge, with parents of firstborn daughters facing significantly higher risks of divorce. The dots indicate the ages at which the divorce risks for couples with firstborn daughters differ significantly from those with firstborn sons. After age 19, the risks are again indistinguishable.

The magnitudes of these disparities are shown in the second panel of Figure 1, which plots the excess age-specific hazard probabilities of divorce for couples with firstborn daughters relative to couples with firstborn sons. From ages 13 to 18, firstborn daughters are associated with 3.8% to 9.2% higher hazard probabilities of divorce, which amounts to 5.5% higher hazard probabilities on average. There are no differences before or after those ages.

Other coefficient estimates, which are listed in Appendix Table A1, reveal that childless couples face the same hazard probabilities as parents whose firstborn children are 12 years old. For couples with children, the risks are higher if the first child was born prior to marriage. Couples in which both spouses immigrated to the Netherlands have lower risks than native-Dutch couples, while couples with mixed immigration backgrounds (e.g., an immigrant married to a native) have higher divorce risks. Divorce risks are negatively associated with the spouses' education levels and positively associated with calendar time.

These estimates help put the magnitudes of the teenage daughter differences into context. The 5.5% increase of hazard probabilities of divorce faced by parents with firstborn daughters aged 13 to 18 is just under half the size of the risk reduction associated with the

husband completing secondary schooling, and it is twice as large as the average year-on-year difference in hazard probabilities over the period of observation.

To provide more interpretable magnitudes of the child-gender associations, we calculate cumulative divorce rates by the time firstborn children reach ages 12, 18, and 26 (Appendix Section 2 describes how these were derived). In line with the coefficients in Figure 1, the cumulative divorce rates through the child's 12^{th} year do not differ significantly by the child's gender. The rate for parents of 12-year-old firstborn girls is 15.63%, which is 0.08 percentage points higher than the rate for parents of firstborn boys. The disparity widens to 0.36 percentage points when the firstborn children reach 18 years of age, with cumulative divorce rates of 20.12% for parents of firstborn sons and 20.48% for parents of firstborn daughters. By age 26, the cumulative divorce rates increase to 25.69% and 26.02%, respectively, with the disparity narrowing slightly to 0.33 percentage points.

3.3 *Divorce Hazards Associated with Higher-Parity Children*

Next, we estimate a model that adds controls for the ages and genders of higher-parity children. The extended specification helps us see whether the firstborn gender associations are mediated through subsequent fertility outcomes, and whether higher-parity children are associated with the same age and gender patterns of divorce risks as firstborns. To control for the ages and genders of higher-parity children, we use count variables instead of dummies. This is to account for twins and closely-spaced siblings. The full specification and additional modelling details are presented in Appendix Section 3. The resulting estimates of the age-specific coefficients for both firstborn and higher-parity children are graphed in Figure 2.

The first two panels of Figure 2 show the gender-specific age profiles of the conditional hazard probabilities for firstborn and higher-parity children (coefficient estimates are reported in Appendix Table A1, column 3). The estimates continue to indicate that firstborn girls are associated with higher divorce risks at ages 13 to 18 and not at other ages,

which tells us that the gender association for firstborn children is not mediated through subsequent fertility. Perhaps of more interest, the age profiles for higher-parity children exhibit strikingly similar patterns to those of firstborn children, both in terms of timing and magnitude (the magnitudes are presented in the latter two panels of Figure 2). Higher-parity daughters aged 13 to 17 are associated with significantly higher hazard probabilities of divorce, and these probabilities are on average 5% higher than those faced by couples with higher-parity sons of the same age. Higher-parity daughters are not associated with higher divorce risks at other ages.

Our age-specific findings can reconcile some of the conflicting findings reported by earlier studies. They may explain why several studies that focused on parents of young children estimated insignificant daughter effects and why estimates of daughter effects tend to increase in studies with older children. The age pattern in the results can also help us distinguish among mechanisms that have been proposed to explain the child-gender association. The absence of gender differences in infancy and early childhood speaks against theories that assume the existence of overarching, time-invariant preferences for sons. It also runs counter to hypotheses that involve forward-looking parents with age-specific preferences for spending time with their sons, and parents facing foreseeable differences in children's costs or other constraints. Sex selection into live birth is also incompatible with the presented patterns, which suggests that the associations between children's genders and divorce can be interpreted causally.

The age patterns and the fact that they appear for both firstborn and higher-parity children strongly point to explanations that involve unexpected changes in parents' preferences, constraints, or family processes during children's adolescence. Family process explanations in the form of gender-differentiated strains during the teenage years seem plausible. Family relations research points to ways that parent-child conflict and other family

stresses can strain spousal relationships (Whiteman *et al*., 2007; Cui *et al*., 2009), and there is evidence that these effects are stronger in families with adolescent daughters than adolescent sons (VanderValk *et al*., 2007; Webb *et al*., 2017). A source of gendered and age-patterned conflict may be differences among family members in their attitudes towards gender roles (Marks *et al*., 2009). These differences may become more salient as girls mature, gain autonomy, become psychologically independent, and improve their argumentation skills (Cui *et al*., 2009), and the resulting conflict may increase the risks of divorce for couples with daughters.[5]

In the next two sections, we look for empirical evidence in support of this explanation, and we also evaluate the plausibility of other candidate explanations that involve unexpected changes of parents' marriage valuations. Alternative explanations that do not involve unexpected changes of parents' marriage valuations are discussed and evaluated in Appendix Section 4.

4. Subgroup analyses

4.1 *Simple subgroup analyses*

To examine whether the excess divorce risks associated with teenage daughters might stem from differences in gender-role attitudes, we estimate hazard models of divorce for various subgroups that are more or less likely to experience such differences. We use a simplified version of the cloglog model without controls for higher-parity children (Equation 2). We keep all the covariates from that specification except for the 27 age-specific interactions of the firstborn-daughter indicator, which are replaced by three age-group

[5] Earlier research indicates that the most common source of child-related conflict during adolescence is parental control over child's personal choices, such as appearance, relationships, and employment (Smetana 1995, Nucci *et al*. 2005, Sorkhabi 2010). Gendered differences in this domain (*e.g.*, daughters being subject to more parental control than sons) can give rise to gendered patterns of conflict and marital instability. Gendered differences in other domains, such as moral judgements (Arutyunova *et al*., 2016; Capraro and Sippel 2017), political values (Pratto *et al*., 1997; Reed, 2006), or sexual attitudes (Oliver and Hyde, 1993) might have the same effect, although they are generally viewed as less potent sources of child-related conflict (Smetana 2005, Sorkhabi 2010).

interactions corresponding to the firstborn daughter being a child (aged 0-12), teenager (aged 13-18), or young adult (aged 19-26). The reduced set of interaction coefficients simplifies comparisons across subgroups. Table 3 lists coefficients and standard errors corresponding to the subgroup models, with the age-group interactions for daughters aged 0-12 and 13-18 listed in the first two columns. We focus on these two coefficients because the young-adult coefficients are estimated imprecisely in several of the smaller subsamples. The third and fourth columns of Table 3 list couples' predicted cumulative probabilities of divorce by the time their firstborn son or daughter is 18 years old. The fifth column lists daughter-son differences in these cumulative probabilities. Detailed results are reported in Appendix Table B1.

As a benchmark, we estimate the simplified cloglog model for the full sample of couples and report the results in the first row of Table 3. In line with our principal model specification, the firstborn daughter interaction for children aged 0-12 is small (0.2%) and insignificant, while the interaction for teenagers is positive (5.5%) and significant. The restricted model also yields the same cumulative divorce rates as our principal specification.

Our first set of subgroup models splits the data by the couples' immigration backgrounds. We distinguish couples in which the spouses/partners are both native Dutch, are both immigrants, or have mixed immigration backgrounds (that is, one spouse is a native and the other is an immigrant). We expect that gender-role disagreements are more likely if daughters are raised in a different cultural context than their parents and that these disagreements may be more pronounced if the parents were raised in different cultural contexts from one another. The results are consistent with these expectations. The firstborn teenage daughter effect is larger among immigrant parents than native-Dutch parents, though not significantly so. The largest effect appears for parents with mixed immigration backgrounds, with the estimated coefficient being significantly different from the coefficients

for couples in which neither or both spouses are immigrants.[6] Among the one-immigrant couples, we do not find significant differences between husbands or wives being the immigrant (see Appendix Table A2).

Next, we examine the birth years and educational attainments of the spouses/partners. Gender-role attitudes have become more egalitarian over time, and so we expect larger gender effects among parents born in earlier cohorts. This is confirmed by the data: fathers born before 1956 have larger teenage first-daughter effects than fathers who were born more recently. Gender-role attitudes also tend to be more egalitarian among people with more schooling (Warner, 1991; Treas *et al*., 2014), and more educated parents might also be better informed about parent-child relationships or more capable of navigating potential difficulties. Consistent with this and previous findings (Spanier and Glick, 1981; Dahl and Moretti, 2008), we find that the magnitudes of the teenage first-daughter effects decrease with fathers' education levels. Analyses that consider the mothers' birth years and education levels produce similar findings (see Appendix Table A2). Corroborating the results for parents with mixed immigration backgrounds, the gender effects are more pronounced among parents who differ in terms of educational attainment (see Appendix Table A2).

To examine the within-couple differences in observable characteristics more comprehensively, we construct a summary index of couple dissimilarity by applying the Mahalanobis distance formula—a weighted, generalised quadratic formula that transforms multi-dimensional distances into a univariate metric—to the spouses' birth years, education levels, and immigration backgrounds (see Kabátek and Ribar [2018] for a previous application). For this analysis, we restrict the sample to couples with non-missing education information, and we split the sample by terciles of the dissimilarity distribution. The results

[6] When we further split the data by the parents' countries of origin, we find that immigrant parents with European backgrounds face similar teenage daughter effects as native-Dutch parents, whereas immigrant parents with non-European background face substantially larger effects. Results are available upon request.

indicate that the firstborn daughter effects rise with couple dissimilarity. The teenage first-daughter coefficient is not significantly different from zero for couples in the lowest tercile of dissimilarity; it is marginally significant for couples in the middle tercile; and it is large and statistically significant (both from zero and from the other coefficients) for couples in the highest tercile.

The final subgroup analyses listed in Table 3 examine whether the firstborn teenage daughter effect varies with the gender composition of the parents' siblings. Growing up with an opposite-sex sibling may provide insights into mixed-gender family relationships. It may also affect the development of gender norms. Crouter *et al*. (2007) found that, among families with traditional parents, firstborn sons who grew up with brothers had more traditional gender-role attitudes at age 19 than firstborn sons who grew up with sisters; however, they did not find differences in households with less traditional parents. In contrast, Healy and Malotra (2013) found that U.S. men who grew up with younger sisters had more traditional gender-role attitudes. We link the records of native-Dutch spouses who were born after 1965 with the records of maternal siblings who were born within 10 years of them, and we estimate the model separately for spouses who did and did not grow up with siblings of the opposite gender. The gender composition of the father's siblings has a striking influence on the teenage daughter effect. The effect for fathers who grew up without sisters is significant and nearly twice the size of the effect for the general cohort born after 1965, whereas the effect for fathers who grew up with one or more sisters is insignificant and effectively zero. Among the mothers, there are no significant differences in the teenage first-daughter effects between mothers who grew up with and without brothers.

4.2. *The range of effect heterogeneity*

The preceding analyses reveal that even though the teenage daughter effect is moderate on average, it is substantially stronger for some subgroups. To investigate the range

of this heterogeneity, we have estimated the subgroup model for 96 distinct subpopulations, splitting the dataset into smaller groups according to the combinations of parental characteristics. We estimate 48 models which are conditioned on combinations of the fathers' characteristics and 48 models conditioned on combinations of the mothers' characteristics.

The results are presented graphically in Figure 3, using specification charts that rank the estimated teenage daughter effects by magnitude and catalogue the composition of the subpopulation corresponding to each estimate. The estimates vary from modest negative to large positive effects. For 25% of couples, the hazard probabilities of divorce associated with a firstborn teenage daughter are at least 7% higher than the risks associated with a firstborn teenage son. None of the negative effects is significantly different from zero.

The effects for parents who are native, educated and relatively young tend to be very close to zero. This is to be expected as these are the parents whose gender-role attitudes are most likely to be aligned with the attitudes of their daughters. The largest effects, on the other hand, occur among less-educated parents in couples with mixed immigration backgrounds.

The ranges of effects corresponding to fathers' and mothers' characteristics are relatively similar in magnitude. This likely reflects the high degree of assortative mating on parental education and age. Nevertheless, the effects associated with fathers' characteristics vary somewhat more than the effects associated with mothers' characteristics.[7]

5. Survey evidence

To look for more direct evidence of the mechanisms responsible for the teenage daughter effect, we turn to the Longitudinal Internet Studies for the Social Sciences survey panel. The LISS followed a representative sample of Dutch households, totalling 11,500 individuals, for nine years and asked household members about the quality of their

[7] A formal variance-comparison test reveals that this difference is insignificant. However, if we expand the analysis to account for parental sibship, the variance of the effects for the father's characteristics is significantly larger than the variance for the mothers' characteristics.

relationships, attitudes, time-use, and expenditures. We analyse these outcomes among married couples whose firstborn is younger than 19 years of age and is the biological child of both spouses.[8] The full set of sample restrictions is described in Appendix Section 5. The resulting sample has approximately 7,200 parent-year observations.

We first estimate multivariate models of the outcomes reported by mothers and fathers. For responses involving relationships and attitudes, we estimate ordered logit models, and for reports of expenditures and time-use, we estimate OLS regressions. The principal variables of interest are dummy variables for firstborn daughters aged 0-12 and 13-18. The reference groups for these two variables consist of parents with firstborn sons who are in the same age groups. The top panel of Table 4 reports estimated coefficients separately for mothers and fathers. The models also control for the parent's age, education, and immigration background; the numbers of higher-parity boys and girls; and wave fixed effects. We adjust the significance levels for multiple hypothesis testing using Anderson's (2008) False Discovery Rate control method.

Several results indicate that households with teenage daughters face more strained marital and parenting relationships than households with teenage sons. Fathers of teenage daughters report significantly more parenting disagreements with their partners than fathers of teenage sons. Fathers of teenage daughters also report significantly worse relationships with their families. Mothers of teenage daughters report significantly more disagreements with their partners over money, more favourable attitudes towards divorce, and significantly lower life satisfaction than mothers of teenage sons. Parents of teenage daughters are also somewhat but not significantly less likely to indicate that married people are generally happier than unmarried people. However, other results are equivocal, providing neither

[8] We would have also liked to have analysed outcomes among a smaller set of couples who were later observed to divorce. Unfortunately, there are not enough observations to support this analysis. Instead, we use a broader sample of married couples with children, with the implicit assumption that these couples represent the risk set for divorce and that some will be subject to relationship processes that will lead to breakdowns.

support nor refutation of the relationship strains explanation.

The LISS provides no evidence of higher time or 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. Mothers of younger daughters, on the other hand, report fewer burdens than mothers of younger boys. Parents also do not report differences in the time spent with daughters, nor do they report differences in expenditures for children aged 0 to 15 if they have daughters.

The LISS also surveyed children who were age 16 or older and asked about their relationships with their parents and behaviour. The second panel of Table 4 reports estimates from ordered logit models of the responses from a sample of firstborn children aged 16 to 18 and a larger sample of all children aged 16 to 18. The models include a dummy variable for the respondent being a girl and controls for the parent, household, and temporal characteristics from our previous specifications. Compared to teenage boys, teenage girls report worse relationships with their fathers, but they do not report worse relationships with their mothers. In terms of risky behaviours, teenage boys are more likely than teenage girls to report drinking alcohol. Teenage boys are also more likely to report being smokers, although this difference is not statistically significant. Teenage girls, on the other hand, are more likely to report having a romantic partner.

The results from our analyses are consistent with relationship strains stemming from differences in family members' gender-role attitudes. As with the subgroup analyses, the results from the LISS survey point to the mutual importance of conflict between the parents and their children, and conflict between the parents themselves. On the one hand, teenage daughters report having worse relationships with their fathers, which supports the existence of parent-child conflict. On the other hand, fathers of teenage daughters report having more disagreements with their spouses over parenting, and mothers report having more

disagreements over expenditures. These results suggest that conflict can emerge both from parent-child and parent-parent interactions. The responses of teenage daughters further point to the importance of father-daughter interactions.[9]

However, other explanations may also be consistent with the results. First, we consider the compensation hypothesis, which postulates that teenage boys build solidarity between their parents because they require more supervision than teenage girls. This solidarity in turn strengthens the parents' marriages.[10] Estimates from the LISS data show that teenage boys are more likely than teenage girls to engage in risky behaviours, which lends empirical support for the need for additional supervision. However, we do not see a corresponding difference in parents' behaviours; parents indicate no differences in expenditures or care burdens conditional on teenagers' genders. A related hypothesis posits that teenage boys may be more vulnerable to negative outcomes from divorce, but these vulnerabilities only become apparent in adolescence. It is unclear, however, why these vulnerabilities would be more pronounced among parents with mixed immigration backgrounds or dissimilar levels of education, or why we would observe more relationship strains between fathers and daughters.

Second, the gender of children may affect the values held by the parents. As children age, parents with daughters may adopt less traditional attitudes than parents with sons, which may lower the stability of their marriages. The LISS results indicate that fathers of teenage daughters become less traditional with respect to gender roles, and mothers of teenage daughters become more open to divorce. However, a value explanation is hard to reconcile

[9] We conducted supplementary analyses that split the LISS sample according to the dissimilarity of gender-role attitudes between (a) the parents and their children and (b) the mothers and the fathers. We found that fathers who are relatively traditional (compared to their children or spouses) are much more likely to report parenting disagreements and dissatisfaction with family relationships conditional on having a teenage daughter. Similarly, teenage daughters are more likely to report worse relationships if their fathers are relatively traditional, but not if their fathers are relatively progressive. Conditioning on mothers' gender-role attitudes does not yield similar results. These findings lend further support to the conflict hypothesis. Results are available upon request.

[10] A countervailing hypothesis is that the supervision requirements for teenage boys increase stresses between the parents and weaken marriages.

with child-gender differences in divorce risks disappearing in early adulthood. Previous research suggests that the changes of parental values induced by the presence of daughters are persistent (Washington, 2008; Borrell-Porta *et al*., 2019). The lack of child-gender differences in the early adulthood suggests that the physical presence of teenage daughters in the household is an important attribute of the true causal mechanism.

Another possible explanation involves parental role modelling. Mothers with teenage daughters may feel a stronger need to leave dysfunctional marriages, so that they would act as better role models for their daughters. This argument is supported by the child-gender differences we see in mothers' attitudes towards divorce. However, it is not clear why would we see differences conditional on fathers' sibling composition or why role modelling would stop being a motivation after the daughters reach age 18. In addition, fathers' reports of parenting disagreements point to explanations that involve child-gender differences in the underlying level of conflict.

Naturally, multiple explanations may be invoked at the same time. A conflict stemming from differences in gender-role attitudes may be more damaging to marital stability because parents' family values are changing, or because mothers want to act as role models for their daughters. Nevertheless, age-specific conflicts involving gender roles fit with multiple patterns in the register and LISS analyses; other explanations do not.

6. External validity and relation to previous findings

We replicate the event-history analysis of divorce from the Netherlands using U.S. data from the 1980, 1985, 1990, and 1995 CPS Marriage and Fertility Supplements extracted from the IPUMS-CPS and NBER CPS archives. The supplements in these years asked adult women living in CPS households about the wedding, separation, and divorce dates for their first three marriages and about the genders and years of birth for their first four children and their most recent child. Supplements in subsequent years have only asked about the youngest

child and current marital status.

In line with other studies that analysed the CPS-MFS data (Morgan *et al*., 1988; Morgan and Pollard, 2001; Dahl and Moretti, 2008), we restrict our sample to women with children. Specifically, we analyse the durations of first marriages of women who were aged 20-65 at the time of the CPS-MFS Survey, and who gave birth to their first child during the course of their first marriage.[11] Each marriage spell is assigned a failure time if the marriage ended in divorce prior to the survey. Spells are right-censored if the woman was still in her first marriage at the time of the survey, if the woman was widowed during the marriage, if the duration of the marriage reached 35 years, or if the youngest child reached age 27. The data structure and censoring conditions are similar to our administrative dataset. Note that the CPS-MFS sample is not a stock sample of marriages that were observed at a specific point in time. It is a retrospective sample of first marriages corresponding to women who were interviewed in one of the four waves of the CPS-MFS survey.

We estimate a discrete-time cloglog hazard model that is similar to the specifications used in the Dutch subgroup analyses. The covariates include dummies for the firstborn's age; interactions of age groups (0-12, 13-18, and 19-26 years) with a firstborn daughter indicator; an indicator for being childless; dummies for marriage duration; controls for maternal age at marriage, education, race, birth cohort, and census division; and cohort and wave fixed effects. The observations are weighted by the CPS-MFS weights. Estimates for the daughter/age-group interactions are reported in the last row of Table 3; further information can be found in Appendix Section 6.

The estimated age pattern for the U.S. resembles the pattern for the Netherlands. We find a small and statistically insignificant difference in divorce risks for firstborn daughters

[11] The CPS-MFS surveys do not ask about paternity; we drop women with children born before their first marriage because the children may have a stepfamily relationship with the husband. We also drop observations that do not indicate the gender of the firstborn child or that have a sampling weight of zero.

aged 0-12 years. In contrast, we find a statistically significant and substantially higher risk of divorce for firstborn daughters aged 13-18 years. The relative teenage daughter coefficient is nearly twice as large as the coefficient for Dutch married couples. We also calculate cumulative divorce probabilities. By the time the firstborn children reach age 19, 24.53% of the U.S. parents of firstborn daughters are predicted to divorce compared to 23.62% of the parents of firstborn sons; a difference of 0.91 percentage points. The CPS-MFS findings are robust to changes of the sample and model specifications, including alternative censoring thresholds, expanding the sample to include childless women, and using fewer controls (detailed results are available upon request).

Dahl and Moretti (2008) also examined data from the CPS-MFS. Because of the difference in methodologies, it is hard to compare the estimated effect sizes exactly. However, the implied divorce differential from our analysis appears to be smaller than theirs. To examine the differences more carefully, we attempted to replicate their CPS-MFS results. We first constructed an analysis sample and estimated a linear probability model of mothers' first marriages ending in divorce, using the sample inclusion criteria and control variables that were described in their article.[12] When we applied their stated inclusion criteria, we obtained an analysis sample that was much larger than the sample they described, and when we estimated their stated model specification, we obtained a firstborn daughter effect that was only 0.5 percentage points (half the size of the estimate they reported).

We investigated alternative samples and specifications and examined an input file provided by Dahl and Moretti. Our analysis indicates that their CPS-MFS estimate of 1.03 percentage points was likely obtained from a model that: (a) omitted data from the 1985 wave of CPS-MFS, (b) constructed the outcome variable as a marriage ending in *widowhood or*

[12] Dahl and Moretti (2008) stated (p. 1089) that their sample included "all ever-married mothers between the ages of 20 and 70", and their explanatory variables included "a cubic in age, and dummies for race (White people, Black people, Asian, other), education (less than high school, high school, college), region of residence (nine regions), and cohort of birth (10-year birth cohorts)".

divorce rather than just ending in *divorce*, (c) omitted census division and education control variables, (d) did not incorporate the CPS-MFS sampling weights, and (e) included mothers whose first children were born after the first marriage ended. The smaller estimates of firstborn daughter effects that we obtain from samples that address these issues are very close to Dahl and Moretti's Census estimates and reconcile the differences between their Census and CPS-MFS results.

Our finding from the CPS-MFS of insignificant gender differences in early childhood is consistent with the null results from the US study by Reichman *et al.* (2004) that focused on families with very young children. However, it contrasts with the Census analysis of Dahl and Moretti (2008), which yielded a significant gender difference in divorce rates among couples with firstborn children who were aged 12 or younger. Although our point estimate of the firstborn daughter effect for 0- to 12-year-olds is small, it is also relatively imprecise, and we cannot rule out the existence of gender effects of the magnitude reported by Dahl and Moretti in their Census analysis.

This implies that American couples may be subject to other causal mechanisms. Dahl and Moretti's (2008) additional analyses of fertility choices, shotgun marriages, and child legitimisation rates provide evidence of a preference for sons among American couples. This mechanism can explain Dahl and Moretti's Census result, although the analyses of Blau *et al*. (2020) suggest that the practical importance of this mechanism has waned in recent years. Furthermore, our finding of a larger firstborn teenage daughter effect in the CPS-MFS indicates that other age-related mechanisms are also at work and that they may rival the effects of preferences for sons.

7. Robustness checks

We subject our analyses of the Dutch data to many robustness checks. Detailed results for these analyses are reported in Appendix Table B2.

7.1 *Different dissolution events and family statuses*

Our principal specifications examine formal divorce and registered partnership dissolution outcomes. As an alternative, we use residential separations (based on changes in the household composition records) to measure the end of relationships. In the data, separations precede formal divorce and dissolution events by 4.5 months on average. Consistent with this, the gender differential in separations appears when firstborn daughters are slightly younger, between ages 12 to 17.

We also investigate whether the age-specific patterns exist among parents who are cohabiting informally (living together but not in a marriage or registered partnership). Our analysis shows that cohabiting couples are also subject to higher separation risks during their daughters' teenage years. The effect for cohabiting couples is twice as large as the effect for married couples. Thus, our findings extend to all cohabiting couple relationships in the Netherlands and not just formal marriages and partnerships. Details corresponding to this model can be found in Appendix Section 7.

In another robustness check, we expand the principal analysis data set to include couples in which the firstborn had a stepchild relationship with one of the spouses. The same age patterns persist.

7.2 *Alternative analysis windows*

Our principal analysis dataset excludes marriages of people who died or emigrated prior to 1995. To see whether this selectivity affects our results, we re-estimate our models with a restricted data set of couples whose first child was born in or after 1995. The disadvantage with this analysis is that we observe no couples with children aged over 20, and so we cannot examine divorce differentials when the children are young adults. Nevertheless, we still find evidence of the teenage daughter effect (see Appendix Table B2, column 3).

7.3 *Alternative functional forms*

Logit specifications of our event-history models produce results that are similar to our cloglog specifications. We further examine the sensitivity of our estimates to the proportionality assumption in the cloglog model by estimating separate specifications of the hazard of divorce at each possible age of the firstborn child (26 specifications). This effectively allows the firstborn-age-duration pattern to vary with all the observed characteristics in our data. The estimated age-specific daughter effects from the separate models are nearly identical to the estimates from the principal specification. The results are available upon request.

8. Conclusion

Our analyses show that Dutch and U.S. parents with teenage daughters are more likely to divorce than parents with teenage sons. In the Netherlands, the effect peaks at age 15, when couples with firstborn daughters face a 9.2% higher hazard probability of divorce compared to couples with firstborn 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 firstborn daughters is 0.36 percentage points (1.8 percent) higher than the rate for couples with firstborn sons. The teenage daughter divorce differential for U.S. couples appears to be even bigger. The same age and gender pattern is found for higher-parity children.

In contrast to these results, we find no differences in divorce risks among couples with children aged 0 to 12 or older than 18. These null results, which are precisely estimated in the Dutch data, are important for narrowing the set of potential explanations for the association between children's gender and divorce. They run counter to explanations based on time-invariant preferences for sons. Our findings of slightly higher fertility in Dutch couples with firstborn sons, and of no gender differences in legitimisation rates of Dutch

children are also inconsistent with overarching preferences for sons. The age patterns of divorce risks are also inconsistent with selection into live birth. The paucity of observable differences in the characteristics of parents prior to the birth of their first child corroborates this interpretation and suggests that the child-gender associations can be interpreted causally. The age patterns also fit poorly with explanations based on forward-looking rational behaviour based on other foreseeable differences in parents' valuations of marriages with sons and daughters.

The isolation of the gender effect in the teenage years and its appearance for both firstborn and higher-parity children are more consistent with age-dependent mechanisms which influence marital stability through unexpected changes in preferences, constraints, or family processes. Although several mechanisms are consistent with the gender-age pattern, we see strained family relationships stemming from gender-role conflicts as a likely mechanism.

This explanation is supported by family relations research which has found links between adolescents' problems and marital stress. It is also supported by further analyses of our data. We find that parents whose children are more likely to hold conflicting beliefs about the gender roles (such as immigrant parents or parents from earlier birth cohorts) are subject to larger teenage divorce differentials. We find even larger differentials when parents themselves are likely to hold dissimilar gender-role attitudes (for example, parents with mixed immigration backgrounds or parents with different levels of education). More strikingly, the teenage divorce differential only appears for fathers who grew up without sisters and disappears completely if the father's sibship includes sisters.

The conflict hypothesis is also supported by our analyses of household survey data, which show that parents of teenage daughters report more disagreements over child rearing, fathers of teenage daughters report worse family relationships, and mothers of teenage

daughters report more favourable attitudes towards divorce and lower life satisfaction. Teenage daughters also report worse relationships than teenage sons with their fathers. Other results from the survey data regarding relationship strains are inconclusive, but there are no significant results that refute the strains explanation. The responses of parents with teenage sons and daughters show no significant difference in levels of expenditures on children, levels of parental time investments, or perceptions of care burdens.

The limitation of our study is that we cannot conclusively rule out other age-related explanations. For example, traits could emerge in adolescence that indicate that boys are more vulnerable to divorce. There could be unexpected age-specific changes in costs that we are not able to measure or control for. It is even possible that parents' preferences could evolve in unexpected ways. We also cannot rule out complex, age-varying interactions in explanations. For example, parents could have time-invariant preferences for sons but be constrained from acting on them until the children reach adolescence.

Finally, the age pattern that we uncover in the Dutch and U.S. data helps to reconcile differences in findings from previous studies. Although several studies have found that daughters are associated with higher divorce risks, others have failed to detect associations. Most studies that have found associations have considered children across a range of ages that includes their teenage years, when the gender effect is present. In contrast, studies with null findings focused predominantly on young children or had other age restrictions, which reduced their ability to find effects.

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Table 1. Previous estimates of the effect of firstborn daughters on marriage ending in divorce and/or separation

Study	Country	Data / Outcome	Child-age or marriage-duration restrictions	Baseline divorce rate	Relative Effect	Absolute Effect
Spanier and Glick (1981)	USA	CPS-MFS 1975 / Divorce or separation	Dur. ≤ 25 years	37.4%	2.4%	0.9 p.p.
Morgan *et al*. (1988)	USA	CPS-MFS 1980 / Divorce or separation	Age ≤ 15 years	10.0%[A]	6.3%	0.6 p.p.[A]
Andersson and Woldmicael (2001)	Sweden	Admin. 1971-1995 / Divorce	Dur. ≤ 15 years	17.3%[A]	1.0% [NS]	0.2 p.p.[A, NS]
Morgan and Pollard (2002)	USA	CPS-MFS 1980-1995 / Divorce or separation	Age ≤ 14 years	-	6% (1960-74) 1% [NS] (1975-94)	-
Diekmann and Schmidheiny (2004)	18 countries	Fertility & Fam. Survey 1988-1996 / Divorce	No restrictions	-	Insignificant for 17 of 18 countries	-
Bedard and Deschênes (2005)	USA	Census 1980 / Divorce	Age ≤ 17 years	20.0%	4.0%	0.8 p.p.
Ananat and Michaels (2008)	USA	Census 1980 / Divorce	Age ≤ 16 years	17.2%	3.7%	0.6 p.p.
Dahl and Moretti (2008)	USA	Census 1960-1980 / Divorce	Age ≤ 12 years	20.6%	2.2%	0.5 p.p.
Dahl and Moretti (2008)	USA	CPS-MFS 1980-1995 / Divorce	No restrictions	32.0%	3.2%	1.0 p.p.
Leigh (2009)	Australia	Census 1981-2001 / Divorce	Age of youngest child ≤ 12 years	11.5%	0.9% [NS]	0.1 p.p.[NS]
Flouri and Malmberg (2010)	UK	Mill. Cohort Study 2003 / Father absence	Age ≤ 3 years	7.5%	-8.0% [NS]	-0.1 p.p.[NS]

Note: The table lists study characteristics and estimated associations between children's gender and marriages ending in divorce.
[A] Own calculations using the results and descriptive statistics provided in the listed manuscript.
[NS] Not statistically significant.

Table 2: Average Characteristics of Couples with Firstborn Sons and Daughters

	Sons	Daughters	Diff.	P-val.
Father's birth year	1962.68	1962.67	0.011	0.36
Mother's birth year	1965.15	1965.14	0.012	0.33
Year of wedding/registration	1989.78	1989.77	0.013	0.40
Father's age at wedding/registration	27.61	27.61	0.001	0.89
Mother's age at wedding/registration	25.14	25.14	0.000	0.94
Mother's age at birth of the firstborn	29.90	29.90	-0.007	0.28
Father's age at birth of the firstborn	27.43	27.43	-0.007	0.21
Marriage/RP duration at birth of the firstborn (in months, cond. on being married/RP at birth)	35.17	35.28	-0.008	0.01
Nativity				
Father and mother born in the Netherlands	72.65%	72.76%	-0.107%	0.05
Father native; mother 1st gen. immigrant	3.35%	3.39%	-0.032%	0.16
Father native; mother 2nd gen. immigrant	3.96%	3.96%	0.003%	0.90
Father 1st gen. immigrant; mother native	2.55%	2.53%	0.020%	0.30
Father and mother 1st generation immigrants	10.96%	10.80%	0.153%	0.00
Father 1st gen. immig.; mother 2nd gen. immig.	0.98%	1.00%	-0.018%	0.16
Father 2nd gen. immigrant; mother native	4.04%	4.06%	-0.020%	0.41
Father 2nd gen. immig.; mother 1st gen. immig.	0.76%	0.76%	0.001%	0.96
Father and mother 2nd generation immigrants	0.75%	0.75%	0.006%	0.59
Father's completed education				
Less than High School	4.18%	4.17%	0.009%	0.72
High-school	23.81%	23.78%	0.034%	0.52
University	17.43%	17.40%	0.033%	0.49
Missing records	54.58%	54.66%	-0.076%	0.22
Mother's completed education				
Less than High School	5.90%	5.53%	0.069%	0.02
High-school	26.46%	26.69%	-0.233%	0.00
University	16.73%	16.64%	0.084%	0.07
Missing records	50.92%	50.84%	0.080%	0.20
Labour supply 1yr prior to first child's birth [a]				
Father employed	85.25%	85.20%	0.051%	0.59
Mother employed	84.57%	84.63%	-0.059%	0.48
Father's annual earnings, in 1000€	30.38	30.35	0.031	0.65
Mother's annual earnings, in 1000€	22.19	22.21	-0.022	0.53
Child born prior to marriage/RP	9.10%	9.07%	0.023%	0.52
Parents are registered partners	1.37%	1.36%	0.018%	0.23
Number of siblings	1.21	1.21	0.003	0.02
Time between first two children, in months	34.61	34.49	0.118	0.00
Parents divorced or ended registered partnership	20.64%	20.91%	-0.270%	0.00
Observations	1,284,277	1,220,755		

Note: The first two columns list means and shares of characteristics of different-sex parent couples in first marriages or registered partnerships estimated conditionally on the gender of the firstborn child. The third column lists differences in the means and shares by the gender of the firstborn, and the fourth column lists *p*-values of two-sided t-tests of the equality of the means and shares. Please see the text for a complete description of the analysis sample.
[a] Employment and earnings are only available for parents of first children born after 1999.

Table 3: Relative and absolute divorce effects of firstborn daughters

Sample	Excess hazard probabilities, firstborn daughters (%)		Cumulative divorce rate by age 18 (%)		
	Age 0-12	Age 13-18	Sons	Daughters	Difference
Full sample	0.21	5.47	20.12	20.48	0.36
(2,950,986 spells)	(0.32)	(0.60)	(0.04)	(0.04)	(0.06)
Immigration background					
Both spouses native	0.17	4.64	19.16	19.46	0.30
(2,359,518 spells)	(0.40)	(0.67)	(0.04)	(0.04)	(0.06)
Both spouses immigrants	0.56	6.80	20.66	21.05	0.40
(338,831 spells)	(1.01)	(2.03)	(0.13)	(0.13)	(0.18)
Mixed imm. background	0.36	11.71‡	31.01	32.11	1.10‡
(252,637 spells)	(1.07)	(2.29)	(0.18)	(0.19)	(0.29)
Husband's birth cohort					
1955 and earlier	-0.06	7.54	14.04	14.38	0.34
(761,890 spells)	(0.81)	(1.24)	(0.06)	(0.06)	(0.09)
1956-1965	0.03	5.01‡	20.65	20.98	0.33
(975,847 spells)	(0.59)	(0.91)	(0.07)	(0.07)	(0.09)
1966 and later	0.56	4.34‡	26.05	26.51	0.46
(1,213,249 spells)	(0.54)	(1.27)	(0.10)	(0.09)	(0.13)
Husband's education					
Less than high school	-0.06	8.37	27.35	28.14	0.78
(122,078 spells)	(1.19)	(2.77)	(0.21)	(0.20)	(0.29)
High school	-0.15	5.12‡	25.01	25.31	0.29‡
(697,964 spells)	(0.55)	(1.10)	(0.08)	(0.08)	(0.11)
University	1.07	3.74‡	18.69	19.01	0.31‡
(501,063 spells)	(0.97)	(1.76)	(0.11)	(0.12)	(0.16)
Missing records	0.11	5.59‡	18.02	18.38	0.36‡
(1,629,881 spells)	(0.44)	(0.86)	(0.05)	(0.05)	(0.07)
Parent's dissimilarity index					
1st tercile	0.34	3.05	21.19	21.43	0.24
(290,494 spells)	(1.16)	(2.06)	(0.13)	(0.13)	(0.17)
2nd tercile	-0.24	3.42	24.07	24.29	0.21
(290,495 spells)	(1.08)	(2.03)	(0.06)	(0.06)	(0.20)
3rd tercile	-0.10	9.26‡	29.59	30.46	0.87‡
(290,495 spells)	(0.98)	(2.07)	(0.16)	(0.17)	(0.23)
Sibship, native husbands born after 1965					
No sisters	1.01	7.87	27.74	28.61	0.87
(368,377 spells)	(0.98)	(2.52)	(0.18)	(0.19)	(0.27)
At least one sister	0.93	0.41‡	25.33	25.57	0.25‡
(524,592 spells)	(0.85)	(1.96)	(0.13)	(0.15)	(0.21)
Sibship, native wives born after 1965					
No brothers	-0.29	3.94	27.93	28.23	0.31
(413,832 spells)	(0.88)	(2.01)	(0.17)	(0.15)	(0.23)
At least one brother	0.70	3.72	25.75	26.21	0.46
(640,062 spells)	(0.73)	(1.59)	(0.11)	(0.12)	(0.17)

Other data					
US married couples (CPS)	2.89	10.05	23.62	24.53	0.91
(101,880 spells)	(1.80)	(4.39)	(0.22)	(0.25)	(0.34)

Note: Estimates of excess hazard probabilities of divorce faced by couples with firstborn daughters aged 0-12 and 13-18, and cumulative divorce probabilities corresponding to the time when the firstborn sons and daughters reach age 18. The excess hazard probabilities are computed using the formula $(\exp(\hat{\beta})-1)*100$, where $\hat{\beta}$ denotes the estimated coefficient for the respective age-group firstborn daughter dummies from cloglog hazard models of divorce risks. The formula for the cumulative divorce probabilities is given in Appendix Section 2. Bootstrapped standard errors are presented in parentheses. With the exception of the last specification, all models were estimated using registry data for Dutch different-sex couples who married after 1971 and did not have children with other partners prior to the marriage. More detailed results can be found in Appendix Table B1. The last specification was estimated using retrospective marital histories of American women aged 20-65, collected in CPS-MFS waves 1980, 1985, 1990 and 1995. More detailed results for this model can be found in the third column of Appendix Table A4. To test whether the coefficient estimates presented in columns 1, 2 and 5 are significantly different across the listed subgroups, we estimate a fully interacted version of our model for each topical group. In this model, the original (not interacted) set of covariates corresponds to the first category in the group, and the interacted sets of covariates correspond to the other categories in the group.
† = different from the coefficient for the first category in the respective group at 0.10 significance level.
‡ = different from the coefficient for the first category in the respective group at 0.05 significance level.

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

	Fathers with firstborn daughters		Mothers with firstborn daughters	
Subjective questions	Age 0-12	Age 13-18	Age 0-12	Age 13-18
1.How satisfied are you with your current relationship?	0.011	0.023	-0.190**	-0.117
2.[Did] you and your partner (have) any differences of opinion regarding money expenditure over the past year?	-0.143	0.173	0.149	0.264**
3.[Did] you and your partner (have) any differences of opinion regarding raising the children over the past year?	-0.140	0.256**	-0.098	0.155
4.A woman is more suited to rearing young children than a man	-0.115	-0.409***	0.042	-0.165
5.Divorce is generally best solution if a married couple cannot solve their marital problems	-0.100	0.063	0.050	0.332***
6.Married people are generally happier than unmarried people	-0.103	-0.177	-0.064	-0.086
7.How would you generally describe the relationship with your family?	-0.200	-0.350**	-0.016	-0.142
8.Caring for my child is not such a burden	0.095	0.121	0.362***	0.070
9.How satisfied are you with the life you lead at the moment?	-0.028	-0.056	-0.137	-0.259**
Expenditures and time-use				
10.How much time did you spend in the last seven days on activities with own child (in hours)	0.252	1.558	-1.994	-0.679
11.Log total expenditure per month for children living at home, children 0-15	0.242	-0.017	0.058	0.102

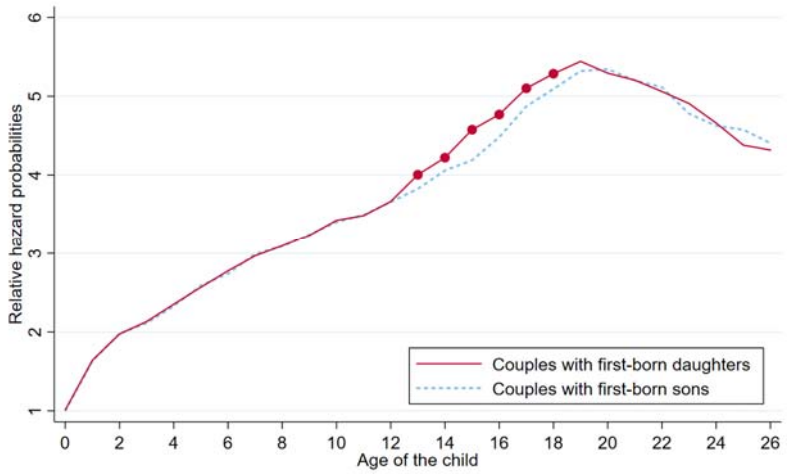
Children's responses	Firstborn daughters aged 16-18	All daughters aged 16-18
12. How would you describe your overall relationship with your father?	-0.461**	-0.268**
13. How would you describe your overall relationship with your mother?	-0.211	-0.018
14. Have you ever smoked?	-0.111	-0.143
15. Do you smoke now?	-0.359	-0.313

16. How often did you have a drink containing alcohol over the last 12 months?	-0.996**	-0.847**
17. Did you have a drink containing alcohol during the last seven days?	-0.763**	-0.800**
18. Do you currently have a partner?	1.168***	0.908***

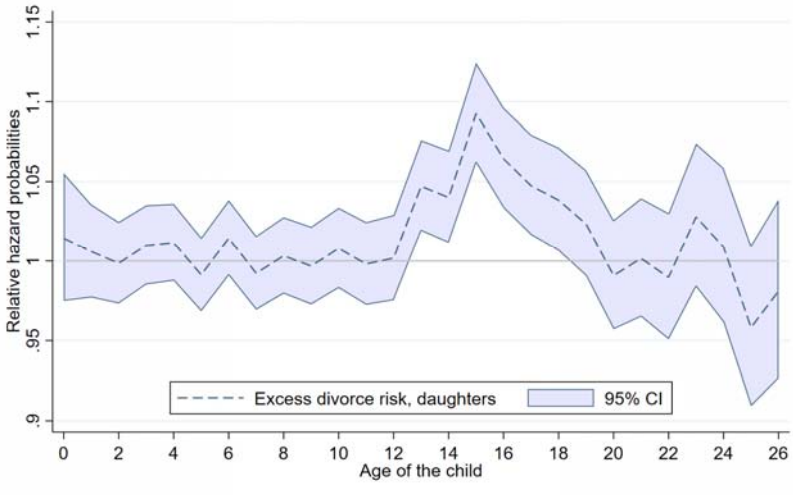
Note: Authors' estimates of coefficients of the daughter-age interactions from ordered logit, binomial logit, and OLS regression models of fathers', mothers', and children's survey responses from the 2008-2016 panels of the LISS. Parental responses correspond to different-sex parents with a firstborn biological child younger than 19 at the time of the survey. Child responses correspond to children aged 16 to 18 who were born to different-sex couples whose firstborn is a biological child. Sample sizes vary depending on the response rates of parents and children. More detailed results can be found in Appendix table B3.
*** = 0.01 significance level, ** = 0.05 significance level, * = 0.1 significance level.

Figure 1: Conditional hazard probabilities of divorce by age and gender of a firstborn child

(a) Hazard probabilities by age and gender of a firstborn child, relative to couples with a firstborn son aged 0



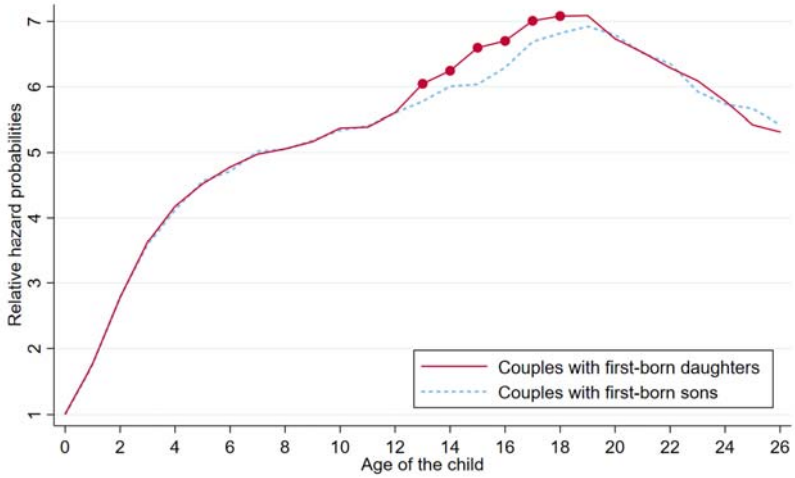
(b) Excess hazard probabilities faced by couples with a firstborn daughter, relative to couples with a firstborn son of the same age



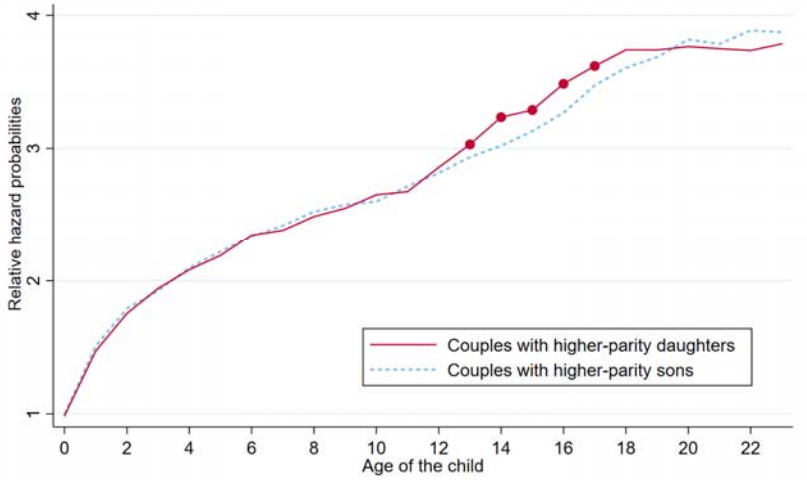
Note: Authors' estimates of exponentiated coefficients from a cloglog hazard model of divorce risks which controls for the presence, age, and gender of a firstborn child, marriage durations and parent's observable characteristics. In panel a, the lines indicate the magnitude of conditional divorce risks faced by the couples with a firstborn of the given gender and age relative to the divorce risks faced by couples with a firstborn son younger than 1 year of age. The dots indicate the ages at which the divorce risks for couples with a firstborn daughter are significantly different ($p<0.05$) from divorce risks for couples with a firstborn son. The model uses Dutch registry data for different-sex couples who married after 1971 and did not have children with other partners prior to the marriage.

Figure 2: Conditional hazard probabilities of divorce associated with firstborn and higher-parity children

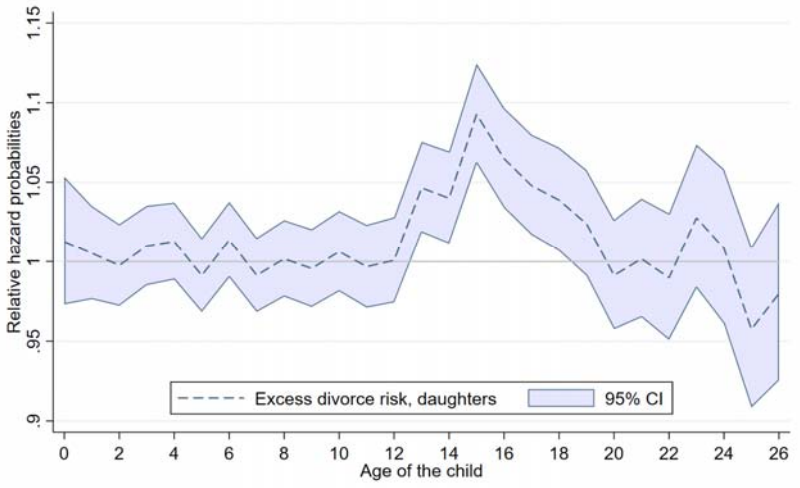
(a) by age and gender of a firstborn child, relative to couples with a firstborn son aged 0 years



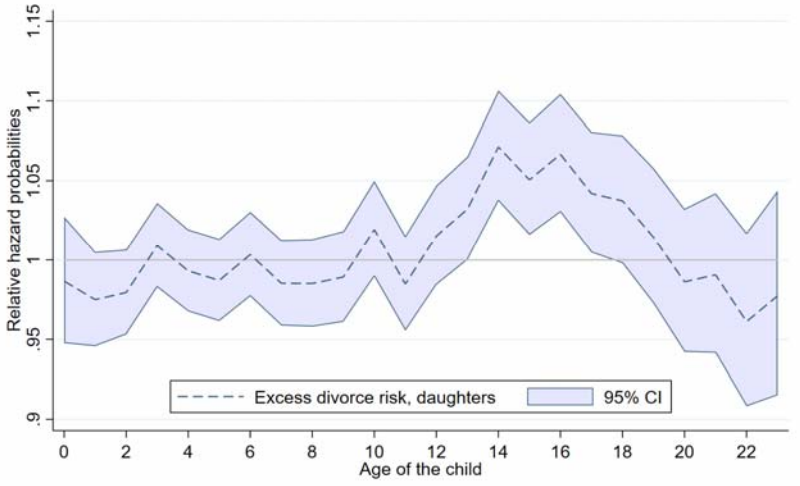
(b) by age and gender of higher-parity children, relative to couples with a higher-parity son aged 0 years



(c) Excess hazard probabilities faced by couples with a firstborn daughter, relative to couples with a firstborn son of the same age



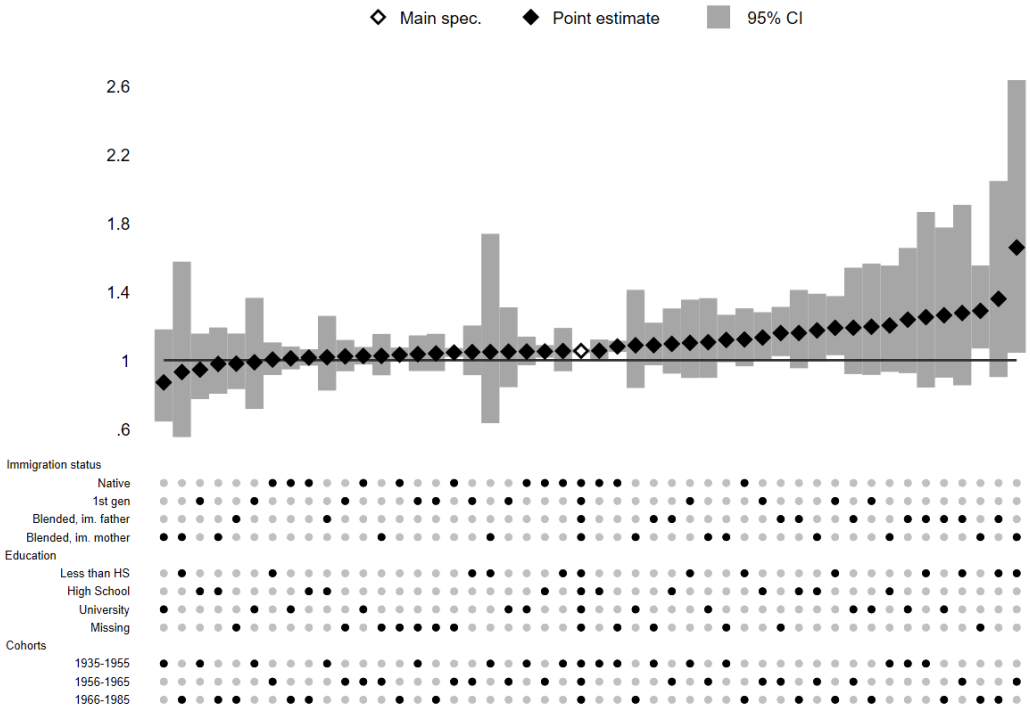
(d) Excess hazard probabilities faced by couples with a higher-parity daughter, relative to couples with a higher-parity son of the same age



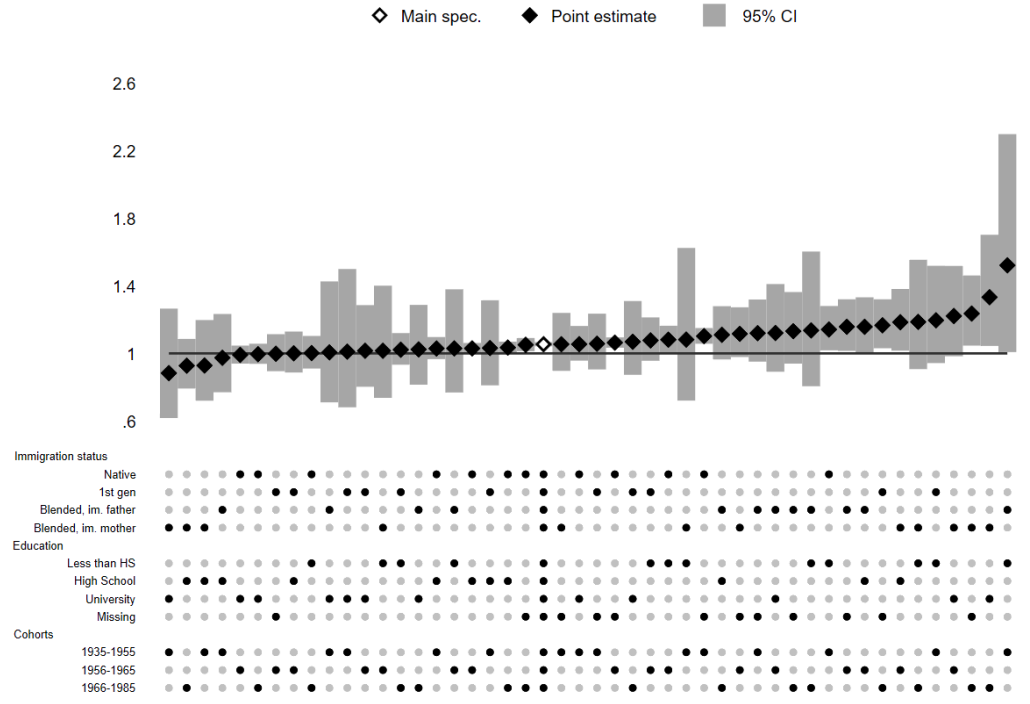
Note: Authors' estimates of exponentiated coefficients from a cloglog hazard model of divorce risks which controls for marriage durations, parent's observable characteristics, and presence, ages, and genders of firstborn and higher-parity children. The dots in the first two panels indicate the ages at which the conditional divorce risks for couples with daughters are significantly different ($p<0.05$) from divorce risks for couples with sons. The model uses Dutch registry data for different-sex couples who married after 1971 and did not have children with other partners prior to the marriage.

Figure 3. Excess hazard probabilities of divorce faced by couples with a firstborn teenage daughter, conditional on combinations of parental characteristics

(a) Split by father's characteristics



(b) Split by mother's characteristics



Note: The top part of each figure plots exponentiated coefficients and 95% confidence intervals corresponding to the firstborn teenage daughter dummies from cloglog hazard models of divorce risks estimated for the subgroups indicated in the lower part of each figure. Each model is specified like the models presented in Table 3. The models use Dutch registry data for different-sex couples who married after 1971 and did not have children with other partners prior to the marriage.

Appendix, Daughters and Divorce
(for online publication only)

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APPENDIX A

1. Log-likelihood function

In our principal analyses, we use a complementary log-log (cloglog) hazard model, which is a discrete analog of the continuous proportional hazards model. The log-likelihood function of the cloglog model is

$$\log(L) = \sum_{i} \sum_{t=t_{1i}}^{t_{2i}} y_{it} \log \lambda_{it} + (1 - y_{it}) \log(1 - \lambda_{it}),$$

where y_{ij} is an indicator function which equals 1 if a marriage spell *i* ends in divorce at duration *t*, and 0 otherwise. λ_{it} is a hazard rate defined as $\lambda_{it} = 1 - \exp\left(-\exp\left(\mathbf{x}'_{it}\boldsymbol{\beta}\right)\right)$, and t_{1i} and t_{2i} are durations of marriage spell *i* at the first and the last point of observation.

2. Calculation of absolute effects

Our event-history models yield child-gender coefficient estimates that can be interpreted as relative effects of the children's gender on divorce risks. To get a better sense of the absolute magnitude of these gender effects, we calculate cumulative probabilities that a marriage with children ends in divorce at different ages of firstborn children, distinguishing between the cumulative probabilities faced by the couples with firstborn boys and firstborn girls. We use the following formula,

$$CP_k^s = \hat{\lambda}_0^s + \sum_{A=1}^{k} \left(\hat{\lambda}_A^s \cdot \prod_{a=0}^{A-1} (1 - \hat{\lambda}_k^s) \right),$$

where $\hat{\lambda}_k^s$ is an estimate of the average age-specific hazard rate of divorce, and corresponds to the average of predicted hazard probabilities of divorce for couples whose firstborn child is of gender $s = \{b,g\}$ and is *k* years old:

$$\hat{\lambda}_k^b = \frac{1}{N_k^b}\sum_{i=1}^{N}\left(\hat{y}_{it}\cdot\left(FB\,age_{it}=k\right)\cdot\left(FB\,daughter_i=0\right)\right)$$
$$\hat{\lambda}_k^g = \frac{1}{N_k^g}\sum_{i=1}^{N}\left(\hat{y}_{it}\cdot\left(FB\,age_{it}=k\right)\cdot\left(FB\,daughter_i=1\right)\right)$$

Using the cumulative divorce probabilities, we can calculate absolute gender effects (defined as the difference between gender-specific cumulative probabilities at a given age), and relative gender effects (defined as the ratio of increments of gender-specific cumulative probabilities over a pre-specified period). Standard errors are bootstrapped using 200 draws from the multivariate distribution of coefficient estimates $\boldsymbol{\beta}_{BS} \sim N(\hat{\boldsymbol{\beta}}, \hat{\Sigma})$.

3. Model for higher-parity children

The functional form corresponding specification is

$$\mathbf{x}_{\mathbf{it}}'\boldsymbol{\beta} = \beta_0 + \sum_{k=0}^{26}\left[\mathbf{1}\left(FB\,age_{it}=k\right)\cdot\left(\beta_{1k}+\beta_{2k}\cdot FB\,daughter_i\right)\right] + \sum_{n=1}^{N}\left[\mathbf{1}\left(nch_{it}\geq n\right)\cdot\beta_{3n}\right]$$
$$+\sum_{j=1}^{40}\beta_{4j}\cdot\mathbf{1}\left(Duration_{it}=j\right)+\beta_5\cdot RegPar_{it}+\sum_{y=1972}^{2015}\beta_{6j}\cdot\mathbf{1}\left(Year_{it}=y\right)+\mathbf{z}_{\mathbf{i}}'\boldsymbol{\beta}_{\mathbf{7}}$$
$$+\sum_{k=0}^{26}\left[\sum_{n=2}^{N}\left[\mathbf{1}\left(HP\,age_{itn}=k\right)\cdot\left(\beta_{8k}+\beta_{9k}\cdot HP\,daughter_{in}\right)\right]\right]$$

The higher-parity controls are included in the last line of Equation 4, with *n* indexing the parity of children, and *N* denoting the total number of children. Besides including these controls, we also replace the *Childless* dummy from Equations 2 and 3 with a set of dummies which control for the number of children in the household. We restrict the higher-parity age coefficients (β_{8k} and β_{9k}) to each have a single value at ages 24 and beyond, because our right-censoring condition for the firstborn's age (26) leads to very few observations of higher-parity children who are older than 24. We also restrict the fertility coefficients (β_{3n}) to a single value if there are five or more children in the family.

In addition to the coefficient β_{10}, the coefficient β_{80} is also set to zero. This means

that the reference group for the higher-parity age coefficients (β_{8k}) consists of couples with higher-parity children that are less than one year old. The reference groups for the higher-parity age-gender interactions (β_{9k}) consist of couples with higher-parity boys of the same age (k). The reference groups for the firstborn age and gender coefficients remain the same.

4. Analyses of mechanisms that do not involve unexpected changes of parents' valuations of marriage

We investigate whether teenage daughter effect could be attributed to mechanisms that do not involve unexpected changes of parents' marriage valuations. We evaluate whether the teenage daughter effect could be explained by a delayed onset of preferences for sons in couples with no sons. In this scenario, the preference for sons only affects marital stability when a couple stops having children and there is no longer a possibility of adding a son. To examine this, we estimate specifications of our model for couples who did and did not have higher-parity sons and find that the firstborn teenage daughter effect is present among both types of couples. The coefficient estimates corresponding to these two models are listed in columns 4 and 5 of Table B2 We also estimate the higher-parity (Equation 4) model conditional on the gender of the firstborn child and find significant higher-parity teenage daughter effects for couples with firstborn boys and firstborn girls (results available upon request).

Another possible explanation is that marital strains in couples with daughters occur at earlier ages, but the mothers cannot act on them immediately because they lack work experience to support a single-earner household. Mothers with younger girls may therefore focus on gaining more work experience and leave the spouse only when they have accumulated enough human capital. However, we find no evidence of such behaviour. Both labour force participation and earnings of mothers with firstborn boys and girls are not

significantly different from each other throughout the first 11 years of the child's life. In the following years, mothers of firstborn girls work marginally more hours than mothers of firstborn boys.

A different work-related hypothesis is that the presence of adolescent girls reduces the costs of mothers working because adolescent girls perform more housework and require less supervision than boys (Kalenkoski *et al*. 2011). The increased labour market activity might make it easier for mothers to divorce. To examine this, we estimate a specification that adds maternal and paternal employment and earnings to the divorce model. The employment data are available as of 1999, which reduces the pool of marriage spells that we can examine. Nevertheless, the age-specific firstborn daughter coefficients remain similarly sized, and they retain statistical significance from ages 13 to 17. Thus, work-related hypotheses do not explain the age pattern that we uncover. The corresponding coefficient estimates are listed in column 6 of Table B2.

5. LISS survey analysis - 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 9 years (2008-2016). Each household member older than 15 years of age is surveyed individually. Children up to the age 15 do not participate actively, meaning that their presence in the household (and a basic set of characteristics) is reported by the parents. Conditional on participating, the response rates of household members average 75-80%.

Our sample is restricted to couples with children who are married and whose firstborn 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 7,182 person-year records of participating parents, and 681 person-year records of participating firstborn teenagers (aged 16-18). The extended sample of all participating teenagers (regardless of birth parity)

contains 1,309 person-year records. The loss of parental observations due to the sample restrictions is documented in Table A3.

The numbers of observations corresponding to individual regression models listed in Table B3 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 which lowers the numbers of observations. Furthermore, the questions regarding expenditures were asked only to the adult household member who is usually responsible for shopping.

The significance levels reported in Table 3 are adjusted for multiple hypothesis testing using the FDR control method of Anderson (2008). Specifically, we have split the survey questions into thematic blocks and tested for statistical significance of the daughter dummies together with other questions in the respective blocks. This adjustment is more suitable for analyses with large number of tested hypotheses which differ with respect to the desirability of the null hypothesis. Further details can be found in the notes of Table B3.

Complete results for the models of fathers' and mothers' survey responses are reported in Table B3a-B3c. Complete results for the models of teenagers' responses are reported in Table B4a-B4b.

6. Analysis of Current Population Survey

We conduct event-history analyses using U.S. data from 1980, 1985, 1990 and 1995 CPS Marriage and Fertility Supplements. We estimate three specifications of the discrete-time hazard model which are modelled on the specifications used in our main analysis. The set of covariates is simplified to account for the smaller sample size of CPS data, and fewer observable characteristics. Akin to Dahl & Moretti (2008), our sample does not contain mothers that are childless, which means that we cannot model the divorce risks faced by

childless women and the risks faced by mothers prior to the child's birth. Table A3 lists the estimated coefficients. The model in column 1 mimics the unconditional model presented in Column 1 of Table A1, the model in column 2 adds marriage duration controls, and the preferred specification in column 3 adds further observable characteristics.

7. Model for cohabiting couples

We investigate whether teenage daughter effects exist among parents who are living out of wedlock. The analysis of cohabiting couples is complicated by the lack of precise information on the starting and ending dates of the relationships, however we are able to construct measures which approximate the relationship durations.

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, doing so is further complicated by left-censoring of the cohabitation data (residential histories were not collected prior to 1995). For this reason, we use the age of the firstborn child to approximate the duration of the relationship. The event of separation is defined as the point when one of the two parents leaves the shared household and does not move back within the period of observation. Apart from the previously discussed censoring events, the cohabitation spells are treated as right-censored at the point when a couple is observed to get married. The specification of the hazard model for cohabiting couples includes dummies for the yearly ages of the firstborn, three gender-age group interaction terms, cubic polynomials of parental age at birth, and year and cohort fixed effects. The coefficient estimates are listed in the last column of Table B1c.

Table A1: Regression results, baseline specifications

	(1) Firstborn's age-gender dummies & childless dummy		(2) Variables in (1), duration dummies & family controls		(3) Variables in (2), & higher-parity children controls	
VARIABLES	exp(Beta)	St.e.	exp(Beta)	St.e.	exp(Beta)	St.e.
Firstborn's age dummies						
age 1	1.851***	0.033	1.634***	0.029	1.760***	0.031
age 2	2.408***	0.041	1.981***	0.034	2.787***	0.047
age 3	2.624***	0.044	2.114***	0.036	3.592***	0.061
age 4	2.858***	0.047	2.328***	0.039	4.127***	0.071
age 5	3.098***	0.051	2.593***	0.043	4.563***	0.078
age 6	3.143***	0.052	2.739***	0.046	4.716***	0.081
age 7	3.275***	0.054	2.993***	0.050	5.021***	0.087
age 8	3.209***	0.053	3.090***	0.053	5.050***	0.089
age 9	3.184***	0.053	3.244***	0.056	5.190***	0.092
age 10	3.142***	0.053	3.401***	0.059	5.342***	0.096
age 11	3.037***	0.051	3.494***	0.061	5.408***	0.099
age 12	3.000***	0.051	3.659***	0.065	5.599***	0.104
age 13	2.963***	0.051	3.826***	0.069	5.777***	0.109
age 14	2.965***	0.051	4.059***	0.074	6.005***	0.115
age 15	2.878***	0.051	4.188***	0.078	6.037***	0.118
age 16	2.889***	0.051	4.481***	0.084	6.293***	0.125
age 17	2.952***	0.053	4.872***	0.092	6.689***	0.135
age 18	2.902***	0.052	5.093***	0.098	6.818***	0.14
age 19	2.853***	0.052	5.320***	0.104	6.924***	0.146
age 20	2.700***	0.051	5.344***	0.108	6.793***	0.148
age 21	2.470***	0.048	5.200***	0.108	6.515***	0.148
age 22	2.274***	0.046	5.116***	0.111	6.355***	0.151
age 23	1.982***	0.042	4.779***	0.109	5.927***	0.149
age 24	1.781***	0.040	4.624***	0.111	5.734***	0.154
age 25	1.647***	0.038	4.575***	0.115	5.665***	0.162
age 26	1.481***	0.037	4.406***	0.118	5.426***	0.172
*Firstborn's age * daughter dummies*						
age 0	1.014	0.021	1.012	0.020	1.012	0.020
age 1	1.006	0.015	1.004	0.015	1.005	0.015
age 2	0.999	0.013	0.998	0.013	0.998	0.013
age 3	1.010	0.013	1.009	0.013	1.010	0.013
age 4	1.011	0.012	1.011	0.012	1.013	0.012
age 5	0.991	0.012	0.991	0.012	0.991	0.012
age 6	1.014	0.012	1.014	0.012	1.014	0.012
age 7	0.992	0.012	0.992	0.012	0.991	0.012
age 8	1.003	0.012	1.002	0.012	1.002	0.012
age 9	0.997	0.012	0.996	0.012	0.996	0.012
age 10	1.008	0.013	1.007	0.013	1.006	0.013

age 11	0.998	0.013	0.997	0.013	0.997	0.013
age 12	1.002	0.014	1.001	0.014	1.001	0.014
age 13	1.047***	0.015	1.047***	0.015	1.046***	0.015
age 14	1.040***	0.015	1.040***	0.015	1.040***	0.015
age 15	1.093***	0.016	1.092***	0.016	1.093***	0.016
age 16	1.064***	0.016	1.064***	0.016	1.065***	0.016
age 17	1.047***	0.016	1.047***	0.016	1.048***	0.016
age 18	1.038**	0.017	1.038**	0.017	1.039**	0.017
age 19	1.023	0.017	1.023	0.017	1.024	0.017
age 20	0.991	0.018	0.990	0.018	0.991	0.018
age 21	1.002	0.019	1.001	0.019	1.002	0.019
age 22	0.990	0.020	0.989	0.020	0.990	0.020
age 23	1.028	0.023	1.027	0.023	1.028	0.023
age 24	1.009	0.025	1.008	0.025	1.008	0.025
age 25	0.958	0.026	0.957	0.026	0.958	0.026
age 26	0.981	0.029	0.980	0.029	0.980	0.029
No Children dummy	3.618***	0.052	5.138***	0.074		
Registered Partnership			1.461***	0.019	1.457***	0.019
Child born prior to marriage			1.636***	0.009	1.638***	0.009
Spousal immigration background						
Husband native, Wife 1st gen.			1.576***	0.009	1.517***	0.009
Husband native, Wife 2nd gen.			1.456***	0.008	1.430***	0.008
Husband 1st gen. Wife native			2.173***	0.013	2.137***	0.012
Husband 1st gen. Wife 1st gen.			0.788***	0.004	0.807***	0.004
Husband 1st gen. Wife 2nd gen.			1.551***	0.016	1.509***	0.015
Husband 2nd gen. Wife native			1.421***	0.008	1.398***	0.008
Husband 2nd gen. Wife 1st gen.			1.343***	0.017	1.283***	0.016
Husband 2nd gen. Wife 2nd gen.			1.588***	0.020	1.550***	0.019
Age at wedding						
Husband, linear			0.793***	0.008	0.806***	0.008
Husband, quadratic			1.007***	0.000	1.006***	0.000
Husband, cubic			1.000***	0.000	1.000***	0.000
Wife, linear			0.921***	0.009	0.929***	0.009
Wife, quadratic			1.002***	0.000	1.001***	0.000
Wife, cubic			1.000***	0.000	1.000***	0.000
Education levels						
Husband, High School			0.901***	0.005	0.890***	0.005
Husband, University			0.697***	0.005	0.712***	0.005
Husband, Missing			0.834***	0.005	0.828***	0.005
Wife, High School			1.037***	0.005	1.017***	0.005
Wife, University			0.882***	0.005	0.896***	0.005
Wife, Missing			0.586***	0.003	0.579***	0.003
Number of children						
1 child					0.192***	0.003
2 children					0.054***	0.001

3 children					0.041***	0.001
4 children					0.034***	0.001
5 children and more					0.023***	0.001
Number of higher-parity daughters by age						
age 0					0.986	0.020
age 1					0.974*	0.015
age 2					0.979	0.014
age 3					1.008	0.014
age 4					0.993	0.013
age 5					0.987	0.013
age 6					1.003	0.014
age 7					0.985	0.014
age 8					0.984	0.014
age 9					0.988	0.015
age 10					1.018	0.015
age 11					0.983	0.015
age 12					1.013	0.016
age 13					1.029**	0.017
age 14					1.068***	0.018
age 15					1.048***	0.018
age 16					1.064***	0.019
age 17					1.040**	0.019
age 18					1.035*	0.020
age 19					1.012	0.022
age 20					0.984	0.023
age 21					0.988	0.026
age 22					0.959	0.028
age 23					0.975	0.033
age 24 and older					0.934*	0.038
Constant	0.004***	0.000	0.005***	0.002	0.004***	0.002
Marriage duration dummies			✓		✓	
Parental birth cohort dummies			✓		✓	
Calendar year dummies			✓		✓	
Count variables for number of higher-parity children by age					✓	
Observations	57,267,483		57,267,483		57,267,483	
Marriage spells	2,950,986		2,950,986		2,950,986	
ln likelihood	-3,742,885		-3,607,353		-3,586,930	

Note: Authors' estimates of exponentiated coefficients from cloglog hazard models of divorce risks. 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.

Table A2: Excess hazard probabilities of firstborn daughters in additional subsample analyses

Model specification	Age 0-12	Age 13-18	Age 19-26	Spells
Baseline, full sample	1.002	1.055***	1.000	2,950,986
	(0.004)	(0.006)	(0.007)	
Immigration background and homogeneity				
Both spouses native	1.002	1.046***	0.987	2,359,518
(incl. 2nd gen. immigrant)	(0.004)	(0.007)	(0.008)	
Both spouses immigrants	1.006	1.068***	1.077***	338,831
	(0.011)	(0.021)	(0.025)	
Father native, mother	0.990	1.118***	1.040	140,702
immigrant	(0.014)	(0.032)	(0.039)	
Mother native, father	1.020	1.112***	1.057	111,935
immigrant	(0.016)	(0.032)	(0.041)	
Birth cohort, husband				
Cohorts 1955 and earlier	0.999	1.075***	1.003	761,890
	(0.008)	(0.012)	(0.012)	
Cohorts 1956-1965	1.000	1.050***	1.002	975,847
	(0.006)	(0.009)	(0.010)	
Cohorts 1966 and later	1.005	1.043***	0.993	1,213,249
	(0.006)	(0.013)	(0.025)	
Birth cohort, wife				
Cohorts 1955 and earlier	0.997	1.082***	1.019	519,039
	(0.011)	(0.016)	(0.015)	
Cohorts 1956-1965	1.001	1.051***	1.003	983,451
	(0.006)	(0.009)	(0.010)	
Cohorts 1966 and later	1.005	1.049***	0.971	1,448,496
	(0.005)	(0.011)	(0.018)	
Education, husband				
Less than High School	0.999	1.084***	1.043	122,078
	(0.014)	(0.028)	(0.032)	
High School	0.999	1.051***	0.981	697,964
	(0.007)	(0.013)	(0.015)	
University	1.011	1.037***	1.032	501,063
	(0.010)	(0.017)	(0.022)	
Missing records[1]	1.001	1.056***	0.997	1,629,881
	(0.005)	(0.009)	(0.010)	
Education, wife				
Less than High School	1.005	1.075***	1.003	167,580
	(0.012)	(0.022)	(0.023)	
High School	0.996	1.039**	0.990	774,145
	(0.006)	(0.010)	(0.012)	
University	1.003	1.016	0.983	492,088
	(0.009)	(0.018)	(0.023)	
Missing records[1]	1.013*	1.075***	1.010	1,517,173
	(0.007)	(0.010)	(0.011)	
Education dissimilarity				
Same level of schooling	0.990	1.033**	0.991	360,762
	(0.008)	(0.015)	(0.019)	

Different schooling levels	1.017* (0.010)	1.085*** (0.019)	0.986 (0.022)	199,993

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

[1] 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.

*** = 0.01 significance level, ** = 0.05 significance level, * = 0.1 significance level .

Table A3: Person-year records of coupled individuals who live with at least some of their children in the same household, LISS panel, years 2008-2016

Sample	Number of person-year records
Coupled adults in LISS households who live with their children	27,180
Out of whom:	
Adults who participate in the survey	20,685
Adults who filled out the family module (necessary to identify biological children)	15,899
Out of whom:	
Adults who share biological firstborn with their current partner	14,246
Adults whose biological firstborn is alive and at most 18 years old	10,761
Adults whose biological firstborn is at most 18 years old and lives in the same household	8,749
Out of whom:	
Adults who are married	7,182
Adults who are married and their firstborn is a teenager	2,786

Appendix Table A4: Regression results, CPS-MFS sample

	(1) Firstborn's age-gender dummies & childless dummy		(2) Variables in (1) & duration dummies		(3) Variables in (2), time dummies & family controls	
VARIABLES	exp(Beta)	St.e.	exp(Beta)	St.e.	exp(Beta)	St.e.
Firstborn daughter aged 0-12	1.028	0.018	1.029	0.018	1.029	0.018
Firstborn daughter aged 13-18	1.096**	0.044	1.096**	0.044	1.101**	0.044
Firstborn daughter aged 19-26	0.929	0.044	0.929	0.044	0.936	0.045
Mother's characteristics						
Age at wedding - linear term					0.811***	0.049
Age at wedding - quad. term					0.916	0.226
Age at wedding - cubic term					2.509***	0.802
Race, black					1.260***	0.036
Race, other					0.702***	0.033
Education, high school					1.156***	0.024
Education, college					1.119***	0.035
Constant	0.006***	0.000	0.003***	0.001	0.029***	0.025
Firstborns' age dummies	✓		✓		✓	
Marriage duration dummies			✓		✓	
Calendar year, birth cohort and geographical dummies					✓	
Observations	1,613,397		1,613,397		1,613,397	
Marriage spells	101,880		101,880		101,880	
ln likelihood	-116,237		-115,920		-111,855	

Note: Authors' estimates of exponentiated coefficients corresponding to the cloglog hazard models of divorce risks. The model uses retrospective marital histories of American women aged 20-65, collected in CPS-MFS waves 1980, 1985, 1990 and 1995. The sample consists of first marriage spells of women whose first child was born within those first marriages.
*** = 0.01 significance level, ** = 0.05 significance level, * = 0.1 significance level.

APPENDIX B – TABLES OF RESULTS CORRESPONDING TO SUPPLEMENTARY ANALYSES

1. Heterogeneity analyses

Table B1a: Estimates corresponding to the heterogeneity analyses

		Immigration background				Education, husband			
VARIABLES	Baseline	Native	Immigrants	Mixed, husband native	Mixed, husband immigrant	Primary school	High school	University	Missing
Firstborn daughter aged 0-12	1.002	1.002	1.006	0.990	1.020	1.003	1.004	1.014	0.996
	(0.004)	(0.004)	(0.011)	(0.014)	(0.016)	(0.014)	(0.007)	(0.010)	(0.005)
Firstborn daughter aged 13-18	1.055***	1.046***	1.068***	1.118***	1.112***	1.082***	1.066***	1.054***	1.075***
	(0.006)	(0.007)	(0.021)	(0.032)	(0.032)	(0.028)	(0.013)	(0.018)	(0.009)
Firstborn daughter aged 19-26	1.000	0.987	1.077***	1.040	1.057	1.048	1.018	1.008	0.990
	(0.007)	(0.008)	(0.025)	(0.039)	(0.041)	(0.032)	(0.015)	(0.022)	(0.010)
No Children dummy	5.112***	5.374***	3.768***	4.086***	6.745***	2.927***	4.487***	8.122***	5.276***
	(0.054)	(0.066)	(0.111)	(0.148)	(0.335)	(0.104)	(0.083)	(0.276)	(0.080)
Registered Partnership	1.461***	1.373***	3.175***	1.312***	2.024***	1.291***	1.372***	1.665***	1.459***
	(0.019)	(0.019)	(0.213)	(0.084)	(0.101)	(0.119)	(0.029)	(0.042)	(0.032)
Child born prior to marriage	1.636***	1.628***	1.496***	1.351***	1.499***	1.465***	1.609***	1.492***	1.645***
	(0.009)	(0.010)	(0.024)	(0.026)	(0.031)	(0.029)	(0.015)	(0.023)	(0.013)
Spousal immigration background									
Husband native, Wife 1st gen.	1.576***				1.062***	1.802***	1.621***	1.326***	1.559***
	(0.009)				(0.016)	(0.051)	(0.018)	(0.019)	(0.013)
Husband native, Wife 2nd gen.	1.456***	1.455***				1.402***	1.473***	1.340***	1.468***
	(0.008)	(0.008)				(0.038)	(0.016)	(0.019)	(0.011)
Husband 1st gen. Wife native	2.173***			1.198***		2.096***	1.972***	1.860***	2.250***
	(0.013)			(0.016)		(0.045)	(0.023)	(0.033)	(0.017)
Husband 1st gen. Wife 1st gen.	0.788***					0.598***	0.792***	0.961**	0.876***
	(0.004)					(0.008)	(0.007)	(0.016)	(0.005)
Husband 1st gen. Wife 2nd gen.	1.551***					0.966	1.252***	1.506***	1.979***
	(0.016)					(0.031)	(0.023)	(0.051)	(0.029)
Husband 2nd gen. Wife native	1.421***	1.420***				1.347***	1.419***	1.356***	1.417***
	(0.008)	(0.008)				(0.038)	(0.015)	(0.019)	(0.011)
Husband 2nd gen. Wife 1st gen.	1.343***					1.328***	1.140***	1.530***	1.413***
	(0.017)					(0.065)	(0.024)	(0.048)	(0.027)
Husband 2nd gen. Wife 2nd gen.	1.588***	1.512***				1.421***	1.539***	1.558***	1.605***
	(0.020)	(0.019)				(0.086)	(0.031)	(0.048)	(0.031)

Age at wedding									
Husband, linear	0.793***	0.748***	0.822***	1.006	0.930***	0.736***	0.823***	0.748***	0.804***
	(0.008)	(0.011)	(0.015)	(0.032)	(0.025)	(0.025)	(0.017)	(0.027)	(0.010)
Husband, quadratic	1.007***	1.009***	1.004***	0.999	1.001	1.009***	1.006***	1.009***	1.007***
	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.000)
Husband, cubic	1.000***	1.000***	1.000***	1.000	1.000	1.000***	1.000***	1.000***	1.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Wife, linear	0.921***	0.582***	1.044**	0.870***	1.107***	0.943**	0.949***	0.974	0.881***
	(0.009)	(0.009)	(0.018)	(0.023)	(0.030)	(0.023)	(0.017)	(0.032)	(0.011)
Wife, quadratic	1.002***	1.018***	0.997***	1.003***	0.996***	0.999	1.001	1.002	1.003***
	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.000)
Wife, cubic	1.000***	1.000***	1.000***	1.000**	1.000***	1.000*	1.000	1.000**	1.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Education levels									
Husband, High School	0.901***	0.785***	1.123***	0.810***	0.734***				
	(0.005)	(0.006)	(0.014)	(0.016)	(0.019)				
Husband, University	0.697***	0.617***	0.840***	0.614***	0.523***				
	(0.005)	(0.005)	(0.015)	(0.014)	(0.014)				
Husband, Missing	0.834***	0.721***	1.103***	0.904***	0.599***				
	(0.005)	(0.005)	(0.013)	(0.017)	(0.015)				
Wife, High School	1.037***	0.861***	1.601***	0.870***	0.940***	1.459***	0.941***	0.954**	0.970***
	(0.005)	(0.005)	(0.017)	(0.017)	(0.018)	(0.021)	(0.008)	(0.022)	(0.007)
Wife, University	0.882***	0.725***	1.479***	0.762***	0.858***	1.611***	0.899***	0.765***	0.863***
	(0.005)	(0.005)	(0.024)	(0.017)	(0.019)	(0.042)	(0.010)	(0.018)	(0.007)
Wife, Missing	0.586***	0.493***	0.639***	0.531***	0.723***	1.145***	0.839***	0.875***	0.396***
	(0.003)	(0.003)	(0.007)	(0.011)	(0.014)	(0.016)	(0.008)	(0.020)	(0.003)
Constant	0.005***	0.262**	0.261	0.012***	0.003***	0.750***	0.012***	0.003***	0.013***
	(0.002)	(0.138)	(0.224)	(0.014)	(0.004)	(0.054)	(0.012)	(0.003)	(0.006)
Firstborns' age dummies	✓	✓	✓	✓	✓	✓	✓	✓	✓
Marriage duration dummies	✓	✓	✓	✓	✓	✓	✓	✓	✓
Parental birth cohort dummies	✓	✓	✓	✓	✓	✓	✓	✓	✓
Calendar year dummies	✓	✓	✓	✓	✓	✓	✓	✓	✓
Observations	57,267,483	47,717,907	5,714,145	1,717,289	2,115,613	2,397,579	12,705,916	8,046,135	34,109,928
Spells	2,950,986	2,359,518	338,831	140,702	111,935	122,078	697,964	501,063	1,629,881
ln likelihood	-3,607,364	-2,859,358	-366,057	-187,745	-182,378	-195,902	-965,756	-481,270	-1,944,811

Table B1b: Estimates corresponding to the heterogeneity analyses

	Education, wife				Education dissimilarity		Cohorts, husband		
VARIABLES	Primary school	High school	University	Missing	Same level	Different levels	1935-1955	1956-1965	1966-1985
Firstborn daughter aged 0-12	1.005	0.996	1.003	1.006	0.990	1.017*	0.999	1.000	1.006
	(0.012)	(0.006)	(0.009)	(0.006)	(0.008)	(0.010)	(0.008)	(0.006)	(0.005)
Firstborn daughter aged 13-18	1.075***	1.039***	1.016	1.075***	1.033**	1.085***	1.075***	1.050***	1.043***
	(0.022)	(0.010)	(0.018)	(0.010)	(0.015)	(0.019)	(0.012)	(0.009)	(0.013)
Firstborn daughter aged 19-26	1.003	0.990	0.983	1.010	0.991	0.986	1.003	1.002	0.993
	(0.023)	(0.012)	(0.023)	(0.011)	(0.019)	(0.022)	(0.012)	(0.010)	(0.025)
No Children dummy	2.954***	4.155***	7.689***	5.863***	4.450***	4.496***	6.323***	4.480***	5.228***
	(0.094)	(0.069)	(0.238)	(0.102)	(0.103)	(0.129)	(0.166)	(0.078)	(0.081)
Registered Partnership	1.184*	1.325***	1.638***	1.695***	1.426***	1.412***	1.723***	1.289***	1.514***
	(0.115)	(0.028)	(0.035)	(0.043)	(0.032)	(0.042)	(0.159)	(0.056)	(0.021)
Child born prior to marriage	1.617***	1.542***	1.408***	1.650***	1.568***	1.555***	1.839***	1.634***	1.537***
	(0.026)	(0.013)	(0.020)	(0.015)	(0.018)	(0.022)	(0.025)	(0.015)	(0.012)
Spousal immigration background									
Husband native, Wife 1st gen.	1.334***	1.385***	1.366***	1.778***	1.407***	1.360***	1.785***	1.616***	1.401***
	(0.028)	(0.015)	(0.020)	(0.015)	(0.021)	(0.023)	(0.020)	(0.015)	(0.014)
Husband native, Wife 2nd gen.	1.365***	1.421***	1.347***	1.529***	1.349***	1.401***	1.483***	1.471***	1.404***
	(0.032)	(0.013)	(0.018)	(0.013)	(0.018)	(0.023)	(0.016)	(0.013)	(0.013)
Husband 1st gen. Wife native	2.232***	2.047***	1.979***	2.265***	1.910***	1.964***	2.133***	2.264***	2.072***
	(0.048)	(0.020)	(0.027)	(0.021)	(0.029)	(0.032)	(0.023)	(0.022)	(0.021)
Husband 1st gen. Wife 1st gen.	0.566***	0.923***	1.079***	0.814***	0.692***	0.669***	1.038***	0.804***	0.673***
	(0.006)	(0.007)	(0.015)	(0.006)	(0.008)	(0.008)	(0.010)	(0.006)	(0.005)
Husband 1st gen. Wife 2nd gen.	1.358***	1.272***	1.649***	1.831***	1.188***	1.115***	2.206***	2.263***	1.290***
	(0.049)	(0.019)	(0.043)	(0.035)	(0.027)	(0.028)	(0.061)	(0.051)	(0.016)
Husband 2nd gen. Wife native	1.470***	1.401***	1.336***	1.448***	1.407***	1.374***	1.400***	1.448***	1.398***
	(0.033)	(0.013)	(0.019)	(0.013)	(0.018)	(0.022)	(0.015)	(0.013)	(0.014)
Husband 2nd gen. Wife 1st gen.	0.879***	1.078***	1.495***	1.758***	1.119***	1.066**	1.803***	1.771***	1.124***
	(0.030)	(0.024)	(0.049)	(0.035)	(0.031)	(0.032)	(0.055)	(0.045)	(0.018)
Husband 2nd gen. Wife 2nd gen.	1.581***	1.471***	1.573***	1.703***	1.479***	1.491***	1.647***	1.548***	1.533***
	(0.085)	(0.028)	(0.045)	(0.037)	(0.036)	(0.046)	(0.049)	(0.036)	(0.026)
Age at wedding									
Husband, linear	0.817***	0.795***	0.871***	0.727***	0.793***	0.849***	0.660***	0.650***	0.724***
	(0.020)	(0.014)	(0.024)	(0.012)	(0.020)	(0.023)	(0.011)	(0.012)	(0.018)
Husband, quadratic	1.005***	1.007***	1.004***	1.010***	1.007***	1.004***	1.012***	1.015***	1.011***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)

Husband, cubic	1.000***	1.000***	1.000***	1.000***	1.000***	1.000***	1.000***	1.000***	1.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Wife, linear	0.914***	0.842***	0.725***	0.826***	0.993	1.044*	0.732***	0.869***	1.116***
	(0.023)	(0.016)	(0.023)	(0.012)	(0.024)	(0.025)	(0.012)	(0.013)	(0.013)
Wife, quadratic	1.002*	1.005***	1.011***	1.005***	1.000	0.997***	1.008***	1.002***	0.995***
	(0.001)	(0.001)	(0.001)	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)	(0.000)
Wife, cubic	1.000	1.000***	1.000***	1.000***	1.000	1.000***	1.000***	1.000***	1.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Education levels									
Husband, High School	1.149***	0.763***	0.730***	0.837***	1.098***	0.630***	0.842***	0.861***	0.948***
	(0.015)	(0.007)	(0.018)	(0.008)	(0.014)	(0.011)	(0.010)	(0.008)	(0.009)
Husband, University	0.920***	0.614***	0.481***	0.742***	0.645***	0.545***	0.860***	0.732***	0.619***
	(0.023)	(0.007)	(0.012)	(0.008)	(0.009)	(0.006)	(0.012)	(0.008)	(0.007)
Husband, Missing	1.448***	0.987	0.792***	0.503***			0.917***	0.793***	0.813***
	(0.018)	(0.009)	(0.019)	(0.005)			(0.010)	(0.007)	(0.008)
Wife, High School						0.811***	0.968***	0.967***	1.119***
						(0.014)	(0.009)	(0.008)	(0.010)
Wife, University						0.872***	1.174***	0.888***	0.791***
						(0.010)	(0.014)	(0.008)	(0.008)
Wife, Missing							0.611***	0.559***	0.592***
							(0.006)	(0.004)	(0.005)
Constant	0.019***	0.017***	0.015***	0.018***	0.028***	0.017***	0.184***	0.618*	0.020***
	(0.025)	(0.017)	(0.019)	(0.010)	(0.022)	(0.007)	(0.085)	(0.180)	(0.006)
Firstborns' age dummies	✓	✓	✓	✓	✓	✓	✓	✓	✓
Marriage duration dummies	✓	✓	✓	✓	✓	✓	✓	✓	✓
Parental birth cohort dummies	✓	✓	✓	✓	✓	✓	✓	✓	✓
Calendar year dummies	✓	✓	✓	✓	✓	✓	✓	✓	✓
Observations	3,390,912	14,138,448	6,968,914	32,767,445	8,956,584	5,238,225	20,176,172	21,981,618	15,108,888
Spells	167,580	774,145	492,088	1,517,173	360,762	199,993	761,890	975,847	1,213,249
Log-likelihood	-280,582	-1,213,481	-495,860	-1,594,188	-631,984	-413,485	-696,775	-625,857	-1,082,133

Table B1c: Estimates corresponding to the heterogeneity analyses

	Cohorts, wife			Dissimilarity index			Sibship, husband	
VARIABLES	1935-1955	1956-1965	1966-1985	1st tercile	2nd tercile	3rd tercile	No sisters	At least one sister
Firstborn daughter aged 0-12	0.997	1.001	1.005	1.006	0.998	0.999	1.012	1.009
	(0.011)	(0.006)	(0.005)	(0.012)	(0.011)	(0.010)	(0.010)	(0.008)
Firstborn daughter aged 13-18	1.082***	1.051***	1.049***	1.030	1.034*	1.093***	1.079***	1.004
	(0.016)	(0.009)	(0.011)	(0.021)	(0.020)	(0.021)	(0.025)	(0.020)
Firstborn daughter aged 19-26	1.019	1.003	0.971	0.988	0.980	0.997	0.932	0.945
	(0.015)	(0.010)	(0.018)	(0.025)	(0.025)	(0.025)	(0.054)	(0.044)
No Children dummy	7.008***	4.594***	5.136***	4.765***	4.367***	4.356***	5.304***	5.464***
	(0.248)	(0.083)	(0.072)	(0.163)	(0.138)	(0.125)	(0.148)	(0.133)
Registered Partnership	1.986***	1.318***	1.515***	1.487***	1.411***	1.393***	1.425***	1.554***
	(0.276)	(0.070)	(0.020)	(0.050)	(0.045)	(0.040)	(0.031)	(0.030)
Child born prior to marriage	1.858***	1.634***	1.548***	1.684***	1.665***	1.464***	1.434***	1.499***
	(0.032)	(0.016)	(0.011)	(0.029)	(0.026)	(0.020)	(0.020)	(0.019)
Spousal immigration background								
Husband native, Wife 1st gen.	1.814***	1.698***	1.381***			1.131***	1.200***	1.310***
	(0.025)	(0.017)	(0.012)			(0.014)	(0.019)	(0.019)
Husband native, Wife 2nd gen.	1.453***	1.476***	1.432***		1.115***	1.134***	1.321***	1.435***
	(0.020)	(0.013)	(0.012)		(0.025)	(0.014)	(0.020)	(0.020)
Husband 1st gen. Wife native	2.269***	2.307***	2.012***			1.559***		
	(0.031)	(0.021)	(0.018)			(0.019)		
Husband 1st gen. Wife 1st gen.	1.270***	0.896***	0.666***	0.679***	0.652***	0.609***		
	(0.017)	(0.007)	(0.004)	(0.011)	(0.009)	(0.008)		
Husband 1st gen. Wife 2nd gen.	2.286***	2.454***	1.313***		1.053	1.003		
	(0.082)	(0.058)	(0.016)		(0.052)	(0.019)		
Husband 2nd gen. Wife native	1.413***	1.433***	1.410***		1.156***	1.135***	1.305***	1.434***
	(0.019)	(0.013)	(0.012)		(0.030)	(0.014)	(0.020)	(0.020)
Husband 2nd gen. Wife 1st gen.	1.799***	1.854***	1.152***		0.875**	0.966	0.838***	0.998
	(0.068)	(0.048)	(0.018)		(0.047)	(0.022)	(0.024)	(0.022)
Husband 2nd gen. Wife 2nd gen.	1.628***	1.576***	1.536***	1.463***	1.453***	1.432***	1.269***	1.591***
	(0.061)	(0.037)	(0.025)	(0.045)	(0.044)	(0.059)	(0.038)	(0.035)
Age at wedding								
Husband, linear	0.705***	0.822***	0.874***	0.762***	0.746***	0.859***	0.721***	0.700***
	(0.018)	(0.012)	(0.012)	(0.065)	(0.034)	(0.019)	(0.045)	(0.036)
Husband, quadratic	1.010***	1.005***	1.004***	1.009***	1.009***	1.003***	1.011***	1.013***

	(0.001)	(0.000)	(0.000)	(0.003)	(0.002)	(0.001)	(0.002)	(0.002)
Husband, cubic	1.000***	1.000***	1.000***	1.000***	1.000***	1.000***	1.000***	1.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Wife, linear	0.682***	0.612***	1.139***	0.856**	1.022	1.125***	0.852***	0.889***
	(0.016)	(0.010)	(0.021)	(0.063)	(0.041)	(0.023)	(0.039)	(0.034)
Wife, quadratic	1.011***	1.017***	0.995***	1.004	0.998	0.995***	1.005***	1.004***
	(0.001)	(0.001)	(0.001)	(0.003)	(0.001)	(0.001)	(0.002)	(0.001)
Wife, cubic	1.000***	1.000***	1.000***	1.000	1.000	1.000***	1.000***	1.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Education levels								
Husband, High School	0.831***	0.838***	0.956***	1.388***	0.992	0.832***	0.804***	0.758***
	(0.014)	(0.008)	(0.008)	(0.040)	(0.015)	(0.008)	(0.018)	(0.014)
Husband, University	0.879***	0.734***	0.655***	1.135***	0.707***	0.569***	0.519***	0.481***
	(0.016)	(0.008)	(0.006)	(0.037)	(0.012)	(0.007)	(0.012)	(0.009)
Husband, Missing	0.915***	0.809***	0.826***				0.649***	0.608***
	(0.014)	(0.007)	(0.007)				(0.015)	(0.011)
Wife, High School	0.928***	0.943***	1.121***	0.827***	1.050***	1.081***	0.912***	0.898***
	(0.012)	(0.007)	(0.009)	(0.021)	(0.015)	(0.011)	(0.018)	(0.015)
Wife, University	1.188***	0.928***	0.822***	0.645***	0.843***	0.912***	0.647***	0.610***
	(0.018)	(0.009)	(0.007)	(0.019)	(0.014)	(0.011)	(0.014)	(0.011)
Wife, Missing	0.593***	0.560***	0.600***				0.509***	0.460***
	(0.007)	(0.004)	(0.005)				(0.010)	(0.008)
Constant	0.164***	0.555**	0.001***	0.082***	0.036***	0.012***	0.348	0.227*
	(0.086)	(0.160)	(0.000)	(0.047)	(0.018)	(0.010)	(0.376)	(0.173)
Firstborns' age dummies	✓	✓	✓	✓	✓	✓	✓	✓
Marriage duration dummies	✓	✓	✓	✓	✓	✓	✓	✓
Parental birth cohort dummies	✓	✓	✓	✓	✓	✓	✓	✓
Calendar year dummies	✓	✓	✓	✓	✓	✓	✓	✓
Observations	14,044,260	23,659,301	19,563,050	5,000,172	4,780,093	4,414,031	4,329,417	6,417,758
Spells	519,039	983,451	1,448,496	290,494	290,495	290,495	368,377	524,592
Log-likelihood	-606,449	-1,428,105	-1,563,789	-316,180	-340,223	-387,754	-364,887	-493,810

Table B1d: Estimates corresponding to the heterogeneity analyses and the model for cohabiting couples

VARIABLES	Sibship, wife		Cohabiting couples
	No brothers	At least one brother	
Firstborn daughter aged 0-12	0.997	1.007	1.002
	(0.009)	(0.007)	(0.007)
Firstborn daughter aged 13-18	1.039**	1.037**	1.095***
	(0.020)	(0.016)	(0.024)
Firstborn daughter aged 19-26	0.955	0.959	1.045
	(0.040)	(0.030)	(0.038)
No Children dummy	5.232***	5.105***	
	(0.133)	(0.107)	
Registered Partnership	1.352***	1.521***	
	(0.029)	(0.029)	
Child born prior to marriage	0.997	1.007	
	(0.009)	(0.007)	
Spousal immigration background			
Husband native, Wife 1st gen.			1.522***
			(0.023)
Husband native, Wife 2nd gen.	1.347***	1.458***	1.325***
	(0.018)	(0.017)	(0.018)
Husband 1st gen. Wife native	1.766***	1.935***	1.865***
	(0.026)	(0.024)	(0.025)
Husband 1st gen. Wife 1st gen.			1.333***
			(0.014)
Husband 1st gen. Wife 2nd gen.	1.111***	1.154***	1.813***
	(0.027)	(0.018)	(0.039)
Husband 2nd gen. Wife native	1.347***	1.422***	1.364***
	(0.018)	(0.017)	(0.018)
Husband 2nd gen. Wife 1st gen.			1.694***
			(0.047)
Husband 2nd gen. Wife 2nd gen.	1.317***	1.564***	1.656***
	(0.039)	(0.032)	(0.038)
Age at wedding			
Husband, linear	0.917**	0.891***	1.170***
	(0.032)	(0.027)	(0.024)

Husband, quadratic	1.003**	1.004***	0.994***
	(0.001)	(0.001)	(0.001)
Husband, cubic	1.000***	1.000***	1.000***
	(0.000)	(0.000)	(0.000)
Wife, linear	0.661***	0.703***	1.595***
	(0.032)	(0.027)	(0.056)
Wife, quadratic	1.014***	1.012***	0.982***
	(0.002)	(0.001)	(0.001)
Wife, cubic	1.000***	1.000***	1.000***
	(0.000)	(0.000)	(0.000)
Education levels			
Husband, High School	0.822***	0.863***	0.953***
	(0.015)	(0.012)	(0.015)
Husband, University	0.578***	0.564***	0.582***
	(0.011)	(0.009)	(0.011)
Husband, Missing	0.696***	0.729***	0.950***
	(0.013)	(0.010)	(0.015)
Wife, High School	0.832***	0.824***	1.118***
	(0.014)	(0.011)	(0.017)
Wife, University	0.596***	0.582***	0.748***
	(0.011)	(0.009)	(0.013)
Wife, Missing	0.452***	0.420***	0.840***
	(0.008)	(0.006)	(0.013)
Constant	1.368***	1.527***	0.001***
	(0.018)	(0.016)	(0.000)
Firstborns' age dummies	✓	✓	✓
Marriage duration dummies	✓	✓	✓
Parental birth cohort dummies	✓	✓	✓
Calendar year dummies	✓	✓	✓
Observations	5,303,059	8,535,195	3,483,763
Spells	413,832	640,062	461,997
ln likelihood	-455,812	-672,814	-429.302

Note: Authors' estimates of exponentiated coefficients corresponding to the cloglog hazard models of divorce risks. The model uses linked marriage, divorce, and other registry data for different-sex couples who married after 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.

2. Robustness checks

Table B2: List of the robustness specifications

Col. No.	Description of the specification
1	Model of residential separation risks, principal sample
2	Model of divorce risks, principal dataset expanded to include couples whose firstborn is a step-child
3	Model of divorce risks, principal dataset restricted to couples with firstborn children born in or after 1995
4	Model of divorce risks, principal dataset restricted to couples with no male children
5	Model of divorce risks, principal dataset restricted to couples with at least one male child
6	Model of divorce risks, covariates expanded to include parental employment and earnings controls

Note: estimates corresponding to the model estimating the separation risks among unwedded couples are included in the last column of table B1.

Table B2: Estimates corresponding to the robustness checks

VARIABLES	(1) Separation risks	(2) Incl. step-children	(3) Born after 1995	(4) No boys	(5) One or more boys	(6) Earnings controls
*Firstborn's age * daughter dummies*						
age 0	1.007	1.003	1.027			0.997
	(0.018)	(0.020)	(0.032)			(0.038)
age 1	1.005	0.996	1.025			1.018
	(0.014)	(0.015)	(0.023)			(0.026)
age 2	1.009	1.003	0.991			0.976
	(0.013)	(0.013)	(0.019)			(0.021)
age 3	1.003	0.997	0.997			1.002
	(0.012)	(0.013)	(0.018)			(0.020)
age 4	1.002	1.009	1.010	1.017	1.051**	1.001
	(0.012)	(0.012)	(0.018)	(0.015)	(0.022)	(0.018)
age 5	0.994	0.988	0.980	1.035**	0.957**	0.985
	(0.012)	(0.012)	(0.017)	(0.016)	(0.019)	(0.017)
age 6	1.008	1.004	1.019	1.031**	1.019	1.023
	(0.012)	(0.012)	(0.018)	(0.016)	(0.019)	(0.017)
age 7	0.999	0.987	1.006	0.998	1.005	1.002
	(0.012)	(0.012)	(0.018)	(0.016)	(0.018)	(0.016)
age 8	0.994	1.005	0.996	0.995	1.028	1.015
	(0.012)	(0.012)	(0.019)	(0.017)	(0.018)	(0.017)
age 9	0.999	0.987	0.981	1.003	1.004	0.980
	(0.012)	(0.012)	(0.019)	(0.017)	(0.018)	(0.016)
age 10	1.003	0.999	1.050**	1.012	1.013	1.034**
	(0.013)	(0.013)	(0.022)	(0.018)	(0.018)	(0.017)
age 11	1.002	0.997	1.004	1.034*	0.975	1.006
	(0.013)	(0.013)	(0.023)	(0.020)	(0.018)	(0.017)
age 12	1.018	1.003	1.019	1.041**	0.975	1.009
	(0.014)	(0.014)	(0.024)	(0.021)	(0.018)	(0.017)
age 13	1.050***	1.050***	1.034	1.064***	1.042**	1.040**
	(0.015)	(0.015)	(0.026)	(0.021)	(0.020)	(0.018)
age 14	1.047***	1.042***	1.024	1.058***	1.033*	1.051***
	(0.015)	(0.015)	(0.028)	(0.022)	(0.020)	(0.018)
age 15	1.083***	1.088***	1.070**	1.111***	1.087***	1.083***
	(0.016)	(0.016)	(0.032)	(0.024)	(0.022)	(0.019)
age 16	1.067***	1.065***	1.074**	1.071***	1.068***	1.075***
	(0.016)	(0.016)	(0.036)	(0.024)	(0.022)	(0.019)
age 17	1.059***	1.054***	1.055	1.063***	1.043**	1.040**
	(0.016)	(0.016)	(0.040)	(0.024)	(0.022)	(0.019)
age 18	1.013	1.034**	0.985	1.045*	1.042*	1.023
	(0.016)	(0.017)	(0.045)	(0.024)	(0.023)	(0.019)
age 19	1.002	1.026*	1.033	1.011	1.041*	1.019
	(0.017)	(0.017)	(0.061)	(0.025)	(0.023)	(0.019)
age 20	0.993	0.986	1.103	0.970	1.013	0.981
	(0.018)	(0.017)	(0.108)	(0.026)	(0.024)	(0.019)
age 21	1.002	0.999		1.052*	0.967	0.990
	(0.020)	(0.019)		(0.030)	(0.024)	(0.021)
age 22	0.996	1.006		0.994	0.988	0.980
	(0.021)	(0.020)		(0.031)	(0.027)	(0.022)
age 23	1.038	1.025		1.026	1.030	1.021
	(0.024)	(0.023)		(0.035)	(0.030)	(0.025)
age 24	1.006	1.009		0.947	1.055*	0.999
	(0.026)	(0.025)		(0.036)	(0.034)	(0.026)
age 25	0.926***	0.969		0.978	0.944*	0.947**
	(0.026)	(0.026)		(0.041)	(0.033)	(0.026)
age 26	0.966	0.998		0.961	0.994	0.971
	(0.030)	(0.029)		(0.044)	(0.038)	(0.029)
No Children dummy	4.062***	4.922***	0.235***	1.660***	10.242***	6.105***
	(0.052)	(0.061)	(0.008)	(0.011)	(0.133)	(0.163)
Registered Partnership	1.514***	1.229***	1.410***	1.406***	1.508***	1.448***

	(0.019)	(0.012)	(0.025)	(0.023)	(0.031)	(0.019)
Child born prior to marriage	0.017***	1.400***	1.315***	1.655***	1.598***	1.384***
	(0.007)	(0.006)	(0.013)	(0.012)	(0.013)	(0.009)
Spousal immigration background						
Husband native, Wife 1st gen.	1.558***	1.556***	1.290***	1.502***	1.603***	1.394***
	(0.009)	(0.007)	(0.016)	(0.012)	(0.015)	(0.011)
Husband native, Wife 2nd gen.	1.470***	1.402***	1.374***	1.419***	1.458***	1.349***
	(0.008)	(0.007)	(0.016)	(0.011)	(0.012)	(0.010)
Husband 1st gen. Wife native	2.183***	2.261***	1.867***	2.172***	2.125***	1.741***
	(0.013)	(0.011)	(0.024)	(0.017)	(0.019)	(0.015)
Husband 1st gen. Wife 1st gen.	0.752***	1.009**	0.616***	0.872***	0.757***	0.691***
	(0.003)	(0.004)	(0.006)	(0.005)	(0.005)	(0.005)
Husband 1st gen. Wife 2nd gen.	1.525***	1.753***	1.223***	1.591***	1.486***	1.228***
	(0.016)	(0.015)	(0.021)	(0.021)	(0.023)	(0.016)
Husband 2nd gen. Wife native	1.432***	1.364***	1.375***	1.389***	1.432***	1.333***
	(0.008)	(0.007)	(0.016)	(0.010)	(0.012)	(0.010)
Husband 2nd gen. Wife 1st gen.	1.315***	1.452***	1.035	1.342***	1.283***	1.196***
	(0.016)	(0.015)	(0.023)	(0.022)	(0.025)	(0.018)
Husband 2nd gen. Wife 2nd gen.	1.611***	1.564***	1.475***	1.609***	1.543***	1.401***
	(0.020)	(0.017)	(0.032)	(0.026)	(0.030)	(0.022)
Age at wedding						
Husband, linear	0.752***	0.921***	0.886***	0.828***	0.763***	0.937***
	(0.007)	(0.005)	(0.021)	(0.011)	(0.011)	(0.012)
Husband, quadratic	1.008***	1.002***	1.003***	1.005***	1.008***	1.002***
	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)
Husband, cubic	1.000***	1.000***	1.000***	1.000***	1.000***	1.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Wife, linear	0.870***	0.937***	1.023	0.857***	0.972**	1.130***
	(0.008)	(0.004)	(0.028)	(0.011)	(0.013)	(0.013)
Wife, quadratic	1.004***	1.001***	0.999	1.004***	1.000	0.996***
	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)
Wife, cubic	1.000***	1.000***	1.000	1.000***	1.000	1.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Education levels						
Husband, High School	0.888***	0.861***	0.827***	0.850***	0.938***	0.990
	(0.005)	(0.004)	(0.012)	(0.007)	(0.008)	(0.008)
Husband, University	0.679***	0.635***	0.525***	0.671***	0.738***	0.792***
	(0.004)	(0.003)	(0.008)	(0.006)	(0.007)	(0.007)
Husband, Missing	0.809***	0.797***	0.740***	0.784***	0.872***	0.910***
	(0.004)	(0.004)	(0.011)	(0.006)	(0.007)	(0.008)
Wife, High School	1.019***	0.964***	0.985	0.979***	1.059***	1.047***
	(0.005)	(0.004)	(0.014)	(0.007)	(0.008)	(0.008)
Wife, University	0.858***	0.751***	0.645***	0.865***	0.902***	0.752***
	(0.005)	(0.004)	(0.010)	(0.007)	(0.008)	(0.006)
Wife, Missing	0.562***	0.536***	0.571***	0.576***	0.577***	0.536***
	(0.003)	(0.002)	(0.008)	(0.004)	(0.004)	(0.004)
Employed in previous year, husband						1.284***
						(0.010)
Employed in previous year, wife						1.152***
						(0.007)
Log earnings, husband						0.765***
						(0.004)
Log earnings, wife						1.117***
						(0.005)
Constant	0.017***	0.001***	0.010***	0.049***	0.002***	0.001***
	(0.007)	(0.000)	(0.003)	(0.024)	(0.001)	(0.000)
Firstborns' age dummies	✓	✓	✓	✓	✓	✓
Marriage duration dummies	✓	✓	✓	✓	✓	✓
Parental birth cohort dummies	✓	✓	✓	✓	✓	✓
Calendar year dummies	✓	✓	✓	✓	✓	✓
Observations	57,597,846	66,968,517	20,191,008	26,882,338	30,385,145	26,751,766
Spells	2,949,415	3,723,060	1,066,945	1,464,874	1,486,112	2,567,219
Log-likelihood	-3,629,853	-4,732,579	-820,204	-1,915,286	-1,664,811	-1,858,790

Note: Authors' estimates of exponentiated coefficients corresponding to the cloglog hazard models of divorce risks. The models use linked marriage, divorce, and other registry data for different-sex couples who married after 1971. In the two specifications which consider the presence of male children, the dummy variables corresponding to the first three years of the firstborn's life have been replaced by a single dummy variable for ages 0-3. The last specification does not use records preceding the year 2000 due to the limited availability of employment records

*** = 0.01 significance level, ** = 0.05 significance level, * = 0.1 significance level.

3. LISS analyses

Table B3: List of dependent variables from LISS panel

Column number	Full description of the variable
1	How satisfied are you with your current relationship? (11-point scale)
2	Can you indicate whether you and your partner had any differences of opinion regarding money expenditure over the past year? (3-point scale)
3	Can you indicate whether you and your partner had any differences of opinion regarding raising the children over the past year? (3-point scale)
4	Do you agree with the following statement: A woman is more suited to rearing young children than a man. (5-point scale)
5	Do you agree with the following statement: A divorce is generally the best solution if a married couple cannot solve their marital problems. (5-point scale)
6	Do you agree with the following statement: Married people are generally happier than unmarried people. (5-point scale)
7	How would you generally describe the relationship with your family? (5-point scale)
8	Do you agree with the following statement: All in all... caring for my child is not such a burden. (5-point scale)
9	How satisfied are you with the life you lead at the moment? (11-point scale)
10	Logarithm of total expenditure per month for children living at home, children 0-15
11	How much time did you spend in the last seven days on: activities with own child. (hours)
12	How would you describe your overall relationship with your father? (4-point scale)
13	How would you describe your overall relationship with your mother? (4-point scale)
14	Have you ever smoked? (yes-no question)
15	Do you smoke now? (yes-no question)
16	How often did you have a drink containing alcohol over the last 12 months? (8-point scale)
17	Did you have a drink containing alcohol during the last seven days? (yes-no question)
18	Do you currently have a partner? (yes-no question)

Note: Thematic blocks used for the FDR control adjustment are: (1, 9), (2, 3, 7), (4, 5, 6), (10, 11), (12, 13), (14, 15, 16, 17, 18).

Table B3a: Regression analysis of parental responses in the LISS panel

	(1)		(2)		(3)		(4)		(5)		(6)	
VARIABLES	father	mother	father	mother	father	mother	father	mother	father	mother	father	mother
FB daughter	0.011	-0.190	-0.143	0.149	-0.140	-0.098	-0.115	0.042	-0.100	0.050	-0.103	-0.064
aged 0-12	(0.085)	(0.075)	(0.093)	(0.083)	(0.093)	(0.083)	(0.091)	(0.083)	(0.091)	(0.084)	(0.091)	(0.083)
FB child	0.195	0.255	-0.409	-0.215	-0.231	-0.448	0.383	0.100	-0.052	-0.040	0.242	0.291
aged 13-18	(0.114)	(0.101)	(0.127)	(0.113)	(0.126)	(0.112)	(0.123)	(0.108)	(0.125)	(0.110)	(0.123)	(0.110)
FB daughter	0.023	-0.117	0.173	0.264	0.256	0.155	-0.409	-0.165	0.063	0.332	-0.177	-0.086
aged 13-18	(0.105)	(0.094)	(0.119)	(0.106)	(0.116)	(0.104)	(0.113)	(0.100)	(0.115)	(0.103)	(0.114)	(0.100)
Number of siblings	-0.071	0.157	-0.003	0.061	0.330	0.217	0.065	0.296	-0.198	-0.175	0.126	0.058
	(0.055)	(0.048)	(0.061)	(0.053)	(0.062)	(0.054)	(0.060)	(0.052)	(0.060)	(0.053)	(0.060)	(0.053)
Number of sisters	0.008	-0.113	-0.104	-0.134	-0.323	-0.186	0.213	-0.009	-0.004	0.001	0.096	0.105
	(0.063)	(0.056)	(0.070)	(0.063)	(0.069)	(0.062)	(0.068)	(0.061)	(0.068)	(0.063)	(0.069)	(0.062)
Age at the time of survey collection												
Husband, linear	0.179	0.109	-0.118	0.042	0.127	0.173	-0.222	-0.025	0.262	0.504	-0.359	-0.298
	(0.181)	(0.168)	(0.198)	(0.184)	(0.202)	(0.187)	(0.197)	(0.181)	(0.194)	(0.187)	(0.197)	(0.184)
Husband, quadratic	-0.005	-0.004	0.003	-0.001	-0.003	-0.004	0.004	0.001	-0.004	-0.009	0.008	0.007
	(0.004)	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Husband, cubic	0.000	0.000	-0.000	0.000	0.000	0.000	-0.000	-0.000	0.000	0.000	-0.000	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Wife, linear	0.076	-0.070	-0.544	-0.439	0.209	-0.469	0.553	-0.208	0.124	0.511	0.192	-0.272
	(0.355)	(0.354)	(0.418)	(0.397)	(0.401)	(0.382)	(0.393)	(0.392)	(0.386)	(0.401)	(0.386)	(0.387)
Wife, quadratic	-0.003	0.002	0.013	0.011	-0.003	0.013	-0.014	0.003	-0.004	-0.014	-0.006	0.005
	(0.009)	(0.009)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)
Wife, cubic	0.000	-0.000	-0.000	-0.000	0.000	-0.000	0.000	-0.000	0.000	0.000	0.000	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Education levels												
Husband, high school	-0.011	0.156	0.079	-0.108	-0.258	-0.296	-0.318	-0.284	0.028	-0.045	0.264	0.120
	(0.100)	(0.082)	(0.109)	(0.091)	(0.108)	(0.092)	(0.104)	(0.087)	(0.107)	(0.090)	(0.104)	(0.087)
Husband, university	-0.175	0.351	-0.152	-0.082	-0.557	-0.567	-0.149	-0.089	0.108	-0.155	0.594	0.612
	(0.105)	(0.087)	(0.114)	(0.095)	(0.114)	(0.096)	(0.111)	(0.092)	(0.113)	(0.095)	(0.111)	(0.093)
Wife, high school	0.212	-0.202	-0.258	-0.289	0.026	0.064	0.187	-0.016	-0.199	-0.150	-0.007	0.075
	(0.096)	(0.089)	(0.106)	(0.097)	(0.106)	(0.099)	(0.103)	(0.093)	(0.105)	(0.098)	(0.104)	(0.094)
Wife, university	0.240	-0.276	-0.452	-0.541	-0.140	-0.094	-0.278	-0.333	-0.225	-0.258	0.278	0.188
	(0.106)	(0.097)	(0.118)	(0.106)	(0.117)	(0.107)	(0.114)	(0.103)	(0.117)	(0.107)	(0.116)	(0.102)

Immigration background												
Husband native,	-0.436	-0.500	0.423	0.994	0.886	1.207	0.295	0.665	-0.157	0.121	0.310	0.436
wife 1st gen.	(0.165)	(0.164)	(0.185)	(0.171)	(0.191)	(0.186)	(0.177)	(0.172)	(0.179)	(0.172)	(0.178)	(0.168)
Husband native,	-0.490	-0.669	-0.073	0.504	-0.067	1.001	-0.548	-0.284	-0.233	0.228	-0.238	0.024
wife 2nd gen.	(0.165)	(0.164)	(0.180)	(0.164)	(0.177)	(0.170)	(0.171)	(0.159)	(0.170)	(0.169)	(0.175)	(0.162)
Husband 1st gen.,	0.126	-0.471	0.569	1.222	1.293	1.805	0.197	0.338	0.492	0.420	0.497	0.675
wife native	(0.209)	(0.192)	(0.219)	(0.224)	(0.224)	(0.224)	(0.212)	(0.210)	(0.224)	(0.211)	(0.210)	(0.212)
Husband 1st gen.,	0.284	-0.039	1.022	1.078	0.511	0.881	1.146	1.502	0.453	0.496	1.234	1.326
wife 1st gen.	(0.202)	(0.198)	(0.223)	(0.206)	(0.217)	(0.210)	(0.214)	(0.217)	(0.202)	(0.208)	(0.214)	(0.208)
Husband 1st gen.,	0.748	-0.148	-0.225	-0.169	0.177	0.728	0.517	1.032	0.159	-0.267	0.432	0.850
wife 2nd gen.	(0.374)	(0.356)	(0.413)	(0.370)	(0.408)	(0.372)	(0.386)	(0.370)	(0.389)	(0.359)	(0.384)	(0.380)
Husband 2nd gen.,	0.307	-0.076	0.350	0.355	0.375	0.428	-0.414	-0.483	0.310	0.425	-0.087	-0.588
wife native	(0.170)	(0.166)	(0.193)	(0.180)	(0.199)	(0.187)	(0.205)	(0.184)	(0.208)	(0.196)	(0.198)	(0.187)
Husband 2nd gen.,	-0.366	1.043	0.510	0.106	0.161	-0.823	0.844	-1.437	-1.085	0.228	0.807	0.994
wife 1st gen.	(0.638)	(0.634)	(0.769)	(0.763)	(0.726)	(0.729)	(0.724)	(1.099)	(0.811)	(0.966)	(0.838)	(0.984)
Husband 2nd gen.,	-1.570	-1.718	1.385	0.769	1.847	1.977	-0.770	0.017	-0.142	0.130	0.505	0.852
wife 2nd gen.	(0.509)	(0.585)	(0.675)	(0.623)	(0.600)	(0.676)	(0.586)	(0.761)	(0.648)	(0.671)	(0.587)	(0.672)
Information missing	-0.147	-0.191	0.644	0.141	0.526	0.332	0.251	0.120	-0.187	0.175	0.277	0.082
	(0.100)	(0.070)	(0.111)	(0.078)	(0.112)	(0.078)	(0.113)	(0.077)	(0.111)	(0.079)	(0.113)	(0.078)
Calendar year dummies	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Observations	3,091	3,870	3,065	3,847	3,071	3,866	2,742	3,355	2,742	3,357	2,742	3,357
ln likelihood	-5,014	-6,248	-2,572	-3,251	-2,681	-3,395	-3,877	-4,685	-3,721	-4,336	-3,692	-4,461

Table B3b: Regression analysis of parental responses in the LISS panel

	(7)		(8)		(9)		(10)		(11)	
VARIABLES	father	mother	father	mother	father	mother	father	mother	father	mother
FB daughter	-0.200	-0.016	0.095	0.362	-0.028	-0.137	0.242	0.058	0.252	-1.994
aged 0-12	(0.129)	(0.111)	(0.149)	(0.134)	(0.094)	(0.084)	(0.324)	(0.118)	(0.792)	(0.991)
FB child	0.129	0.125	-0.219	0.254	0.167	0.117	0.015	-0.204	-4.086	-9.574
aged 13-18	(0.169)	(0.145)	(0.209)	(0.183)	(0.126)	(0.111)	(0.397)	(0.166)	(1.043)	(1.308)
FB daughter	-0.350	-0.142	0.121	0.070	-0.056	-0.259	-0.017	0.102	1.558	-0.729
aged 13-18	(0.151)	(0.133)	(0.193)	(0.168)	(0.115)	(0.102)	(0.389)	(0.155)	(0.949)	(1.193)
Number of siblings	-0.131	0.169	-0.340	-0.244	-0.106	0.107	0.506	0.146	0.819	2.506
	(0.082)	(0.069)	(0.101)	(0.088)	(0.061)	(0.054)	(0.236)	(0.077)	(0.515)	(0.632)
Number of sisters	0.091	-0.286	-0.024	0.083	0.012	-0.042	-0.296	0.114	-1.230	0.363
	(0.092)	(0.082)	(0.112)	(0.102)	(0.069)	(0.063)	(0.257)	(0.089)	(0.582)	(0.737)
Age at the time of survey collection										
Husband, linear	0.317	0.464	-0.237	-0.029	-0.093	0.151	3.287	0.663	-2.123	-5.256
	(0.481)	(0.504)	(0.297)	(0.285)	(0.248)	(0.218)	(1.289)	(0.282)	(1.623)	(2.192)
Husband, quadratic	-0.008	-0.012	0.006	0.000	-0.001	-0.005	-0.070	-0.012	0.039	0.096
	(0.011)	(0.012)	(0.006)	(0.006)	(0.005)	(0.005)	(0.029)	(0.006)	(0.032)	(0.044)
Husband, cubic	0.000	0.000	-0.000	-0.000	0.000	0.000	0.000	0.000	-0.000	-0.001
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Wife, linear	-0.109	-0.155	-0.886	0.268	-0.075	0.544	-4.735	-0.150	-4.623	-0.191
	(0.631)	(0.636)	(0.701)	(0.606)	(0.443)	(0.405)	(2.572)	(0.619)	(3.269)	(4.522)
Wife, quadratic	0.000	0.004	0.025	-0.006	0.006	-0.011	0.117	0.003	0.099	-0.005
	(0.015)	(0.016)	(0.018)	(0.015)	(0.011)	(0.010)	(0.063)	(0.016)	(0.081)	(0.113)
Wife, cubic	0.000	-0.000	-0.000	0.000	-0.000	0.000	-0.001	-0.000	-0.001	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)	(0.001)	(0.001)
Education levels										
Husband, high school	-0.081	0.282	0.089	0.297	-0.035	0.228	-0.202	0.009	-0.604	-0.179
	(0.153)	(0.125)	(0.172)	(0.146)	(0.110)	(0.091)	(0.461)	(0.129)	(0.899)	(1.066)
Husband, university	-0.022	0.375	0.129	0.153	0.034	0.367	-0.104	-0.085	-0.131	1.633
	(0.159)	(0.131)	(0.180)	(0.153)	(0.114)	(0.096)	(0.449)	(0.133)	(0.926)	(1.110)
Wife, high school	0.257	0.250	-0.037	-0.122	0.330	-0.011	1.001	-0.085	1.243	-0.800
	(0.151)	(0.139)	(0.170)	(0.156)	(0.106)	(0.097)	(0.412)	(0.137)	(0.869)	(1.129)
Wife, university	0.485	0.064	-0.002	-0.166	0.335	-0.014	0.559	0.053	2.893	2.432
	(0.165)	(0.147)	(0.193)	(0.170)	(0.116)	(0.106)	(0.416)	(0.152)	(0.967)	(1.235)

Immigration background										
Husband native,	-0.121	-0.473	-1.175	-0.158	-0.533	-0.507	0.290	-0.154	1.492	2.907
wife 1st gen.	(0.233)	(0.225)	(0.313)	(0.289)	(0.182)	(0.179)	(0.476)	(0.306)	(1.513)	(2.120)
Husband native,	0.174	-0.347	-0.039	-0.656	-0.430	-0.605	0.251	0.376	-0.524	0.525
wife 2nd gen.	(0.270)	(0.222)	(0.269)	(0.248)	(0.180)	(0.169)	(0.894)	(0.240)	(1.514)	(1.897)
Husband 1st gen.,	0.281	0.209	-0.633	-1.207	-0.306	-0.539	0.412	-0.073	0.970	9.818
wife native	(0.308)	(0.273)	(0.370)	(0.332)	(0.218)	(0.204)	(0.622)	(0.391)	(1.842)	(2.777)
Husband 1st gen.,	1.061	0.947	0.946	0.916	-0.357	-0.613	0.647	0.634	1.709	-3.329
wife 1st gen.	(0.296)	(0.260)	(0.360)	(0.338)	(0.219)	(0.214)	(0.575)	(0.371)	(1.698)	(2.353)
Husband 1st gen.,	0.361	-0.898	0.504	1.259	-0.717	-1.138	1.102	-1.148	-4.424	4.940
wife 2nd gen.	(0.541)	(0.463)	(0.700)	(0.748)	(0.502)	(0.423)	(1.033)	(0.619)	(3.334)	(4.244)
Husband 2nd gen.,	0.218	0.436	0.224	0.368	-0.048	0.270	-1.723	0.094	2.124	2.913
wife native	(0.258)	(0.231)	(0.333)	(0.293)	(0.197)	(0.185)	(1.028)	(0.251)	(1.691)	(2.071)
Husband 2nd gen.,	0.737				-2.427	-1.784				
wife 1st gen.	(0.859)				(0.739)	(0.684)				
Husband 2nd gen.,	-0.585	0.286	2.037	-2.167	-0.522	-1.409		-1.050	4.913	-12.789
wife 2nd gen.	(0.756)	(0.770)	(1.326)	(1.070)	(0.573)	(0.823)		(1.059)	(5.712)	(9.840)
Information missing	0.218	-0.025	0.334	-0.260	-0.207	-0.058	-0.138	0.039	2.307	-0.335
	(0.148)	(0.106)	(0.193)	(0.130)	(0.117)	(0.079)	(0.376)	(0.111)	(0.978)	(0.931)
Constant							16.153	-4.640	120.510	126.252
							(33.820)	(7.027)	(39.549)	(53.534)
Calendar year dummies	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Observations	1,794	2,217	996	1,262	2,657	3,305	207	1,025	1,079	1,354
ln likelihood	-1,764	-2,196	-1,439	-1,805	-3,858	-4,896	-385.8	-1,842	-3,977	-5,461

Table B3c: Regression analysis of responses of teenage children in the LISS panel

	(12)		(13)		(14)		(15)		(16)	
VARIABLES	firstborn	all children	firstborn	all children	firstborn	all children	firstborn	all children	firstborn	all children
Daughter	-0.211	-0.018	-0.461	-0.268	-0.111	-0.143	-0.359	-0.313	-0.996	-0.847
	(0.165)	(0.115)	(0.161)	(0.112)	(0.203)	(0.140)	(0.254)	(0.179)	(0.160)	(0.112)
Number of siblings	0.001	0.016	-0.194	-0.112	-0.112	0.094	-0.086	0.069	0.068	0.138
	(0.128)	(0.082)	(0.122)	(0.079)	(0.156)	(0.092)	(0.197)	(0.118)	(0.119)	(0.076)
Number of sisters	0.104	-0.068	0.258	0.045	-0.049	-0.110	-0.347	-0.148	0.238	-0.028
	(0.153)	(0.103)	(0.146)	(0.100)	(0.186)	(0.121)	(0.239)	(0.155)	(0.137)	(0.093)
Age at the time of survey collection										
Husband, linear	0.187	2.199	0.493	1.829	-1.980	-0.045	-1.366	0.197	0.887	1.089
	(1.109)	(0.747)	(1.117)	(0.748)	(1.382)	(0.961)	(1.600)	(1.096)	(1.150)	(0.800)
Husband, quadratic	0.001	-0.043	-0.006	-0.035	0.044	0.001	0.030	-0.005	-0.014	-0.017
	(0.024)	(0.015)	(0.024)	(0.015)	(0.030)	(0.020)	(0.035)	(0.023)	(0.024)	(0.016)
Husband, cubic	-0.000	0.000	0.000	0.000	-0.000	-0.000	-0.000	0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Wife, linear	1.161	-0.535	1.122	-0.393	0.325	-0.624	0.443	-0.738	-0.400	-0.388
	(0.775)	(0.344)	(0.778)	(0.326)	(0.982)	(0.633)	(1.134)	(0.589)	(0.753)	(0.305)
Wife, quadratic	-0.024	0.015	-0.022	0.011	-0.015	0.012	-0.015	0.015	0.011	0.011
	(0.018)	(0.008)	(0.018)	(0.008)	(0.023)	(0.014)	(0.027)	(0.014)	(0.018)	(0.008)
Wife, cubic	0.000	-0.000	0.000	-0.000	0.000	-0.000	0.000	-0.000	-0.000	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Education levels										
Husband, high school	-0.168	-0.270	0.137	-0.110	-0.266	0.105	-0.105	0.344	0.276	0.482
	(0.221)	(0.156)	(0.213)	(0.150)	(0.268)	(0.187)	(0.327)	(0.232)	(0.210)	(0.147)
Husband, university	-0.243	-0.414	0.128	-0.235	-0.638	-0.303	-0.723	-0.317	-0.042	0.154
	(0.233)	(0.160)	(0.224)	(0.154)	(0.287)	(0.194)	(0.363)	(0.254)	(0.216)	(0.150)
Wife, high school	-0.009	-0.074	-0.462	-0.172	-0.033	-0.249	-0.164	-0.463	0.084	-0.003
	(0.215)	(0.151)	(0.210)	(0.146)	(0.274)	(0.186)	(0.345)	(0.231)	(0.203)	(0.143)
Wife, university	0.201	0.223	-0.194	0.180	0.612	0.133	0.625	-0.134	0.250	0.013
	(0.253)	(0.177)	(0.248)	(0.173)	(0.309)	(0.212)	(0.381)	(0.265)	(0.231)	(0.166)
Immigration background										
Husband native, wife 1st gen.	-0.121	0.158	0.751	0.753	0.454	-0.745	0.167	-0.798	-0.545	-1.112

	(0.573)	(0.403)	(0.630)	(0.420)	(0.656)	(0.552)	(0.822)	(0.752)	(0.568)	(0.336)
Husband native, wife 2nd gen.	0.067	-0.053	-0.103	0.047	0.752	0.702	0.188	0.677	-0.128	0.241
	(0.490)	(0.326)	(0.484)	(0.335)	(0.502)	(0.354)	(0.673)	(0.426)	(0.416)	(0.312)
Husband 1st gen., wife native	0.456	0.130	-0.029	0.067		-1.336		-0.622	-1.012	-1.234
	(0.620)	(0.456)	(0.553)	(0.433)		(0.754)		(0.758)	(0.611)	(0.433)
Husband 1st gen., wife 1st gen.	1.012	0.625	-0.130	-0.327	-1.320	-1.614	-1.021	-1.267	-1.312	-1.407
	(0.411)	(0.322)	(0.365)	(0.302)	(0.642)	(0.619)	(0.776)	(0.751)	(0.335)	(0.288)
Husband 1st gen., wife 2nd gen.										
Husband 2nd gen., wife native	0.010	0.012	-0.347	-0.313	-0.974	0.004	-1.510	-1.308	-0.669	-0.093
	(0.431)	(0.369)	(0.430)	(0.362)	(0.658)	(0.461)	(1.066)	(1.033)	(0.405)	(0.364)
Husband 2nd gen., wife 1st gen.										
Husband 2nd gen., wife 2nd gen.	-1.410	-1.209	12.656	11.832						
	(1.709)	(1.664)	(620.498)	(401.765)						
Immigration information missing	0.884	0.370	0.614	-0.036	-0.742	-0.033	-0.440	0.277	0.018	0.244
	(0.372)	(0.238)	(0.340)	(0.222)	(0.507)	(0.290)	(0.599)	(0.350)	(0.290)	(0.211)
Calendar year dummies	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Observations	666	1,264	662	1,252	607	1,191	606	1,190	618	1,189
ll	-563.9	-1131	-617.8	-1235	-318.1	-644.4	-227.5	-447.3	-958.5	-1864

Table B3d: Regression analysis of responses of teenage children in the LISS panel

	(17)		(18)	
VARIABLES	firstborn	all children	firstborn	all children
Daughter	-0.763	-0.800	1.168	0.908
	(0.201)	(0.139)	(0.213)	(0.152)
Number of siblings	-0.079	0.203	-0.509	-0.450
	(0.156)	(0.098)	(0.165)	(0.112)
Number of sisters	0.394	-0.001	0.203	0.246
	(0.182)	(0.120)	(0.187)	(0.134)
Age at the time of survey collection				
Husband, linear	4.765	0.372	2.067	-0.179
	(3.338)	(2.052)	(2.012)	(1.189)
Husband, quadratic	-0.088	-0.003	-0.048	0.001
	(0.066)	(0.040)	(0.041)	(0.024)
Husband, cubic	0.001	-0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)
Wife, linear	2.785	1.392	1.337	0.939
	(1.819)	(1.280)	(1.673)	(0.901)
Wife, quadratic	-0.070	-0.031	-0.030	-0.023
	(0.041)	(0.028)	(0.038)	(0.020)
Wife, cubic	0.001	0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)
Education levels				
Husband, high school	0.264	0.483	0.375	0.296
	(0.277)	(0.189)	(0.274)	(0.200)
Husband, university	-0.347	-0.001	0.409	0.283
	(0.280)	(0.193)	(0.290)	(0.209)
Wife, high school	0.328	0.166	0.082	0.063
	(0.261)	(0.181)	(0.266)	(0.191)
Wife, university	0.948	0.596	0.206	-0.060
	(0.310)	(0.217)	(0.316)	(0.229)
Immigration background				
Husband native, wife 1st gen.	-0.515	-0.859	-0.236	-0.067
	(0.757)	(0.480)	(0.653)	(0.496)
Husband native, wife 2nd gen.	-0.850	-0.159	-0.474	-0.060
	(0.548)	(0.375)	(0.664)	(0.424)
Husband 1st gen., wife native	-0.594	-1.111	-0.613	-0.470
	(0.760)	(0.554)	(0.820)	(0.648)
Husband 1st gen., wife 1st gen.	-2.609	-2.423	-0.520	-0.482
	(0.795)	(0.652)	(0.447)	(0.391)
Husband 1st gen., wife 2nd gen.				
Husband 2nd gen., wife native	-0.478	-0.253	-2.115	-2.229
	(0.548)	(0.454)	(1.067)	(1.028)
Husband 2nd gen., wife 1st gen.				
Husband 2nd gen., wife 2nd gen.				
Immigration information missing	0.076	0.469	-0.270	-0.206
	(0.398)	(0.287)	(0.389)	(0.285)
Calendar year dummies	✓	✓	✓	✓
Observations	529	1,028	669	1,271
ll	-315.7	-634.8	-326.9	-608.0

Note: Authors' estimates of coefficients corresponding to the ordered logit and regression models of responses of parents and children in different-sex couples with a firstborn biological child younger than 19 at the time of the survey. LISS data 2008-2016. Standard errors in parentheses.