# Do Parents of Girls Have a Higher Risk of Divorce? An Eighteen-Country Study*

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March 11, 2004

* We gratefully acknowledge the help and advice of Anatol Rapoport and Tom DiPrete. We thank S. Philipp Morgan for a fair review and valuable suggestions. We also thank two anonymous reviewers for valuable hints and the Advisory Group of the Fertility and Family Survey programme of comparative research for its permission, granted under identification number 69, to use the Fertility and Family Survey data on which this study is based. While working on this research, Andreas Diekmann was at the Institute of Sociology and Kurt Schmidheiny at the Economics Department of the University of Bern.

*Abstract*

Using data from June 1980 Current Population Survey, Morgan, Lye, and Condran (1988) reported that families with a daughter have a higher divorce risk than families with a son. They attribute this finding to the higher involvement of fathers in raising a son, which, in turn, promotes marital stability. We investigate the relation between gender composition of children and parents' divorce risk with cross-national data from the Fertility and Family Survey. These data, which cover 16 European countries, Canada, and the U.S, do not support a general hypothesis that sons contribute more to marital stability than daughters.

*Key Words*: Children and Divorce, Cross-national, Divorce, Gender of Children

Are families with daughters more divorce-prone than families with sons? Although in general couples with children have a lower risk of divorce than childless couples, the size of the stabilizing effect may depend on the children's gender. In a study using data from the 1980 Current Population Survey, Morgan, Lye, and Condran (1988) concluded that the risk of marital disruption in one-child families is moderately higher when the child is a daughter. They estimated that for families with one or two children each girl increases the risk of divorce by 9%, all other things being equal. Unsurprisingly, these findings captured the attention of family researchers. In a review article on the determinants of divorce, White (1990, p. 907) went so far as to comment: "Perhaps the most interesting finding of the decade is Morgan, Lye, and Condran's (1988) finding that parents of sons are less likely to divorce than parents with daughters."

To explain their finding, Morgan et al. put forward a *father-involvement hypothesis*. They claimed that, on average, fathers play a greater role in raising sons and therefore spend more time with them than with daughters. The higher degree of fathers’ involvement in childrearing lowers the divorce risk. The supposed reason for this effect is that fathers’ involvement contributes to a change in the household division of labor, which leads to less likelihood of divorce. In a recent paper, Morgan and Pollard (2002) outline this reasoning. They believe that fathers’ involvement increases “solidarity based on similarity, especially shared experience and values” (p. 3). In the context of “the contemporary companionate marriage” this type of solidarity or “companionship” creates a barrier to divorce (p. 3). Thus, because of the increased involvement of fathers in raising a son, these marriages enjoy more companionship than those in which children are daughters. This in turn leads to a reduced risk of divorce.

Evidence for the differential involvement of fathers is also provided by cross-cultural studies (White, 1990). Using data from the 1981 National Survey of Children, Morgan et al. further demonstrated that mothers report a closer relationship between fathers and sons than between fathers and daughters. Moreover, fathers and sons share more activities than fathers and daughters. In terms of the family economics, one might also hypothesize that fathers' specific investments are higher in marriages with sons than in marriages with daughters. If this assumption is met and because divorce rates are inversely related to marriage-specific investments (Becker, Michael, & Landes, 1977), marriages with sons are expected to have a lower divorce risk than marriages with daughters.

Somewhat surprisingly, Morgan et al. failed to mention the issue of gender preference (Andersson & Woldemicael, 2001). If one gender of offspring is preferred, marital-specific investment should be perceived to be higher for the preferred gender. In this case one might argue that divorce rates that are gender-of-child-specific reflect differential preferences for children's gender. Andersson and Woldemicael (2001) showed this relationship for two-child families in Sweden. They found both that the divorce risk is slightly reduced if a woman has one child of each gender and that this gender composition is preferred by the majority of two-child parents. Hank and Kohler (2000) investigated gender preferences in 17 European countries using data from the Fertility and Family Survey. Data are from the period 1988 to 1996, and preferences are determined indirectly by estimating the probability of having or desiring a third child dependent on the gender composition of the first two children. With East Germany and West Germany analyzed separately, they found preferences for mixed gender compositions in 10 countries, preferences for girls in three countries, and no indication for gender preferences in five countries. If gender preferences for children were related to parents' divorce rates, the above cross-country pattern should be mirrored in the relative divorce rates. This cross-country comparison of gender preference and gender-specific divorce risk was thus waiting to be investigated.

There might be another alternative that is often neglected by researchers; the possibility that the reported effect is very weak or does not exist at all and that Morgan et al. were the victims of a Type I ($\alpha$) error. Andersson and Woldemicael (2001, p. 5), after reviewing the few existing studies, conclude: "Since very little support has materialized for the finding of Morgan et al. (1988), it might be plausible to suspect that their finding was mostly a result of random variation showing up in their data." Andersson and Woldemicael themselves conducted a very careful investigation using Swedish register data that excluded sources of random variation as far as possible. The database consisted of all first marriages with children in the years 1971 to 1995, of which about 100,000 ended in divorce. Estimation of a piece-wise constant hazard-rate model revealed that in one-child marriages the child’s gender has no effect on the divorce rate. Furthermore, and contrary to Morgan et al., in two-child marriages the divorce risk is slightly higher (by 4 percent) if both children are of the same gender. For the relatively small number of families with three children, however, boys moderately reduce the divorce risk. Similar studies have been carried out in Australia (Bracher et al., 1993), Germany (Wagner, 1997), and Switzerland (Diekmann & Schmidheiny, 2001) using much smaller population samples. No significant effect of gender composition showed up in either the Australian, Swiss, or East German sample. Only for West Germany did Wagner (1997) report a marriage-stabilizing effect if families with one child had a boy. All in all, very few studies explore how gender of offspring affects the risk of divorce in highly industrialized countries. Although findings are mixed, most of these studies are not congruent with the Morgan-et-al. hypothesis.

In a recent paper Morgan and Pollard (2002) provide evidence from the U.S. Current Population Survey (CPS) for the attenuation of the effect. They argue that the gender composition of children was related to divorce for marriages in the time span from 1960 to 1979. Thereafter, the effect has declined because changing family roles have led to a more egalitarian distribution of fathers' attention to sons and daughters. We deal with the *attenuation hypothesis* later in this article.

In our study we investigate possible effects of gender composition of children on parents' divorce rate using the *Fertility and Family Survey*. To estimate these effects we apply event-history methods. Whether a daughter or a son is born in a marriage is by nature a *random experiment* (Morgan et al. 1988). Thus in principle, a multivariate model is not necessary in estimating gender-composition effects. To be cautious, however, we included other independent variables as controls. Moreover, estimates of control variable effects support the validity of the data and of the model estimation. Doubts concerning data validity would arise if we were unable to reproduce well known results from a large mass of previous research. It is known, for example, that divorce rates vary by cohort, that children reduce the likelihood of divorce, and that age at marriage is negatively correlated with divorce risk (White, 1990). We expect these relations to show up in our data as well.

The 19 samples from 18 countries allow us to study cross-cultural variations in the effects on the divorce rate. Our main aim was to determine whether the findings of Morgan et al. could be replicated in the U.S. and other highly industrialized countries. Second, we examined cultural differences, as it is conceivable that, for example, Scandinavian countries might exhibit a different pattern of gender-composition effects than the U.S. or Canada. Third, we turned our attention to the preference hypothesis. Our cross-cultural comparison shows whether there is a consistent pattern of correlation between dominant preferences for the gender composition of children and marital stability.

## METHOD

### *Data*

The Fertility and Family Survey comprises completed surveys from 20 countries, but the necessary information on the duration of marriages or about children who have left the respondent's household is lacking for two of them. Thus, with West and East Germany analyzed separately, our estimates are based on 19 data sets collected from 16 European countries, Canada, and the U.S. in the early 1990s. Note, however, that the data are retrospective and include marriage cohorts from the 1970s (in some countries from the 1960s) up to the early 1990s. Samples were drawn from the total population within certain age limits. The Belgian sample covers only Flanders and the region of Brussels. For more information on the Fertility and Family Survey see Schoenmaeckers and Lodewijcky (1999) and Festy and Prioux (2002). For information on samples and descriptive statistics, see Table 1. In several countries participating in the Fertility and Family Survey, men were excluded. Hence, for cross-national comparisons, we used only samples from female respondents and we confined our analysis to those women who were or had previously been married. Families with adopted children or that experienced the death of a child and cases with missing data for one or more variables in the estimated equation were excluded. With these restrictions, net sample sizes varied from 1,219 (West-Germany) to 5,396 (U.S.).

[Table 1 about here]

### *Dependent Variable*

The variable of main concern is *duration of first marriage* in months. We consider a marriage as terminated when it ends in divorce or permanent separation, as the Fertility and Family Survey does not distinguish between these two occurrences. The dissolution of a common household, or, in the case of intact marriage, the date of the interview is taken as the terminal date. This definition leaves marriage duration unaffected by the time between the end of coresidence and the date of the legal divorce, which varies substantially across the different jurisdictions. (See Festy & Prioux , 2002, p. 32, for a discussion of the comparability of Fertility and Family Survey partnership data.) We use duration data as well as information on the gender composition of the children and other covariates to estimate a model linking the divorce risk to covariate effects (see below).

### *Independent Variables*

*Gender composition of children.* The independent variable of interest was the gender composition of children. Children's gender pattern was determined by seven dichotomous variables: (a) reference group of childless marriages; (b) marriages with a girl as first child; (c) marriages with a boy as first child; (d) marriages with a child of each gender; (e) marriages with two girls; (f) marriages with two boys; (g) marriages with three or more children. The gender of the third child is not considered because of the very low number of cases for the various types of gender composition in families with three children.

*Control variables.* In addition to gender composition, we included as independent variables marriage cohorts, age at marriage, and education. We used five-year *marriage cohorts* (marriages before 1969, 1970 – 1974, 1975 – 1979, 1980 – 1984, 1985 – 1989, 1990 and later) with the cohort of marriages between 1975 and 1979 as the reference group. We also control for the *marriage age* of the respondent, that is, the wife. *Education* of the female respondents was measured in accordance with the International Standard Classification of Education. This scale covers seven educational levels from preprimary (0) to the second stage of tertiary education (6). The other levels of the scale are: (1) Primary education or first stage of basic education, (2) lower secondary or second stage of basic education, (3) (upper) secondary education, (4) postsecondary nontertiary education, and (5) first stage of tertiary education. (See UNESCO 1997 for more details.) We summarized education in three dummy variables: *lower* (valued 0, 1, or 2 by International Standard Classification of Education), *medium* (3 or 4), and *higher* (5 or 6). Table 1 displays the variables and their means.

### *Statistical Model and Estimation*

The effects of gender composition of children and control variables on divorce risk were estimated using event history analysis. The multivariate estimation of effects was based on the parametric *sickle model* (see Figure 1). We chose this model because it is well known that divorce risk increases with the duration of marriage to a maximum value and decreases thereafter. This sickle-shaped time dependency of the hazard rate of divorce can be modelled by the following function (Diekmann & Mitter, 1984):

$$r(t) = a\, t\, \exp(-t/\lambda),$$

where $a = a_0 a_1^{x_1} \ldots a_k^{x_k} \ldots a_m^{x_m}$ and $\alpha_k > 0$ for all $k = 1, \ldots, m$.

[Figure 1 about here]

In this model, $x_1, \ldots, x_m$ are covariates and $a_0$ as well as $\lambda$ denote parameters to be estimated empirically. $(\alpha_k - 1)\cdot 100$ can be interpreted as the percentage effect of covariate $k$ on the risk of divorce $r(t)$. If $\alpha_k > 1$, there is a positive effect of a covariate on the risk of divorce; if $\alpha_k < 1$, the effect is negative. The parameter $\lambda$ is interpreted as the marriage duration up to the maximum risk. A further feature of the model is that it allows for *immunity*. In the present context, this means that the model allows for a certain proportion of marriages to last indefinitely.

We used the maximum likelihood method to estimate the $a$-parameters of covariate effects and the $\lambda$-parameter. The independent variables, marriage cohort, age of marriage, and education, were treated as time constant. The dichotomous gender-composition variables were time-dependent covariates that could vary within a marriage. A gender-composition variable was treated as zero in the beginning, switched to 1 when a child with the gender of interest was born, and returned to zero again when another child was born. We estimated the parameters of time-dependent covariates in the likelihood function using the method of episode-splitting (see, for example, Blossfeld & Rohwer, 1995). Roughly speaking, episode-splitting is a method for decomposing an episode such as marriage duration into subintervals. Within subintervals covariates remain constant, and the likelihood function can be rewritten as a product of the subinterval-specific likelihoods.

The complete length of the episode can be observed only in marriages ending in divorce before the interview. Marriages still existing at the time of the survey or those ended by the death of a spouse were treated as censored data. The complete episodes as well as the censored ones were used for estimating the $\alpha$-effects and the $\lambda$-parameter. In the presence of censored data, the maximum likelihood method provided consistent and (asymptotically) normally distributed estimates of the parameters.

## Results and Discussion

### *One-Child Families*

We begin by considering gender effects in marriages with one child. Figure 2 shows divorce risks in one-child marriages with a son relative to those with a daughter for all 19 samples. (Table 2 shows complete estimation results.) In Austria, for example, the point estimate is 1.14 compared with 0.65 in Canada. This means that Austrian families with a son show a 14% higher divorce risk than Austrian families with a daughter. In Canadian families with a son, the divorce risk is 35% lower than in Canadian families with a daughter. We found more countries with a relative risk in favor of families with boys (11 samples) than countries with no gender difference (relative risk in the range 0.99 to 1.01 in four samples) or countries with a gender difference in favor of families with girls (four samples).

[Table 2 about here]

Although there is a tendency for the sign of the effects to be in accordance with the Morgan et al. hypothesis, the estimates are not statistically significant at the $p < .05$ level for any of the countries except Canada. In other words, only one sample out of 19 confirms the Morgan-et-al. findings. Note that by assuming that the probability of a Type I ($\alpha$) error is 0.05, we would expect one coefficient to be significant, even if the null hypothesis of no gender effect is valid for all 19 samples. Even if we increase the probability of a Type I error to $p < .10$ or if we perform a one-sided test, there is no significant gender effect except for Canada.

[Figure 2 about here]

*Two-Child Families*

Evidence for the Morgan-et-al. hypothesis is weaker if we inspect the estimates for two-child families. Figure 3 shows divorce risks in two-girl families and those in two-boy families relative to families with mixed-gender composition. According to the Morgan-et-al. hypothesis, we would expect the lowest relative risk in two-boy families, a medium risk in famlies with mixed-gender composition, and a higher risk in families with two girls. We observe this pattern in only three of the 19 samples, however (Estonia, France, USA). Two-boy families have the lowest risk in six samples (Canada, Estonia, France, Hungary, Poland, USA). Comparing two-girl families with two-boy families, the risk for the latter is lower in ten samples, higher in seven samples, and there is no risk difference in two samples (Norway and Spain). This is not much better than tossing a coin. Also, none of the estimated coefficients is statistically significant at $p < .05$. Note, however, that the point estimates for the U.S. sample – significance of the coefficients not being considered – are perfectly in line with the Morgan et al. study, both for one-child families as well as for those with two children.

[Figure 3 about here]

### *The Attenuation Hypothesis*

Is the gender-composition effect confined to the 1960 to 1979 cohorts and has it diminished since then? This attenuation hypothesis was put forward by Morgan and Pollard (2002). To test this hypothesis we incorporated in our model an interaction effect with an additional dichotomous cohort variable (marriages up to 1979 versus marriages contracted after 1979, table not shown). Only one sample supports the attenuation hypothesis: In Switzerland we observe a significant ($p < .05$) gender-composition effect for older but not for younger cohorts, as predicted by Morgan and Pollard. Contrary to the prediction of the attenuation hypothesis, there is a significant effect for younger cohorts but not for older ones in Italy. The gender-composition effects for all other countries are not significant. Ignoring significance, in only 6 samples the initially negative effect becomes weaker (East and West Germany, Slovenia) or even positive (Belgium, Slovenia, Switzerland). The other 13 countries are not in accordance with the attenuation hypothesis. For the U.S. sample, the relative risk is 0.93 in the older cohorts and 0.85 in the younger cohorts. In sum, our findings do not support the attenuation hypothesis.

### *The Effect of the Control Variables*

In contrast, we find robust estimates of other well known factors contributing to divorce risk or promoting family stability (see Table 2). In almost all countries, a child acts as a barrier to divorce, and two children inhibit divorce even more, East Germany being a notable exception. In almost all countries, age at marriage is significantly and inversely related to the probability of marriage dissolution. Although these findings are not new, clear and consistent replication of these effects confirm the validity of the data.

### *Country Differences*

As discussed in detail below, one can argue that the effects of children's gender on parents' divorce are small and difficult to detect by survey data of moderate sample size. Yet Andersson and Woldemicael's (2001) analysis of Swedish register data for the total population does not support the Morgan-et-al. hypothesis either. One might argue that Sweden is not a perfect counterexample, as the magnitude and direction of the effect probably vary in different cultures, and Sweden is more egalitarian than, say, North America. Thus, it comes as no surprise that the effect of gender composition of children cannot be observed in Sweden. The line of argumentation related to cultural variability would be more convincing if we could demonstrate that countries with similar cultural traits showed similar patterns of effect. Unfortunately there is no consistent pattern of coefficient signs for Catholic southern Europe, for Protestant Scandinavia, for the Baltic states, or for Eastern Europe (see Figure 2).

### *Testing the Preference Hypothesis*

We now turn our attention to the preference hypothesis. Three studies (Hank & Kohler, 2000; Marleau & Saucier, 1996; Yamaguchi & Ferguson, 1995) provide data on gender preferences of parents for all countries in our sample except Estonia. No gender preference for boys was detected in any of the Fertility and Family Survey countries. In the majority of countries, if there were any preferences at all, they were for mixed gender. If gender preferences were related to divorce rates and if there was a preference for mixed gender in a specific country, we would expect the lowest divorce rates in couples with a child of each gender. This is observed in Belgium, Italy, Spain, Sweden, and Switzerland (Hank & Kohler, 1999, Figure 3). Mixed gender preference is also observed in Austria, Canada, East Germany, Hungary, Latvia, Slovenia, and the U.S., however, although in these countries marriages with a son-daughter composition do not have lower divorce risks than those with homogeneous compositions. A simple cross-tabulation of the two variables, gender preference in country (mixed vs. other) and divorce risk (lowest for mixed composition vs. lowest for two boys or two girls) shows that there is no substantial relation between the two variables ($\phi = 0.08$, $p = 0.73$, $n = 18$).

### *Fathers' Involvement*

The Fertility and Family Survey data provide information on fathers' involvement in raising children. In eight countries participating in the survey, respondents were asked to indicate which parent took care of the child. Five children's needs were examined: (a) being fed, (b) getting dressed, (c) care during illness, (d) played with, and (e) assistance in homework. To obtain meaningful comparisons, we confined our samples to married female respondents who were raising one child under age 15 at the time of the survey. Using a five-point scale, respondents reported their husbands' involvement in caring for the child in the situations mentioned above. The two upper categories were combined. For each country and category, the percentage of involvement in raising a son was compared to that in raising a daughter. The countries included were Austria, Belgium, Germany, Hungary, Italy, Lithuania, Spain, and Switzerland. Because categories (d) and (e) above were not included in the Belgian survey, we performed 38 comparisons (seven countries with five categories plus three categories for Belgium). We are cautious here because these are not hard data and may be biased by wishful thinking or social desirability. Surprisingly, only one significant difference (at $p < .05$) was observed: 48% of Spanish fathers cared for a son during illness, whereas only 37% cared for a daughter, according to their wives' reports. As we detect no systematic gender-specific involvement of fathers, the investigation of its possible relationship with divorce rates becomes pointless.

In most Fertility and Family Survey countries, data on the fathers' involvement were collected in the 1990s. Thus, it might be possible that there has been a trend toward egalitarian behavior of fathers in respect to a child's gender, as suggested by Morgan and Pollard (2000). At least the data do not contradict the hypothesis of an egalitarian trend. Except for one item in Spain, we found no difference in reported involvement with respect to child's gender in the countries included in the Fertility and Family Survey. From these data alone, therefore, we do not know whether there was a significant difference in the past, so that a trend toward more egalitarian behavior cannot be proven.

### *Sample Size*

There is one additional question concerning the validity of our findings. Morgan et al. reported that the effect of gender composition of children on divorce is quite small. To detect a small effect, one needs a large sample size. Of course, there is not only a Type I error but also a Type II error. With a small sample, a real effect might fail to reach significance. In our analysis, sample size is in the range of 1,219 (West Germany) to 5,396 (U.S.) and the number of divorces varied from 118 (Slovenia) to 1,800 (U.S.). Yet, for the U.S. the point estimate of 12% reduction in risk for families with a boy, which is in accordance with the Morgan et al. hypothesis, is still not significant. Therefore, one might object that by design our test of the hypothesis is overly strict and biased against a positive outcome. Only a large effect, which was not postulated by Morgan et al., can meet the $p < 0.05$ level. Even if we ignore inferential statistics and simply concentrate on point-estimates, however, the cross-national comparisons do not support the hypothesis that boys stabilize a marriage. For families with one child, the postulated direction of the effect was observed in 11 of 18 samples. The picture is even clearer for families with two children. Here, in only 3 of 18 samples, the prediction is in accordance with the rank order of effects.

### *Conclusion*

In sum, neither cultural diversity nor gender preference explains differences in the size and direction of gender-composition effects on divorce. If there is a gender-composition effect in a country at all, it is relatively weak. The least that one can say is that the relation of gender composition of children to divorce is not universal. Evidence from 19 samples in the Fertility and Family Survey supports neither the hypothesis that in general families with sons have a lower divorce risk than families with daughters, nor that any other type of gender composition is systematically related to the risk of divorce in one- or two-child families.

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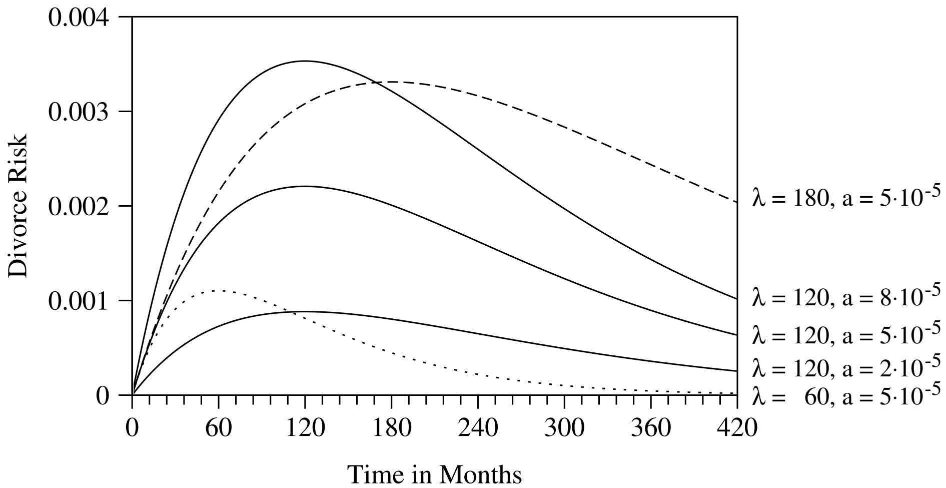
# LEGENDS

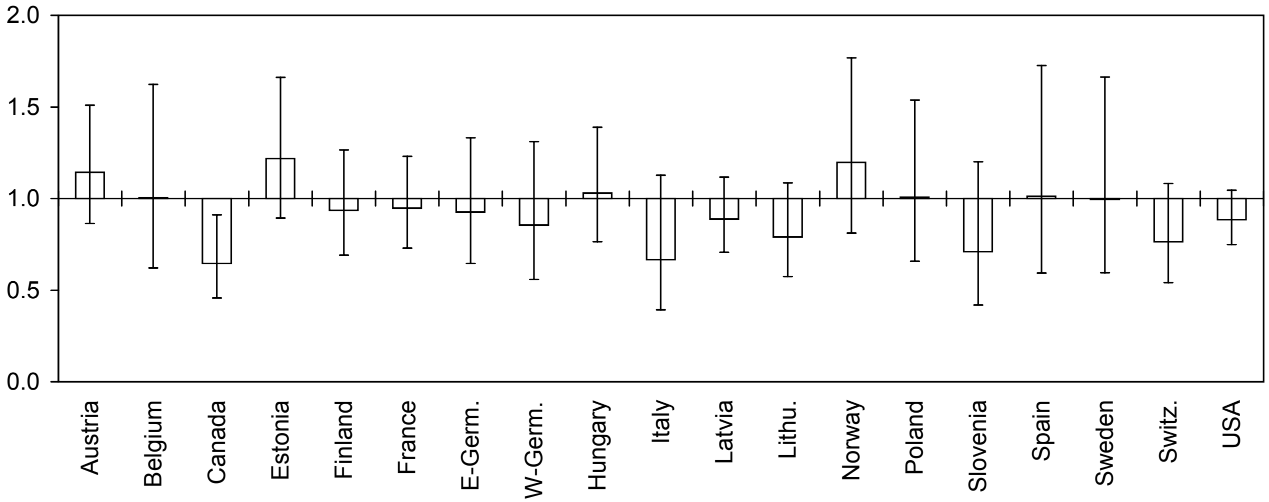
Figure 1: The divorce risk function of the sickle model for different parameter values. Values of $\lambda$ and a are indicated to the right on the corresponding line.

Figure 2: Risk of divorce in a one-child family. Risk relative to one girl. 95% confidence interval marked by error bars. Reading the effect (e.g., Austria): Families with one boy have a 14% (i.e., 1.14 - 1) higher divorce risk than families with one girl.

Figure 3: Risk of divorce in a two-child family. Risk relative to mixed-gender composition. Reading the effect (e.g., Austria): Families with two boys have a 22% (i.e., 1.22 - 1) higher divorce risk than families with mixed-gender composition.



0.004
0.003
0.002
0.001
0
Divorce Risk
0
60
120
180
240
300
360
420
Time in Months
λ = 180, a = 5·10-5
λ = 120, a = 8·10-5
λ = 120, a = 5·10-5
λ = 120, a = 2·10-5
λ = 60, a = 5·10-5



2.0
1.5
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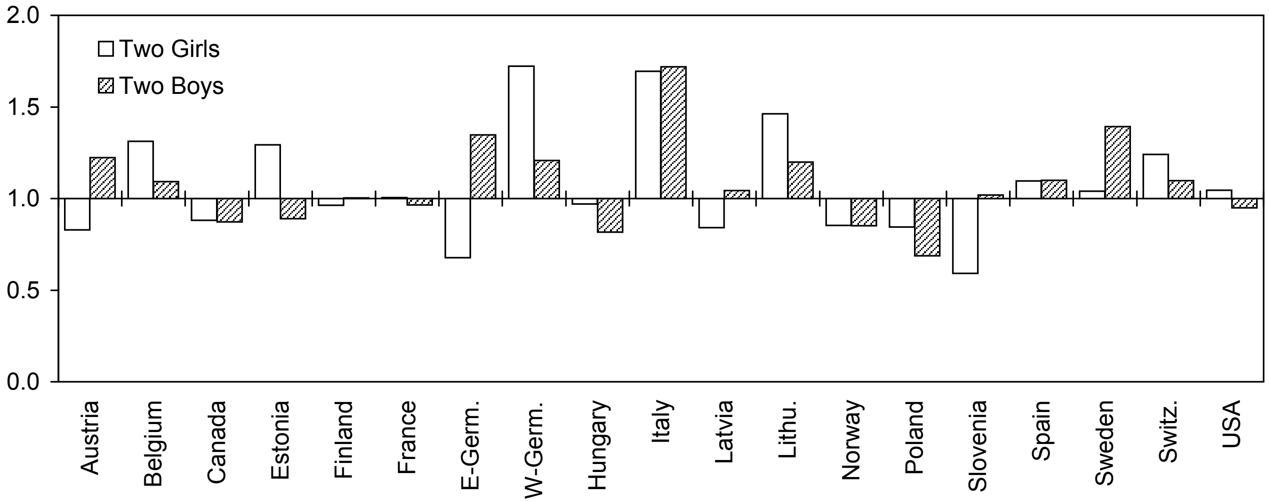
0.5
0.0
Austria
Belgium
Canada
Estonia
Finland
France
E-Germ.
W-Germ.
Hungary
Italy
Latvia
Lithu.
Norway
Poland
Slovenia
Spain
Sweden
Switz.
USA


2.0
1.5
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0.0
Two Girls
Two Boys
Austria
Belgium
Canada
Estonia
Finland
France
E-Germ.
W-Germ.
Hungary
Italy
Latvia
Lithu.
Norway
Poland
Slovenia
Spain
Sweden
Switz.
USA

Table 1: Descriptive Statistics and Information on Sample

| | Austria | Belgium | Canada | Estonia | Finland | France | E-Germ. | W-Germ. | Hungary | Italy | Latvia | Lithu. | Norway | Poland | Slovenia | Spain | Sweden | Switz. | USA |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| *Gender pattern* | | | | | | | | | | | | | | | | | | | |
| No children | 0.08 | 0.21 | 0.18 | 0.05 | 0.11 | 0.06 | 0.08 | 0.21 | 0.09 | 0.13 | 0.07 | 0.09 | 0.11 | 0.07 | 0.04 | 0.11 | 0.11 | 0.17 | 0.18 |
| One girl | 0.45 | 0.38 | 0.38 | 0.47 | 0.44 | 0.46 | 0.43 | 0.38 | 0.45 | 0.41 | 0.46 | 0.44 | 0.44 | 0.44 | 0.40 | 0.38 | 0.41 | 0.40 | 0.38 |
| One boy | 0.47 | 0.41 | 0.43 | 0.47 | 0.46 | 0.47 | 0.49 | 0.41 | 0.47 | 0.46 | 0.47 | 0.47 | 0.44 | 0.49 | 0.56 | 0.51 | 0.48 | 0.44 | 0.43 |
| One girl & one boy | 0.33 | 0.27 | 0.29 | 0.34 | 0.33 | 0.34 | 0.29 | 0.23 | 0.30 | 0.28 | 0.27 | 0.28 | 0.34 | 0.32 | 0.32 | 0.30 | 0.37 | 0.32 | 0.30 |
| Two girls | 0.17 | 0.12 | 0.15 | 0.16 | 0.15 | 0.17 | 0.12 | 0.10 | 0.14 | 0.13 | 0.15 | 0.12 | 0.16 | 0.16 | 0.13 | 0.13 | 0.14 | 0.14 | 0.13 |
| Twoboys | 0.18 | 0.14 | 0.17 | 0.18 | 0.18 | 0.18 | 0.14 | 0.11 | 0.17 | 0.16 | 0.15 | 0.16 | 0.18 | 0.19 | 0.24 | 0.21 | 0.19 | 0.18 | 0.17 |
| Three children | 0.22 | 0.15 | 0.22 | 0.24 | 0.23 | 0.30 | 0.09 | 0.11 | 0.13 | 0.15 | 0.15 | 0.11 | 0.22 | 0.25 | 0.12 | 0.19 | 0.26 | 0.20 | 0.25 |
| *Marriage cohorts* | | | | | | | | | | | | | | | | | | | |
| Cohort - 1969 | 0.19 | | 0.16 | 0.14 | 0.33 | 0.15 | | | | 0.08 | 0.08 | 0.07 | 0.20 | 0.18 | 0.02 | 0.05 | | 0.08 | |
| Cohort 1970 - 1974 | 0.16 | 0.15 | 0.12 | 0.18 | 0.19 | 0.22 | 0.13 | 0.13 | 0.18 | 0.18 | 0.16 | 0.15 | 0.23 | 0.20 | 0.15 | 0.13 | 0.14 | 0.15 | 0.17 |
| Cohort 1975 - 1979 | 0.15 | 0.28 | 0.16 | 0.21 | 0.18 | 0.24 | 0.27 | 0.22 | 0.27 | 0.19 | 0.21 | 0.17 | 0.22 | 0.21 | 0.24 | 0.22 | 0.20 | 0.17 | 0.20 |
| Cohort 1980 - 1984 | 0.15 | 0.27 | 0.20 | 0.18 | 0.17 | 0.18 | 0.28 | 0.29 | 0.23 | 0.18 | 0.21 | 0.19 | 0.19 | 0.18 | 0.21 | 0.21 | 0.20 | 0.18 | 0.22 |
| Cohort 1985 - 1989 | 0.18 | 0.31 | 0.17 | 0.19 | 0.14 | 0.12 | 0.32 | 0.37 | 0.23 | 0.18 | 0.19 | 0.21 | 0.15 | 0.23 | 0.23 | 0.22 | 0.33 | 0.23 | 0.22 |
| Cohort 1990 - | 0.18 | | 0.17 | 0.11 | | 0.08 | | | 0.10 | 0.20 | 0.15 | 0.23 | | | 0.16 | 0.19 | 0.14 | 0.20 | 0.19 |
| *Age at marriage* | | | | | | | | | | | | | | | | | | | |
| Age women | 22.0 | 21.6 | 22.6 | 21.7 | 22.5 | 21.4 | 21.3 | 22.6 | 20.2 | 22.8 | 21.3 | 21.6 | 22.2 | 21.5 | 21.3 | 22.6 | 25.1 | 24.2 | 21.7 |
| *Highest level of education* | | | | | | | | | | | | | | | | | | | |
| Low | 0.27 | 0.31 | 0.15 | 0.11 | 0.24 | 0.44 | 0.13 | 0.50 | 0.48 | 0.54 | 0.07 | 0.05 | 0.16 | 0.51 | 0.26 | 0.75 | 0.13 | 0.14 | 0.16 |
| Middle | 0.55 | 0.64 | 0.65 | 0.68 | 0.68 | 0.38 | 0.63 | 0.43 | 0.52 | 0.37 | 0.71 | 0.67 | 0.61 | 0.41 | 0.59 | 0.20 | 0.47 | 0.80 | 0.60 |
| High | 0.18 | 0.05 | 0.20 | 0.21 | 0.09 | 0.18 | 0.24 | 0.07 | | 0.09 | 0.22 | 0.28 | 0.23 | 0.08 | 0.15 | 0.06 | 0.41 | 0.06 | 0.24 |
| Year of Survey | 95/96 | 91/92 | 95 | 94 | 89/90 | 94 | 92 | 92 | 92/93 | 95/96 | 95 | 94/95 | 88/89 | 91 | 94/95 | 94/95 | 92/93 | 94/95 | 95 |
| Age Range | 20 - 54 | 21 - 40 | 15 - 54 | 20 - 69 | 22 - 51 | 20 - 49 | 20 - 39 | 20 - 39 | 18 - 41 | 20 - 49 | 18 - 49 | 18 - 49 | 20 - 43 | 18 - 49 | 15 - 45 | 18 - 49 | 23 - 43 | 20 - 49 | 15 - 44 |
| Female Respondents | 4581 | 3235 | 4166 | 1918 | 4155 | 2944 | 2984 | 3012 | 3554 | 4824 | 2699 | 3000 | 4019 | 4211 | 2798 | 4021 | 3318 | 3881 | 10847 |
| Married | 3377 | 2437 | 2668 | 1424 | 3094 | 1734 | 1996 | 1626 | 2829 | 3260 | 2146 | 2311 | 2359 | 3249 | 2005 | 2693 | 1880 | 3087 | 6844 |
| Dropped Cases | 597 | 156 | 719 | 135 | 271 | 141 | 620 | 407 | 211 | 229 | 221 | 209 | 303 | 323 | 273 | 151 | 446 | 229 | 1448 |
| Marriages in Sample | 2780 | 2281 | 1949 | 1289 | 2823 | 1593 | 1376 | 1219 | 2618 | 3031 | 1925 | 2102 | 2056 | 2926 | 1732 | 2542 | 1434 | 2858 | 5396 |
| Divorced | 529 | 220 | 511 | 340 | 490 | 643 | 249 | 206 | 422 | 147 | 541 | 310 | 312 | 186 | 118 | 126 | 199 | 418 | 1800 |

Note: The descriptive statistics for the gender pattern counts every occurence of a gender pattern during the course of a marriage as one observation. Families with more than one child appear thus in several categories.

Table 2: Relative Risk of Divorce and Gender-Composition Effects of Children

| | Austria | Belgium | Canada | Estonia | Finland | France | E-Germ. | W-Germ. | Hungary | Italy | Latvia | Lithu. | Norway | Poland | Slovenia | Spain | Sweden | Switz. | USA |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| *Gender pattern* | | | | | | | | | | | | | | | | | | | |
| No children | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 |
| One child | 0.69 ** | 0.45 *** | 0.78 | 0.56 ** | 0.57 *** | 0.88 | 1.16 | 0.87 | 0.38 *** | 0.40 *** | 0.53 *** | 0.59 ** | 0.58 ** | 0.43 *** | 0.47 * | 0.58 | 0.38 *** | 0.72 * | 0.72 *** |
| One child * one boy | 1.14 | 1.00 | 0.65 * | 1.22 | 0.94 | 0.95 | 0.93 | 0.86 | 1.03 | 0.67 | 0.89 | 0.79 | 1.20 | 1.01 | 0.71 | 1.01 | 1.00 | 0.76 | 0.88 |
| Two children | 0.39 *** | 0.36 *** | 0.51 *** | 0.30 *** | 0.39 *** | 0.65 ** | 0.91 | 0.48 ** | 0.19 *** | 0.11 *** | 0.23 *** | 0.18 *** | 0.43 *** | 0.19 *** | 0.19 *** | 0.19 *** | 0.15 *** | 0.27 *** | 0.51 *** |
| Two children * two girls | 0.83 | 1.31 | 0.88 | 1.29 | 0.96 | 1.01 | 0.68 | 1.72 | 0.97 | 1.69 | 0.84 | 1.46 | 0.85 | 0.84 | 0.59 | 1.10 | 1.04 | 1.24 | 1.05 |
| Two children * two boys | 1.22 | 1.09 | 0.87 | 0.89 | 1.00 | 0.97 | 1.35 | 1.21 | 0.82 | 1.72 | 1.04 | 1.20 | 0.85 | 0.69 | 1.02 | 1.10 | 1.39 | 1.10 | 0.95 |
| Three children | 0.39 *** | 0.31 *** | 0.35 *** | 0.14 *** | 0.25 *** | 0.62 ** | 0.93 | 0.21 ** | 0.21 *** | 0.08 *** | 0.21 *** | 0.36 *** | 0.28 *** | 0.11 *** | 0.20 ** | 0.28 *** | 0.19 *** | 0.18 *** | 0.58 *** |
| *Marriage cohorts* | | | | | | | | | | | | | | | | | | | |
| Cohort - 1969 | 0.58 *** | na | 1.00 | 0.92 | 0.76 | 0.52 *** | na | na | na | 0.43 * | 0.73 | 0.66 * | 0.72 | 1.10 | 0.65 | 0.73 | na | 0.84 | na |
| Cohort 1970 - 1974 | 0.81 | 0.87 | 0.97 | 0.80 | 0.70 * | 0.77 * | 0.56 ** | 1.14 | 1.11 | 0.60 * | 0.93 | 0.63 ** | 0.79 | 0.80 | 1.13 | 1.02 | 0.67 * | 0.75 | 0.95 |
| Cohort 1975 - 1979 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 |
| Cohort 1980 - 1984 | 1.13 | 1.26 | 1.58 ** | 1.01 | 1.61 ** | 1.99 *** | 0.98 | 1.52 * | 1.13 | 0.91 | 0.83 | 0.87 | 1.55 * | 1.35 | 1.06 | 1.17 | 1.03 | 1.10 | 1.06 |
| Cohort 1985 - 1989 | 1.73 *** | 1.48 | 1.91 *** | 0.86 | 2.01 * | 2.71 *** | 1.33 | 2.83 *** | 1.09 | 1.48 | 1.14 | 0.97 | 2.89 ** | 1.98 * | 1.02 | 1.66 | 1.35 | 1.11 | 1.27 ** |
| Cohort 1990 - | 2.57 *** | na | 2.68 *** | 0.34 | na | 6.48 *** | na | na | 1.95 | 1.54 | 1.97 *** | 1.37 | na | na | 0.88 | 3.18 ** | 1.06 | 1.22 | 1.80 *** |
| *Age at marriage* | | | | | | | | | | | | | | | | | | | |
| Age women | 0.89 *** | 0.87 *** | 0.94 *** | 0.88 *** | 0.88 *** | 0.95 *** | 0.94 | 0.94 * | 0.91 *** | 0.86 *** | 0.89 *** | 0.94 ** | 0.81 *** | 0.95 | 0.82 *** | 0.85 *** | 0.93 ** | 0.95 ** | 0.88 *** |
| *Highest level of education* | | | | | | | | | | | | | | | | | | | |
| Low | 1.18 | 1.25 | 1.18 | 1.19 | 1.07 | 0.85 | 1.39 | 1.07 | 1.06 | 0.59 ** | 1.32 | 1.76 ** | 0.87 | 0.68 * | 1.06 | 0.45 *** | 0.85 | 1.12 | 1.11 |
| Middle | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 |
| High | 0.93 | 0.93 | 0.99 | 0.92 | 1.34 | 1.08 | 1.07 | 1.23 | na | 2.06 ** | 1.04 | 1.05 | 1.25 | 0.76 | 2.27 ** | 1.07 | 0.95 | 1.28 | 0.82 ** |
| *Constants* | | | | | | | | | | | | | | | | | | | |
| a0 ×1000 | 0.77 *** | 0.70 *** | 0.22 *** | 4.15 *** | 1.15 *** | 0.25 *** | 0.28 *** | 0.28 *** | 0.93 *** | 0.65 *** | 2.49 *** | 0.46 *** | 5.82 *** | 0.12 *** | 3.75 *** | 0.94 *** | 0.95 *** | 0.24 *** | 3.50 *** |
| ? | 121.1 *** | 120.5 *** | 151.2 *** | 69.9 *** | 125.6 *** | 182.9 *** | 68.0 *** | 74.0 *** | 93.7 *** | 256.6 *** | 86.2 *** | 102.6 *** | 143.3 *** | 148.5 *** | 72.8 *** | 195.8 *** | 95.3 *** | 124.0 *** | 58.7 *** |
| -log Likelihood | 4040 | 1749 | 3727 | 2385 | 3767 | 4374 | 1836 | 1521 | 3185 | 1293 | 3802 | 2363 | 2376 | 1613 | 1004 | 1100 | 1476 | 3207 | 12251 |
| *N* | 2780 | 2281 | 1949 | 1289 | 2823 | 1593 | 1376 | 1219 | 2618 | 3031 | 1925 | 2102 | 2056 | 2926 | 1732 | 2542 | 1434 | 2858 | 5396 |
| % divorced | 19% | 10% | 26% | 26% | 17% | 40% | 18% | 17% | 16% | 5% | 28% | 15% | 15% | 6% | 7% | 5% | 14% | 15% | 33% |

Note: Reported are the a-parameters of the maximum likelihood-estimation of the sickle model, a is the risk of divorce relative to the reference group indicated by a 1. Parameters with (***,**,*) are significantly different from 1 at $p < .001$ resp. $p < .01$, $p < .05$. The dummy variable for children and their gender are time-dependent covariates. The parameters of the gender variables (One Child * One Boy, Two Children * Two Girls, and Two Children * Two Boys) report the interaction effects relative to One Girl or Mixed Gender Composition, respectively. *N* is the number of marriages included in the estimation.

Reading the gender effects (e.g., Austria): The a-effect of 0.69 means that families with one girl have a - 31% (i.e., 0.69 - 1) lower risk of divorce than families with no child. The a-effect of 1.14 means that families with one boy have a 14% (i.e., 1.14 - 1) higher divorce risk than families with one girl and a - 21% (i.e., 0.69 · 1.14 - 1) lower divorce risk than families with no child.