Working Paper Series 765
(ISSN 2788-0443)

The Impact of the Firstborn Gender on Family Formation and Dissolution: Evidence from Russia

Sergii Maksymovych

CERGE-EI
Prague, October 2023

ISBN 978-80-7343-572-1 (Univerzita Karlova, Centrum pro ekonomický výzkum a doktorské studium)
ISBN 978-80-7344-696-3 (Národohospodářský ústav AV ČR, v. v. i.)

The Impact of the Firstborn Gender on Family Formation and Dissolution: Evidence from Russia *

Sergii Maksymovych †

September 27, 2023

Abstract

In this paper, I obtain three findings regarding the impact of the first-born child's gender on family stability. First, couples who have a first-born daughter aged 6-18 are more likely to divorce than those who have a son of that age. Second, single mothers with first-born daughters are less likely to marry. Third, couples who have a first-born daughter aged 0-5 are less likely to divorce than those who have a son of that age. The first two findings are in accord with findings in the literature. The third finding is specific to the Russian context. My analysis is based on the Russia Longitudinal Monitoring Survey (RLMS-HSE) data for the 1994-2018 period. I estimate complementary log-log (cloglog) regressions of divorce and marriage (for single mothers) on firstborn gender, age, and a set of household socio-demographic characteristics. My findings support the conclusion that the impact of children's gender on family living arrangements depends on family socioeconomic conditions and thus has a different character and magnitude in different contexts.

JEL codes: J13, J15, J16
Keywords: marriage, divorce, children, gender, living arrangements

*I am grateful to Štěpán Jurajda for insightful and essential comments and suggestions. Also, my thanks go to Jan Kabátek, Nikolas Mittag, Ján Palguta, and Mariola Pytliková for providing feedback on my manuscript. Further, I appreciate useful remarks by committee members on several DP workshops, in particular, Randall Filer, Daniel Münich, Jan Švejnar, Jan Hanousek, Gérard Roland, Krešimir Žigić, Filip Matějka, Christian Ochsner, Jan Zápal, Vasily Korovkin, and Sebastian Ottinger. Moreover, I owe a debt of gratitude to participants of the 12th CES conference for generous and helpful recommendations and, especially, to Sofia Trommlerová. Last but not least, I want to thank Andrea Downing and Grayson Krueger for proofreading and valuable advice.

†CERGE-EI, a joint workplace of Charles University in Prague and the Economics Institute of the Academy of Sciences of the Czech Republic, Politickych veznu 7, 111 21 Prague, Czech Republic. Address: CERGE-EI, P.O. Box 882, Politickych veznu 7, 111 21 Prague 1, Czech Republic. Email: Sergii.Maksymovych@cerge-ei.cz.

1 Introduction

Historically in many societies around the globe, a child's gender affects the level of education he or she is likely to receive, the occupation he or she will choose, and the wages he or she will be paid (e.g., Blau (1997), Exley and Kessler (2022), Blau and Kahn (2017)[1]). A growing body of research examines how a child's gender may be associated with differential treatment by parents from birth, which could contribute to gendered differences in adult market outcomes (e.g., Lundberg (2005) among others). One strand of this literature has found associations between child gender and parental marriage and separation, with implications for the living arrangements of children. Fathers are more likely to be present in the home if a child is male (Dahl and Moretti, 2008); the presence of sons decreases the probability of divorce (Mott, 1994; Katzev et al., 1994); and a birth outside marriage is more likely to be followed by marriage if the child is a son (Lundberg and Rose, 2003). These facts may have serious consequences for the well-being of children, because family structure is an important predictor of children's later life outcomes. Research on children's well-being broadly supports the idea that children who grow up with only one parent, most often the mother, fare worse than those who grow up with two married biological parents (McLanahan et al., 2005).

Recent research has moved beyond documenting the associations and establishing causality between children's gender and family living arrangements, and has aimed to find causes behind the patterns observed (Dahl and Moretti, 2008; Kabátek and Ribar, 2020). Many authors have proposed that parental, and especially fathers' preference for sons may offer an explanation. This explanation has been supported by systematic evidence from several US survey data sets by Dahl and Moretti (2008).

[1]The lists of references in these papers contain studies investigating various aspects of the subject. The citation list accompanying Blau and Kahn (2017) contains almost seven hundred related papers published over the previous five years.

However, a more recent study by Kabátek and Ribar (2020) concluded that strained relationships in families with teenage daughters likely lie behind the higher divorce rate of couples who have daughters. This explanation is compatible with the finding of Kabátek and Ribar (2020) that single Dutch mothers of firstborn girls are just as likely to marry as single Dutch mothers of firstborn boys and they take about the same amount of time to do so. However, it would not have accounted for the finding of Dahl and Moretti (2008) that single mothers of first-born daughters in US were less likely to marry.

My paper, like the two aforementioned studies, aims to simultaneously estimate the impact of the firstborn gender on family formation and dissolution, conditional on firstborn age.[2] In contrast with Dahl and Moretti (2008), who also look into family formation and dissolution in their paper, I consider an event-history model that not only allows the risks of separation to change with children's ages, but also incorporates the length of cohabitation of spouses and the right-censoring in marriage spells. [3] Compared to Kabátek and Ribar (2020), my work employs data on multiple socioeconomic characteristics of households, which allows me to check whether the effect in question is confined to a particular group of the population. Moreover, I look at actual cohabitation and not only at the officially registered relationships, as Kabátek and Ribar (2020) do.

My study examines the impact of the firstborn child's gender on living arrangements in the Russian setting. The case of Russia deserves particular attention because it has been one of several countries with the highest reported divorce rates for decades.

I focus on children of different ages, more specifically on pre-school children (0-5 years) and school-age children (6-18). I group children's ages this way because I

[2] Plausibility of the assumption of the firstborn gender randomness is discussed in the Method section.

[3] Moreover, I use data from one longitudinal survey while Dahl and Moretti (2008) use pooled data from US Current Population Surveys (CPS) over a period 1960-2000.

expect differential effects of preschool children. First, events before five years old can have large long term impacts on adult outcomes (see e.g., Currie and Almond (2011)). Second, having daughters of this age might have an impact on such personal traits of fathers as neuroticism and extraversion (van Lent, 2020), which in turn are related to a higher chance of divorce (Diederik and Mortelmans, 2018; Fani and Kheirabadi, 2011; Zare et al., 2013). Third, single mothers in my data set most often have young children.

Each of the two proposed mechanisms - the son bias and tensions between teenage daughters and their fathers - implies different predictions for my results. If the overarching son preference holds, lower marriage rates of single mothers of daughters should hold simultaneously with higher separation rates of married mothers of daughters. If the explanation through strained relations with teenage children holds, there should be no effect for either marriage or separation. [4] My results show that the effect on probability of separation [5] is negative, and the effect on probability of partnership formation is also negative. Additionally, I find an indication that having teenage daughters increases the probability of separation in line with Kabátek and Ribar (2020).[6] My results, based on the Russia Longitudinal Monitoring Survey (RLMS-HSE) data, suggest that while son preference might play a notable role in marriage decisions, it is outweighed by other determinants of the family process, leading to a negative observed effect of preschool firstborn daughters on separation. Investigation of the precise nature of these determinants is currently beyond the scope of this paper. [7]

[4] Kabátek and Ribar (2020) argue that parents do not foresee difficulties in relationship with teenage daughters when daughters are of a younger age.

[5] When I use the word "divorce" I actually mean divorce together with separation rather than the legal divorce on its own. I will discuss this aspect in more detail in Section 3.3.

[6] This result is less warranted because I cannot observe many teenage children because of the attrition of households over survey waves

[7] I suggest six possible causes for the negative effect of first-born daughters on separation in Russia. First, parents of spouses might be more supportive of a marriage in which daughters are born. Second, losses of marriage surplus due to separation are higher when children are female. Third, higher marriage rates among single mothers of first-born sons reduces the estimated cost

My results contribute to the literature by adding a novel fact about the negative relation between the presence of preschool daughters and divorce, and also by supporting already established patterns mentioned above in the Russian context. Overall, gender-related attitudes and practices are highly culturally dependent. Hence, it is important to examine the same research question in different cultural contexts. My paper, in which I replicate some of the results for other countries (teenage daughters) but some of my results are new (young daughters), confirms the need to examine this topic in different countries. Moreover, better understanding of how the gender of children influences living arrangements could be of use to policy makers. Measuring the magnitude and character of the impact of first-born child's gender on family living arrangement potentially could become standardized and make its way into routine practice of international organizations, as is already the case for the gender preference measures.

2 Data

RLMS-HSE data

The source of data for my analysis is the Russia Longitudinal Monitoring Survey, RLMS-HSE, for the 1994-2018 period. The data set covers 23 yearly rounds of a national representative survey on the social, health and economic situation in Russia. Two years are missing, 1997 and 1999. The survey is scheduled annually during fall and winter months exact dates varying from year to year (i.e., one household could be surveyed twice during the same calendar year). The RLMS-HSE is conducted

of separation for them as they perceive their remarriage prospects to be more favorable. Fourth, mothers of younger sons should be able to remarry faster than mothers of older sons because sons, compared to daughters, are more likely to come to terms with stepfathers when they are younger. Fifth, public policy is more oriented towards women, because women constitute the majority of voters demographically and vote more actively. Sixth, gender of children induces changes in parents' (especially fathers') personality, which in turn influence marriage stability.

by several organizations including the National Research University Higher School of Economics and the Carolina Population Center, University of North Carolina at Chapel Hill. [8]

The RLMS is a survey representative at the national level. The sampling was designed to obtain a replicated three-stage, stratified cluster sample of residential addresses excluding military, penal, and other institutionalized populations. Households participating in the survey were selected through a multi-stage probability sampling procedure in order to guarantee cross-sectional national representativeness. Within each selected primary sample unit, the population was stratified into urban and rural substrata in order to guarantee the representativeness of the sample in both areas. The data covers approximately 5000 households (dwelling units), with 12,000 adults and 2000 children per wave. [9]

The RLMS-HSE was established to create a nationally representative survey to monitor the economic and health impact of the reforms in the Russian Federation. Throughout the entire set of surveys, detailed basic household and individual data have been collected. The major data components are: economic (detailed income, assets, expenditures and labor force behavior data, including type of employment, earnings, hours and ownership form, i.e., public, private or joint), demographic/sociological (household structure and age-gender composition, background, education and school behavior); and health (24-hour dietary recall, nutrient intake levels, smoking, drinking activity, BMI direct measurement). All in all, there are more than 3,000 variables. [10]

[8] These are two organizations which provide access to the data. A more comprehensive list of people and agencies involved in conducting the survey is available at this link: https://rlms-hse.cpc.unc.edu/team/.

[9] The target sample size was set at 4 000 households (Kozyreva et al., 2016b). Details of the sampling design, including specification of primary and secondary statistical units along with targeted sample sizes for households and individuals, can be found, e.g., in Kozyreva et al. (2016a) and also here: https://www.hse.ru/rlms/sample (in Russian).

[10] In many aspects the design of the RLMS-HSE, which was established in 1992, mirrored the design of the China Health and Nutrition Survey (CHNS) (Kozyreva et al., 2016a) initiated in 1989.

The RLMS-HSE is a panel survey with a longitudinal component. A point of concern is that of attrition in the panel. [11] Some households are inevitably lost from the panel as a result of moving house, splitting up, or other common causes of attrition. The size of attrition across years is reported in, e.g., Kozyreva et al. (2016a), and Heeringa (1997), along with reports by organizations that administer the data. The rate of household attrition is gradually decreasing as households are observed over consecutive survey waves, being 13% in the second wave, 5% in the fourth wave, and about 2% in the twentieth wave. The rate of individual attrition is a little higher. All in all, for the first 18 rounds (1994-2014), only about 29% of households and 19% of individuals continued to participate but, for the first 9 rounds (1994-2004), the results were about 60% and 51%. From 1996, the RLMS-HSE followed households in the panel even if they moved away from the sample address or split into several households, each of which is inducted in the panel. However, households that moved out of primary sampling units (the entire country is divided into 35 primary sampling units) were not tracked in subsequent surveys. Heeringa (1997) finds that attrition does not notably change the distribution of demographic characteristics of households (number of children, total number and employment status of members). Still, households that move out of their original residences or decline to participate tend to have higher median incomes and expenditures. Gerry and Papadopoulos (2015) investigate patterns of the RLMS attrition and how it is related to demographic, health and other socio-economic characteristics of participants. The authors confirm the presence of non-random attrition for the RLMS. At the same time, they also conclude that the non-random attrition does not significantly distort estimates of statistical models. [12]

[11] Researchers who are not interested in exploiting the longitudinal element of the data, can still use the univariate statistics from individual cross-sections, which are unbiased because of the annual replenishing undertaken to restore representativeness.

[12] Specifically, Gerry and Papadopoulos (2015) consider a case of the dynamic Probit model. The methods applied to estimation and analysis of the Probit model are also applied to the cloglog model. Thus, their conclusions should also hold for the cloglog model.

The household response rate was about 40% during 2006-2013. It increased to 60% in 2014, when the target sample size was reduced from 6,000 to 4,800 households. Since then, it gradually decreased to 56% in 2019. In urban areas the response rate is lower. The individual response rate, conditional on a household responding, has constantly been around 96-98%. The imbalances caused by differences in response rates across regions and socio-demographic strata of the population were corrected for by the survey design so that the actual proportion of completed household interviews compares well to the proportion of the population in each stratum. All in all, the longitudinal sample consists of 16,789 households, of which 73 percent were observed for at least 2 consecutive years, and 25 percent for at least 7 consecutive years

Selected variables and descriptive statistics

I identify 1,788 firstborn children whose mothers participated in the RLMS-HSE survey in the year of their birth, i.e. before they turned 1. Of these 1,788 firstborn children, 1,431 were born to partnered women (either in a married couple or in a consensual union) [13] and 1,367 were observed in more than one survey wave.[14] Therefore, the firstborns from the analyzed sample are born to already partnered couples and their gender cannot influence the matching or selection of their parents into partnership. Correspondingly, 357 firstborns were born to single women, [15] of whom 255 were observed in more than one survey wave (of them 73 report being married in the individual file, 102 are observed starting partnership of which 55

[13]Of these partnerships, 1,231 have a known start date. Estimation of models in which the time under risk starts from the year of partnership formation rather than from the year of firstborn arrival yields estimates close to my reported results. I use terms "family formation" and "partnership formation" interchangeably.

[14]These women are partnered according to the household questionnaire responses. According to the individual questionnaire responses, their marital statuses are: never married (9), in marriage (1,155), cohabiting but not married (200), divorced and not married (2), no answer (1). Thus, 11 responses from the individual file are in conflict with responses from the household file.

[15]Of these, 230 firstborns were born to women who had never been married before. The results of the estimations run on the sub-sample of women who never married before are in line with the results for the whole sample.

start marriages and 47 start cohabiting). Therefore, I have two main samples for analysis: a sample of 1,367 firstborns with two parents present and a sample of 255 firstborns born to single mothers.

The variables used for analysis are described in Table 3.B.1 in Appendix 3.B. Descriptive statistics are presented in Table 1. Differences in means of selected characteristics between households with first-born sons and first-born daughters are not statistically significant at the 10% level in most cases.

Table 1: Average Characteristics of Couples with Firstborn Sons and Daughters

	Sons	Daughters	Diff	p-val.
Mother's age at birth of the first-born	24.21	24.39	-0.18	0.45
Father's age at birth of the first-born	26.91	27.04	-0.13	0.63
Father's employment status	88.45%	88.44%	0.01%	0.99
Father is Orthodox	47.08%	50.30%	1.73%	0.52
Mother is Orthodox	49.36%	48.42%	-0.94%	0.73
Father is Muslim	3.71%	2.40%	1.31%	0.16
Mother is Muslim	3.71%	1.50%	2.21%	0.01
Mother reports other religious affiliation	1.4%	1%	0.4%	0.36
Mother reports no religious affiliation	4%	3%	1%	0.22
Urban area	74.32%	76.73%	-2.40%	0.30

Father is Russian	89.24%	87.98%	1.27%	0.46
Mother is Russian	91.61%	89.26%	2.35%	0.14
Father has vocational or tertiary education	52.00%	48.05%	3.95%	0.14
Mother has vocational or tertiary education	62.34%	61.26%	1.08%	0.68
Number of family members	3.96	4.01	-0.05	0.54
Mother reporting satisfactory life	64.25%	62.44%	1.82%	0.49
Father reporting satisfactory life	60.54%	58.13%	2.41%	0.37
Reporting satisfactory job	33.5%	33.5%	0.04%	0.99
Having subordinates at work	9.1%	10.5%	-1.4%	0.39
Concerned about losing job	26.7%	28.4%	-1.7%	0.48
Well-being improved in the last year	22.7%	22.4%	0.3%	0.89
Expecting economic improvement in 12 months	40.4%	41.9%	-1.5%	0.57
Relative economic standing	15.3%	14%	1.3%	0.50
Feeling empowered	13.7%	12.3%	1.4%	0.45
Feeling respected	61.9%	62.8%	-0.9%	0.75
Satisfactory material condition	23.3%	20.1%	3.1%	0.16
Understanding between generations is possible	28.8%	28.1%	0.7%	0.76
Knowing foreign language	22.3%	23.3%	-1.0%	0.65
Having disability	0.6%	0.5%	0.1%	0.76

Smoking now	12.3%	11.0%	1.7%	0.32
Ever having smoked	18.8%	20.7%	-1.9%	0.38
Drinking alcohol recently	25.4%	27.2%	-1.8%	0.45
Drinking alcohol sometimes	35.7%	37.5%	-1.9%	0.47
Ability to have three meals daily	33.0%	34.5%	-1.6%	0.54
[a] Living in regional center	47.6%	44.4%	3.2%	0.24
Family size	3.6	3.7	-0.1	0.32
Owning accommodation	74.8%	77.3%	%	
Number of rooms	2.13	2.10	0.03	0.48
Area of accommodation (m^2)	33.31	33.71	-0.40	0.67
Running water	89.4%	88.4%	1%	0.56
Owning country house	14.6%	17.4%	-2.8%	0.17
Saving money recently	17.6%	13.1%	4.5%	0.02
Receiving economic support	49.7%	47.1%	2.6%	0.33
Having credit debt	29.1%	28.0%	1%	0.67
Owing money to individuals	4.3%	6.1%	-1.8%	0.13

Notes. The results are based on 1,367 observations, 701 boys and 666 girls.

[a] The following variables are taken from the household file and descriptive statistics based on 1,354 observations, 698 boys and 656 girls.

The only exception is that the mothers of first-born sons more frequently report being Muslim. This might mean that they are more likely to follow prescriptions of tradition in family life and have stronger reservations about divorce. However, not including firstborns with mothers who report as Muslim does not have a notable impact on estimates.

Descriptive statistics for single mothers are presented in Table 2. There are three main differences between single mothers of first-born sons and daughters. First, single mothers of first-born sons tend to be about one year younger than those of first-born daughters. This is compatible with my finding that single mothers of first-born sons form partnerships faster.[16] Second, mothers of sons appear to have lower educational attainment that mothers of daughters. This could be partially explained by their younger average age. Another possible reason is a lower response rate from mothers of first-born daughters, which could be even lower for those with lower educational attainment.

Table 2: Average Characteristics of Single Mothers with Firstborn Sons and Daughters

	Sons	Daughters	Diff	p-val.
Mother's age at birth of the first-born	24.22	25.20	-0.98	0.07
Mother is Orthodox	47.09%	49.04%	-0.20%	0.72
Mother is Muslim	1.59%	2.55%	-0.96%	0.53
Urban area	68.25%	70.06%	-1.81%	0.72
Mother is Russian	90.11%	92.05%	-1.94%	0.54
Mother has vocational or tertiary education	47.01%	61.78%	-14.69%	0.01

[16]More precisely, this would be the case for children born out of wedlock on January 1, the boy's mother is already partnered by the time of the survey (October-December), while the girl's mother is still single. This statement could be supported by the results of a back-of-an-envelope numerical simulation with the probability of partnership formation following a Bernoulli distribution. The difference in the age of mothers appear to be noticeable even during the first year after a firstborn birth (the period for which the numbers in Table 2 are calculated). This result is even stronger when younger mothers are more likely to marry (which again makes sense because the observed mean age difference accumulates only during the first year after a birth).

Number of family members	4.28	4.24	0.04	0.81
Satisfactory life	40%	49%	-9%	0.08

Notes. The calculations are based on 346 observations, 189 boys and 157 girls.

Such an explanation is consistent with a higher proportion of mothers of first born sons in the sample (1.2) than the sex-ratio at birth in the population. It is also consistent with first-born daughters' mothers more frequently dropping out of the survey after separation, which I observe in the data and which implies that single mothers of daughters are less willing to participate in the survey. Moreover, the fact that more educated mothers are more likely to participate in the survey is in line with their reported higher life satisfaction. Further, single mothers are less likely to live in urban areas and have a less satisfactory life than partnered mothers.

The numbers of first-born boys and girls of different ages living with partnered and single mothers can be seen in Table 3. The share of boys among children living with single mothers is higher than for children living with partnered mothers when the children are younger. This could be explained by two simultaneous tendencies. First, married mothers of first-born sons are more likely to separate. Second, mothers of daughters more frequently drop out of the survey after separation, [17] while single mothers of sons more frequently drop out of the survey after partnership formation.[18] The latter fact could explain why the share of first-born boys remaining in the survey increases over time.

[17] This could happen because mothers with daughters tend to separate when children are older, as the results in Table 3.4 show. Moving to a different location is easier with older children.

[18] The possible sex-specific attrition that I mention occurs after partnership formation for single mothers or after separation for partnered mothers. Therefore, it does not have an effect on the results of the baseline statistical analysis in my paper because the latter is based on observations of either single mothers before partnership formation or partnered mothers before separation. I do not find evidence that sex-selective attrition occurs during partnership spells or during single-motherhood spells.

Table 3: Numbers and shares of first-born children living with partnered and single mothers by age and gender

	1 wave before birth[a]		The year of birth		1-year-olds	
	P-ed	S-le	P-ed	S-le	P-ed	S-le
All firstborns	774	245	1,367	255	1,085	177
Boys	390	131	701	139	559	98
Girls	384	114	666	116	526	79
Share of boys	0.504	0.535	0.512	0.545	0.515	0.554

2-year-olds		3-year-olds		4-year-olds*		5-year-olds	
P-ed	S-le	P-ed	S-le	P-ed	S-le	P-ed	S-le
944	141	792	107	659	78	573	69
493	80	414	63	330	47	298	40
451	61	378	44	329	31	275	29
0.522	0.567	0.523	0.589	0.501	0.603	0.520	0.580

6-year-olds		7-year-olds		8-year-olds		9-year-olds	
P-ed	S-le	P-ed	S-le	P-ed	S-le	P-ed	S-le
495	55	442	41	361	35	309	29
250	32	221	22	189	17	161	15
245	23	221	19	172	18	148	14
0.505	0.582	0.500	0.537	0.524	0.486	0.521	0.517

10-year-olds		11-year-olds		12-year-olds		13-year-olds	
P-ed	S-le	P-ed	S-le	P-ed	S-le	P-ed	S-le
249	21	221	17	189	16	152	13
128	11	118	10	97	10	80	9
121	10	103	7	92	6	72	4
0.514	0.524	0.534	0.588	0.513	0.625	0.526	0.692

[a] These numbers include firstborn's mothers who were observed one wave before the firstborn birth out of 1,367 referred to in the analysis of marriage dissolution.

* The difference is statistically significant at a 0.1 level.

To take into account the last-mentioned fact, I calculate the differences between shares of boys living with single and partnered mothers from Table 3 for firstborns aged 0-10 and show them on Figure 1.

The differences between shares of sons living with single and partnered mothers tend first to increase and then to decrease. This tendency is compatible with more

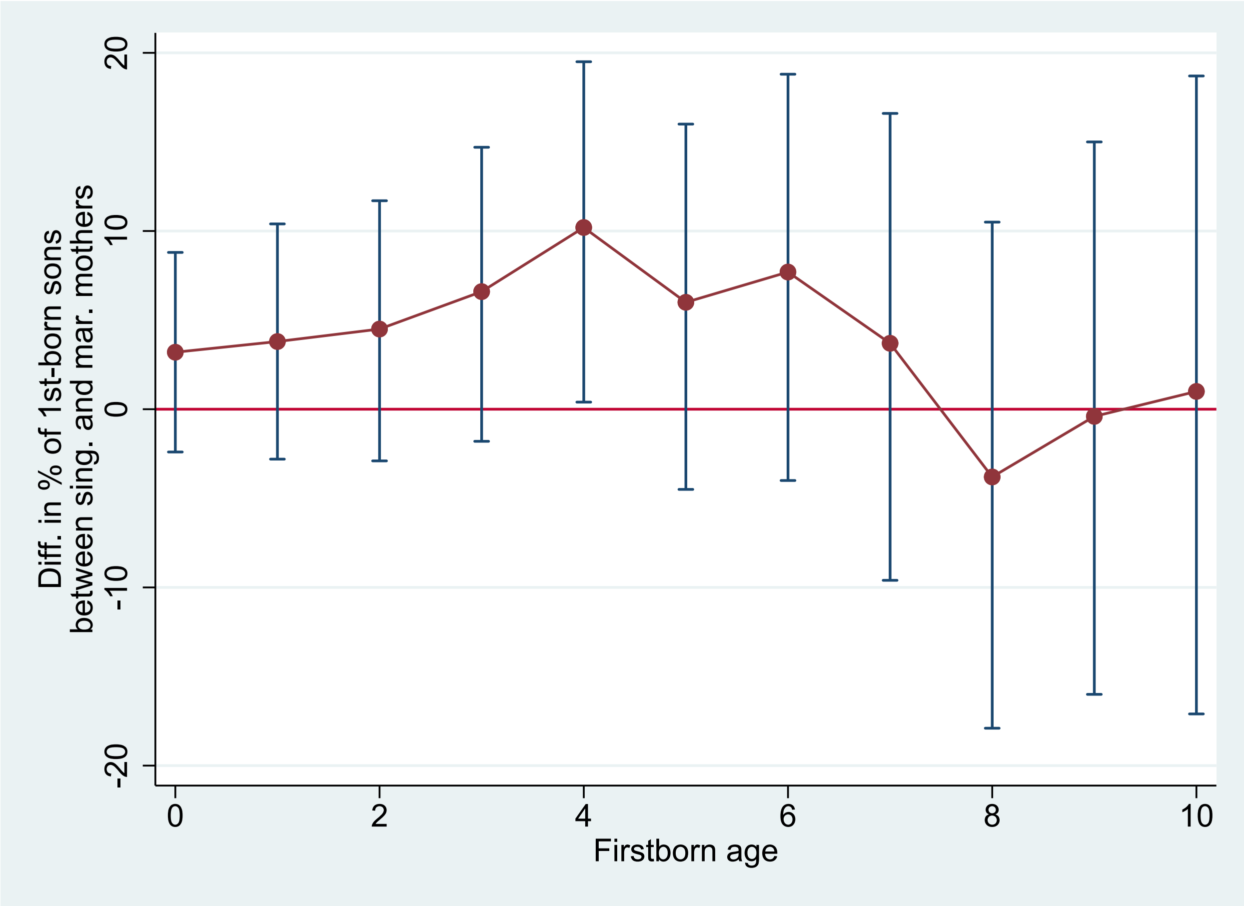

Figure 1: Difference between percentages of firstborn sons among single and partnered mothers conditional on firstborn age

Note: Capped spikes show 90% confidence intervals.

separations among mothers of younger first-born sons and fewer separations among mothers of older first-born sons.

3 Method

Following (Kabátek and Ribar, 2020), I estimate a complementary log-log (cloglog) discrete-time hazard model of family separation and family formation. This model has a number of advantages. First, the results are straightforward to interpret: exponentiated estimated coefficients can be interpreted as approximate odds ratios. Second, it is widely used in the literature, being a discrete analog of the continuous proportional hazards model. Third, the underlying link function more closely approximates the distribution of observed partnership durations (left-tailed) than alternative link functions (logistic, Gaussian, etc.) (Simonoff, 2003, p. 396). In other words, the cloglog model is suitable when one of the outcomes is rare relative to the other. This applies to the partnership duration data, in which the separation is relatively rare and hence most partnership spells are long which leads to a left-tailed (or left-skewed) distribution of baseline separation hazards.[19] Fourth, the assumption of proportional hazards is intuitively plausible in the current setting. The Cox PH model assumes that predictors act multiplicatively on the hazard function. The baseline hazard is common to all units in the population; individual hazard functions differ proportionately depending on values of observed covariates (see, e.g., (Hurrell, 2015, p. 475) or (Wooldridge, 2002, p. 690)).

The functional form of the cloglog model is:

$$\Pr[y_{it} = 1|\mathbf{x}_i] = 1 - \exp(-\exp(\mathbf{x}_i'\boldsymbol{\beta})), \tag{1a}$$

where the hazard probability of y_i, a separation for a couple i in year t, is defined as a function of covariates $\mathbf{x}_i$ that are specific to the given couple (not time variant).

The corresponding specification of the predictor (also called *index*) $\mathbf{x}_i'\boldsymbol{\beta}$ in the model of separation with conditioning on firstborn age is:

[19] The cloglog model is a discrete time analog of the Cox proportional hazards (PH) model. In the Cox PH model the exact form of the baseline hazard function is not of interest. Still, when the assumed properties of the baseline hazard function (the negative skew) mirror the actual ones, the precision of the estimates is higher.

$$\mathbf{x}_i^{'}\boldsymbol{\beta} = \beta_0 + \mathbf{1}(FB\ age\_range_{it} = 0-5)(\beta_{0_{0-5}} + \beta_{1_{0-5}} \cdot FB\ daughter_i)$$
$$+\mathbf{1}(FB\ age\_range_{it} = 6-18)(\beta_{0_{6-18}} + \beta_{1_{6-18}} \cdot FB\ daughter_i)$$
$$+\sum_{k=1}^{18} \beta_{3j} \cdot \mathbf{1}(FB\ years\_obs_{it} = j) + \mathbf{z}_i^{'}\boldsymbol{\beta}_4 \qquad \text{(1b)}$$

where the base category is all firstborns older than 18. In this case, the exponentiated coefficients on the firstborn age dummies show which factor the separation hazard increases by over the separation hazard in families with firstborns older than 18 in any given year. In this regression specification the constant terms is suppressed. I also present an estimate without age ranges that includes only the first-born daughter dummy. Then, I present two sets of estimates similar to (1.b) each with a dummy included for only one age range, with the other age range being a base category. The vector $\mathbf{z}$ includes employment status, age, religious affiliation, living in an urban area, being of Russian ethnicity, and educational accomplishment of both spouses (see Table 3.1).

The functional form for the model of family formation is the same as for the model of separation:

$$\Pr[y_{it} = 1|\mathbf{x}_i] = 1 - \exp(-\exp(\mathbf{l}_i^{'}\boldsymbol{\alpha})) \qquad \text{(2a)}$$

The specification of the predictor in equation (2.a), which I now denote $\mathbf{l}_i^{'}\boldsymbol{\alpha}$, with conditioning on firstborn age is:

$$\mathbf{l}_i^{'}\boldsymbol{\alpha} = \alpha_0 + \mathbf{1}(FB\ age\_range_{it} = 0-5)(\alpha_{0_{0-5}} + \alpha_{1_{0-5}} \cdot FB\ daughter_i) + \sum_{k=1}^{18} \alpha_{3j} \cdot \mathbf{1}(FB\ years\_obs_{it} = j) + \mathbf{w}_i^{'}\boldsymbol{\alpha}_4 \quad (2b)$$

I also estimate specifications analogous to those of the model of separation. The vector $\mathbf{w}$ includes employment status, age, religious affiliation, living in an urban area, being of Russian ethnicity, and educational accomplishment for single mothers. Vectors of parameters $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$ are estimated by maximum likelihood. Selection of covariates follows previous studies, but also takes into account the numbers of missing observations and results of likelihood ratio tests. [20]

For the model of couple's separation, I test two hypotheses: a) first-born teen- and school age daughters cause a different parental separation rate than their peer first-born sons, i.e. $H_0 : \beta_{1_{6-18}} = 0$ versus $H_A : \beta_{1_{6-18}} \neq 0$; b), and first-born daughters of preschool age (0-5 years old) cause a different parental separation rate than their peer first-born sons, i.e. $H_0 : \beta_{1_{0-5}} = 0$ versus $H_A : \beta_{1_{0-5}} \neq 0$. In other words, the first hypothesis states that parents who have a first-born daughter aged 6-18, are either more or less likely to break up their union in a particular year than parents with otherwise similar characteristics other than having a first-born son aged 6-18. In the same way, the second hypothesis states that parents who have a first-born daughter aged 0-5, are either more or less likely to break up their union in a particular year than parents with otherwise similar characteristics who have a first-born son in that age range. For the family formation model, I test $H_0 : \alpha_{1_{0-5}} = 0$ versus $H_A : \alpha_{1_{0-5}} \neq 0$. That is, single mothers with preschool (0-5 years old) first-born daughters are either more or less likely to form a union than single mothers with first-born sons aged 0-5 who have otherwise similar characteristics. I expect,

[20]More details on the estimation procedure are included, e.g., in the Online Appendix Section of Kabátek and Ribar (2020)

in line with previous studies, the coefficient $\beta_{1_{6-18}}$ to have a positive value and the coefficient $\alpha_{1_{0-5}}$ to have a negative value.

For identification, I rely upon the assumption that the firstborn's gender is random. This assumption would be violated if there were sex-selective abortions. At the level of the entire sample this assumption appears to be warranted because the sex ratio does not notably differ from that in the population, and the average age of women who give birth to their first child does not notably differ by the gender of the first child. This fact, however, does not rule out a possibility that sex-selective abortions could be biased either towards sons or towards daughters in different subgroups of the population. In this case, the effects of sex-selective abortions across different subgroups could cancel out, so the sex-ratio at birth at the level of the entire population would be close to the natural one. To the best of my knowledge, no evidence, however, has been reported on sex selective abortions biased towards different sexes and confined to particular subgroups of the population in Russia. [21]

While there are few reasons for concern about reverse causality and unobserved heterogeneity, there are several issues regarding the empirical framework that should be pointed out. First is the measurement error in age ranges of firstborns. Specifically, I measure firstborn ages as differences between the year of observation and the year of birth. Therefore, when two consecutive survey waves are less then one year apart, some children have the same age at these two consecutive waves. In such cases, I add one year to their age in the second wave. [22] Second is construction of

[21] As for specific groups of population in other countries, some studies indicate this assumption might not hold. These studies, however, have not been frequently replicated so far, the effect they found is small, and it is not clear how characteristics causing a shift in sex ratio of children are related to marriage stability. More detailed discussion of the firstborn gender randomness assumption plausibility is presented in Abramishvili et al. (2019). Kim and Shapiro (2021) explicitly deny the presence of sex-selective abortion in Russia as a whole at a statistically noticeable scale (recorded online presentation of their paper can be accessed at this URL: `https://youtu.be/f1_qHepWozU`).

[22] This happens only for children born between September and December, i.e. during the period of the year when the interviewing takes place. Thus, my measure of children's age might overstate the actual age by up to four months.

the dependent variable, the indicator of family dissolution. [23] While Kabátek and Ribar (2020) focus on formal marriage status,[24] I also take into account information about the actual cohabitation of spouses. I do this for two reasons. First, the related literature tends to focus on the actual absence of fathers rather than on official marriage status, as in the paper by Dahl and Moretti (2008). Second, women who appear to single according to the individual-level data quite often have a husband according to the household-level data. That is why my dependent variable takes a value of 1 when a couple stops cohabiting according to the household-level data file and if this couple is not officially married in the individual-level data (as reported by a wife).[25] In other words, in the basic family dissolution model I consider those women who appear to be not married based on individual data and separated based on household-level data as actually separated. Women who fulfill only one of these conditions or none, are regarded as being in partnership. At the same time, my marriage formation indicator takes a value of 1 when a couple starts cohabiting according to the household-level data file. Finally, the character of association between covariates and the dependant variable, along with estimation results with alternative errors specifications suggest that concerns about non-monotonicity and heteroskedasticity are not justified.[26]

[23] I use terms "family dissolution" and "separation" interchangeably.

[24] This is the only information on family living arrangements contained in their administrative data set.

[25] Other living arrangements recorded in the individual-level data file, besides being officially married, are: never having been married, cohabiting but not officially registered, divorced, widowed, and officially married but not cohabiting. I use alternative measures for the family dissolution for robustness checks. Table 3.C.1 in Appendix 3.C shows results for the family dissolution model with the dependent variable being cohabitation termination according to the household data file.

[26] Parametric methods (e.g., Probit and Logit) assume strict monotonicity and homoscedasticity (Jurajda, 2021).

4 Results and discussion

4.1 Results for partnership dissolution

Estimates of the model (1.a) with eight specifications of predictor (1.b) are presented in Table 3.4. [27] The presence of older first-born daughters predicts a substantially higher likelihood of divorce than the presence of sons in the same age range. This finding is in line with the result in Kabátek and Ribar (2020), but the effect size is much higher than they find. Nevertheless, their estimated values lie within the 95% confidence interval of my estimated effect. [28] Moreover, it is necessary to take into account the fact that the divorce rate in Russia has been high on the global scale. Specifically, if similar mechanisms to those underlying the results of earlier research are at work in the Russian setting, my estimate of the impact of firstborn's gender on marriage dissolution will be higher than in previous studies.

The estimated effect of young (0-5 years old) first-born daughters on the probability of marriage dissolution is negative. It, however, is not statistically significant. Still, when sons aged 6-18 years become a reference category, the impact of first-born daughters aged 0-5 becomes statistically significant as can bee seen from columns (1) and (2) of Table 5.

Table 5 also contains results for separation regressions with interaction terms

[27]Results for cohabitation termination as a dependent variable are presented in Table 3.C.1. Specifically, the dependent variable takes a value of 1 when cohabitation is terminated according to the household file without conditioning on family status in the individual data file.

[28]Using a relatively low number of observations could explain the high standard errors of my estimates. Moreover, the age range of 6-18 includes not only teens (for whom Kabátek and Ribar (2020) observe higher hazard of divorce of spouses with first-born daughters), but also school-age children aged 6-12. When the age range of 13-18 instead of 6-18 is included in the model, the estimates do not differ notably, but the null hypothesis cannot be rejected. Overall, there are 418 firstborns in the age range of 6-12 observed on average for 4.37 years (1,826 family-years observations in total) and 119 firstborns in the age range of 13-18 observed on average for 3.13 years (372 family-years observations in total). The numbers 418 and 119 include only firstborns who remain in families in which they were born. These numbers are lower than corresponding numbers in Table 3, which include all firstborns that remain in the survey.

Table 4: The impact of the firstborn gender on family dissolution from complementary log-log model estimation.

Explanatory var-s:	Separation and marriage termination (1)	Separation and marriage termination (2)	Separation and marriage termination (3)	Separation and marriage termination (4)	Separation and marriage termination (5)	Separation and marriage termination (6)	Separation and marriage termination (7)	Separation and marriage termination (8)
Firstborn is daughter 0-18 years	1.03	0.87						
	(0.18)	(0.17)						
Firstborn 0-5 years old is daughter			0.93	0.86	0.92	0.82		
			(0.19)	(0.18)	(0.19)	(0.17)		
Firstborn 6-18 years old is daughter			2.32	2.00			2.29	2.03
			(0.99)**	(0.86)*			(0.97)**	(0.85)*
Firstborn 0-5 years old			0.019	0.012	1.36	1.43		
			(0.005)***	(0.006)***	(0.46)	(0.49)		
Firstborn 6-18 years old			0.011	0.008			0.48	0.54
			(0.006)***	(0.005)***			(0.21)*	(0.24)
Year	1.05	1.09	1.09	1.10	1.07	1.12	1.07	1.08
	(0.08)	(0.08)	(0.06)*	(0.09)	(0.08)	(0.09)	(0.08)	(0.09)
Year^2	0.99	0.99	0.99	0.98	0.99	0.99	0.99	0.99
	(0.005)	(0.004)*	(0.004)*	(0.004)**	(0.005)*	(0.005)*	(0.005)*	(0.005)
N of marriage spells	1,367	1,367	1,367	1,367	1,367	1,367	1,367	1,367
N of marriage-year observations	7,163	7,064	7,163	7,064	7,163	7,064	7,163	7,087
Log-likelihood	-633.15	-575.45	-633.75	-585.09	-632.74	-575.25	-630.79	-586.94
Socio-demographic controls	No	Yes	No	Yes	No	Yes	No	Yes

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Notes: Socio-demographic controls include employment status, mother's age, religious affiliation, living in an urban area, being of Russian ethnicity, the number of children in the family, and educational accomplishment. In model specifications corresponding to columns (3) and (4), dummies for both age ranges are included and the base category is all firstborns older than 18. In columns (5)-(8) the base category also includes first-born sons and daughters aged either 6-18 (columns (5)-(6)) or 0-5 (columns (7)-(8)). In all columns, however, numbers corresponding to first-born daughter dummies indicate the factor by which the hazard of union break up in families with first-born daughters is higher than in families with first-born sons in the same age range. That is, coefficients in the table become statistically significant when they are different enough from 1 (and not 0). Constant terms for estimating specifications (3) and (4) are suppressed. Columns' headings are "Separation and marriage termination" because the dependent variable takes value 1 if a separation happens according to the household-level data and an official marriage does not take place according to the individual-level data.

between the dummy for a first-born daughter aged 0-5 and moderator variables, observed socio-economic characteristics of mothers (being Orthodox, Russian, employed; having secondary education; living in an urban area; having above average well-being). Adding these interaction terms makes the coefficient on the dummy for a first-born daughter aged 0-5 comparable in magnitude and statistical significance to that on the dummy for a first-born daughter aged 6-18 in Table 4. Therefore, the lower separation among parents of first-born daughters aged 0-5 is driven by belonging to groups of the population with the foregoing characteristics (most notably, by

Table 5: The impact of preschool first-born daughters on separation after conditioning on interaction with socio-demographic characteristics.

Explanatory var-s:	Separation and marriage termination (1)	Separation and marriage termination (2)	Separation and marriage termination (3)	Separation and marriage termination (4)
Firstborn is daughter 0-18 years	2.30	2.03		
	(0.98)**	(0.86)*		
Firstborn 0-5 years old is daughter	0.40	0.41	1.27	2.48
	(0.19)**	(0.19)**	(0.56)	(1.21)**
Firstborn 0-5 years old	2.21	2.23	1.36	1.41
	(1.01)*	(0.99)*	(0.39)	(0.34)
Firstborn 0-5 years old is daughter			0.52	0.46
*Orthodox			(0.20)*	(0.18)**
Firstborn 0-5 years old is daughter			1.17	0.99
*Russian			(0.43)*	(0.41)
Firstborn 0-5 years old is daughter			0.76	0.61
*Secondary education			(0.23)	(0.20)
Firstborn 0-5 years old is daughter			0.73	0.70
*Urban			(0.27)	(0.26)
Firstborn 0-5 years old is daughter			0.69	0.46
*High well-being			(0.38)	(0.32)
Firstborn 0-5 years old is daughter			1.53	0.94
*Employed			(0.52)	(0.34)
Year	1.07	1.11	1.07	1.12
	(0.08)	(0.09)	(0.08)	(0.09)
Year ^{2}	0.99	0.99	0.99	0.99
	(0.04)*	(0.04)*	(0.04)*	(0.04)*
N of spells	1,367	1,367	1,367	1,367
N of marriage-year observations	7,163	7,096	7,103	7,048
Log-likelihood	-630.65	-583.98	-620.63	-562.54
Socio-demographic controls	No	Yes	No	Yes

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Notes: Socio-demographic controls are as in Table 4. In the first two columns, the base category is a first-born son aged 6-18 and in the other two columns the base category is all firstborns aged 6 and above. In all columns, as in Table 3.4, numbers corresponding to first-born daughter dummies indicate the factor by which the hazard of union breakup in families with first-born daughters is higher than in families with first-born sons in the same age range. That is, coefficients in the table become statistically significant when they are different enough from 1 (and not 0).

being an Orthodox).

Results of the estimation with cohabitation termination as a dependent variable in Table 3.C.1 in Appendix 3.C are in line with the results in Table 3.4.

As for the effect direction for the preschool daughters, I can put forth six explanations for $\beta_{1_{0-5}}$ to be negative. The first is the relationship with parents of spouses (esp. mothers of husbands), who may be more supportive of the marriage when a child is a daughter (Duflo, 2003; Aduškina, 2015; Aivazova, 2015; Mkchtrian, 2017).[29] The second is the loss in marriage surplus due to loss in the human capital of daughters. In other words, if investment in the human capital of children brings higher returns in terms of marriage surplus for daughters than for sons (Abramishvili et al., 2019), the divorce of spouses with a first-born daughter could cause an especially high loss in the marriage surplus (Currie and Almond, 2011; Myck et al., 2021; Kim and Shapiro, 2021).

The third possible reason is that a higher marriage rate among single mothers of first-born sons reduces the cost of separation for them as they perceive their remarriage prospects to be more favorable.

The fourth reason is that older children and boys fare worse in "stepfamilies" than younger children and girls (Brown, 2004). Thus, mothers of younger boys should try not to delay separation, once they opt for it, because remarriage is more challenging when sons get older. This conjecture is also supported by the comparison of baseline results with results of the robustness check. Namely, the fact that the observed effect for younger daughters becomes more statistically significant when separations with the preservation of the official marriage are not counted means that mothers who separate while remaining in official marriage are more likely to

[29]First, research by Duflo (2003) estimated the effect of starting pension payments in South Africa on grandchildren co-residing with pension recipients. The most pronounced effect on children's health (catching up with boys) was observed for granddaughters when pension recipients were grandmothers. The author does not investigate whether those grandmothers were paternal or maternal, but they are more likely to be paternal because a wife is likely to come to the household of her husband's parents. Thus, paternal grandmothers might support grandchildren more when they are daughters. Second, people might value potential old-age care when they approach their dotage. For example, a source in an Eastern European periodical (Mkchtrian, 2017) reports colloquial evidence that daughters pay more attention to old parents dwelling in rest homes than sons.

be mothers of daughters. [30] This might suggest that mothers of young sons are less willing to preserve a formal marriage because, for instance, the remarriage becomes more problematic when sons get older.

The fifth reason is a cumulative effect of existing policies, especially social policies. The fact that policies can be not neutral with respect to gender is discussed in Washington (2008). For example, Cygan-Rehm et al. (2018) show that increasing child benefit in Germany leads to higher cohabitation rates of couples having first-born daughters. In Russia, compared to other countries, social policies are likely to favor women more than men because women constitute a larger share of voters demographically and, at the same time, vote more actively (Goncharenko, 2018). Examples of such policies might include mother's exclusive entitlement to maternity capital introduced in 2007, [31] generous public pensions (relative to average salary) that disproportionately benefit women due to a significant gender gap in life expectancy, [32] indexing public sector salaries, which is likely to be more beneficial to women, who constitute a majority of public sector employees (including those in education and health care), or "gender asymmetry" in family law (Klimashevskaya, 2021). [33] These circumstances enhance the chances of daughters to be employed, financially secure, and, ultimately, capable of supporting their parents in their dotage. Thus, fathers are likely to have fewer reasons for leaving a family

[30] This is true if some of the observations of separations without official divorce correctly reflect actual living arrangements. I do not count such separations in baseline results because they might signal an error in a response. A separation when someone disappears from a household rooster is likely to be short-term one. It could hardly be the case when a person remains in an official marriage because in that case a person has legal rights in the household. That is why I do not count separations with preservation of official marriage, suspecting that the household file contains erroneous information in this regard.

[31] Except in the case when a man adopts a child and is the only parent.

[32] In Post-Soviet countries this gap is most pronounced globally.

[33] In addition, divorce in Russia has two characteristics that are at odds with conventional understanding of divorce causes in the literature. First, divorce is mostly initiated by women, which is at odds with the skewed sex ratio in Russia. Second, the main reason for divorce is "financial difficulties" (Antonov and Smagin, 2021), which is at odds with positive returns to scale from living together. The latter also needs explanation in view of the fact that the first divorce in Russia is not correlated with income and the second one is positively correlated (Laktjuhina and Antonov, 2017).

when their firstborn is a daughter.

Finally, the sixth reason is changes in personalities of parents induced by the gender of the firstborn. Specifically, van Lent (2020) found that fathers of first-born daughters have higher scores on neuroticism and extraversion. A more detailed discussion of this cause is provided in Appendix 3.A.

The six proposed explanations are based on the related academic literature and narratives in common contemporary socioeconomic discourse. However, they are not likely to exhaust the list of all possible explanations.

As for the baseline hazard, the exponentiated coefficient on the log(time) is significantly less than 1 in all specifications. That is, the estimated baseline hazard decreases with elapsed survival time. This result seems plausible because the divorce hazard is falling during most of a typical marriage after a relatively short period of rising following the start of a marriage. When the quadratic polynomial of time is used in the regression instead of the log(time), the coefficient on the squared year is negative and mostly statistically significant, corresponding to a bell-shaped form of the empirical divorce hazard.

4.2 Results for partnership formation

Estimates of model (2.a) with four specifications of the predictor (2.b) are presented in Table 3.5. The first-born daughter delays marriage, even without conditioning on age, as can be seen in column (2). Conditioning on age does not change the result substantially. This could also be partially explained by the fact that, in most observations, children living with single mothers are younger. The impact of first-born daughters on marriage of single mothers is close in magnitude (but opposite in direction) to the impact of 0-5-year-old first-born daughters on separation. Thus, the higher marriage rate among single mothers of first-born sons is outweighed to some extent by a higher separation rate among married mothers of first-born sons.

The possibility that the estimates obtained are driven by family background of

Table 6: The impact of the firstborn gender on family formation from a complementary log-log model estimation.

Explanatory var-s:	Partnership formation (1)	Partnership formation (2)	Partnership formation (3)	Partnership formation (4)
Firstborn is daughter	0.73	0.62		
	(0.18)	(0.15)**		
Firstborn 0-5 years old is daughter			0.59	0.63
			(0.16)**	(0.18)*
Firstborn 0-5 years old			0.81	0.86
			(0.31)	(0.47)
Year	0.75	0.81	0.71	0.78
	(0.07)***	(0.08)**	(0.08)***	(0.09)**
Year ^{2}	1.01	1.01	1.01	1.01
	(0.006)	(0.007)	(0.006)*	(0.007)
N of single mothers observed	255	255	255	255
N of marriage-year observations	1,051	1,051	1,051	1,051
Log-likelihood	-316.52	-280.34	-315.27	-288.93
Socio-demographic controls	No	Yes	No	Yes

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Notes: Socio-demographic controls include employment status, mother's age, religious affiliation, living in an urban area, being of Russian ethnicity, the number of children in the family, and educational accomplishment. In the first two columns, the base category is a first-born son and in the other two columns the base category is all firstborns aged 6 and above. In all columns, as in Table 3.4, numbers corresponding to first-born daughter dummies indicate the factor by which the hazard of union breakup in families with first-born daughters is higher than in families with first-born sons in the same age range. That is, coefficients in the table become statistically significant when they are different enough from 1 (and not 0).

mothers is examined in Appnedix 3.C

5 Conclusion

I confirm the finding from the previous literature that having daughters delays marriage of single mothers. However, I do not confirm the finding that having daughters of 0-18 years old causes separation. At the same time, I find that the effect of daughters on parental living arrangements depends on a daughter's age. In particular, having daughters aged 0-5 is related to a lower chance of parental separation, while having daughters aged 6-18 predicts a higher chance of parental separation.

The two effects seem to cancel each other in the pooled sample of firstborns aged 0 to 18. The latter effect accords with Kabátek and Ribar (2020) who report the negative impact of teenage daughters (aged 13-18) on the duration of the parental marriage. The former fact is a novel contribution to the extant literature.

Therefore, my findings give a reason to believe that the impact of the children's gender on family living arrangements likely depends on family socioeconomic conditions and thus has a different character and magnitude in different contexts. In this regard, it is worth mentioning that most studies finding that daughters cause separation use data from countries for which a son preference has been reported.[34] Son preference has not been established for Russia so far, [35] and, hence, in the Russian context it might have a lower impact than in other contexts. Moreover, some features peculiar to the Russian socioeconomic landscape are likely to mediate the relationship between the gender of children and parental marriage stability. At least six such features may be pointed out: relationship between spouses and their parents, higher returns to children human capital investment for daughters than for sons, lower subjective cost of separation among mothers of sons, women constituting the majority of voters and having relatively high electoral activity, older sons having harder time in "stepfamilies", and changes in parental personalities induced by having children of a particular gender. Examining the plausibility of these possible mechanisms will be a focus of my further research.

[34]Despite it not being found for the Netherlands, the country studied by Kabátek and Ribar (2020), it was reported for US (Dahl and Moretti, 2008), India (Barcellos et al., 2014), Australia (Kippen et al., 2006)

[35]Kim and Shapiro (2021) report indicative evidence supporting the daughter preference in Russia.

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Appendix A

Impact of changes in extraversion and neuroticism of fathers on marriage stability

According to van Lent (2020), neuroticism increases only among fathers of daughters aged 0-5 while extraversion remains higher for fathers of older daughters as well. Each of these two personality traits is related to marriage duration according to previous studies.

Most evidence in the literature supports a positive relation between higher scores on neuroticism of spouses and likelihood of divorce. This is, in turn, attributed to the negative relation between neuroticism and marital satisfaction, which was confirmed on data from the US (Claxton et al., 2012; Heaven et al., 2006; Boertien and Mortelmans, 2018), Great Britain(Boertien and Mortelmans, 2018) and Germany (Boertien and Mortelmans, 2018; Lundberg, 2012).[36] A similar relationship was confirmed for Iranian (Fani and Kheirabadi, 2011), Malaysian (Zare et al., 2013), and Russian (Kornienko and Silina, 2020; Nikolaieva, 2018) local questionnaire survey data. Therefore, if increased neuroticism among fathers of first-born 0-5-year-old daughters occurs in Russia, as is reported for the Netherlands (van Lent, 2020), they should be more likely to divorce. Still, Kabátek and Ribar (2020) do not observe a different divorce rate for fathers of young first-born daughters in the Netherlands. This means that the neuroticism effect should be sufficiently strong to have an impact on divorce. Hardly any research has been conducted so far to examine this question in the Russian context.

[36]Lundberg (2012), using data from German Socio-economic Panel Study, finds that neuroticism (as well as extraversion) is statistically significant only for women. According to Boertien and Mortelmans (2018), neuroticism appears to be related to a smaller likelihood of coping well with stressful events, as negative emotions appear to impede the ability to choose appropriate coping strategies

As for extraversion, its possible impact is less clear-cut. On the one hand, extraversion is assumed to be related to positive emotions (Donellan et al., 2004; Heaven et al., 2006), but on the other hand this trait increases the productivity of searching for partners, along with the arrival and assessment of marriage alternatives (Lundberg, 2012) since this trait is related to the ease of socializing and building social networks (Asendorpf and Wilpers, 1998). Accordingly, the available empirical evidence on the relationship between extraversion and divorce is inconclusive. While Lundberg (2012), Boertien et al. (2017) and Boertien and Mortelmans (2018), using German and UK data, report a positive association between extraversion and divorce risk; Solomon and Jackson (2014) do not find such an association in an Australian nationally representative sample. Moreover, Solomon and Jackson (2014) report a positive relationship between extraversion and marital satisfaction. The latter, according to Hirschberger et al. (2009), negatively predicts prospective marriage dissolution for men.[37] Therefore, the direction and magnitude of relation between extraversion and divorce likely depends on the relative strengths of influential factors in a specific socio-economic context.

In particular, in the Russian context, Kornienko and Silina (2020) find that, in the first 10 years of marriage, spouses with higher extraversion are focused on active development of a family relationship at the stage of its formation. Moreover, the authors admit that open expression of feelings is valued in young families: they are ready to address conflicts and express discontent with the spouse because the organization of rules and norms within the family is important for them.[38] Zelenskaja and Liders (2015) say that extraversion is associated with the presence of "communicative resources" needed for discussing the role structure of the family. [39] Nevertheless, it remains unclear how the role of extraversion at the onset of a

[37]Marital satisfaction around the first child's transition to school is the strongest predictor.

[38]In addition, Švecova and Kondraševa (2015) report that young husbands assign relatively high value to friend networks (compared to wives and older spouses).

[39]This fact resonates with the conclusion by Somville (2019) that having a daughter reduces

marriage compares to that in other societal settings. A study comparing attitudes among Lithuanian and Russian high school students towards family (Voroncova and Ermolaev, 2016) finds that Russian youth aim for more control over building relations in the family and rely less on norms and conventions than their Lithuanian peers. Exercising more control over family relations would apparently require more communication which, in turn, is facilitated by extraversion. The reason behind a higher reliance on intra-family negotiation along with lower reliance on norms and conventions among Russian youth could lie in the history of family and marriage in Russia. In particular, Brainerd 2008; 2016 finds that pronounced sex-ratio imbalances caused by World War II had a lasting effect on family structure in Russia, including lower rates of marriage and fertility, higher non-marital births and reduced bargaining power within marriage for women most affected by war deaths. Moreover, the author argues that the mentioned effects were likely magnified by family policies in the former USSR. This is why the effect of increased extraversion of young husbands is likely to be more pronounced than in other socio-cultural contexts.

Appendix B

Table 3.B.1: Data Description for Selected Variables

Separation	Dummy for whether a person who cohabited with a partner in a previous wave is not cohabiting with the same partner and reports being divorced or cohabits with a different partner in a current wave
Partnership formation	Dummy for whether a person who did not cohabit with a partner and reported being single in a previous wave is cohabiting with a partner in a current wave

male violence against a partner.

First-born child age	How many years have passed since firstborn birth
First-born child gender	Dummy for the first-born child being a girl
First-born child age 0-5	Dummy for the first-born child being 0-5 years old
First-born daughter age 0-5	Dummy for the first-born child being 0-5 years old and being a girl
First-born child age 6-18	Dummy for the first-born child being 6-18 years old
First-born daughter age 6-18	Dummy for the first-born child being 6-18 years old and being a girl
Father employment status	Dummy for whether a father is in registered employment
Mother employment status	Dummy for whether a mother is in registered employment
Father age (in years)	Age of a father at the time of an interview
Mother age (in years)	Age of a mother at the time of an interview
Father is Orthodox	Dummy for whether a father reports being an Orthodox Christian
Mother is Orthodox	Dummy for whether a mother reports being an Orthodox Christian
Father is Muslim	Dummy for whether a father reports being a Muslim
Mother is Muslim	Dummy for whether a mother reports being a Muslim
Mother reports other religion	Dummy for whether a mother reports adherence to another religion than Orthodox Christianity or Islam
Mother reports	Dummy for whether a mother reports adherence

no religion	to no religion
Number of children in the household	How many children below 18 live in the household
Urban area	Dummy for whether an interviewed household is located in an urban area
Mother is Russian	Dummy for whether a mother reports being of Russian ethnicity
Father is Russian	Dummy for whether a father reports being of Russian ethnicity
Mother has vocational or tertiary education	Dummy for whether a mother reports attaining vocational or tertiary education
Father has vocational or tertiary education	Dummy for whether a father reports attaining vocational or tertiary education
Number of family members	How many people live in the household
Mother reporting satisfactory life	Dummy for whether a mother is fully satisfied or rather satisfied with life at the current moment
Father reporting satisfactory life	Dummy for whether a father is fully satisfied or rather satisfied with life at the current moment
Having three meals per day	Dummy for being able regularly to have three meals per day
Consuming alcohol	Dummy for consuming alcoholic drinks at least sometimes
Drinking alcohol last month	Dummy for consuming alcoholic drinks during previous 30 days

Ever smoked	Dummy for being a smoker at some time
Smoking now	Dummy for being a smoker at the present moment
Foreign language	Dummy for knowing other language than languages of former USSR republics
Understanding between generations	Dummy for agreeing that understanding between generations is possible
Satisfied with material condition	Dummy for being currently satisfied with own material condition
Respectability	Dummy for feeling oneself sufficiently respected
Empowerment	Dummy for feeling oneself sufficiently empowered
Relative economic standing	Dummy for feeling oneself better off than others
Expecting economic improvement	Dummy for expecting improvement in economic situation of a family in next 12 month
Well-being improved in the last year	Dummy for improvement of well-being during previous 12 months
Unemployment concern	Dummy for being concerned about loosing a job
Having subordinates	Dummy for having subordinates at work
Satisfied with job	Dummy for being satisfied with one's own job
[a] Living in a regional center	Dummy for living in a regional center
Dwelling area	Area of an occupied dwelling in square meters
Population	Population of a municipality of residence
Family size	Number of people residing in a household
Rooms	Number of rooms in an occupied dwelling
Running water	Dummy for availability of running water in a dwelling

Running sewerage	Dummy for availability of running sewerage in a dwelling
Country house	Dummy for a family having a country house
Saved money last month	Dummy for a family saving money during previous 30 days
Help from others	Dummy for a family receiving financial or in kind help from others
Credit debt	Dummy for a family having credit debt
Owing money to others	Dummy for a family being in debt to other individuals

Notes. Covariates were selected with the minimum number of missing observations and based on the previous literature.

[a] The following covariates are taken from the household file.

Appendix C

Robustness check for marriage dissolution estimate

Table 3.C.1: The impact of the firstborn gender on cohabitation termination from complementary log-log model estimation.

Explanatory var-s:	Cohabitation termination (1)	Cohabitation termination (2)	Cohabitation termination (3)	Cohabitation termination (4)	Cohabitation termination (5)	Cohabitation termination (6)	Cohabitation termination (7)	Cohabitation termination (8)
Firstborn is daughter	1.09 (0.15)	0.95 (0.14)						
Firstborn 0-5 years old is daughter			1.04 (0.17)	1.04 (0.17)	1.03 (0.16)	0.95 (0.16)		
Firstborn 6-18 years old is daughter			1.51 (0.43)	1.50 (0.43)			1.49 (0.42)	1.58 (0.45)*
Firstborn 0-5 years old			0.02 (0.007)***	0.02 (0.03)***	1.02 (0.26)	1.10 (0.29)		
Firstborn 6-18 years old			0.03 (0.01)***	0.02 (0.001)***			0.79 (0.23)	0.72 (0.21)
Year	1.01 (0.06)	1.06 (0.06)	1.03 (0.06)	1.00 (0.07)	1.02 (0.02)	1.07 (0.07)	1.02 (0.06)	1.05 (0.07)
Year ^{2}	1.00 (0.004)	0.99 (0.004)	0.99 (0.003)**	0.99 (0.004)**	0.99 (0.004)	0.99 (0.004)	1.00 (0.004)	0.99 (0.004)
N of marriage spells	1,367	1,367	1,367	1,367	1,367	1,367	1,367	1,367
N of marriage-year observations	7,163	7,064	7,163	7,064	7,163	7,064	7,163	7,040
Log-likelihood	-954.18	-892.02	-957.54	-926.67	-954.36	-894.19	-953.37	-894.54
Socio-demographic controls	No	Yes	No	Yes	No	Yes	No	Yes

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Notes: Socio-demographic controls include employment status, mother's age, religious affiliation, living in an urban area, being of Russian ethnicity, the number of children in the family and educational accomplishment. The dependent variable is termination of cohabitation according to the household questionnaire data. That is, a husband stops residing in the household and and a wife does not report widowhood. Explanation of base categories and interpretation of estimates are provided in the notes under Table 3.4.

Controlling for family background of mothers

The family background of firstborn mothers can lie behind the observed effects of first-born daughters on marriage dissolution and formation. For example, Brainerd (2016) says that growing up in an incomplete family might in its turn lead to a higher chance of a woman being divorced or unmarried.[40] The estimates of the

[40] The author focuses her analysis on the situation of Soviet women in the wake of WWII. Thus the author's claim is limited to women. It might apply to men as well. Nevertheless, it appears intuitively plausible that women who grew up with single mothers might be more confident about bringing up their daughters on their own than mothers who have sons. Thus, in the following analysis I use only dummies for women's family background and not their interactions with the firstborn gender.

first-born daughters' impact on marriage dissolution and formation after controlling for women's family background are presented in Table 3.8. Variables characterizing the family background are four dummies for not being able to answer about father's and mother's year of birth, and occupation at the time when a respondent was 15 years old. Each dummy takes a value of 1 if a respondent finds it hard to answer about the four mentioned characteristics of the parental family background, and 0 in other cases (provides an answer, declines to answer, or there is no answer).

For parental occupation, respondents can specify an option that they are not able to provide an answer because they did not cohabit with a certain parent when they were 15. The latter should be correlated to some extent with parental divorce or separation. Not reporting a year of birth of a parent likely correlates with parental divorce or separation too (albeit to a lesser extent than in the case with parental occupation because not reporting a parental year of birth might be caused either by not knowing it or by unwillingness to respond for some reason). The estimates of the first-born gender impact in Table 3.8 appear to accord with the results in Tables 3.4 and 3.5 in direction and magnitude, but have a lower statistical significance, which could be explained by a smaller sample size (due to missing observations on the family background). Not knowing the father's occupation predicts a higher separation hazard, in line with expectations. The level of statistical significance for this effect is somewhat lower than 0.1, but it might increase after new waves are added into the sample. Regarding marriage formation, not knowing the parental occupation does not have a sizeable effect. Interestingly, not reporting a father's year of birth notably accelerates the marriage of single mothers. This might be related to a possible correlation between not knowing or not being willing to report a father's year of birth, and less-demanding expectations of a potential husband.

Table 3.C.2: The impact of the firstborn gender on cohabitation termination and marriage formation from complementary log-log model estimation with regressors for family background of mothers.

Explanatory var-s:	Separation (1)	Separation (2)	Separation (3)	Partnership formation (4)	Partnership formation (5)
Firstborn 0-5 years old is daughter	0.93 (0.06)	0.90 (0.21)			0.58 (0.19)*
Firstborn 6-18 years old is daughter	1.91 (0.70)**		2.49 (1.07)**		
Firstborn 0-5 years old	0.005 (0.002)***	1.41 (0.45)			0.79 (0.50)
Firstborn 6-18 years old	0.003 (0.001)***		0.44 (0.20)*		
Firstborn daughter				0.78 (0.1)**	
Father's occupation not known	1.55 (0.11)***	1.73 (0.43)**	1.98 (0.58)**	0.71 (0.05)***	0.46 (0.23)*
Mother's occupation not known	1.24 (1.67)	1.22 (1.78)	0.87 (0.38)	1.73 (0.99)	2.07 (0.99)
Father's birth year not known	0.40 (0.19)*	0.43 (0.19)**	0.52 (0.24)	1.90 (0.32)***	2.04 (0.58)***
Mother's birth year not known	1.46 (0.47)	1.52 (1.10)	1.22 (1.08)	0.66 (0.02)***	0.64 (0.40)
N of (marriage) spells	900	900	900	197	197
N of (marriage-)year observations	5,882	5,882	5,882	903	893
Log-likelihood	-476.69	-472.33	-470.48	-257.34	-243.80
Socio-demographic controls	Yes	Yes	Yes	No	Yes

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Notes: Socio-demographic controls include employment status, mother's age, religious affiliation, living in an urban area, being of Russian ethnicity, the number of children in the family, and educational accomplishment. These estimates capture the effect of including the family background of partners or single mothers on the results on previous estimations. Descriptions of base categories and interpretations of estimates are provided in Table 3.4 for columns(1)-(3) and Table 3.5 for columns (4)-(5).

Abstrakt

V tomto článku docházím ke třem hlavním zjištěním týkajícím se vlivu pohlaví prvorozeného dítěte na stabilitu rodiny. Za prvé, páry, které mají prvorozenou dceru ve věku 6-18 let, se častěji rozvedou než ty, které mají syna v tomto věku. Za druhé, svobodné matky s prvorozenými dcerami se méně často vdávají. Za třetí, u párů, které mají prvorozenou dceru ve věku 0-5 let, je menší pravděpodobnost rozvodu než u těch, které mají syna v tomto věku. První dva poznatky jsou v souladu s poznatky v literatuře. Třetí zjištění je specifické pro ruský kontext. Moje analýza je založena na datech Ruského Longitudinálního Monitorovacího Průzkumu (RLMS-HSE) za období 1994-2018. Odhaduji komplementární log-log (cloglog) regresní model závislosti rozvodu a manželství (u svobodných matek) na pohlaví prvorozených, věku a souboru sociodemografických charakteristik domácnosti. Moje zjištění podporují závěr, že vliv pohlaví dětí na uspořádání rodinného života závisí na socioekonomických podmínkách rodiny, a má tedy různý charakter a velikost v různých kontextech.

Working Paper Series
ISSN 2788-0443

Individual researchers, as well as the on-line version of the CERGE-EI Working Papers (including their dissemination) were supported from institutional support RVO 67985998 from Economics Institute of the CAS, v. v. i.

Specific research support and/or other grants the researchers/publications benefited from are acknowledged at the beginning of the Paper.

Published by
Charles University, Center for Economic Research and Graduate Education (CERGE)
and
Economics Institute of the CAS, v. v. i. (EI)
CERGE-EI, Politických vězňů 7, 111 21 Prague 1, tel.: +420 224 005 153, Czech Republic.
Phone: + 420 224 005 153
Email: office@cerge-ei.cz
Web: https://www.cerge-ei.cz/

Editor: Byeongju Jeong

The paper is available online at https://www.cerge-ei.cz/working-papers/.

ISBN 978-80-7343-572-1 (Univerzita Karlova, Centrum pro ekonomický výzkum a doktorské studium)
ISBN 978-80-7344-696-3 (Národohospodářský ústav AV ČR, v. v. i.)