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Dissertation Thesis

2023 **Sergii Maksymovych**

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Parenting of Sons and Daughters, Household Decision Making and Family Characteristics

Dissertation Thesis

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Abstract

The first chapter examines how household living conditions are related to alternative allocations of control over decision-making in the household. This study has three main findings. First, more equally shared decision-making in a household is closely connected to better household living conditions. Second, while predominant decision-control accrued to any of partners is correlated with worse living conditions, this is more pronounced for women rather than men. Finally, the distribution of the mode of decision-making in households does not strongly predict the regime of family finances.

The second chapter contributes to the body of research indicating the presence of a parental preference for a particular gender of children. The main objective of this paper is to test between the two main explanations for the existence of such preference, namely differences in the costs of raising sons and daughters versus the gender bias (corresponding to parental utility derived from a child's gender or from characteristics exclusive to that gender). Our evidence corroborates the cost difference explanation in countries exhibiting daughter preference.

In the third chapter, 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.

Abstrakt

První kapitola zkoumá, jak jsou životní podmínky domácností spojeny s alternativními alokacemi kontroly nad rozhodovacím procesem v domácnostech. Tato studie dochází ke třem hlavním zjištěním. Za prvé, více rovnostářský styl rozhodování uvnitř domácností je spojen s lepšími životními podmínkami. Za druhé, korelace mezi převládající kontrolou jednoho z partnerů nad rozhodovacím procesem uvnitř domácností a horšími životními podmínkami domácnosti je silnější pro ženy než pro muže. Za třetí, typ rozhodovacího procesu uvnitř domácností nijak silně nepředpovídá režim správy rodinných financí.

Druhá kapitola přispívá k výzkumu, který naznačuje přítomnost rodičovských preferencí ohledně určitého pohlaví dětí. Cílem kapitoly je otestovat dva hlavní vysvětlení existence těchto preferencí. Konkrétně se jedná o rozdílné náklady na výchovu synů a dcer na straně jedné a upřednostňování jednoho pohlaví na straně druhé (dané rozdílným užitkem rodičů z určitého pohlaví dítěte nebo charakteristik výhradně spojených s jedním z pohlaví). Naše důkazy podporují vysvětlení rozdílnými náklady na výchovu v zemích s preferencí dcer.

Ve třetí kapitole 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.

Keywords

Material deprivation; intra-household allocation; female empowerment; parental gender preference; parental inputs; gender gap; fertility; union formation; divorce; living arrangements

Klíčová slova

Materiální deprivace; alokace zdrojů domácností; posílení postavení žen; preference pohlaví dítěte; rodičovské investice; genderová propast; plodnost; uzavření manželství; utváření životního společenství; rozvod; rodinný stav

Length of the work: 153876 characters (including spaces)

Declaration

1. I hereby declare that I have compiled this thesis using the listed literature and resources only.
2. I hereby declare that my thesis has not been used to gain any other academic title.
3. I fully agree to my work being used for study and scientific purposes.

In Prague on May 14, 2023

Sergii Maksymovych

Acknowledgement

“…When as a youth I waxed more bold,
Time *strolled*
When I became a full-grown man,
Time RAN…”
- ***Henry Twells, Time's Paces***

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Introduction

"We must be careful not to
confuse data with abstractions
we use to analyze them."
- William James

"...economic analysis is really journalism...
Fiction and journalism"
- Ed Leamer, EconTalk episode with Russ Roberts

My study uses observational data to better understand how characteristics of families are shaped by constraints on decisions made by household members. I find certain uniformity in the way how collective nature of decision making in households mediates the relationship between constraints and family characteristics. I consider two constraints on decisions of household members: the assignment of decision control for different spheres in the household and predominance of either sons or daughters among children. In particular, I investigate how the assignment of decision control determines the depth of household material deprivation and how the predominance of sons or daughters among children determines family size, spending on children, and living arrangements.

The first chapter contributes to the recent work that focuses on the nexus between the extent of women's control and household living conditions. I employ direct measures of household decision control, instead of income-based proxies. I include direct measures of household decision control simultaneously for husbands and wives, whereas most of studies on the subject look at correlates of income transfers only to women. Further, unlike the existing work on household living conditions, which focuses on shares of specific goods in total household expenditures, I study direct measures of material deprivation and ask about their relationship to the nature of household decision making.

My analysis in the first chapter yields two main findings. First, not only female but also male control are concurrently correlated with household material conditions. Second, predominant male control over decisions and, even more so, predominant female control is associated with worse material conditions of the household compared to the balanced control of household decision making. The first fact, together with results in the related literature, indicates that using income transfers targeted to women as a proxy variable for women control might not be plausible. The second fact indicates that income transfers might not affect household outcomes only through female empowerment and their effects are actually driven by an increased collaboration of household members, i.e., by increase in the balanced control. Beyond that, I also observe that the degree to which household members pool their incomes is not closely related to the allocation of control over spending decisions which is in line with the foregoing interpretation of the second fact. All in all, my results question validity of not only the unitary model of the household but also the collective model which represents household decision making as the tug-of-war game. More realistic model of the household, which could better account for observed effects of income transfers, has to reflect the collective nature of decisions in the household.

In the second chapter I and co-authors aim to identify which of the two possible causes of gender preference is more prevailent in Balkan and Scandinavian countries. The two causes are parental bias in favor of one or another gender or different costs of raising sons and daughters. In the process of this research, my responsibility was to clearly formulate a specific substantive research question, propose a research design allowing for defensible model identification, conduct econometric analysis of large micro-level data sets (applying linear and Probit models using primarily Stata software), write up and present the results at academic conferences (e.g, CES, SEAM, and MiC conferences), and confirm listed steps with co-authors. We verify son preference using the parity-three progression method applied to a pooled EU-SILC 2004-2015 cross-sectional sample from four Balkan countries: Bulgaria, Croatia, Slovenia, and the Republic of Serbia. We also verify daughter preference for three Scandinavian countries, i.e. Denmark, Norway, and Sweden. Next, we find out which of the two aforementioned explanations - gender bias or differential costs - is more prevailent in Balkan and Scandinavian countries. Each explanation implies a distinctive relationship between the gender of children and the allocation of household resources. We test between the two explanations by checking which relationships hold for the household-level data.

We find that Balkan households with more female children replace furniture less frequently than households with fewer female children. Moreover, in households with more female children, mothers report a lower ability to spend on themselves. Additionally, for Balkan countries we find no difference in parental investment in male and female children and no impact of the gender composition of children on the ability to make ends meet or the minimum amount of money needed to make ends meet. We argue, based on earlier studies, that these findings are consistent with the gender bias explanation and not with the differential expenses explanation. For Scandinavian countries we find no impact of the gender composition of children on replacing furniture or on consumption of other household public goods, and we find significantly larger parental investment in households with more female children. Moreover, we do not find a systematic impact of the gender of their children on parental consumption. We argue based on conclusions in previous studies, that these findings are not consistent with the gender bias explanation but are in line with the differential expenses explanation. Supplementary analyses of the top-income-decile sub-sample and of cross-country relationships between gender preference, parental investment, and conventional measures of gender equality support our argument. Like my first chapter, the second chapter shows that considering the household as a single decision-making entity is not adequate to explain observed family characteristics.

The third chapter aims to estimate the impact of the gender of the first-born child on family formation and dissolution simultaneously conditional on the child age. I improve upon earlier studies in several ways. First, I use data from extensive panel survey, while most studies on the subject use cross-section data and do not control for the age of children. Second, a study, which uses longitudinal data and controls for the age of children, looks only at family dissolution (not looking at formation) and uses fewer covariates (while using many more observations) to see how they mediate the connection in question. Third, my research examines the impact of the firstborn gender on living arrangements in the Russian setting. The case of Russia deserves particular attention because earlier similar studies on Russia are lacking and because for decades it has been among several countries with highest reported divorce rates. I focus on children of in two different age ranges, pre-school children (0-5 years) and school-age children (6-18). I group children ages this way because I expect, based on peculiarities of the Russian context, there to be differential effects of preschool sons and daughters for marriage formation and dissolution.

I find that having daughters aged 0-5 years is related to lower chance of parental divorce while having daughters aged 6-18 years predicts higher chance of parental divorce. The latter effect accords with a previous study while the former one is specific to the Russian context. I point out several possible causes for these findings. Expressing these causes formally in terms of preferences and constraints could be less straightforward than formulating them conceptually in terms of psychological factors, broader family network characteristics, and specifics of institutional environment. The third chapter complements other two by showing that explanation of household choices might require taking into account not only multilateral modality of intrahousehold decision making but also peculiarities of the socioeconomic context.

In the end, complexity and diversity of social phenomena might limit applicability of deductive reasoning and corroborative techniques. As A. Deaton put it (Deaton, 2010), "Technique is never a substitute for the business of doing economics." That is why accumulation of useful knowledge and understanding might require "abduction" (a term introduced by J. Heckman (Heckman & Singer, 2017)). This means the process of looking for consilience across bodies of evidence and across studies. This also means revising models, hypotheses, and data analyzed to provide defensible explanations for surprising phenomena. In my work, I have followed this approach to show that choices of households, elementary units in the economy, are determined to a large extent by continually evolving conditions rather than by particular generic principles.

1. Decision-Making in the Household and Material Deprivation

1.1 Introduction

Multiple studies find that providing income transfers to women has a stronger impact on particular household outcomes than providing income transfers[1] to men. These outcomes include child health and higher household expenditure on nutrition, health, and housing (e.g., Bobonis, 2009; Duflo, 2003; Lee and Pocock, 2007; Lundberg et al., 1997; Lundberg and Ward-Batts, 2000; Hoddinott and Haddad, 1995). The mechanisms underlying this impact remain an active research area {Lundberg, 2007, The American Family and Family Economics}. Specifically, the mechanism underscored in much of the recent work is that transfers to women enhance women's empowerment so that household expenditures become more in line with women's preferences which are more pro-family and pro-child than men's (e.g., Duflo, 2012; Bobonis, 2009). Nevertheless, several studies do not find a strong positive association between women's income and household living conditions (e.g., Braido et al. (2012); Haushofer and Shapiro (2016); Thomas (1990)). Thus, there is some doubt regarding the conventional mechanism behind the effects of women-targeting income transfers targeted to women. This paper provides additional evidence that questions the logic of the conventional mechanism and supports an alternative interpretation.

The conventional explanation that the positive effects of targeting income transfers to women result from women empowerment would be correct under two assumptions. First, the income transfers targeted to women should be correlated with women's control in the household. Second, conditional on women's control, these transfers should not be correlated with other determinants of household outcomes (Cameron and Triverdi, 2005)[2]. Based on evidence in the literature, I conclude that the first assumption holds. However, the findings of this study combined

[1] Income transfer mainly refers to cash transfer programs (e.g., Bolsa Alimentação or PROGRESA), but also to transfers during field experiments or incomes brought by price shocks in markets of female-specific crops (cultivated only by women).

[2] In other words, the studies estimating the impact of income transfers targeted to women use presence of these transfers as a proxy variable for women control. The two mentioned assumptions that presence of the transfers should be correlated with women empowerment and not correlated with other determinants of the household outcomes are correspondingly the relevance and the redundancy assumptions made about the proxy variable.

with findings in the literature suggest that the second assumption might not hold. Specifically, in my analysis of Eurostat data I find that male control is also correlated with household material conditions. Few other studies also find that the income transfers to women also change the extent of male control in the household. For example, Ashraf (2009) suggests that windfall income transfers also impact men's behavior: men negotiate the use of transfers (more or less intensively, depending on who receives the transfer and on the allocation of control in the household before transfers[3]) and men may adjust their contributions to the household budget depending on how the transfer is spent. Iversen et al. (2011) find that men and women contribute the largest shares of individual windfall incomes to the common pool in a trust game when a woman is ultimately assigned the control of the final allocation and that men contribute more than women to the common pool independently of who is assigned the control of the final allocation. In view of these last findings, the authors and Jackson (2013) conclude that women's preferences are not necessarily more pro-family than men's and that cash contributions to the common pool do not determine the bargaining power which is largely context-dependent. Natali et al. (2016) find that a cash transfer increases balanced control, i.e., control by both partners, over a number of household decisions, but does not increase solely female control. Another study (Haushofer and Shapiro, 2016) finds that cash transfers do not affect female empowerment in households, but their measure of female empowerment is based on reported instances and attitudes to domestic violence. Therefore, income transfers to women affect male control and female control simultaneously while each is independently related to household material conditions. Thus, it may not be accurate to ascribe the effects of women-targeting income transfers to women empowerment alone[4].

In this research, I employ *direct* measures of household decision control, instead of income-based proxies. Further, unlike the existing work on household living conditions, which focuses on shares of specific goods on total household expenditures, I study direct measures of material deprivation and ask about their relationship to the nature of household decision making. The analysis is based on EU-SILC data from 2010, covering 18 EU member countries. The ability to measure both household control structure and relevant household outcomes directly affects the

[3] The author assumes that the allocation of control is determined before the marriage and does not change thereafter.

[4] In addition, Tommasi (2017) finds that positive effects of cash transfers to women in the Mexican PROGRESA program are mostly driven by a subsample of women who already controlled most of household resources before the transfer. However, if the conventional explanation was accurate, these results would be driven instead by a subsample of women who do not possess significant control over household resources before the transfer.

results qualitatively. As in most existing studies, women's relative income is strongly associated with better household outcomes and this association turns out to be robust to controlling for direct measures of decision-making control in the household. However, predominant male control over decisions and, even more so, predominant female control is associated with worse material conditions of the household compared to the balanced control of household decision making[5]. Most of the existing work implicitly assumes that changes in female control are necessarily accompanied by equivalent and opposite changes in male control[6]. However, the EU-SILC data suggest that female and male control can increase simultaneously when both partners start deciding on more items, i.e., when the overlap of their spheres of responsibility expands. Hence, without further evidence, the effects of women's windfall income on household outcomes should not be interpreted as corresponding only to female control, which is crucial from a policy perspective. An alternative explanation is that the observed effects are driven by an increased collaboration of household members, i.e., by increase in the balanced control. In this paper, I attempt to verify this conjecture.

The EU-SILC data does not allow me to study the association of *windfall* income and female and male control. Moreover, my OLS results could be affected by measurement error or reverse causality. Thus, to shed more light on the negative association of unbalanced control and household income, I instrument for the observed mode of decision-making using the share of 4-year-olds in formal childcare, the gender-gap in unemployment, and the gender-gap in weekly work-hours. All three variables are measured at the NUTS 2 regional level. These are region-based instrumental variables that affect the women's position in the household, but are arguably unrelated to household-specific outcomes. The results of the IV estimation are in accord with the baseline OLS results supporting the notion that the mode of decision-making affects material status of households.

To further support the main argument that income transfers do not affect household outcomes only through female empowerment, I show that the degree to which household members pool their incomes is not closely related to the allocation of control over spending decisions. However, for the conventional interpretation to be correct, households which do not pool income

[5] In Section 4, I indirectly test for the possibility that this association is a consequence of reverse causality and reject this notion.
[6] This is the case in the framework of collective models; see, e.g., Almås (2015).

should be much more likely to make decisions separately. This is because individual transfers to households which pool income should not make a difference as it is not important who exactly receives income when it is pooled. Moreover, households which do not pool incomes should be much more likely to make decisions separately because if they make decisions collectively, it matters less if they pool income or not[7]. Nevertheless, the share of households making decisions collectively among those who do not pool incomes is about the same as among those who do. This last finding also contributes to the two strands of empirical literature: the literature that tests the unitary model of the household, and in the socio-economic literature that explores the management of finances in families.

In the context of the unitary model, the concept of income pooling is used in two different senses (Lundberg and Pollak, 2007; Lundberg and Pollak, 2008). First, this means that the budget constraint of the household contains the sum of the individual income of partners, i.e. the entire income of partners is "pooled". In other words, all income is spent on maximization of a single household utility function. In the second sense, income pooling means that the income recipient is irrelevant to the allocation of family resources. This is true if and only if income pooling in the first sense holds and the individual income of partners does not enter the household utility function. Turning to the literature on family finances, it understands income pooling in its first meaning, i.e. as partners contributing individual income to a "common pot", "kitty", etc. (Pahl, 2005) and drawing it down at their individual discretion. If such income pooling takes place, individuals who do not make decisions on important household matters are still insulated from being much poorer than other household members. This is because non-primary decision-makers can satisfy their needs by taking money from the common pool if the predominantly deciding household member does not does not take their needs into account in budgeting decisions. When income is not pooled, individuals who do not make decisions may be more deprived than other household members as they do not have access to pooled resources. To evaluate how likely such a situation is to arise, I check if individualized decision making is also accompanied by individualized family finances (i.e. no income pooling). The absence of the strong relationship between the individualized decision-making and income pooling belies the previously mentioned concerns in the literature. (Jackson, 2013)

[7] Some studies, however, mention the "labelling effect" or "spending inertia" (Lundberg and Pollak, 2007) which characterize income transfers. However, these effects are not large and the evidence is scarce.

The EU-SILC 2010 data contains responses about the shares of partners' individual income contributed to a common pool. I find households that pool income are more likely to use a more balanced decision-making mode. This finding is important to the literature on the unitary model of the household and the family finance literature. In particular, it underlies concerns expressed in the family finances literature that households with no income pooling are likely to end up in a situation in which there is significant inequality between household members (Elizabeth, 2001; Pahl, 2005). Moreover, it is clear that not all households pool income, thus violating the assumption of the unitary model[8].

1.2 Data

The data is obtained from the 2010 round of the European Union Survey of Income and Living Conditions (EU-SILC). The EU-SILC is based on a national representative probability samples. It collects a comprehensive dataset containing information on income, poverty, social exclusion, and living conditions. The reference population includes all households and their current members residing in the countries at the time of data collection. Bases of sampling differ from country to country. In most cases, it is either the population register or the population census. In 2010, EU-SILC was implemented in the EU-27, Croatia, Montenegro, Iceland, Turkey, Norway, and Switzerland. The data used in this research covers only the EU-27 countries, and except Austria, Cyprus, Belgium, Denmark, Finland, Hungary, the Netherlands, Slovenia, and Sweden, because these countries use a sample of individuals and all persons in a household are not interviewed. In the rest of the EU-27 countries, all household members aged 16 and up are surveyed.

The survey collects primary and secondary variables. The primary variables are collected annually. They characterize a household as a whole or as its individual members. The household-level variables are divided into four domains: basic data, housing, material deprivation, and income. The individual-level variables are divided in five domains: basic/demographic data, education, health, labour, and income. The material deprivation variables from the household-level

[8] The fact that partners pool income implies that they behave as if they have common arguments in their utility functions or even a common utility function. If they did not, none would pool individual income because a partner could take it all for personal use.

domain are of particular interest for our research. Its primary focus is the relationship between these variables and the mode of decision-making.

Secondary variables are collected in the so-called ad-hoc modules every five years or less frequently. In year an ad-hoc module on intra-household sharing of resources was implemented. Its objective was to look into the decision-making process and the allocation of resources within the household. The 2010 ad-hoc module supplements primary poverty risk indicators by providing information on differences in living standards between household members. The questionnaire includes questions on participation of household members in important financial decisions. Specifically, each adult household member is asked to evaluate the degree of his/her participation in decisions about common savings, borrowing money, everyday spending, spending on durables, and important purchases for children. In the questionnaire, for all the above questions there are offered three possible answers about the degree of participation: “More me”, “Balanced”, “More my partner”. The reference period is three months preceding an interview.

All except three countries (the United Kingdom, Cyprus, and Ireland) achieved the minimum effective sample size (the sample size stipulated by the EU). The difference between the actual sample size (the number of actually selected households) and the achieved sample size (the number of actually completed interviews) lies between 5.43 % (Bulgaria) and 37.61 % (Belgium) of the actual sample size. The first most common reason for interview non-completion is refusal of a household to cooperate. The second is a household not contacted. The achieved sample size varies from 2,148 households for Cyprus to 8,768 for France, 8,962 for Germany, and 13,318 for Italy. Individual non-response rates vary for different questions. Usually they are low. For most of questions on decision control in most countries, non-response rates do not exceed 1 % (Eurostat, 2012). There are, however, consistently high nonresponse rates in France (between 17.5% and 18%) and Poland (always 25.1%). In addition there are high non-response rates in Belgium for the question about decisions on durables (26.8%) and in Ireland for questions about decisions on everyday shopping (25.2%) and decisions on purchases for children (33.4%). Non-response rate for the primary variables is about 1% (see also Eurostat (2016)).

1.3 Empirical specification

This study explores the correspondence between material deprivation conditions and the mode of decision-making in households. Measures of material deprivation and the mode of decision-making are constructed from responses of household members to the EU-SILC-2010 questionnaire. The household-level responses to questions on material conditions are used to construct the measure of household material deprivation. At the same time, individual-level responses about involvement in making decisions are used to construct an indicator of the mode of decision- making in the household.

I construct one composite measure of material deprivation. This measure is similar but not identical to the Eurostat material deprivation criterion, according to which a household is materially deprived if it fits 3 of 11 material deprivation criteria (Fusco et al., 2010). I do not use all 11 criteria, but only 6. My measure is equal to the sum of six binary variables taking a value of 1 if a household satisfies the corresponding criterion and 0 otherwise. The conditions are: arrears on mortgage payment, arrears on utility bills, arrears on hire installments, inability to afford one-week holiday away from home, inability to face unexpected financial expenses, inability to make ends meet, and inability to afford a meal with meat, fish, or chicken every second day[9]. The distribution of households by the sum of dummies is shown in Figure 5. If the sum of these dummies for a given household exceeds 2, the composite measure of material deprivation takes value 1[10].

Regarding the indicator of the mode of decision-making, it is constructed from individual responses about how much a given person is involved in making specific decisions. The approach is based on the one adopted by (Li and Wu, 2011). Each spouse is offered three alternative options to characterize their involvement in making decisions: “More me”, “Balanced”, and “More my partner”[11]. In my analysis I consider only households in which partners give consistent answers to most of questions; when one partner answers to a given question “More me” and other answers “More my partner” or both answer “Balanced”. I consider only households with mostly consistent responses because it is necessary for constructing my indicator of the mode of decision-making

[9] There are also questions about five more material deprivation conditions. They are disregarded in the analysis because of a very small variation in responses (only 7% of households don't have access to a car because they cannot afford it and literally all have the other three items). They are about the ability to afford the following items: keeping the home adequately warm, having a washing machine, having a colour TV, having a telephone, having a personal car.

[10] The results do not change much neither qualitatively nor quantitatively if the threshold for the sum of component dummies is equal to 3.

[11] When the question is about deciding on common savings, there are also other alternatives including “Never arisen" or “No common savings".

described below. Consistent responses constitute more than 90% of all responses to any considered question.

I focus on responses about the following five decisions[12]: ‘decision-making on everyday shopping’, ‘decision-making on expensive purchases of consumer durables and furniture’, ‘decision-making on borrowing money’, ‘decision-making on use of savings’, ‘decision-making in general’. Based on these responses I, distinguish five different modes of household decision-making: ‘man-led’, ‘primarily man-led’, ‘woman-led’, ‘primarily woman-led’, and ‘balanced’. Specifically, if a man makes four or five decisions, the household is labeled “man-led”. If a man makes three decisions, the household is labeled “primarily man led". Similarly, if women make three decisions or more than three decisions, households are labeled “primarily women-led” or “women-led” correspondingly. The rest of the households are labeled “Balanced” and constitute the reference group. Table 2 illustrates how the indicator of the mode of decision-making is constructed.

It is worthwhile to discuss the intuition behind my measure of the mode of decision-making in more detail. In the literature, indicators similar to the one constructed in the current study are called “bargaining power” (Li and Wu, 2011), “measures of empowerment” (Almås et al., 2015), “decision-making index” (Natali et al., 2016). These indicators are meant to show how much influence a woman has in the household. (Almås et al., 2015), however, assume that the female empowerment measure should be proportional to the share of household income that is spent as if a woman were the sole decision-maker. In the current study I shall stick to the above mentioned terms labeling allocation of control. However, it should be understood that the time and effort committed by either spouse to working out the best possible allocation of household resources underlies the notion of 'control'[13].

[12] Two other decisions are excluded from analysis because they are not pertinent to the condition of the entire household. Namely, ‘ability to decide about expenses for your own personal consumption, your leisure activities and hobbies’, ‘ability to decide about purchases for children's needs (including giving them pocket money)’. Also, not all households have children, so the ‘decision making on important expenses for the child(-ren)’ is also excluded.

[13] Such understanding allows for simultaneous increase in the control of both partners. In other words, both partners can become more involved in working out a specific decision. At the same time, the income sharing-rule interpretation of intra-household control mentioned above (Almås et al., 2015) is based on the collective household model. This model implies that an increase in one partner's control is necessary accompanied by a decrease in another partner's control. That is why our proposed understanding of intra-household control is better captured by the bargaining household model (Lundberg and Pollak, 1993). This model features cooperative and non-cooperative equilibria. In our framework, more balanced decision-making corresponds to the theoretical concept of cooperative equilibrium. Non-cooperative equilibria might be not Pareto optimal. This is in line with later empirical findings (Udry, 1996).

Table 2 Construction of decision-making dummies

Number of decisions made by men	Number of decisions made by women	The mode of decision-making
5	0	Man-led
4	0	Man-led
4	1	Man-led
3	0	Primarily man-led
3	1	Primarily man-led
3	2	Primarily man-led
2	0	Balanced
2	1	Balanced
2	2	Balanced
2	3	Primarily woman-led
1	0	Balanced
1	1	Balanced
1	2	Balanced
1	3	Primarily woman-led
1	4	Woman-led
0	0	Balanced
0	1	Balanced
0	2	Balanced
0	3	Primarily woman-led
0	4	Woman-led
0	5	Woman-led

Turning to the empirical specification, let Dec_i be a vector of four dummy variables for the modes of decision-making (the reference category is “balanced”). The outcome of interest is a value of a material deprivation indicator j for a household i, y_{ij}. I estimate the following empirical model:

$$y_{ij} = \beta_j Dec_i + \alpha_j X_i + \varepsilon_{ij} \tag{1}$$

where X_i is a vector of respondent, spousal, and household characteristics, and ε_{ij} is the error term. The null-hypothesis is $H_0: \beta_j = 0$. The methods used for estimation are OLS and 2SLS. Concerning the covariates X_i, the main ones included are: family income, number of children of specified age, number of daughters of specified age, length of cohabitation of spouses, living in a rural area, being unemployed, employment status, hours spent on job market work, hours spent on housework, hours spent on leisure, education level, and occupation (a more detailed list appears in the next part). Besides being intuitively relevant, these controls are among those most frequently encountered in the literature. A possible theoretical reasoning behind use of specification (1) is contained in the corresponding working paper and could be provided by request.

The Equation 1 incorporates three specifications. The baseline specification contains RHS dummies for modes of decision-making along with the controls listed above. The second includes, in addition to all previously used regressors, interaction variables in between the decision controls and household characteristics. These characteristics are: educational attainment of each partner, unemployment during the preceding six months, length of cohabitation of partners, and pooling or non-pooling of individual incomes. The third specification has four regional characteristics on the RHS in addition to controls in the first specification, and uses instrumental variables for the mode of decision-making.

1.4 Results

The analysis sub sample includes only households composed of one couple of cohabiting partners with or without children. Most households report balanced decision making (Table 3). The sample includes only households in which couples give consistent responses, i.e., if a man responds about his role in some decision “More Me”, then the woman responds “More my partner”. From Table 3 it is clear that about 90% of responses are consistent.

Table 3 Percentages of responses of different types

Type of response			Decision-making measure					
A woman's response	A man's response	Agreement of responses	Decision-making in general	Decision-making on use of savings	Decision-making on borrowing	Decision-making on purchases of durables	Decision-making on everyday shopping	Decision-making on important purchases for children
More me	More me	Disagree	0.7	0.4	0.4	0.4	1.4	0.7
More me	Balanced	Disagree	2.0	1.1	1.4	0.8	3.6	4.7
More me	More my partner	Agree	7.1	4.8	7.6	3.6	47.0	21.3
Balanced	More me	Disagree	2.0	1.3	1.3	1.5	0.8	0.8
Balanced	Balanced	Agree	77.5	84.3	81.6	84.0	39.2	66.6
Balanced	More my partner	Disagree	1.5	1.2	1.7	0.8	2.1	3.3
More my partner	More me	Agree	7.4	5.3	4.7	7.4	4.5	1.5
More my partner	Balanced	Disagree	1.4	1.3	1.1	1.3	0.5	0.6
More my partner	More my partner	Disagree	0.4	0.2	0.3	0.3	0.8	0.4
Total percentage			100	100	100	100	100	100
N of households (non-weighted)			82,459	68,016	77,677	60,518	82,626	27,155

Source: 2010 European Union Survey of Income and Living Conditions and author's calculations.

Table 4 shows absolute frequencies of households in the sample by response consistency. When households are divided into two groups by share of consistent responses being 66%[14], the means of selected household characteristics tend to differ very little between the two groups.

Table 4 Frequencies of responses by consistency when at least one question is answered by both partners

Number of consistent responses	Number of households	Percentages
0	287	0.4
1	717	0.9
2	1,879	2.3
3	4,963	6.0
4	14,364	17.4
5	60,516	73.2
Total	82,726	100

Source: 2010 European Union Survey of Income and Living Conditions.

[14] Despite the vast majority of households give consistent responses to all five questions, the analysis incorporates households who reply at least three questions and give at most one inconsistent response. This reduces the sample selection problem.

This could be seen in Table 5. There are only minor differences in mean household disposable income and earnings: those responding inconsistently tend to earn a little more, despite literally no difference in the hours worked. It might mean that people who give inconsistent responses are more likely to have higher earnings and to be more focused on work-related rather than home-centered activities. Also, partners more frequently report primary decision-making in households with more inconsistent responses. This is a mechanical relationship: an inconsistent response can happen only if one partner reports primary decision-making. The described similarities between the two groups make it possible to focus on households which gave predominantly consistent responses[15].

Table 5 Sub-sample weighted means of selected household characteristics by response consistency

Household characteristics	More than 66% of responses are consistent	Less than 66% of responses are consistent
	Means	Means
Number of persons in a household	3.02	3.06
Age of a woman	50.31	49.51
Age of a man	53.12	52.43
A number of children	0.65	0.70
A woman having tertiary education	0.24	0.25
A man having tertiary education	0.27	0.27
A woman having secondary education	0.42	0.40
A man having secondary education	0.41	0.40
A woman being full-time employed	0.30	0.29
A man being full-time employed	0.48	0.46
A woman being part-time employed	0.14	0.16
A man being part-time employed	0.02	0.03
Yearly earnings of a woman (gross), euros	8,289.43	9,335.15

[15] Several variables statistically significantly predict consistent responses: employment of a man, ownership of accommodation, living in an urban area, and a woman doing more housework. But, the largest associated change in the likelihood of a consistent response is 0.05 for ownership of accommodation and around 0.01 for remaining three variables.

Yearly earnings of a man (gross), euros	15,284.62	16,599.72
Hours worked per week by a woman	16.75	16.45
Hours worked per week by a man	25.91	25.99
Household disposable income	32,416.57	35,913.32
Ownership of a dwelling	0.33	0.28
Living in highly urbanized area	0.46	0.48
Lowest monthly income to make ends meet, euros	2,706.95	1,992.91
Having arrears on mortgage payments during the previous month	0.18	0.11
Having arrears on utility bills during the previous month	0.20	0.13
Having arrears on hire purchase installments during the previous month	0.12	0.13
Inability to afford a two-week holiday once in a year	0.33	0.34
Inability to afford meat-containing diet every second day	0.06	0.07
Inability to face unexpected financial expenditures	0.30	0.32
Inability to make ends meet	0.52	0.54
A woman responds the household questionnaire	0.42	0.39
Incomes are pooled	0.78	0.69
A woman reports primary decision-making in general	0.09	0.21
A man reports primary decision-making in general	0.09	0.23
A woman reports primary decision-making on savings	0.05	0.14
A man reports primary decision-making on savings	0.06	0.21
A woman reports primary decision-making on durables and furniture	0.07	0.20
A man reports primary decision-making on durables and furniture	0.05	0.18
N of households (weighted)	75,102	7,624

Source: 2010 European Union Survey of Income and Living Conditions.

1.4.1 Baseline results

The results of estimating the baseline specification of equation 1 are presented in column 1 of Table 1[16]. Higher reported personal control by either spouse is associated with more frequent instances of any material deprivation measure. Analysis by separate countries yields qualitatively similar results, which, however, are not always statistically significant and not uniform in scale. Therefore, we analyze a pooled data sample while controlling for country specific effects (this approach is quite common in the literature).

The baseline results exhibit three noticeable features. First, the share of women's income is negatively associated with material deprivation[17]. This result is in line with the findings of other studies[18]. Second, the higher the degree of individualization in household decision-making, the higher the frequency of any kind of material deprivation. Third, predominant control by women is connected to higher frequencies of all kinds of material deprivation than individual control by men (these differences are also statistically significant at 10% level for all outcomes except hire purchase installments and mortgage payments). Therefore, the direct measure of female control corresponds to better household material conditions only when women are in control of household decisions together with men. When women are sole decision-makers in the households, household material conditions are worse. The use of women's income share as proxy variable for women's control, however, would suggest that more women's control unconditionally corresponds to better material conditions.

Thus, the interpretation of the share of female income as the female empowerment might be misleading. Moreover, the fact that men's control is also related to household material conditions, combined with the finding that windfall incomes accrued to women change control of both men and women is important in two ways for interpreting the effects of windfall incomes handed to women on household outcomes.

First, a conventional interpretation assigns these effects to increased female control. Still, if men’s control changes simultaneously and is related to household outcomes, the effects in question cannot be assigned only to female control. To reinforce this claim we conduct a series of estimates to check whether the correlations obtained are driven by some confounders. Second, it

[16] Results for six constituent indicators are presented in Table 6.
[17] Also, replacement of women's income share by women's relative earnings yields similar results, but woman's relative earnings are available only for about half of observations. Both woman's income share and relative earnings are used in the literature as a proxy variable for female control.
[18] This is result is not driven by the presence of unemployed women in the sample. It also holds for the subsample of employed women.

is quite possible that the increased balanced control partly drives those positive effects reported in the literature. If causation from balanced control to better material conditions were established, it would support this notion. That is why I also conduct instrumental variables estimation of a modified Equation 1. The results of controlling for potential confounders as well as of 2SLS estimation are reported next.

Table 1 Mode of Decision-making and Composite Material Deprivation

Explanatory var-s:	Dep. Var.: Material Deprivation		
	(1) OLS	(2) OLS	(3) IV
Woman takes control over 4-5[a] decisions	0.068 (0.010)***	0.085 (0.035)**	
Man takes control over 4-5 decisions	0.025 (0.012)**	0.132 (0.038)***	
Woman takes control over 2-3 decisions	0.026 (0.005)***	0.003 (0.018)	
Man takes control over 2-3 decisions	0.006 (0.008)	-0.005 (0.022)	
Control balanced between partners			-0.311 (0.176)*
Woman takes control over 4-5 decisions*Men's unemployment		0.255 (0.036)***	
Regional gender gap in unemployment			0.003 (0.029)
Regional share of employment in hi-tech industries			-0.143 (0.019)***
Share of population having access to broadband internet connection in a region			-0.003 (0.001)***
Regional rate of long-term unemployment			0.019 (0.002)***
Share of woman's income in total household income	-0.024 (0.014)*	-0.028 (0.014)**	-0.037 (0.016)**
Woman responds the questionnaire	0.034 (0.004)***	0.034 (0.004)***	0.022 (0.008)***
Number of children	0.044 (0.002)***	0.043 (0.002)***	0.044 (0.002)***
Woman's age	-0.003 (0.000)***	-0.003 (0.000)***	-0.002 (0.000)***
Man's age	-0.002 (0.000)***	-0.002 (0.000)***	-0.003 (0.000)***
Woman has tertiary education	-0.156 (0.006)***	-0.159 (0.007)***	-0.157 (0.008)***

Table 1 continued

Explanatory var-s:	Dep. Var.: Material Deprivation		
	(1)	(2)	(3)
	OLS	OLS	IV
Man has tertiary education	-0.162 (0.006)***	-0.165 (0.006)***	-0.150 (0.014)***
Woman has secondary education	-0.064 (0.005)***	-0.065 (0.005)***	-0.060 (0.005)***
Man has secondary education	-0.075 (0.005)***	-0.077 (0.005)***	-0.056 (0.009)***
Woman is employed full-time	-0.108 (0.007)***	-0.107 (0.007)***	-0.096 (0.008)***
Man is employed full-time	-0.102 (0.005)***	-0.095 (0.006)***	-0.096 (0.008)***
Woman is employed part-time	-0.057 (0.008)***	-0.060 (0.008)***	-0.033 (0.009)***
Man is employed part-time	0.006 (0.015)	0.011 (0.015)	0.007 (0.016)
Woman is self-employed	-0.064 (0.007)***	-0.066 (0.007)***	-0.056 (0.009)***
Man is self-employed	-0.119 (0.006)***	-0.114 (0.006)***	-0.129 (0.008)***
Woman's earnings	0.000 (0.000)***	0.000 (0.000)***	0.000 (0.000)***
Man's earnings	-0.000 (0.000)	-0.000 (0.000)	-0.000 (0.000)
Household disposable income	-0.000 (0.000)***	-0.000 (0.000)***	-0.000 (0.000)***
Own accommodation	-0.133 (0.005)***	-0.129 (0.005)***	-0.128 (0.005)***
Densely populated area	-0.011 (0.004)***	-0.010 (0.004)***	-0.018 (0.005)***
Country dummies	Yes	Yes	Yes
R^2	0.22	0.22	0.16
N	64,082	62, 358	64,660

[a] Out of 5 or 4 decisions consistently reported by a household

Notes: Robust standard errors are in parentheses. The sample contains households consisting of a couple with or without children. Households with inconsistent responses and with female income higher than total household income were excluded. The dependent variable is the *Material deprivation index*. It is a binary variable taking value 1 if more than 2 of 7 considered material deprivation conditions occur in an observed household. Numbers in the first two columns represent estimations of two specifications of Equation 1. The second specification differs from the first in the presence of interactions between the modes of decision-making on several household characteristics. The RHS in specifications (1) and (2) contains 4 decision-making dummies with the balanced mode being a reference category. The third column contains results of IV estimation explained in Subsection 4.1. In specification (3) the RHS contains only one decision-making dummy which is for the balanced mode, while other modes constitute the reference category. The two instruments used for the balanced mode of decision-making are: the regional rate of involvement of 4-year-olds in formal childcare and the regional gender gap in weekly work hours.

1.4.2 Controlling for potential confounders

There are several variables that could confound the baseline results and are present in the data set. These are: educational attainment of spouses, hours of job market work and of housework, man's long-term unemployment, length of cohabitation of partners, and the regime of family finances. These variables and interactions between them and the mode-of-control dummies are included in the RHS of Equation 1. The results of the estimations are shown in column 2 of Table 1. The main conclusion is that the established relationship between the household mode of decision-making and material deprivation still holds and is not driven by the suggested confounders.

Among the confounders considered, the regime of family finances deserves special attention. If the mode of decision-making is closely related to the regime of family finances, it will support the concerns in the literature about within-family consumption inequality due to individualization of family finances (Pahl, 2008; Vogler and Pahl, 1994). Figure 1 shows distributions of households by mode of decision-making conditional on the regime of family finances. The share of households reporting balanced decision-making decreases when individual incomes are treated autonomously.

The character of the relationship between decision-making and material deprivation, however, does not change compared to the baseline case. This can be seen from column 2 of Table 1. Thus, the concerns in the literature about possible intra-familial consumption inequality are partially warranted due to the fact that families with pooled incomes more frequently make decisions in a balanced way. Nevertheless, the allocation rules probably do not change since the relationship between decision-making and deprivation is similar for both income pooling and non-pooling regimes.

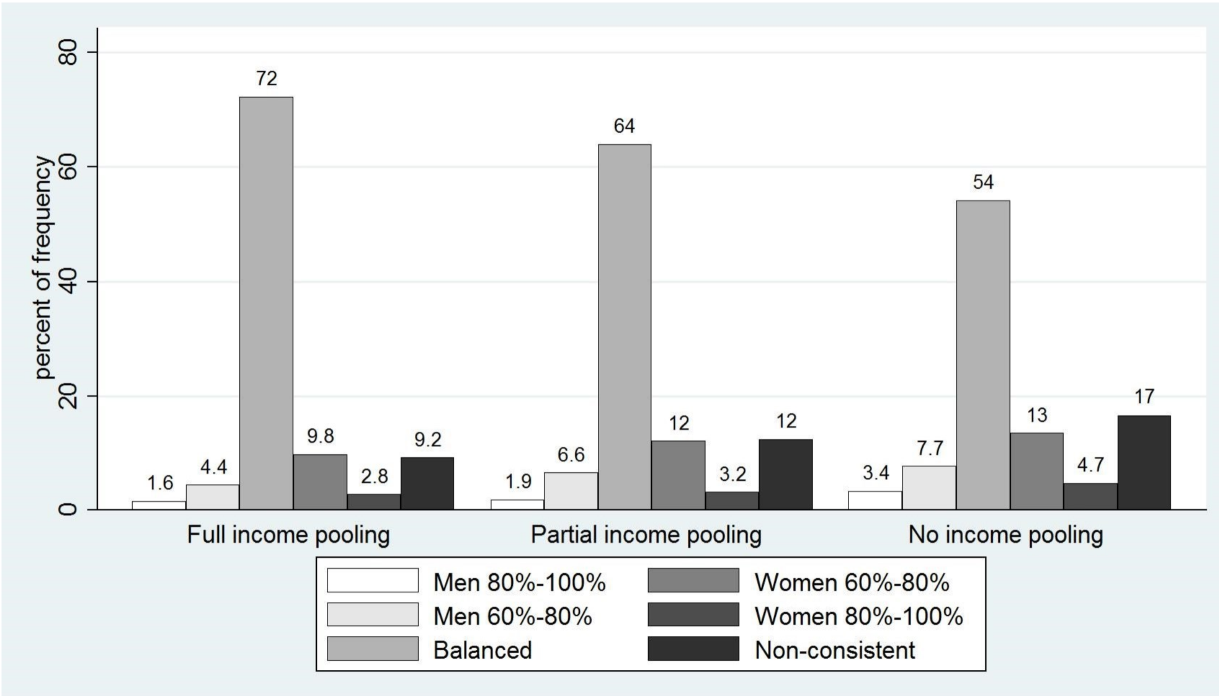

Fig. 1: Shares of families by decision-making mode across the regimes of family finances

1.4.3. Instrumental variables estimation

The robustness of the established correlations might indicate that the established relationship between the household mode of decision-making and material deprivation is actually causal. Accordingly, to make this claim with more certainty, a more refined estimation technique is needed. The ideal way to proceed would be using data from randomized control trial when treatment changes the mode of decision-making. The sample that I analyze, however, does not contain such data. One possibility in this case is to use regional-level variables as instruments for the balanced mode of decision-making. Despite the fact that this requires strong identification assumptions, it is widely used in the literature (Moffit, 2005). To conduct the instrumental variables estimation, I modify Equation 1 so that the RHS contains only one dummy for balanced decision-making rather than four dummies for remaining modes. The OLS estimation has at least two potential problems. First, the estimate can be biased towards zero due to errors in reporting the control allocation. Second, the estimate can be biased due to reverse causality. The direction of the bias depends on the specific mechanism of the reverse causality. For example, if one spouse takes predominant control when the other partner has made a decision with negative consequences, the OLS estimation would yield too a large coefficient in absolute value. If, however, one of spouses joins the decision-making process when the partner makes an error of judgment so that they start deciding in a more balanced way, the OLS estimation would yield too small a coefficient in absolute value. Regarding the instrumental variables, I use two: the rate of involvement of 4-year-olds in formal childcare and the gender differences in weekly work hours.

These instruments can be warranted by previous findings in the literature. First, accessibility of institutional pre-school childcare has been found to have significant consequences for female activity status. A couple of recent studies mention this (Bičáková, 2016; Bičáková and Kalíšková, 2016; Kalíšková, 2016). In turn, the activity statuses of household members are related to the allocation of control in households. This is confirmed empirically (Schneebaum and Mader, 2013; Yusof and Duasa, 2010) and is a departure point in theoretical research (Lundberg and Pollak, 1993). Specifically, when a woman is not employed outside the household, her contribution to the family income is not likely to be high. That is why she does not have enough control over family finances, in particular, over decisions on use of savings and borrowing money. At the same time, she often has more control over purchases for children and everyday shopping (Schneebaum and Mader, 2013). This resembles a theoretical separate spheres equilibrium demonstrated by (Lundberg and Pollak, 1993). Further, the longer a woman stays on maternity leave due to inaccessible childcare the more likely such allocation of control is to arise, and once established, it tends to carry on (Basu, 2006). Therefore, the accessibility of pre-school childcare (which does not appear to change quickly) should influence the intra-household allocation of decision-making. It is hard to see any other channel through which it could influence household material deprivation measures once activity status and incomes of household members are controlled for[19].

Second, besides the activity status, employment opportunities also matter for allocation of control in family. For example, (Morrill and Pabilonia, 2015) show that increasing national unemployment rates reduce time spent together also in households with both partners working due

[19] It might also happen that accessibility of childcare influences women's employment which, in turn, influences household material deprivation. This fact could threaten identification if either the correlation between childcare accessibility and women's employment is too high or when the women's employment is itself endogenous. The former situation is not the case while the latter is likely to attenuate the estimates and not undermine the conclusions. Specifically, the most likely unobservable household-level variable affecting women's employment is women's household productivity. That is, women who are more productive at home are less likely be employed. When there are more employment opportunities due to accessible childcare, more such women become employed. That is why the negative effect of employment on household material deprivation for a subsample of women having access to childcare will be lower than for the entire population (because when they become employed their household loses more in terms of the household production). This difference will translate into smaller estimated effect of balanced decision making under the proposed IV estimation. In other words, the IV estimated coefficient on balanced control will be a lower bound estimate. Therefore, the identification assumption about exogeneity of childcare accessibility comes at no cost for the conclusions of this study.

to rearrangement of working schedules (workers accept less convenient hours to preserve the job). In turn, less time spent together leads to greater specialization of partners (Mansour and McKinnish, 2013) so that they do not decide together, but rather individually on matters of their responsibility. In this case, gender difference in weekly workhours will reflect the difference in spouses' ability to arrange their schedules in order to participate in home-focused activity. The one for whom it is more difficult is likely to be more preoccupied with his or her job and to be less able to participate in household decisions. Moreover, employment perspectives influence spouses' expectation of income in the case of divorce. This is an important factor in intra-household bargaining (Lundberg and Pollak, 1993). Those who have worse employment perspectives will be less likely to resort to divorce in the case of household conflict and, thus, more likely to concede more control to their spouses. The gender gap in weekly work hours will reflect the difference in spouses' outside-marriage options and willingness to concede control.

Both variables are measured across NUTS 2 regions. The results of 2SLS estimation are shown in column 3 of Table 1. The 2SLS coefficient is statistically significant and has the same sign as the OLS coefficient, but its absolute value is much larger. This is consistent with correcting for the attenuation bias and the reverse causality from worse outcomes to more balanced control, as in the second of the two mechanisms explained above. The two instruments used stood up to several tests. First, the value of F-statistic for a test of their joint significance in the first-stage equation is 20. Second, the Hausman test shows that the balanced-control variable is not exogenous at a 10\% confidence level. Third, the overidentifying restriction test statistic is not significant (p-value is 0.97). Fourth, when the reduced form model is estimated on the sub-sample of households with balanced decision-making only, the proposed instruments lose their statistical significance as expected. Fifth, when the reduced form model is estimated on the subsample of single-headed households, the instruments also lose statistical significance. Thus, the 2SLS estimation result supports the claim that balanced decision-making in households reduces material deprivation.

1.5 Conclusions

In this paper, we establish strong and robust correlations between direct measures of each adult household member's control over specific decisions and household-level measures of material deprivation. More individualized control by either partner is closely related to more frequently reported material deprivation if other conditions remain constant. This result holds for primary control by both men and women. This fact questions the interpretation of a positive impact of increases in female income as outcomes of increase in female control, because increases in female incomes change male control as well. In such cases, effects of windfall incomes handed to women documented in the literature could be actually driven by increases in balanced decision-making. This notion is supported by several studies which report increases in balanced decision making as a result of windfall incomes going to women. The negative relationship persists when we control for a number of possible confounders and use IV estimation, suggesting that it is likely to be causal. Possible detailed mechanisms at work behind the observed pattern are partially accounted for by an autonomous regime of family finance management when one partner cannot afford to cover an agreed upon part of common expenditures. However, more research is needed to understand precisely how this relationship works. Detailed information on the routines of managing household finances would be helpful in this case. It is possible that joint expenditures are akin to joint projects (Evertsson and Nyman, 2014). In this case, a lack of cooperation on household decisions could be interpreted as a lack of cooperation on a joint project, which is known to be a very common cause of projects failure. As for the policy implications of the established results, it turns out that individual-specific (usually female-specific) targeting of social assistance, which is frequently highlighted in the literature (Attanasio and Lechene, 2010; Schady and Rosero, 2007; De la Briére and Rawlings, 2006), perhaps should not be unambiguously preferred to household-specific targeting.

1.6 Appendix

Table 6 The Mode of Decision-making and Household Economic Outcomes

Modes of decision-making relative to balanced	Household outcomes						
	Arrears on mortgage payments	Arrears on utility bills	Arrears on hire purchase installments	Inability to afford one week annual holiday	Inability to afford a proper diet	Inability to face unexpected financial expenses	Inability to make ends meet
A woman takes control over 4-5 decisions	0.056 (0.012)***	0.046 (0.007)***	0.053 (0.017)***	0.066 (0.010)***	0.034 (0.007)***	0.070 (0.010)***	0.049 (0.009)***
A man takes control over 4-5 decisions	0.026 (0.014)**	0.031 (0.008)***	0.029 (0.022)	0.039 (0.012)***	0.013 (0.008)*	0.021 (0.012)*	-0.002 (0.012)

Table 6 continued

Modes of decision-making relative to balanced	Household outcomes						
	Arrears on mortgage payments	Arrears on utility bills	Arrears on hire purchase installments	Inability to afford one week annual holiday	Inability to afford a proper diet	Inability to face unexpected financial expenses	Inability to make ends meet
A woman takes control over 2-3 decisions	0.011 (0.005)***	0.024 (0.003)***	0.024 (0.008)***	0.023 (0.005)***	0.012 (0.003)***	0.024 (0.005)***	0.011 (0.005)**
A man takes control over 2-3 decisions	0.014 (0.007)**	0.025 (0.005)***	0.030 (0.012)***	0.005 (0.007)	0.005 (0.005)	-0.007 (0.007)	-0.037 (0.007)***
A woman responds the household questionnaire	-0.003 (0.003)	0.006 (0.002)***	-0.002 (0.005)	0.041 (0.003)***	0.002 (0.002)	0.041 (0.003)***	0.029 (0.003)***
Woman's income share	-.004 (0.007)	-.006 (0.004)	-.015 (0.008)*	-.006 (0.007)	.000 (0.004)	.000 (0.007)	.010 (0.006)*
N of hhds.	27,987	74,733	16,379	75,073	75,089	75,063	75,059
R^2 adj.	0.67	0.58	0.14	0.26	0.14	0.20	0.30

Notes: Each column corresponds to one regression. Each of the four decision-control variables takes value 1 or 0. A decision-control variable takes value 1 if both partners report one of them having a dominant role in making a corresponding decision. A decision-control variable takes value 0 if both partners report balanced participation in making a corresponding decision. Robust standard errors are in parentheses. * denotes statistical significance at 10 percent, ** at 5 percent, and *** at 1 percent.

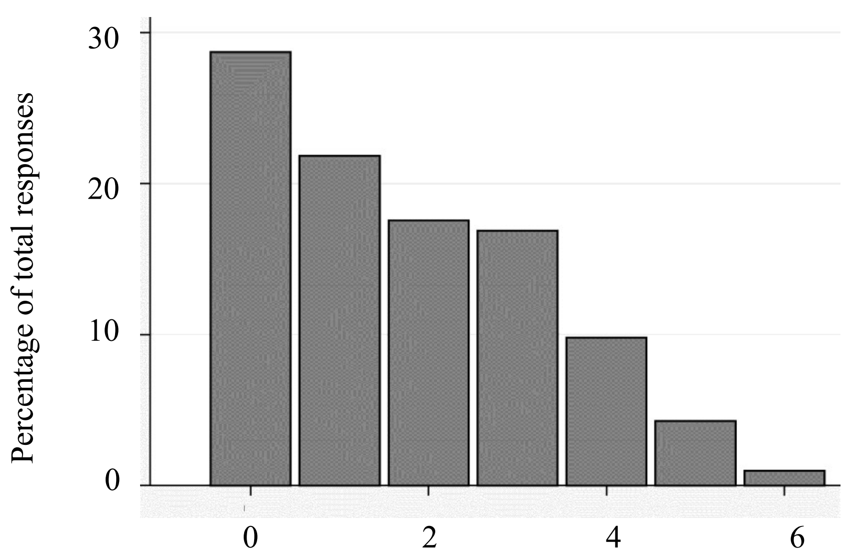

Fig. 2 Percentages of households by reported number of occurring material deprivation conditions

2. Parental Gender Preference in the Balkans and Scandinavia: Gender Bias or Differential Costs?

2.1 Introduction

The impact of the gender of the first-born child on the number of children in a family has been repeatedly observed in many countries. We confirm son preference using the parity-three progression method applied to a pooled EU-SILC 2004-2015 cross-sectional sample from four Balkan countries: Bulgaria, Croatia, Slovenia, and the Republic of Serbia.[20] We also confirm daughter preference for three Scandinavian countries, i.e. Denmark, Norway, and Sweden, which had been identified previously by (Andersson, Hank, Ronsen, & Vikat, 2006) and (Hank & Kohler, 2000). Two possible causes of gender preference considered in the literature are parental bias in favor of one or another gender and different costs of raising sons and daughters (Ben-Porath & Welch, 1976; Lundberg S. , 2005). This paper aims to identify which of the two is more prevailent in Balkan and Scandinavian countries. Each explanation implies a distinctive relationship between the gender of children and the allocation of household resources. We test between the two explanations by checking which relationships hold for the household-level data.

We find that Balkan households with more female children replace furniture less frequently than households with fewer female children. Moreover, in households with more female children, mothers report a lower ability to spend on themselves. Additionally, for Balkan countries we find no difference in parental investment in male and female children and no impact of the gender composition of children on the ability to make ends meet or the minimum amount of money needed to make ends meet. We argue, based on earlier studies, that these findings are consistent with the gender bias explanation and not with the differential expenses explanation. For Scandinavian countries we find no impact of the gender composition of children on replacing furniture or on consumption of other household public goods, and we find significantly larger parental investment in households with more female children. Moreover, we do not find a systematic impact of the gender of their children on parental consumption. We argue based on conclusions in (Lundberg S. , 2005) and (Lundberg & Rose, 2003), that these findings are not

[20] These countries are covered by EU-SILC and had the highest SIGI son bias component in Europe according to OECD: https://www.genderindex.org/ranking/sonbias/.

consistent with the gender bias explanation, but are in line with the differential expenses explanation. Supplementary analyses of the top-income-decile sub-sample and of cross-country relationships between gender preference, parental investment, and conventional measures of gender equality support our argument.

2.2 Literature Review

The evidence on the impact of parental gender preference pertains to developing economies (Barcellos, Carvalho, & Lleras-Muney, 2014; Jiang, Li, & Sanchez-Barricarte, 2016; Altindag, 2015) and developed economies (Dahl & Moretti, 2008; Andersson, Hank, Ronsen, & Vikat, 2006; Michael S & Morgan, 2002). Authors attribute this impact to parental preference for one gender of children. In developing economies, parents usually have more children (progress to higher parities) when their firstborn is a daughter (Filmer, Friedman, & Schady, 2009; Arnold, 1992). The interpretation of such behavior is that they have a son preference, so they continue producing children until they reach a desired number of sons or the upper limit of the desired family size. At the same time, in some developed economies, parents also exhibit son preference (Dahl & Moretti, 2008; Choi & Hwang, 2015), but daughter preference in others (Andersson, Hank, Ronsen, & Vikat, 2006; Brockman, 2001).[21] Consequences of parental gender preference have mostly been researched for developing economies. The main consequence is that girls, on average, have more siblings and receive a lower share of household resources (Vogl, 2013; Jensen, 2003; Basu & De Jong, 2010). Consequences include shorter breastfeeding period for girls (Jayachandan & Kuziemko, 2011), worse health and nutritional status of girls (Arnold, Sex Preference and Its Demographic and Health Implications, 1992), and biased sex ratios (e.g. (Jayachandran S. , 2017; Guilmoto & Duthe, 2013)). In more developed economies, (Kippen, Evans, & Gray, 2007) and (Dahl & Moretti, The Demand for Sons, 2008) argue that a son preference increases fertility in Australia and the US. (Edlund, 1999) demonstrates theoretically that gender preference combined with availability of gender selection technology[22] could lead to a female "under-class", because poorer parents would prefer daughters and richer ones prefer sons (Trivers & Willard, 1973). Other

[21] Sandstrom and Vikstrom (2015) provide evidence for the existence of son preference in Germany in the second half of the 19th century, which faded later, while Outram (2015) finds evidence for son preference in Edwardian England.
[22] Such technologies may include infanticide, sex-selective abortion, or poorer health care.

possible consequences in the setting developed by (Edlund, 1999) are existence of a “backlog” of unmarried men (Gupta, 2014) with ensuing consequences, such as polygamy (Economist, 2018; Seidl, 1995). That is because changes in socio-demographic structure lead to “adoption of adequate institutions” (Seidl, 1995), which is evident, e.g., in the falling marriage-market value of young men in across commuting zones in the US (Autor, Dorn, & Hanson, 2017) accompanied by rising acceptance of polygamy in the US recorded by Gallup pollster (Economist, 2018). Any policy that mitigates the effects of gender preferences would need to take into account the causes behind the observed behavior (Lundberg S. , 2005). Two possible causes considered in the literature are parental bias in favor of some gender and different costs of raising sons and daughters (Ben-Porath & Welch, 1976; Lundberg S. , 2005). This paper studies which of the two is more prevalent across selected European countries. Each explanation implies a distinctive relationship between the gender of children and the allocation of household resources. We test between the two explanations by checking which relationships hold for the household-level data.

Regarding parental gender bias, there are several definitions in the economic literature. The first is that some gender brings more direct utility or has a utility premium. This definition is used in most papers on the subject (e.g., (Jayachandan & Kuziemko, 2011; Dahl & Moretti, 2008; Yoon, 2006)). Authors either forgo explaining possible mechanisms behind the gender bias and take the gender-biased fertility behavior as their starting point (Jayachandran & Kuziemko, 2011) or explain it by a predilection (Dahl & Moretti, 2008) or cultural and biological factors (Yoon, 2006). Scholars in demographic and sociological literature elaborate more and offer further explanations for gender bias, such as expansion of the self, affiliation, stimulation, accomplishment or social comparison (Hank, 2007), as well as the emotional value of children (Sandstrom & Vikstrom, 2015). Moreover, mothers and fathers can perceive the extent to which sons and daughters fulfill these expectations differently (Hank, 2007). Finally, the definition proposed in (Lundberg S. , 2005) encompasses the aforementioned elements, stating that ‘parents have child-gender preferences if the marginal value of an additional male child differs, *ceteris paribus*, from the marginal value of an additional female child, or if the marginal utility of increments in boy quality is not equal to the marginal utility of girl quality.’ Here ‘quality’ means child outcomes that are outputs of a household production process in which inputs are parental time and market goods and services. This definition incorporates two different cases. In the first case, parental valuation of the gender of children or accompanying outcomes does not relate to parental outlays on children (beyond providing for a minimal subsistence level). In the second case children outcomes are closely dependent on parental inputs until these inputs reach significant values. The second case is not consistent with previous definitions since the gender is not preferred *per se*, but because it makes the technology of producing a certain quality cheaper, i.e. it is only one means of reaching a specific discrete end. In this paper we understand gender bias as in the first case, as the predilection for such gender-intrinsic characteristics of children that depend neither in extent nor intensity on parental outlays. Therefore, the gender bias does not mean that parents prefer a son or daughter because s/he will bring higher returns to their investments. Instead, it means that they want a child of a particular gender because of its predetermined characteristics.[23] If gender bias, as we understand it, were the only determinant of the family size connected to the gender of children, two relationships for household outcomes would likely hold. First, parents who desire boys but have a girl or vice versa anticipate having more children in the future and might start saving or work more to support a larger family (Barcellos, Carvalho, & Lleras-Muney, 2014). Second, parents who have children of a preferred gender should spend more on household public goods, because their marriage is more stable, as the preferred gender child generates higher surplus (Lundberg S. , Sons, Daughters, and Parental Behavior, 2005). Therefore, in countries where firstborns of the preferred gender have, on average, fewer siblings, parents of firstborns of this gender should work less, save less, and spend more on household public goods. Moreover, if sons directly increase the utility of fathers, then a standard bargaining model of the household predicts a shift of household resources from fathers to mothers. This redistribution could be observable as increased leisure among mothers of sons, or increased consumption of private commodities typically consumed by women (Lundberg & Rose, 2003).

Turning to the difference in costs of raising sons and daughters, the literature considers two cases.[24] First, when sons and daughters have constant, albeit not necessarily equal, cost. An assumption of constant costs of children is taken in much, if not most, of the applied studies on the

[23] Figures A4 and A5 in the Appendix illustrate a more detailed explanation of the difference between the gender bias and the cost difference.

[24] While we test for the difference in costs of children, it is actually the difference in "prices" of sons and daughters in which we are primarily interested. The price of a child is the commitment of resources required to raise a child of given 'quality'. At the same time, the cost of a child is a measure of the actual amount of resources committed to child-raising (Bradbury, 2004). Thus, the cost of children is deliberately chosen by parents and, in principle, is measurable. In most theoretical models related to the subject, which do not allow for variable quality of children (Dahl & Moretti, 2008; Leung, 1991), the price of children is constant and equals cost, because parents are assumed to pay the full life-time prices of children once they are born or the per-period price every period.

topic (van Praag & Warnaar, 1997), which frequently calculate so-called *normative budgets*.[25] Nominal expenditures or normative budgets, however, do not equal total expenditures on children. The latter also include time costs of childcare and exclude the value of children's contribution to household production. Still, the monetary outlays per se do not fully reflect the quality of inputs. Another issue is whether parents take into account net flow of future transfers from children (Blacklow, 2002; Adda, Dustmann, & Stevens, 2016). Available empirical evidence suggests that parental expectations are important for parental spending (Hao & Yeung, 2015). These assumptions describe a case when parents rely upon some rules of thumb when deciding about outlays on children. These rules of thumb, in turn, are based on perceptions about optimal living arrangements in a given society in a given time (Kornrich & Furstenberg, 2007). Then, to calculate the gender difference in costs of children, studies in the literature employs two methods. The first, the Rothbarth method, measures the adult-good equivalent of children cost. This method, unlike normative budgets or discretionary equivalence scales (van Praag & Warnaar, 1997), is theoretically plausible (Deaton & Muellbauer, 1986). This method estimates a difference in consumption of private adult goods or leisure time (Bradbury, 2004) between parents having first-born sons and first-born daughters. The second method measures gender difference in costs of children relying upon the subjective scales method (Leyden approach) proposed and substantiated in (van Praag & Warnaar, 1997).

The second case considered in the literature regarding the difference in costs of sons and daughters is when the cost consists of fixed and variable components. This case is captured by models like those in, e.g., (Galor, 2011; de la Croix & Doepke, Inequality and Growth: Why Differential Fertility Matters, 2003); and (Hazan & Zoabi, 2015). In this case, either fixed (one-time costs) or variable components (price of human capital) of the child cost could differ. Differences in fixed costs are revealed by parental outlays during the early childhood years. At the same time, differences in the variable component are revealed by the differences in availability of parental investment items. Children with lower human capital costs will receive higher outlays and

[25] For example, the U.S. Department of Agriculture (USDA) has provided estimates of expenditures on children since 1960. Forensic economists use these figures in wrongful death and birth cases, as well as in child support cases (Lino & Carlson, 2010). The constant cost of children is also assumed in, e.g., (Dahl & Moretti, 2008; Hazan & Zoabi, 2015; Leung, 1991; Sienaert, 2008; Bojer, 2002); and (Raurich & Seegmuller, 2017).

have less siblings due to substitution of quality for quantity (Galor, 2011; Aaronson, Lange, & Mazumder, 2014).[26]

We use a set of home items as measures of parental investment (Cunha, Heckman, & Schennach, 2010), as proxy variables for parental outlays on children. Parents buy more of such items when they bring more parental utility per unit of expense for a gender and will have fewer children after having a firstborn of that gender. In our analysis, we assume the costs of children per the latter case, when the costs include of fixed and variable components, so that it is consistent with economic theory. Thus, if the differential cost explanation is true, parents of a child of the more expensive gender should have fewer children thereafter, spend less on themselves (both parents), spend less on adult public goods, and spend more on children. Moreover, parents of a ``more expensive'' child should report higher sums needed to make ends meet. However, if the gender bias explanation specified above is correct, parents will report lower sums, because they should spend more on household public goods which exhibit economies of scale in consumption. The restriction on child age applied in our analysis ensures that child's financial contribution to a household does not confound the estimates obtained.

2.3 Data and Sample Statistics

We use a data set from the European Union Survey of Income and Living Conditions (EU-SILC) for 2004 - 2015. The data set is collected annually by national statistical offices in cooperation with Eurostat from nationally representative samples, which covered the EU-28 and several non-EU countries in 2015. In 2004, only 15 countries were covered by the survey. Our analysis is based on data from four Balkan countries and three Scandinavian countries. The Balkan countries are Bulgaria, Croatia, Slovenia, and the Republic of Serbia.[27] The Scandinavian countries are Denmark, Norway, and Sweden.[28] A primary goal of EU-SILC is to collect cross-sectional and

[26] It could be that either items for some gender are cheaper or produce more parental utility through child human capital. One more case is possible when items generate little human capital and thus, more of them are bought (i.e., the demand for them is inelastic). However, it is unlikely that this effect would be stronger in countries with more gender-equivalent attitudes as Figure A3 in the Appendix shows.

[27] These are Slavic-speaking Balkan countries covered by EU-SILC survey. When we extend the set of Balkan countries to include Greece and Romania, the estimates of gender preference do not change qualitatively.

[28] These groupings of countries have been frequently used in previous studies. For instance, (Estrin & Uvalic, 2014) use a similar grouping of Balkan countries and conduct regression analyses on the pooled sample of data under the assumption that regression parameters do not differ between these countries. Similarly, (Baranowska-Rataj, 2016)

longitudinal microdata using a rotational four-year panel scheme on income, poverty, social exclusion, and living conditions (Eurostat, 2017). The longitudinal component is not used in our research. The reference population in EU-SILC includes all private households and their current members residing in the territory of the respective countries at the time of data collection. All household members are surveyed, but only those aged 16 and older are interviewed. The data set for each year after 2004 consists of two groups of variables: primary and secondary. Primary variables are collected annually. Secondary variables are collected approximately every five years in so-called ad-hoc modules. A variable may include information at the household or personal level about specific topics. The primary variables convey information on household demographic composition, incomes, living conditions, and labor market activity. The secondary variables used in the current research were collected in 2009, 2010, and 2013-2015 in ad-hoc modules on material deprivation. These secondary variables contain more in-depth information on material deprivation in the household than the annual primary variables. Eurostat calculates cross-sectional household and individual weights to correct for non-random sampling and non-responses (Eurostat, 2015).[29]

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and (Ragan, 2013) use the mentioned grouping of Scandinavian countries. Both studies assume that the considered characteristics of those economies (model parameters) are similar across Scandinavian countries. In a similar vein, (Filmer, Friedman, & Schady, 2009) pool HNS data into six sub-samples by parts of the world and assume no difference in parameters between countries within groups.

[29] More detailed information on the dataset is available at the following link http://ec.europa.eu/eurostat/web/microdata/overview

Table 1: Descriptive statistics - demographics and labor market information.

	Balkan countries				Scandinavian countries			
	All families		Married couples		All families		Marr red couples	
Selected household characteristics	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference
Living without	0.114	-0.005			0.106	0.003		
father	(0.318)	(0.003)			(0.308)	(0.003)		
Number of children	1.855	0.047	1.872	0.046	1.996	0.004	2.016	0.005
	(1.047)	(0.010)***	(0.996)	(0.010)***	(0.839)	(0.007)	(0.832)	(0.007)
First-born girl	0.481		0.484		0.487		0.487	
	(0.500)		(0.500)		(0.500)		(0.500)	
Age of mother	26.44	0.035	27.06	0.03	28.91	0.04	29.09	0.07
at first birth[a]	(7.35)	(0.07)	(5.36)	(0.06)	(5.36)	(0.04)	(4.79)	(0.04)
Age of mother	34.68	0.001	35.4	0.0009	37.44	0.002	3T.56	0.05
	(7.40)	(0.07)	(6.12)	(0.06)	(6.26)	(0.05)	(5.80)	(0.05)
Mother having	0.178	-0.005	0.195	-0.007	0.363	0.002	0.402	0.004
tertiary erlucation	(0.382)	(0.003)	(0.396)	(0.004)	(0.481)	(0.004)	(0.490)	(0.004)
Mother employed	0.606	0.000	0.650	-0.003	0.746	-0.001	0.821	0.001
	(0.489)	(0.004)	(0.477)	(0.005)	(0.435)	(0.003)	(0.383)	(0.003)
Mother's weekly	28.100	-0.106	28.738	-0.159	27.985	0.341	28.001	0.340
hours of work	(19.424)	(0.183)	(19.141)	(0.186)	(14.82)	(0.122)*	(14.851)	(0.123)**
Father employed			0.805	0.004			0.924	-0.005
			(0.396)	(0.004)			(0.264)	(0.002)
Father's weekly			37.156	0.082			37.810	-0.165
hoursofwork			(16.689)	(0.162)			(12.762)	(0.106)
Household disposable	20,469	265.421	20,982732	214.079	64,070.609	325.596	65,957259	450.271
income(euros)	(15,431.683)	(141.036)	(15,550.905)	(150.036)	(57,680.462)	(44L583)	(59,032.599)	(483.734)
Living in urban area	0.137	0.003	0.131	0.002	0.347	0.000	0.341	0.002
	(0.344)	(0.003)	(0.337)	(0.003)	(0.476)	(0.004)	(0.474)	(0.004)
Ownership of	0.767	-0.003	0.763	-0.004	0.920	-0.005	0.929	-0.004
accomodation	(0.423)	(0.004)	(0.425)	(0.004)	(0.271)	(0.002)**	(0.257)	(0.002)**
N of hhds	24,951		22,027		28,352		25,294	

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

Note: The statistics were calculated for the subsample of intact families with children. Columns one and three show means and standard deviations while columns two and four show differences between mean values for girls versus boys. Values in parentheses in even numbered columns correspond to t-test standard errors
[a] These statistics were calculated only for families in which the mother is younger than 41 and older than 17 and had her first child at the age of 16 or older and child ages are in the range 0-12.

Table 2: Availability of selected items in the home environment for girls and boys

	Balkan countries				Scandinavian countries			
	All families		Married couples		All families		Marrred couples	
Dependent variables	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference
Household-level material condition characteristics[a]								
Amount of money needed to make ends meet	1,486.629 (830.649)	11.577 (7.705)	1,507.179 (831.329)	8.119 (8.141)	4,725.007 (13,992.615)	44.569 (115.330)	4,823.201 (14,112.091)	84.074 (122.862)
Ability to make ends meet	0.215 (0.411)	0.004 (0.004)	0.225 (0.418)	0.002 (0.004)	0.776 (0.417)	0.002 (0.003)	0.798 (0.402)	0.006 (0.003)**
Replacing worn-out furniture	0.278 (0.448)	-0.008 (0.007)	0.290 (0.454)	-0.005 (0.007)	0.888 (0.316)	-0.006 (0.005)	0.905 (0.293)	-0.004 (0.005)
Adult-specific material condition characteristics[b]								
Ability to spend a small amount of money on oneself (women)	0.522 (0.500)	0.000 (0.007)	0.533 (0.499)	0.000 (0.007)	0.399 (0.490)	0.017 (0.007)**	0.381 (0.486)	0.016 (0.007)**
Ability to spend a small amount of money on oneself (men)	0.540 (0.498)	0.003 (0.007)	0.573 (0.495)	0.005 (0.007)	0.383 (0.486)	-0.013* (0.007)	0.408 (0.492)	-0.014** (0.007)
Availability of two pairs of properly fitting shoes (women)	0.615 (0.487)	-0.003 (0.007)	0.627 (0.484)	-0.001 (0.007)	0.437 (0.496)	0.017 (0.007)**	0.411 (0.492)	0.015 (0.007)**
Availability of two pairs of properly fitting shoes (men)	0.597	-0.000	0.634	0.002	0.408	-0.012*	0.435	-0.012*
Replace worn-out	(0.490)	(0.007)	(0.482)	(0.007)	(0.492)	(0.007)	(0.496)	(0.007)
clothes (women)	0.540	0.003	0.555	0.004	0.415	0.013	0.393	0.011
	(0.498)	(0.007)	(0.497)	(0.007)	(0.493)	(0.007)*	(0.488)	(0.007)
Replace worn-out								

clothes (men)	0.535	0.002	0.571	0.003	0.396	-0.012	0.422	-0.013
	(0.499)	(0.007)	(0.495)	(0.007)	(0.489)	(0.007)*	(0.494)	(0.007)*
Get together with friends/family	0.552	0.004	0.565	0.005	0.429	0.018	0.405	0.017
at least once a month (women)	(0.497)	(0.007)	(0.496)	(0.007)	(0.495)	(0.007)**	(0.491)	(0.007)**
Get together with friends/family at least once a month (men)	0.551	-0.002	0.586	-0.001	0.401	-0.016**	0.426	-0.016**
	(0.497)	(0.007)	(0.493)	(0.007)	(0.490)	(0.007)	(0.495)	(0.007)
Regularly participate in a leisure activity (women)	0.233	-0.006	0.244	-0.006	0.322	0.010	0.307	0.010
	(0.423)	(0.005)	(0.430)	(0.006)	(0.468)	(0.007)	(0.462)	(0.007)
Regularly participate in	0.254	-0.005	0.276	-0.006	0.317	-0.009	0.338	-0.009
a leisure activity (men)	(0.435)	(0.006)	(0.447)	(0.006)	(0.465)	(0.007)	(0.473)	(0.007)
Children home environment items [d]								
Replacing worn-out clothes	0.822	-0.007	0.843	-0.005	0.986	0.000	0.987	0.001
Two pairs of	(0.382)	(0.007)	(0.363)	(0.007)	(0.118)	(0.003)	(0.113)	(0.003)
properly fitting shoes	0.845	0.006	0.867	0.007	0.983	0.000	0.986	-0.002
Fresh fruits and	(0.362)	(0.006)	(0.340)	(0.006)	(0.128)	(0.003)	(0.118)	(0.003)
vegetables once a day	0.866	-0.010	0.885	-0.006	0.982	-0.003	0.983	-0.003
One meal with	(0.341)	(0.006)	(0.319)	(0.006)	(0.134)	(0.003)	(0.127)	(0.003)
fish, chicken or meat (or vegetarian equivalent) at least once a day	0.842	-0.003	0.862	-0.001	0.988	0.003	0.989	0.002
Books at home suitable	(0.365)	(0.006)	(0.345)	(0.006)	(0.108)	(0.002)	(0.103)	(0.002)
for children's ages	0.844	0.006	0.863	0.009	0.983	0.006	0.984	0.005
Outdoor leisure	(0.363)	(0.006)	(0.344)	(0.006)	(0.131)	(0.003)	(0.126)	(0.003)
equipment	0.821	-0.001	0.841	0.004	0.987	-0.002	0.990	-0.003
	(0.383)	(0.007)	(0.366)	(0.007)	(0.112)	(0.002)	(0.102)	(0.002)
Indoor games	0.875	-0.002	0.891	0.000	0.995	-0.000	0.996	-0.001
	(0.331)	(0.006)	(0.312)	(0.006)	(0.072)	(0.001)	(0.066)	(0.001)
Regular leisure activity	0.503	0.010	0.518	0.009	0.776	0.017	0.779	0.019
	(0.500)	(0.009)	(0.500)	(0.009)	(0.417)	(0.008)**	(0.415)	(0.009)**
Celebrations on	0.867	-0.002	0.884	0.000	0.981	0.001	0.983	0.002
special occasions	(0.339)	(0.006)	(0.320)	(0.006)	(0.137)	(0.003)	(0.129)	(0.003)
Invite friends								

over to play	0.790 (0.408)	0.002 (0.007)	0.807 (0.395)	0.005 (0.007)	0.959 (0.198)	0.002 (0.004)	0.959 (0.198)	0.002 (0.004)

tistics were calculated for the subsample of intact families with children. Columns one and three provide means and standard deviations while columns two and four provide differences between mean values for girls versus boys. Values in parentheses in even numbered columns correspond to t test standard errors.

a The amount of money needed to make ends meet and the ability to make ends meet are primary variable collected annually while replacing worn-out furniture was collected in ad-hoc modules in years 2009 and 2013-2015.

b Adult-specific material condition characteristics were collected in ad-hoc modules in years 2009 and 2013-2015.

c This variable and the three next variables were collected in 2010.

d Children’s home environment items were collected in ad-hoc modules in 2009 and 2013-2015.

Two main advantages of this data set are important for our analysis. First, it contains information on the age and gender of all adults and their children living in the household. Second, the ad-hoc modules from 2009, 2010 and 2013-2015 contain detailed information on material deprivation of adults and children in the household. There are also two significant drawbacks. First, not all children might be present in the household at the time of the survey for some reason (e.g., because they study or work elsewhere). We cannot be sure that the firstborn child lives in the household. Second, the information on material deprivation of children is available only for all children in the household together and not for each child separately.[30] To correct for the first drawback, we limit our sample to data where we can claim with high certainty that the firstborn child is still in the household. Specifically, following other studies in the literature (Dahl & Moretti, 2008; Karbownik & Myck, 2017; Ananat & Michaels, 2008), we limit the analysis to mothers aged between 18 and 40 who had their first child at the age of 16 or older. The limit for the age of the oldest child is set at 12 years.[31] Our calculated sex-ratio for firstborns is 1.057, close to the commonly accepted value of 1.06 (Grech, Savona-Ventura, & Vassallo-Agius, 2002).[32] To correct for the second drawback, we connect the material condition of children in the household to the gender composition of children (i.e., the share and presence of daughters among children are instrumented with a dummy for the first child being a girl).

Since the gender of children influences household composition, we limit our analysis, for the most part, to married and cohabiting couples. Table 1 contains descriptive statistics for selected household socio-demographic characteristics separately for all families and for cohabiting couples. Table 2 presents descriptive statistics on variables characterizing different aspects of the household material condition. We use variables in Table 2 as dependent variables

[30] For example, an answer to a question: "Do children have books at home suitable for their age?" should be "Yes" if all children have books and "No" if at least one child does not have books.

[31] The sample bias is likely to be very small because the minimal age of leaving school in all European countries is above 16. Other studies (Dahl & Moretti, 2008; Karbownik & Myck, 2017) use the threshold of 12 years. (Karbownik & Myck, 2017) use this threshold since it corresponds to the grouping of expenditure information on clothing. We need broader range of ages because we aim to control for the age of children (which was not done in other studies). (Dahl & Moretti, The Demand for Sons, 2008) find the 12-year cutoff conservative while (Ichino, Lindström, & Viviano, 2011) and (Ananat & Michaels, 2008) use 15-year and 17-year cutoffs respectively. Importantly, our chosen threshold ensures that child earnings do not confound our results because this threshold is below the compulsory schooling age in all European countries. At the same time, when we estimate our models on the entire sample, the estimates preserve signs and statistical significance but reduce in size.

[32] This fact also suggests that gender-selective abortion or gender difference in early childhood treatment should be too rare to show up in the data.

and variables in Table 1 as covariates. Amongst adult and household material deprivation characteristics, Table 2 also presents the average frequency of the ten home environment items for children along with girl-boy differences. One can readily see that girls are more likely to have books, have an opportunity to invite friends, and to host celebrations. These differences are small, however, and hover around one percent of the standard deviation of the corresponding items. This is less than reported by (Xu, 2016). The largest differences between all families and intact, i.e. married and cohabiting, families appear to be in food and clothing. Specifically, the girl-boy difference is significant for all families but disappears for intact families. This could be explained by more limited resources of non-intact families.[33] Otherwise, the intact families do not appear to differ systematically from all families along the considered characteristics. That supports our decision to focus the analysis on intact families.

2.4 Empirical Analysis

Our paper tests between two alternative explanations for parental gender preference. Each has different implications for household economic behavior. The gender bias hypothesis implies that households with a first-born child of the desired gender save less (Barcellos, Carvalho, & Lleras-Muney, 2014) and spend more on household public goods (Lundberg S. , Sons, Daughters, and Parental Behavior, 2005). We do not have a direct measure of household savings, so we use the capacity to face unexpected financial expenditures as a proxy variable. Here we rely on the intuitively appealing assumption that greater savings mean higher capacity to deal with unexpected expenditures. Regarding the measure of household public goods, we use replacing worn-out furniture. Other measures, like good nutrition and quality of leisure or availability of appliances and cars, are more likely to have a direct impact on child well-being and thus might be not invariant to the gender of children. Moreover, more household public goods available should also result in greater ability to make ends meet and less money needed to make ends meet, because the consumption of household public goods exhibit returns to scale. At the same time, the differential costs hypothesis implies that parents of a child of the preferred gender (i.e., of the more expensive gender, resulting in fewer additional or total births) work

[33] This result is consistent with the Trivers-Willard hypothesis. Further exploration of this question is beyond the scope of this study.

more, save less, and spend less on adult public goods. Parents of more expensive child should report lower ability to make ends meet together with higher sums needed to make ends meet.

One possible way to test our hypotheses is to compare families with different child gender composition. This is the approach taken by (Bogan, 2013), who explores the relationship between household financial assets market participation and the gender of children. Specifically, Bogan estimates a regression in which the dependent variable is stock or bond ownership while the explanatory variables are dummies for only female and only male children or a proportion of female children in the household. However, since the explanatory variable in both specifications (the dummies for same-gender children and share of daughters) might be decided by households and, thus, may be endogenous, therefore, such estimates cannot be taken as evidence of a causal relationship between the variables in question.[34] Similarly, in the case of our analysis, more daughter-preferring parents could also derive more utility from the well-being of their children and, thus, tend to create better material conditions for them. To address these concerns, we use the gender of the firstborn as the explanatory variable. Our identification strategy is to assume that the gender of the firstborn is randomly determined. This assumption has been made in other studies that use the gender of firstborns as an instrument for household characteristics. Some of these characteristics are: the bargaining power of women in China (Li & Wu, 2011), the number of children in a family (Dahl & Moretti, 2008), the occurrence of divorce (Bedard & Deschenes, 2005; Ananat & Michaels, 2008), and the area of accommodations (Dujardin & Goffette-Nagot, 2009).[35]

To test our hypotheses, we proceed in three steps. First, we estimate gender preference across European countries using the third-parity method. Second, we verify the validity of the gender bias explanation by testing its aforementioned implications in daughter-preferring countries and son-preferring countries respectively. That is, in countries where we observe daughter preference, parents of a first-born daughter should be less capable of dealing with unexpected financial expenditures (because they save less), spend less on themselves, be more likely to replace worn-out furniture, be more able to make ends meet, and need less money to make ends meet. The same predictions should hold for parents of first-born sons in son-

[34] More daughter-preferring families, for instance, are more likely to have all daughters: they self-select into having all daughters because son-preferring families who have only daughters are more likely to continue having more children until they have a son. At the same time, daughter-preferring families could be less risk-averse and, consequently, more inclined to participation in financial assets market.

[35] The second Appendix subsection describes additional considerations and reservations about using this instrument.

preferring countries. Third, we verify the validity of the differential costs explanation by testing its implications in daughter- and son-preferring countries. We do this in two stages. In the first stage, we assume constant costs (prices) of sons and daughters(e.g., (Dahl & Moretti, 2008; Jayachandran & Kuziemko, 2011; Leung, 1991)). In the second stage, we relax this assumption and, instead, assume the cost of children to consist of two components, fixed and variable (Galor, 2011; Aaronson, Lange, & Mazumder, 2014; de la Croix & Doepke, 2003). In the latter case, we determine whether the difference is driven by the fixed or the variable component.

The baseline specification of the regression model takes the following form:

$$y_i = \beta(\mathrm{Firstchildgirl})_i + \alpha X_i + \epsilon_i \quad (1)$$

where y_i stands for either the progressing to parity three (having three children) or a children's material conditions indicator for a household i and X_i is a vector of household i socio-demographic and economic characteristics. The *First child girl* indicator takes value 1 if the first-born child was a girl and 0 if a boy. Within a given country, the residual values, ϵ_i, can be correlated. The specific set of variables that make up X depends on the particular regression equation specification. We use this form at each of the three steps of the hypothesis testing.

To test for gender preference, we put the third parity progression on the left-hand side. Progression to the third parity has been the most widely used indicator in the literature to test for gender preference. There are two main reasons it is better to use parity-three progression rather than parity-two progression to measure the gender preference. First, it is likely that the desire for a gender-mix of children (to have at least one son and one daughter) coexists with the gender bias towards one gender (Dahl & Moretti, 2008). In that case, parents who have bias towards any gender will progress to parity two independently of the gender of their firstborn. That is why the causal effect of the gender of the firstborn on the progression to parity two is not likely to be significant. The second reason is that first-born twins would distort the estimates for parity two progression. Still, we also report second parity progression and total number of children. We choose covariates that have been used in similar studies: gender of the first two children, cubic polynomial of mother's age, squared polynomial of mother's age at first birth, length of cohabitation of spouses, mother's education, father's education, mother's employment, father's employment, household disposable income, and living in an urban area (Dahl & Moretti, 2008; Hank & Kohler, 2000; Haughton & Haughton, 1998; Larsen, Chung, & Das Gupta, 1998; Basu & De Jong, 2010). We include higher degree polynomials in the mother's age to account for the conclusions reached by (Yamaguchi & Ferguson, 1995), who argue that the probability of giving birth for women is lower at a younger age, then increases, and then again decreases. Such a relationship is best fit by the third-degree polynomial in age. Finally, we include the family's occupied accommodation tenure along with year and country dummies. We estimate the models with OLS as do most other studies on the subject, because this method yields consistent estimates of the coefficient on the dummy for the gender of the firstborn. The linear probability model may be an especially good choice because right-hand side variables are mostly dummies (of 23 covariates only 7 are continuous variables) and the unboundedness problem is less acute in this case (Wooldridge, 2002, p. 456). Nevertheless, we also run Probit estimations to check for consistency with the OLS-based results.[36] Since we expect observations not to be iid, but correlated within countries, we cluster the standard errors at the country level.

In regard to testing for differential costs of sons and daughters, we assume that the cost of children consists of two components, constant (one-time cost) and the variable (outlays on human capital). Researchers commonly use this assumption in models featuring parental investment in children. The fixed component of child cost primarily represents the time cost of rearing children during infancy, whereas the variable component represents parental expenditures on child human capital. Thus, if our analysis finds that parental outlays on children of one gender are larger, there could be two causes: larger one-time costs or lower price of human capital (parental discounted utility derived from child human capital). The mechanism behind the second cause is that of substitution of quality for quantity of children. For example, parents may spend more on daughter "quality" and have fewer children after daughters. If this explanation is true, daughters in daughter-preferring countries should receive more parental investments. One measure of parental investments used in the literature[37] is the availability of conditions and items at home which are necessary for normal child development (Cunha, Heckman, & Schennach, 2010; Todd & Wolpin, 2007; Juhn, Rubinstein, & Zuppann, 2015).[38] The expected effects of the first-born daughter are systematically presented in Table A7. We use the 2009/2010 and 2013-2015 EU-SILC data on availability of such items in households to test if daughters tend to have better material conditions in daughter-preferring

[36] The Probit estimates correspond to OLS estimates in terms of impact direction and statistical significance.
[37] The most common measure is the years of schooling conditional on household income.
[38] These variables are described in more detail in the Appendix subsection

countries and sons, respectively, in son-preferring countries. Under this assumption, parents having a child of the more expensive gender, in addition to having a lower progression ratio, should also have lower expenditures on private consumption and household public goods, be less able to deal with unexpected financial expenditures, be less able to make ends meet, and need more money to make ends meet. The ability to make ends meet is measured by a binary variable taking value 1 when a household is able to make ends meet. The aforementioned predictions follow from the fact that they have less of financial means left after making outlays on children than parents with a child of the cheaper gender. The method of measuring the cost of children through comparing the amount of money needed to make ends meet reported by families having children of different gender was proposed and used by (van Praag & Warnaar, 1997).[39]

2.5 Results

2.5.1 Estimates of parental gender preferences for children

Table 3 presents coefficients on the gender of the firstborn for different specifications of the dependent variable in Equation 1 estimated on data from Balkan countries. These results resemble those obtained by (Dahl & Moretti, 2008) in the US. The first column indicates that families in which the first child is a girl end up having more children than families in which the first child is a boy, although the difference is not significant. In line with the expectations discussed above, the impact of the gender of the firstborn on progression to parity two in column (2) is much less statistically significant and lower than the impact on progression to

[39] One way to conceptually unify the aforementioned gender differences in costs of raising children is to interpret them as differences in constraints associated with raising sons and daughters (Lundberg S. , 2005). In that case, intact families have comparative advantage in raising a child of a preferred gender provided that, in vast majority of cases, mothers have custody of children (Dahl & Moretti, 2008). Specifically, in the case of father's comparative advantage in raising sons, intact families have a comparative advantage in raising sons over single-mother headed families. In the case of differential costs, an intact family also has a comparative advantage in raising a child with lower price of human capital because it has more resources at its disposal thanks to economy of scale, even if the total nominal incomes of family members remain the same whether it is intact or not. Here the economy of scale means that the opportunity cost of raising a child of a gender with more costly human capital (in terms of utility forgone if the child were gender with lower cost of human capital) increases with family income. This is true, for instance, when a marginal return to parental investment in children is constantly higher for one gender. The proposed unification of child gender differences in costs of children together with the previous reasoning has several implications for household allocation, which are presented in Table A5.

parity three and has much lower percent effect. The numbers in column (3) show the probability of having three or more children is 1.3 percent higher when the first child is a girl, which is an order of magnitude higher than the result obtained by (Dahl & Moretti, 2008) in the US. In other words, first-born girl families are 17% more likely to have three or more children compared to first-born boy families. We also find significant positive effects for the probability of four or more and five or more children when the first-born child is a girl. The positive effect of the first-born daughter on progression to parity three has also been found by (Filmer, Friedman, & Schady, 2009) in Central Asia, South Asia, Middle East, and North Africa. It is this result which is most commonly interpreted in the literature as a manifestation of son preference.

Table 3: The firstborn-child gender and family size in the Balkans.

	Breakdown by number of children				
	(1) Total number of children	(2) Two or more children	(3) Three or more children	(4) Four or more children	(5) Five or more children
First-born child being a girl	0.030 (0.010)***	-0.001 (0.008)	0.013 (0.005)***	0.011 (0.002)***	.003 (0.001)***
Controls	Yes	Yes	Yes	Yes	Yes
First boy baseline	1.57	0.483	0.077	0.011	0.002
Percent effect	0.019	-0.002	0.17	0.18	0.50
R-sq	0.26	0.39	0.13	.04	.02
Observations	19,807	-	-	-	-

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

Notes: S.E. are given in parentheses and are clustered at the country level. Estimates are based on the 2004-2015 EU-SILC samples for Bulgaria, Croatia, Serbian Republic, and Slovenia. The sample consists of households formed by one cohabiting couple, their children, and, occasionally, other relatives. The mother of children in the household is younger than 41 and older than 17 and had her first child at the age of 16 or older, and children's ages are in the range 0–12. The estimation method used is weighted OLS with probability weights reflecting non-random sampling within and between countries. The table presents estimated effects of the firstborn being a daughter compared with the baseline case of the firstborn being a son. The effect is a ratio of the estimated OLS coefficient on the firstborn's gender dummy to the baseline value of the dependent variable. The dependent variables are the total number of children and a set of binary indicators for specific numbers of children. The control variables, besides the gender of the firstborn, are: the dummy for a first-born daughter, gender of the first two children, cubic polynomial of mother's age, squared polynomial of mother's age at first birth, length of cohabitation of spouses, mother's education, father's education, mother's employment, father's employment, household disposable income, living in urban area, tenure status, year and country dummies.

Table 4: The firstborn-child gender and family size in Scandinavia.

	Breakdown by number of children				
	(1)	(2)	(3)	(4)	(5)
	Total number of children	Two or more children	Three or more children	Four or more children	Five or more children
First-born child being a girl	-0.009	0.002	-0.013	0.002	0.0002
	(0.010)	(0.006)	(0.005)***	(0.002)	(0.0002)
Controls	Yes	Yes	Yes	Yes	Yes
First boy baseline	1.82	0.64	0.16	0.02	0.003
Percent effect	0.005	0.003	0.08	0.1	0.07
R-sq	0.29	0.38	0.22	0.05	0.01
Observations	25,227	-	-	-	-

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

Notes: Estimates are based on the 2004-2015 EU-SILC samples for Denmark, Norway, and Sweden. For details about sampling and estimates presentation, see the notes under Table 3

Table 4 presents estimates analogous to those in Table 3, but for Scandinavian countries. These results are notably different from the results for Balkan countries. First, the impact of a first-born daughter on progression to parity three in column (3) is negative and statistically significant. Despite having a similar absolute value, the effect is half of the Balkan effect, because a larger share of Scandinavian families progresses to parity three. Second, impacts of a first-born daughter on the total number of children and on progression to other parities have small absolute magnitudes and are not statistically significant. The parity three progression results in column (3) are in line with those obtained by (Andersson, Hank, Ronsen, & Vikat, 2006), for each of the Scandinavian countries separately. This alone suggests that gender bias is probably not the only mechanism behind these results, because they would then also be similar for progressions to higher parities.

In Appendix Figure A6 and Table A3, we present gender preferences across EU countries. Our results are broadly consistent with those obtained in previous literature (Hank & Kohler, 2000). We also attempt to evaluate how our results would differ if there were no family disruptions caused by child gender, which is frequently reported in the literature (see, e.g., (Lundberg S. , 2005) for a review). Estimates obtained for that counterfactual scenario, however, do not differ qualitatively and do not differ much quantitatively from those reported here. Absence of rank correlations between the country-level impacts of the firstborn's gender on progression to parity two and parity three suggests different driving causes behind these impacts.

2.5.2 Testing between the gender bias and differential cost explanations

The gender bias explanation implies two patterns in household-level allocations.[40] First, expenditures on household public goods should be higher when the firstborn is of the preferred gender (Lundberg S. , 2005). Specifically, if a son increases marital surplus more than a daughter, then the birth of a son reduces the probability of divorce and increases the incentive of partners to invest further in the marriage, i.e. the family as a whole (Lundberg & Rose, 2003). Second, saving should be less, because parents anticipate fewer births in the future (Barcellos, Carvalho, & Lleras-Muney, 2014). To test the first implication, we estimate the impact of a first-born daughter on the frequency of replacing furniture in the household. (Lundberg & Rose, 2003) consider furniture an important household public good along with automobiles and housing conditions as proxies for housing expenditures. Spending on automobiles and housing, however, can be directly influenced by child gender composition. As (Lundberg & Rose, 2003) note, observed differences in housing spending could be influenced the need for greater space to accommodate the size and activity of sons or the desire for a higher quality neighborhood to reduce the probability of risky behavior by boys or probability of crimes against girls. Concerning automobiles, having one might make more sense when a couple has sons, who are possibly expected to be more handy with cars and, for whose socialization, access to an automobile can often be more important. Meanwhile, expenditures on furniture do not appear to be directly influenced by the gender of children. Column (1) of Table 5 contains estimates of the firstborn's gender impact on replacement of worn-out furniture in the household. The negative and statistically significant estimate for Balkan countries confirms the prediction from

[40] Table A6 shows the results of testing between the gender bias and the gender-specific constraints explanations for the Balkans and Scandinavia separately. The rounded cells in Table A6 indicate that data corroborate the gender bias explanation for the Balkans and the differential constraints explanation for Scandinavia. The ensuing discussion clarifies which specific form the differential constraints are most likely to take. The current section further explains that it is the gender difference in price of children's human capital.

the son bias explanation of the observed gender preference. To support the daughter bias explanation for Scandinavian countries, the estimate would need to be positive, which is not the case. Regarding the prediction that savings should be less in families with a firstborn of the preferred gender, we test that by estimating the impact of the firstborn's gender on the ability to deal with unexpected expenditures, assuming that households with higher savings are more likely to respond positively to this question, the estimate should be positive in Balkan countries and negative in Scandinavian countries. The estimates obtained in column (2), however, are small in magnitude and not statistically significant. For Balkan countries, this result could be reconciled with son preference by the fact that common savings are also a household public good and respond positively to the arrival of a child of the preferred gender, countering the negative effect of reduction in expected number of children.

Table 5: Impact of a first-born girl on availability of household public goods across countries grouped by observed gender preference

Countries	(1) Replacing worn-out furniture	(2) Capacity to deal with unexpected expenditures	(3) Ability to make ends meet	(4) Lowest monthly income to make ends meet	(5) Availability of home items
Balkan	-0.020	0.0019	0.008	-0.671	0.017
	(0.011) *	(0.007)	(0.006)	(9.848)	(.015)
Scandinavian	-0.006	0.005	0.005	152.7	0.035
	(0.007)	(0.005)	(0.004)	(142.2)	(0.018) **

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

Notes: The standard errors of estimates on sub-samples for Balkan and Scandinavian countries are clustered at the country level. Estimates in the columns (2), (6), and (7) are based on the 2009 and 2013-2015 EU-SILC ad-hoc modules, while the estimates in the remaining columns are based on the 2004-2015 EU-SILC primary modules. The sample consists of households formed by one cohabiting couple, their children, and, occasionally, other relatives. The mother of children in the household is younger than 41 and older than 17 and had her first child at the age of 16 or older, and children's ages are in the range 0–12. The estimation method used is weighted OLS with probability weights reflecting non-random sampling within and between countries. Dependent variables for columns (1) and (3)-(7) are binary indicators taking value 1 when a household has the indicated condition and value 0 otherwise. The table presents estimated effects of the firstborn being a daughter compared with the baseline case of the firstborn being a son. Other control variables are: the dummy for a first-born daughter, gender of the first two children, cubic polynomial of mother's age, squared polynomial of mother's age at first birth, length of cohabitation of spouses, mother's education, father's education, mother's employment, father's employment, household disposable income, living in urban area, tenure status, year and country dummies.

Table 6: Impact of a first-born girl on employment and consumption of mothers and fathers in the Balkans

	(1) Being employed	(2) Weekly hours of work	(3) Ability to spend on oneself	(4) Two pairs of shoes	(5) Replacing clothes	(6) Get together with friends	(7) Regualr leisure activity
Mothers	-0.011	-0.369	-0.0233	-0.007	-0.007	0.006	0.032
	(0.006) *	(0.265)	(0.0117) **	(0.01)	(0.01)	(0.01)	(0.024)
Fathers	-0.006	-0.328	0.004	-0.002	-0.003	0.002	0.049
	(0.005)	(0.228)	0.011	(0.01)	(0.01)	(0.01)	(0.024) **

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

Notes: The standard errors of estimates on the sub-sample for Balkan countries are clustered at the country level. For details on sampling and estimation see the note under Table 5.

Higher expenditure on household public goods may also be consistent with the comparative advantage a father has in raising sons, i.e. the so called "technology" explanation, according to (Dahl & Moretti, The Demand for Sons, 2008). The gender bias and technology explanations have different implications for consumption patterns of fathers and mothers. The gender bias explanation suggests lower consumption of mothers of daughters while the technology explanation implies it to be higher. Specifically, if sons directly increase the utility of fathers, then a standard bargaining model of the household predicts a shift of household resources from fathers to mothers. This redistribution could be observable through lower consumption of private commodities by mothers of daughters. The negative impact of the mother's ability to spend on herself in Balkan countries in column (3) of Table 6 is in line with the gender bias explanation. In addition, two more facts hold for intrahousehold allocations in Balkan countries. First, mothers of daughters are less likely to be employed. Second, fathers of daughters report more time spent on leisure. The first could be explained by self-selection into unemployment of mothers whose comparative advantage in raising daughters results in an even greater opportunity cost than for similar mothers of sons (otherwise, first-born daughters would also negatively impact the intensive margin of mother's employment). Still, such self-selection of mothers into employment would not undermine our results, because the ``technology'' explanation implies lower progression to parity three when fathers have a sufficiently high comparative advantage in raising sons and a sufficiently wide wage gap in favor of men (Gugl & Welling, 2012). Despite the existence of a wide gender wage gap, our estimates do not support the existence of a sizable comparative advantage of fathers in raising sons in the Balkans, which would be evident from fewer hours of work and higher personal consumption reported by fathers with first-born sons, as explained earlier. Finally, the fact that fathers have more leisure could be explained by longer hours of housework done by daughters.[41] Thus, the obtained results are consistent with the gender bias explanation for Balkan countries.

Table 7: Impact of a first-born girl on employment and consumption of mothers and fathers in Scandinavia

	(1) Being employed	(2) Weekly hours of work	(3) Ability to spend on oneself	(4) Two pairs of shoes	(5) Replacing clothes	(6) Get together with friends	(7) Regualr leisure activity
Mothers	0.005	0.439	-0.002	0.005	0.001	0.013	-0.060
	(0.005)	(0.185) **	(0.008)	(0.006)	(0.007)	(0.007) *	(0.031) **
Fathers	-0.007	-0.357	0.004	(0.004)	0.005	0.0003	-0.032
	(0.003)	(0.156) **	(0.008)	(0.006)	(0.007)	(0.007)	(0.030)

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

Notes: The standard errors of estimates on the sub-sample for Scandinavian countries are clustered at the country level. For details on sampling and estimation see the note under Table 5.

For Scandinavian countries, there is no firstborn gender effect on either furniture replacement nor the ability to deal with unexpected expenditures (the first two columns of Table 5). Moreover, the estimates of the firstborn's gender impact on parental consumption in Table 7 do not differ between fathers and mothers, similar to the unconditional means in Table 3, which would be in line with parental comparative advantage. In other words, the parental consumption difference between fathers and mothers, as the difference between fathers and mothers of the unconditional means in Table 3, both point to the parental comparative advantage explanation.[42] That is because mothers of sons should redirect household resources

[41] This is true for the 2010 ad-hoc sample from Romania and Bulgaria. The question about hours of housework was included in the 2010 EU-SILC ad-hoc module. However, since this was an optional question, and national statistical agencies chose whether or not to include it in the survey presented to their residents, this data is available only for 10 EU countries.
[42] It cannot be the main driving cause for the observed gender preference in Scandinavia, because the gender wage gap should be in favor of women (Gugl & Welling, 2012) and that is not the case. Still, this result is consistent with a comparative advantage of intact families with daughters in producing “child quality”.

to fathers to keep them in the family due to their important role in raising sons (Lundberg S. , 2005). At the same time, estimates of the impacts on ability of mothers to meet with friends and family and to have regular leisure activity do not contradict the gender bias explanation *per se*. However, the estimated impacts on father's consumption should be positive according to the gender bias explanation and it is not. Fathers of daughters work fewer hours but they do not redirect that time to leisure. Moreover, the fewer hours worked by fathers of daughters is not likely to drive the observed daughter preference because similar effects were found in the US and West Germany (Lundber & Rose, 2002; Choi, Joesch, & Lundberg, 2008), which exhibit son preference (Dahl & Moretti, 2008; Hank & Kohler, 2000). All in all, the data does not support the gender bias explanation for Scandinavian countries.

The differential cost hypothesis is not confirmed by household-level estimates for Balkan countries. There are no statistically significant results in the last three columns Table 3 for Balkan countries. Moreover, if expenditures on sons were higher, explaining the lower progression after a first-born son, parents of daughters would have more resources to spend on themselves. This is in contrast with the negative impact of the first-born daughter on private expenditure of mothers in column (3) of Table 5.

The Scandinavian results do show the expected higher outlays on daughters consistent with the differential cost explanation. Households with first-born daughters are more likely to have the entire set of ten important children consumption items. However, neither ability to make ends meet nor the minimum amount of money to make ends meet depend on the gender of the firstborn. Nevertheless, for the top income decile, the minimum amount of money needed to make ends meet is larger for families with a first-born daughter.[43] Mothers of daughters appear to more frequently forgo regular leisure activity and substitute it with apparently less costly socialization through meeting with friends and family. Moreover, more hours worked by mothers of daughters suggest that they are willing to substitute leisure for outlays on daughters. At the same time, fathers of daughters tend to work less than fathers of sons. When (Lundberg & Rose, 2003) reported a similar effect for fathers from the US, they offered an explanation based on the son bias idea but did not formally test it. Our testing, however, does not support the son bias explanation. Furthermore, Norwegian data indicates that paternal leave has more pronounced positive effects for daughters than sons (Cools, Fiva, & Kirkeboen,

[43] The argument why this should be true is developed in the Appendix (the Figure A4 illustrates this idea).

2015). That could be a reason fathers in Scandinavian countries substitute time spent on work for time spent on children (rather than leisure).[44] All in all, the differential expenses explanation of daughter preference in Scandinavian countries is supported by the data.

Figure A2 and Figure A3 show cross-country relationships between the gender preference, the gender gap in parental investment, and conventional measures of gender equality. These relationships are in line with our previous points.[45]

2.6 Conclusion

We find evidence that parental gender preferences in different countries are caused by different reasons. In Balkan countries, the observed son preference is likely driven by gender bias towards sons, that plays the major role. In Scandinavian countries, the observed daughter-preference is likely driven by a lower cost of daughter quality (which incorporates gender-specific personal characteristics and their usefulness for parents). To measure the effect of the gender difference in the cost of children precisely we would need to observe its random variation. Evidence of a lower price for female human capital is most pronounced in more gender-equal societies in line with trends of institutional change in modern societies in favor of women (Baumeister, 2011). If this is not compensated by policies that reduces the price of human capital for sons in less well-off families, the consequences mentioned in (Edlund, 1999) and (Seidl, 1995) might be realized.

2.7 Appendix

2.7.1 The distinction between the gender bias and differential costs concepts

In the literature, there is neither a clear-cut definition of what we have designated as gender bias nor a conventional term for labeling it. In some cases, gender bias is readily recognizable.

[44] Examining data from detailed time-use surveys could shed more light on this issue.
[45] Specifically, Figure A3 shows that daughters tend to receive greater parental investment in countries with higher indicators of gender equality. This suggests that child household items for daughters are either cheaper or more useful in more gender-equal countries. Both situations are consistent with a lower price of child human capital for daughters in countries with greater gender equality. Meanwhile, if the gender equality indicators at hand reflect a degree of gender bias and gender bias drives parental gender preference, Figure A2 should show negative relationships, which is not the case.

For example, (Arnold, Choe, & Roy, 1998) assert that some Indian parents prefer sons for reasons connected with religious beliefs and kinship descent, whereas (Jacobsen, Møller, & Engholm, 1999) argue that women's need of companionship leads to daughter preference in Denmark. Characteristics, like continuing the family name or providing the same-gender companionship to parents, are intrinsically pertinent to the gender of a child and their utility does not directly depend on the parental outlays on children. Preferences for such characteristics are captured by the first part of Lundberg's (2005) definition, because a son has a greater marginal value in the first case and a daughter in the second. This understanding is consistent with other previously provided definitions. In other situations, the gender bias is less recognizable. One possible example is the case of a man who wants a son because the boy may be a player in his favorite soccer team. Yet, the father cannot do much to bring this about beyond encouraging him or taking him to a local soccer academy. Had this man had a daughter instead of a son, he would likely have done not much less for her physical development. Similarly, parents might want a daughter, because she can become a soprano singer. These examples are captured by the second part of the aforementioned definition. That is, the man values a son's soccer skills more than a daughter's, because they increase his chances of him becoming a player in a father's favorite team. While in the second example, parents value a daughter's singing skills more than a son's, because the son's soprano will eventually disappear. In both cases, parents would not need to invest much provided the children have sufficient aptitude (parental time and tuition at a soccer academy or music school). A common feature of the these examples is the absence of a close relationship between the parental investment of time and market goods on one side and child quality (desired characteristics) on the other side beyond some relatively low level of investment.

An alternative example could be parents that want a household member to know a foreign language. One way to proceed is to have a child that would learn that language. On average, it would be cheaper with a daughter because girls are known to be better at picking up foreign languages (Burman, Bitan, & Booth, 2008). Here, the more parents invest in a child's language learning, the better the result (hours with tutors, educational trips abroad, etc.). Keeping other things equal, these parents are likely to invest significantly more in the daughter's language learning, because of greater marginal returns on their investment. We understand such situations as cases of differences in costs of children.

2.7.2 Considerations about using the gender of the firstborn as the instrumental variable

Some authors claim that the gender of the firstborn is not random. For example, (Norberg, 2004) reports that children who were conceived when their mother was living with a partner were 14 per cent more likely to be boys than siblings conceived when the parents were living apart. This finding aligns with the falling gender ratio in a set of industrialized countries (Davis, Gottlieb, & Stampnitzky, 1998). One possible explanation for these findings is the evolutionary advantage of species that can adjust the gender ratio of offspring in response to changes in conditions affecting the relative reproductive success of males and females (Trivers & Willard, 1973). Furthermore, the wealthiest individuals in societies tend to have sons born more frequently (Cameron & Dalerum, 2009). To address these concerns we repeat our analysis on the sample of partners cohabiting at the time when the firstborn arrived, control for the country fixed-effects, and repeat the analysis after dropping the top 1% of wealthiest households in each country from the sample.[46]

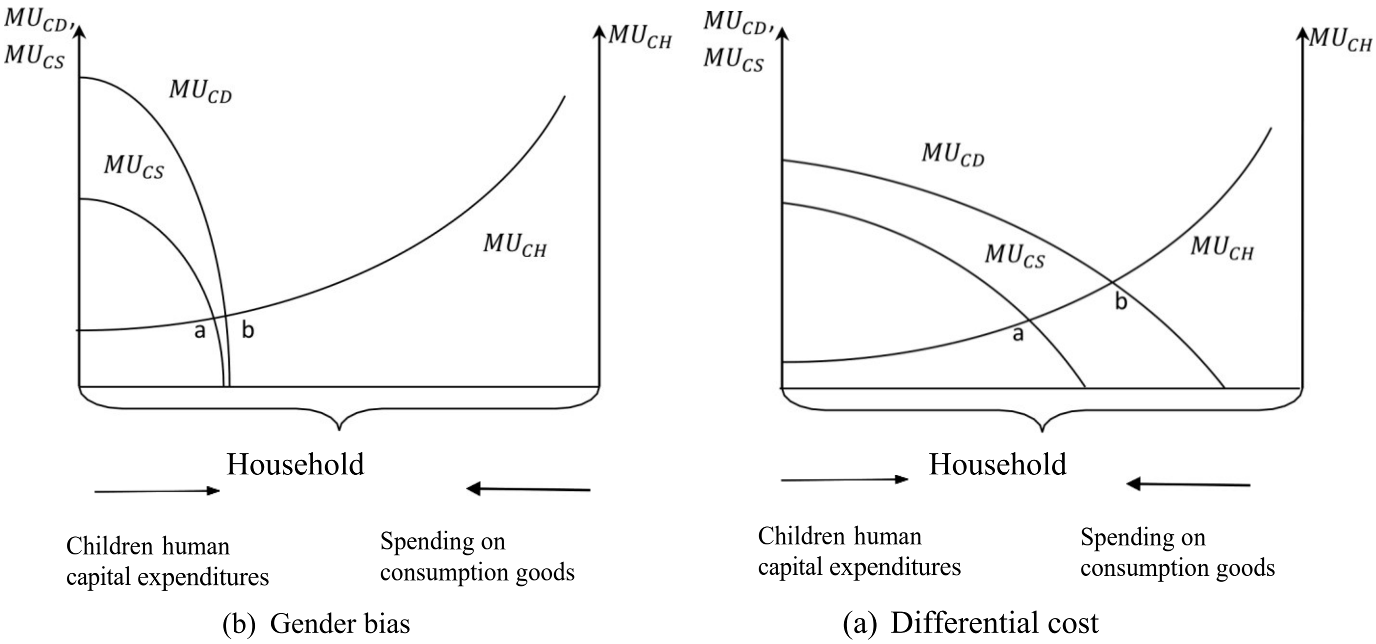


Figure A1: Graphical distinction between cases of gender bias and differential costs

Notes: The graphs show marginal parental utilities of human capital expenditures on children, MU_{CS} and MU_{CD}, together with accompanying marginal utility of household consumption expenditures, MU_{CH}. The underlying

[46] One study (Kanazawa, 2007) reports that physically more attractive parents are significantly more likely to have a daughter. We are not aware of other studies confirming this finding.

unitary household model is assumed. On the horizontal axis is human capital expenditure, household marginal utility is on the vertical axis. Marginal utility of household consumption increases as expenditures on household consumption decrease, which occurs along the horizontal axis as human capital expenditures on children increase. On the left graph, marginal utilities of human capital expenditures on children plummet quickly and parental investments are low and do not differ significantly between genders. At the same time, the difference in parental utility derived from children of different genders is significant. This is a graphically depicted example of gender bias. On the right graph, the marginal utility of investment in a child of some gender is notably larger along a broad range of possible investment volumes. The optimal volumes of investment differ considerably between children of different genders. This is a graphically depicted example of differential cost.

At the same time, the gender of the firstborn might impact marital stability (Lundberg & Rose, 2003; Mammen, 2008; Lundberg, McLanahan, & Rose, Child Gender and Father Involvement in Fragile Families, 2007), family size (Hank & Kohler, 2000; Angrist & Evans, 1998), and parental time allocation (Lundber & Rose, 2002; Lindstrom, 2012; Choi, Joesch, & Lundberg, 2008). This makes "exclusion restrictions a priori unpersuasive" (Lundberg S. , Sons, Daughters, and Parental Behavior, 2005). To solve this problem, we focus our analysis on the sample of intact families, instrument the number of children with twin-births, and argue that the impact of the gender of the firstborn on parental employment does not notably alter our estimates or their statistical significance.

The documented impact of the gender of firstborns on parental employment differs across countries. For example, a first-born son increases a father's work hours in the US by 3% of the mean male work hours more than for fathers with a first-born daughter (Lundber & Rose, 2002). However, (Pabilonia & Ward-Batts, 2007) find $\frac{1}{3}$ of the same effect and not at a statistically significant level. An even larger effect, almost 5% of mean annual male work hours, was found in West Germany (Choi, Joesch, & Lundberg, 2008). Meanwhile, (Ichino, Lindström, & Viviano, 2011) find a negative impact of a first-born son on a mother's working hours and employment in the US, UK, and Italy. This is still smaller than the previously mentioned effect for fathers and hovers across the countries at around 1% of the mean. (Lindstrom, 2012) finds that a first-born son increases parental leave by 0.6 days (1.5 %) and decreases maternity leave by a similar amount.

In our analysis, we do find that the gender of a firstborn affects the employment status of mothers. However, we do not find an effect on their work hours or on father's employment status or work hours. The negative effect of a first-born son on a mother's employment is approximately 1% of mean female employment. This is in line with previously reported estimates from the literature. However, when we multiply this effect on employment with its coefficient, the final effect on the variable of interest is by an order of magnitude smaller than the direct effect of the first-born gender variable. That is why, following (Karbownik & Myck, 2017), we believe that the impact on employment does not undermine our estimates of interest and se we keep the employment status and workload of parents as covariates.

2.7.3 A description of the material deprivation measures

The EU-SILC ad-hoc modules on material deprivation from 2009 and 2014 each contain thirteen questions about the availability of child items and amenities (the module from 2009 contained questions on 22 items, but the recent module was reduced). Each question corresponds to a variable that indicates the presence of a specific item or amenity. Specifically, the variables are: replace worn-out clothes; two pairs of properly fitting shoes; fresh fruit and vegetables once a day; one meal with fish, chicken, or meat (or vegetarian equivalent) at least once a day; books at home suitable for children's age; outdoor leisure equipment; indoor games; regular leisure activity; celebrations on special occasions; invite friends home to play and eat from time to time; participate in school trips and school events that cost money; suitable place to study or do homework; and go on holiday away from home at least 1 week per year. In our analysis, we primarily only use the first ten questions, because they are available for nearly all children in the sample, while the last three are available only for school-age children. These questions do not completely correspond to the questions from other surveys on material conditions of children that have been analyzed in the literature, e.g., NLSY79-CS HOME-SF module (Cunha, Heckman, & Schennach, 2010; Todd & Wolpin, 2007; Juhn, Rubinstein, & Zuppann, 2015) and PISA-2000 (Xu, 2016). Those surveys are more extensive. Instead, the ten questions we consider largely overlap with the resources-spent and time-with-child subcomponents defined by (Juhn, Rubinstein, & Zuppann, 2015) based on the NLSY79 survey. For instance, all questions in the resources-spent and some questions from the time-with-child subcomponents of (Juhn, Rubinstein, & Zuppann, 2015) are contained in EU-SILC ad-hoc modules from 2009 and 2014. All in all, the EU-SILC ad-hoc modules considered here could be seen as extended versions of the two subcomponents mentioned above, and since elements in these two subcomponents were highly correlated with child development (Bradley & Caldwell, 1980; Bradley & Caldwell, 1981; Bradley & Caldwell, 1984) and strongly influencing it (Cunha, Heckman, & Schennach, 2010), the raw score of the EU-SILC ad-hoc modules should also be correlated with and have an impact upon child development. Furthermore, the responses from the PISA-2000 survey analyzed by (Xu, 2016) contain more detailed information, but correspond directly with the EU-SILC questions on participating in regular leisure activity, availability of a suitable place to study, and having books at home. Xu argues that precisely those items are important for a child's adult outcomes and supports the point by referring to multiple related studies.

To test for a gender-gap in children's material conditions at home, we use five alternative dependent variables in equation 1 for measuring material condition. The first is a pure sum of the binary indicators of the presence of the first ten material conditions listed in the previous paragraph. This sum corresponds to the so-called HOME index used in the literature. One problem with this variable is susceptibility to monotonous transformations, also known as the scaling problem (Bond & Lang, 2013). Another problem is that all the items in that dependent variable are assigned equal weights in summation, which means that those with larger variance contribute more to the estimated effect. We attempt to overcome these problems by constructing four other measures of material condition. First, we conduct the principal component analysis (PCA), where the first principal component (the one with the most variance) obtained from this analysis is used as an alternative dependent variable. In this way we follow (Cools & Patacchini, 2017), who also construct a measure for material conditions of children albeit based on a different data set, using different indicators, and addressing a different research question. The rationale behind the method is elaborated, for example, by (McKenzie, 2005). He applies this method to measuring household wealth inequality based on responses about availability of different items. Importantly, he demonstrates that there is invariance of this measure across linear transformations. Additionally, we use ordered probit and Poisson models with the raw sum of ten indicators as the dependent variable. In this case, however, we assume that households acquire the most necessary child items first. The probit and the Poisson regressions measure the probabilities of acquiring the next most necessary items. Finally, the frequency histogram of the raw sum of indicators (Figure A1) shows that around one-half of the households possess all ten items. Therefore, we introduce one more binary alternative dependent variable. It takes a value of 1 for households which possess all specified items and a value of 0 for the other households. This specification of the dependent variable is the most intuitively appealing to us and we rely upon it in the main analysis. Nevertheless, under all specifications of the dependent variable the results of the analysis are qualitatively similar and the estimated coefficients of primary interest are statistically significant.

2.7.4 Cross-country comparison of gender preference and parental investment

Table A3 displays the results of estimating gender preference by country. The geographical pattern of the gender preference at birth is depicted in Figure A6. Our results are broadly consistent with those previously obtained in the literature. As did (Hank & Kohler, 2000), we find son preference in Italy and France and daughter preference in Portugal and Lithuania. Similar to (Andersson, Hank, Ronsen, & Vikat, 2006), we also find daughter preference in Norway, though not in Sweden.[47] We also attempt to evaluate how our results would differ if there were no family disruptions caused by gender of children, which is frequently reported in the literature (see, e.g., Lundberg (2005), for a review). The results are presented in Table A3. Son preference becomes statistically significant in Slovenia and stops being statistically significant in Croatia. However, the estimates obtained after including Slovenia and excluding Croatia from son preferring countries do not differ qualitatively and do not differ much quantitatively from those reported here. The rank correlations between the country-level impacts of the firstborn's gender on the selected household fertility outcomes are presented in Table 4. The absence of a strong correlation between estimated impacts on progression to parity two and parity three suggests different factors driving these impacts as we expected above.

Two measures of the same variable should be correlated, yet the correlations between second-parity coefficients and third-parity coefficients is quite low (Table A1). Still, the last two sets of coefficients are strongly correlated with coefficients for the total number of children. This might spur an examination of whether or not it is proper to use third parity progression for measuring gender preference, a frequent practice in the literature.

[47] Still, our estimates are correlated with (ρ=0.6) and statistically significantly predict comparable estimates to Hank and Kohler(2000).

To rationalize the estimates obtained, we plot the coefficients against several existing measures of gender inequality. As Figure A2 shows, the estimates do not exhibit a strong relationship with those measures. Only the coefficients from the third-parity equation exhibit a negative relationship with our gender equality score based on Eurobarometer data and with the proportion of households reporting balanced decision-making. At the same time, neither the coefficients for the total-number nor the second-parity equations exhibit any such relationship. This fact once more suggests that second parity progression and third parity progression actually measure different kinds of preferences. This is why we use third parity progression results in Figure 1 and beyond.

In addition, the fact that parents tend to invest more in daughters as measured by the presence of the home items[48] hold for the pooled EU-SILC sample. To test for the gender gap in parental investment we estimate Equation 1 with several alternative measures of child material conditions on the LHS. We primarily focus on the specification with the binary home indicator (the dummy variable for all 10 items) on the LHS. Table A8 displays estimates for this specification on a pooled sample. The results suggest that daughters, on average, receive more parental investments in terms of home items. For example, the number in column 1 means that families with first-born girls are 1.5% more likely to have all 10 items. This estimate is robust to the alternative sets of covariates, as can be seen in the rest of Table A8. Still, this effect is not large, remaining between 1,7% and 2% of the standard deviation of the binary home indicator. Results of this scale are typical in the literature on gender effects. Meanwhile, the gender preference pattern established before holds for the sub-sample of households from the highest income decile. These results might suggest that society as a whole is attaching increasingly positive significance to female children, an idea that has appeared in previous studies, such as (Brockman, 2001) and (Andersson, Hank, Ronsen, & Vikat, 2006). A daughter may assume both the role of a breadwinner and that of a caregiver.[49] As Brockmann (2001) puts it, "in the future, the average girl may well wish to become the mother of a one-daughter family."

[48] Availability of these indicators has been frequently used in the literature as a measurement of parental investment. More detailed discussion is presented in Section 3.
[49] In this regard some authors speak about the "boy crisis" (Husain & Millimet, 2009; Sadowski, 2010).

As with the estimates of the preference for gender of children at birth, we relate the estimates of the gender gap in parental treatment to specific country-level measures of gender inequality. The impact of the gender of the firstborn on material conditions exhibits a much stronger relationship with conventional measures of gender inequality than the impact on parity progression. Figure A3 displays the three strongest relationships. Most importantly, there is a strong relationship with the Global Gender Gap (GGG) score, calculated by the World Economic Forum (we used the most recent 2016 data). This index is also strongly related to the gender gap in PISA math achievement (Guiso, Monte, Sapienza, & Zingales, 2008).

However, Xu (2016) did not find any strong relationship between the gender gap in the home environment measure (similar to ours) and the GGG, though he measured the gender gap by difference in the unconditional mean between genders. Moreover, as explained earlier, our measure is preferable to the one used in Xu (2016). Therefore, the gender gap in child material conditions more closely corresponds to conventional gender-inequality measures than the gender gap in the number of younger siblings.[50] Nevertheless, the latter is commonly used as a measure of parental gender preference.

2.7.5 Tables and Figures

Table A1: Coefficients corrected for selection bias

Cntrs.	Coefs.	Cntrs.	Coefs.	Cntrs.	Coefs.	Cntrs.	Coefs.
AT	0.006	EE	-0.0007	IS	-0.003	PL	-0.003
BE	0.0003	EL	-0.006	IT	0.011***	PT	-0.017***
BG	0.0217***	ES	-0.001	LT	-0.006	RO	0.024***
CH	0.002	FI	0.004	LU	0.003	RS	0.029**
CY	-0.016*	FR	0.007	LV	-0.002	SE	0.010
CZ	0.002	HR	0.027*	MT	-0.010	SI	0.012**
DE	0.006	HU	-0.008*	NL	-0.004	SK	0.010
DK	-0.017**	IE	0.007	NO	-0.018**	UK	0.0007

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

Notes: The estimates contained in this table do not differ from those in the third column of the Table A5 except in the sample characteristics and omission of father-related control variables (which have little explanatory power). The sample also includes incomplete families with simulated numbers of additional children—simulated under the assumption that those divorced because of the gender of children are characterized by bias towards that gender and do not stop producing more children until they a child of the desired gender.

[50] A similar and statistically significant relationship also holds between the first-daughter coefficient in the material-conditions regression and two other indexes: the GDI (it highly correlates with the GGG) and the SIGI (though it is available only for seven countries from our sample).

Table A2: Effects of firstborn gender on selected measures of fertility

Explanatory var-s	(1) Total number of children	(2) Two or more children	(3) Three or more children	(4) Four or more children	(5) Five or more children
First child a girl	-0.0050**	-0.0073***	0.0011	0.0004	0.0005*
	(0.0025)	(0.0017)	(0.0012)	(0.0006)	(0.0003)
Controls	Yes	Yes	Yes	Yes	Yes
First boy baseline	1.54	.406	.106	.0248	.00462
Percent effect	-.00323	-.0179	.0102	.018	.109
R-sq	.27	.235	.137	.0491	.0163
Observations	265,507	265,507	265,507	265,507	265,507

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

Notes: S.E. are given in parentheses and are clustered at the country level. Estimates are based on the 2004-2015 EU-SILC samples for Bulgaria, Croatia, Serbian Republic, and Slovenia. The sample consists of households formed by one cohabiting couple, their children, and, occasionally, other relatives. The mother of children in the household is younger than 41 and older than 17 and had her first child at the age of 16 or older, and children's ages are in the range 0–12. The estimation method used is weighted OLS with probability weights reflecting non-random sampling within and between countries. The table presents estimated effects of the firstborn being a daughter compared with the baseline case of the firstborn being a son. The effect is a ratio of the estimated OLS coefficient on the firstborn's gender dummy to the baseline value of the dependent variable. The dependent variables are the total number of children and a set of binary indicators for specific numbers of children. The control variables, besides the gender of the firstborn, are: the dummy for a first-born daughter, gender of the first two children, cubic polynomial of mother's age, squared polynomial of mother's age at first birth, length of cohabitation of spouses, mother's education, father's education, mother's employment, father's employment, household disposable income, living in urban area, tenure status, year and country dummies.

Table A3: Effects of firstborn gender on selected measures of fertility

Countries [a]	(1) Total number of children	(2) Two or more children	(3) Three or more children	(4) Four or more children	(5) Five or more children	Obs
AT	-0.0181	-0.0245*	0.0083	-0.0050	0.0015	6,574
BE	-0.0074	-0.0139	0.0054	0.0007	0.0004	7,694
BG	0.0206	-0.0112	0.0222**	0.0096*	0.0011	3,509
CH	0.0353	0.0364**	0.0013	0.0013	-0.0017	4,461
CY	-0.0422*	-0.0330**	-0.0125	0.0032	0.0002	5,675
CZ	-0.0123	-0.0167*	0.0037	-0.0002	0.0001	10,329
DE	-0.0141	-0.0179*	0.0060	-0.0012	-0.0010	9,790
DK	-0.0183	-0.0023	-0.0178*	0.0012	0.0007	7,889
EE	-0.0147	-0.0091	-0.0032	0.0027	-0.0017	6,594
EL	-0.0040	-0.0075	-0.0065	0.0045	0.0041***	8,147
ES	-0.0292**	-0.0277***	-0.0030	0.0003	0.0008	16,054
FI	-0.0027	-0.0031	0.0070	-0.0000	-0.0011	13,145
FR	0.0209*	0.0102	0.0072	0.0005	0.0029**	14,496
HR	0.0878**	0.0507*	0.026**	0.0127	0.0031	1,742

HU	-0.0082	0.0057	-0.0137**	-0.0027	0.0015	11,281
IE	0.0002	0.0094	0.0030	-0.0074	-0.0007	5,636
IS	-0.0059	0.0009	-0.0022	-0.0028	-0.0014	5,711
IT	0.0091	-0.0032	0.0121***	-0.0004	0.0002	21,486
LT	-0.0352	-0.0096	-0.0090	-0.0098**	-0.0040*	3,742
LU	-0.0068	-0.0069	0.0022	0.0020	-0.0029*	8,084
LV	-0.0172	-0.0204	-0.0020	0.0028	0.0008	5,102
MT	-0.0170	-0.0013	-0.0118	-0.0019	-0.0013	2,872
NL	0.0021	0.0039	-0.0033	-0.0001	0.0001	11,942
NO	-0.0385**	-0.0210*	-0.0191*	0.0006	0.0007	8,108
PL	0.0049	-0.0037	-0.0008	0.0023	0.0035**	18,374
PT	-0.0794	-0.0486***	-0.0216***	-0.0074**	-0.0008	6,044
RO	0.0293	0.0028	0.0218**	0.0075*	-0.0027	4,948
RS	0.0619	0.0378	0.0214	0.0044	-0.0017	1,221
SE	0.0240	0.0112	0.0114*	0.0019	-0.0006	9,228
SI	0.0140	-0.0147	0.0113	0.0093***	0.0036***	10,544
SK	0.0191	-0.0025	0.0093	0.0072*	0.0018	5,802
UK	-0.0155	-0.0104	0.0034	-0.0085*	-0.0012	9,288

* p < 0.1; ** p < 0.05; *** p < 0.01
Notes: See notes for Table 3 for data samples, variable definitions, and included control variables. The columns contain estimated country-level effects of firstborn daughters on the corresponding variables in the column headings.
[a] Table A4 contains names of countries corresponding to the abbreviations.

Table A4: Abbreviations for countries

Abbrev.	Countries	Abbrev.	Countries	Abbrev.	Countries	Abbrev.	Countries
AT	Austria	EE	Estonia	IS	Iceland	PL	Poland
BE	Belgium	EL	Greece	IT	Italy	PT	Portugal
BG	Bulgaria	ES	Spain	LT	Lithuania	RO	Romania
CH	Switzerland	FI	Finland	LU	Luxembourg	RS	Republic of Serbia
CY	Cyprus	FR	France	LV	Latvia	SE	Sweden
CZ	Czech Republic	HR	Croatia	MT	Malta	SI	Slovenia
DE	Germany	HU	Hungary	NL	Netherlands	SK	Slovak Republic
DK	Denmark	IE	Ireland	NO	Norway	UK	The United Kingdom

Source: Eurostat

Table A5: Impact of the first-born daughter on selected household allocation decisions under two alternative explanations of the parental gender preference

Allocation decisions	Bias		Intact family advantage	
	towards sons	towards daughters	in raising sons	in raising daughters
Household public goods expenditure	-	+	.	.

Savings	+	-	.	.
Personal well-being of a father	+	-	-	+
Personal well-being of a mother	-	+	+	-

Note: The sign "+" means a positive impact and the sign"-" means a negative impact. The rationale behind the predictions is explained primarily in the Introduction and also in Sections 3 and 4.

Table A6: Impact of the first-born daughter on selected household allocation decisions under two alternative explanations of the parental gender preference

Allocation decisions	Balkan countries		Scandinavian countries	
	Bias towards sons	Intact family comparative advantage in raising sons	Bias towards daughters	Intact family comparative advantage in raising daughters
Household public goods expenditure	-	+	.	.
Savings	+	-	.	.
Personal well-being of a father	+	-	-	+
Personal well-being of a mother	-	+	+	-

Note: The sign "+" means a positive impact and the sign"-" means a negative impact. The rationale behind the predictions is explained primarily in the Introduction and also in Sections 3 and 4.

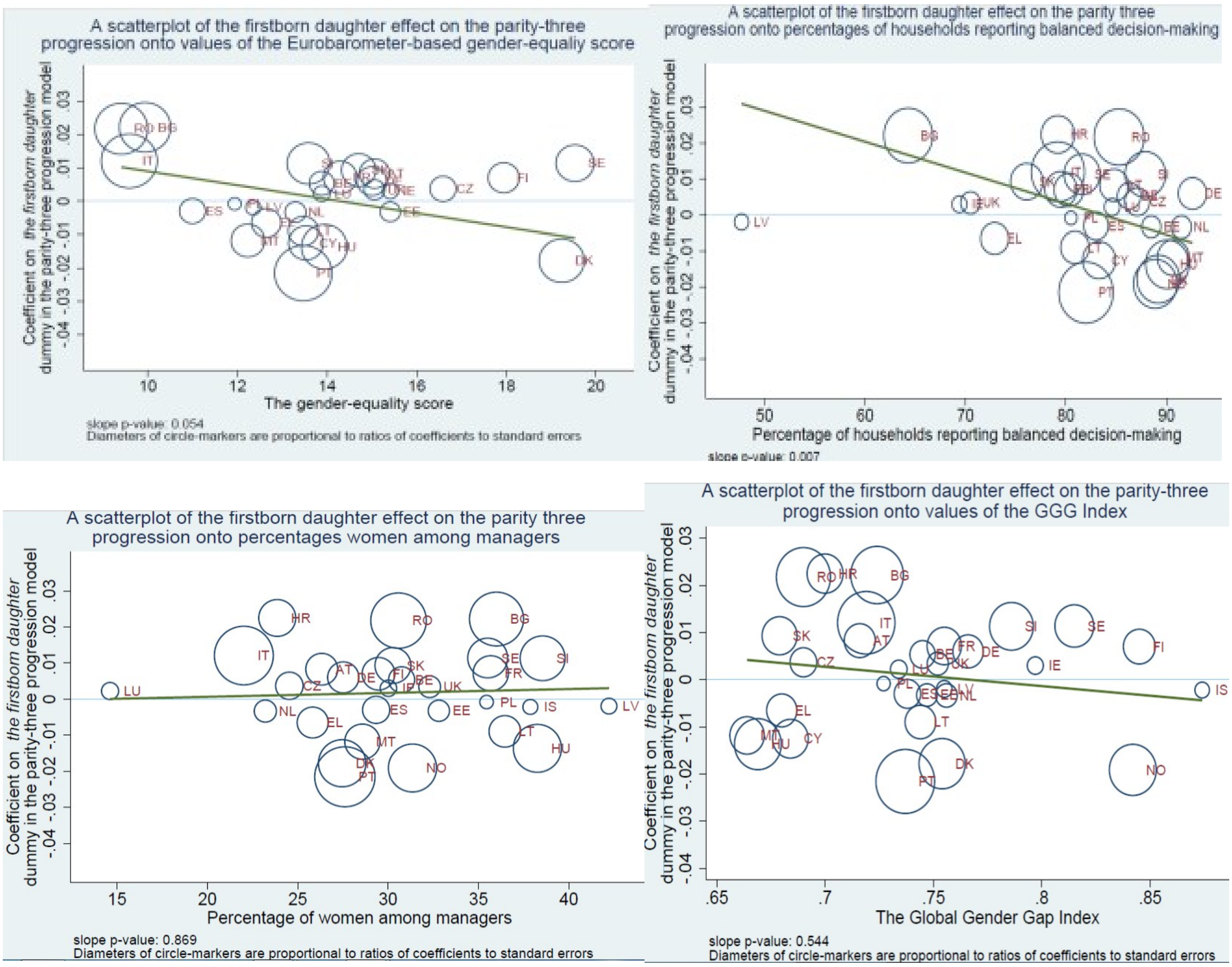


Figure A2: The relationship between the effect of first-born daughters on third parity progression and specific gender-equality measures across countries. We calculate the Eurobarometer-based gender equality score for a particular country as a sum of the country's ranks in responses to questions about attitudes towards gender equality. These responses were collected in the 2009 Eurobarometer special survey (European Commission, 2010). For each question, countries were ordered according to shares of respondents who report an existence/wish to exist in gender-egalitarian conditions in a specified realm of life. The country with the highest share of such respondents was assigned rank 1 for the corresponding question. Then, we calculated the sums of such ranks across all 13 pertinent questions and our gender-equality score. Please note that we do not have scores for Switzerland, Croatia, Iceland, Norway, and the Republic of Serbia, because the Eurobarometer survey was not conducted in those countries. Percentages of households reporting balanced decision-making were taken from the data of Health and Demographic Survey collected by the World Bank in multiple years and from the Survey of Income and Living Conditions collected by Eurostat in 2010. The percentage of women managers was obtained from the data of the Enterprise Surveys, conducted by the World Bank in multiple years. The Global Gender Gap Index was calculated by the World Economic Forum in 2016.

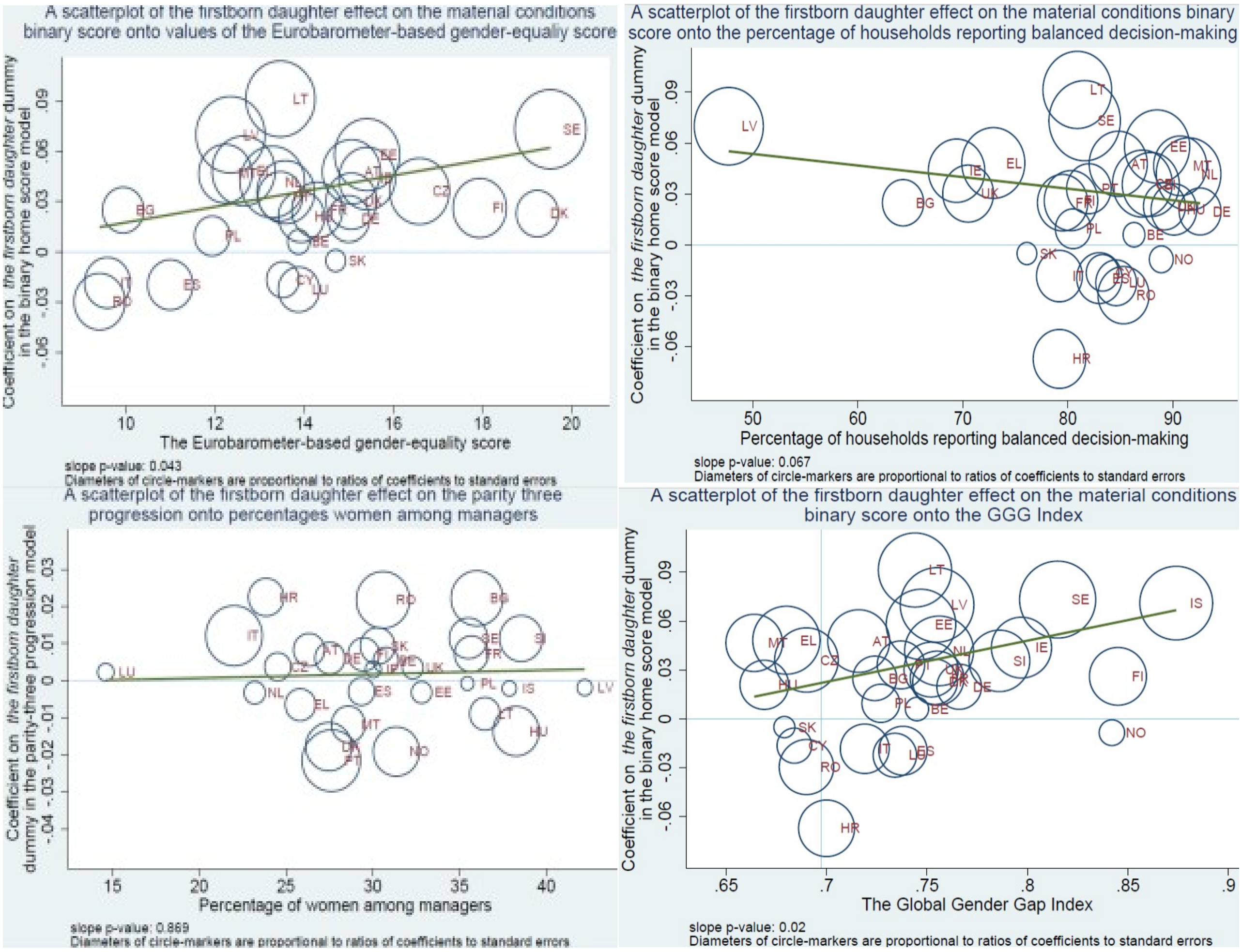
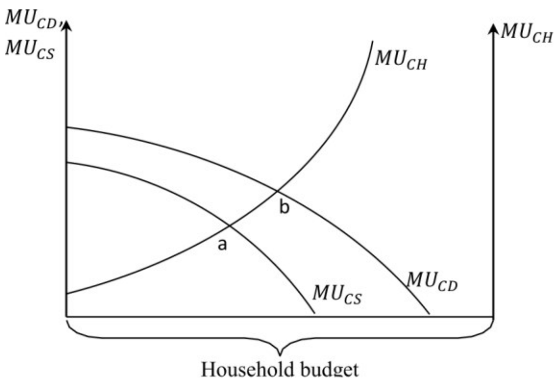
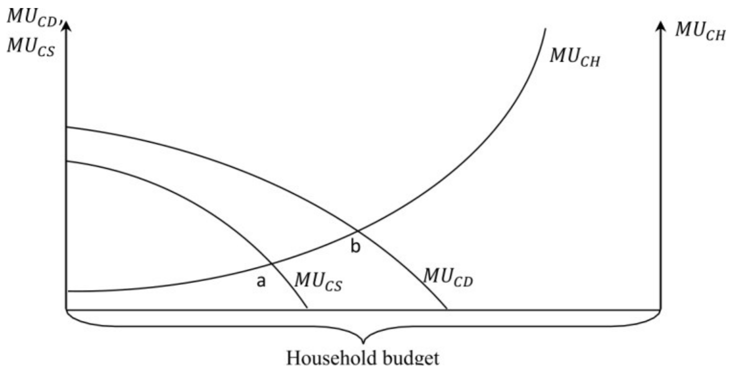


Figure A3: The relationship between the effect of first-born daughters on child material conditions and specific gender-equality measures across countries.



(c) Low-income



(d) High-income families

Figure A4: Differences in expenditures on children between low-income and high-income households *Notes:* See the note to Figure A1 for explantion.

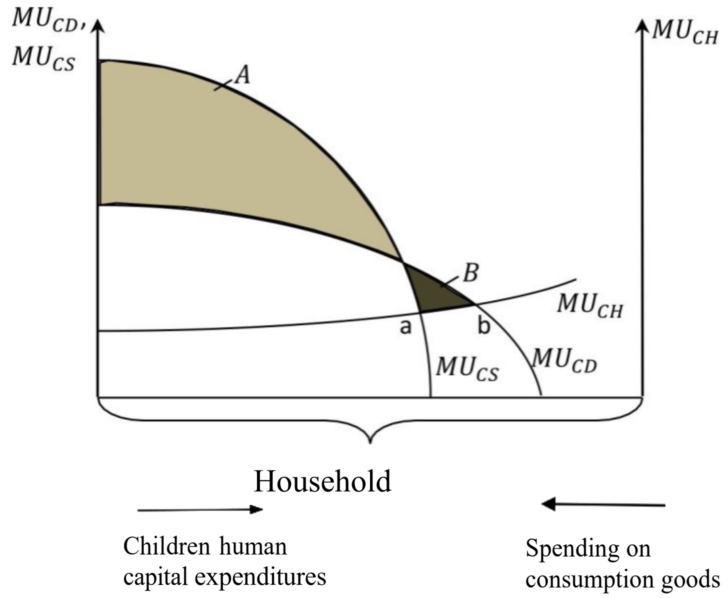


Figure A5: Coexistence of gender (son) bias and differential cost with the gender bias effect on fertility prevailing. *Notes:* See the note to Figure A1 for explanation.

Table A7: Spearman's rank correlations between country-level effects of first-born daughters on selected measures of fertility

	Total number of children	Progression to parity two	Progression to parity three	Progression to parity four	Progression to parity five
Total number of children	1				
Progression to parity two	0.8380***	1			
Progression to parity three	0.7878***	0.4765***	1		
Progression to parity four	0.4758 ***	0.2753	0.3680**	1	
Progression to parity five	0.0037	-0.1334	-0.0169	0.2834*	1

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

Notes: Spearman's rank correlations are based on estimates for 32 European countries covered in the EU-SILC survey during 2004-2015. The estimates are contained in Table A5.

Table A8: The impact of the firstborn gender on the binary material deprivation indicator.

	The binary material deprivation indicator on the LHS			
Explanatory var-s	(1)	(2)	(3)	(4)
	OLS	OLS	OLS	IV
First child a girl	.015***	.0148***	.0168***	.0172***
Number of children		.0896***	.0797***	-.0231*
Covariates	No	No	Yes	Yes
R-Square	.000225	.0191	.168	.146
N obs	51,087	51,087	49,922	49,922

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

Notes: The standard errors of estimates on pooled EU-SILC sample are clustered at the country level. The table presents estimated effects of the firstborn being a daughter compared with the baseline case of the firstborn being a son. Estimates are based on the 2009 and 2013-2015 EU-SILC ad-hoc modules, while the estimates in the remaining columns are based on the 2004-2015 EU-SILC primary modules. The sample consists of households formed by one cohabiting couple, their children, and, occasionally, other relatives. The mother of children in the household is younger than 41 and older than 17 and had her first child at the age of 16 or older, and children's ages are in the range 0–12. The estimation method used is weighted OLS with probability weights reflecting non-random sampling within and between countries. The dependent variable is the binary indicator taking value 1 when a household has all 10 items listed in the Table [Table2]2 and takes value 0 otherwise. Other control variables are: the dummy for a first-born daughter, gender of the first two children, cubic polynomial of mother's age, squared polynomial of mother's age at first birth, length of cohabitation of spouses, mother's education, father's education, mother's employment, father's employment, household disposable income, living in urban area, tenure status, year and country dummies. The estimates in the fourth column are obtained using the 2SLS method from a regression-model in which the number of children is instrumented with twin-birth. The first stage F-statistic value for this model is above two thousand.

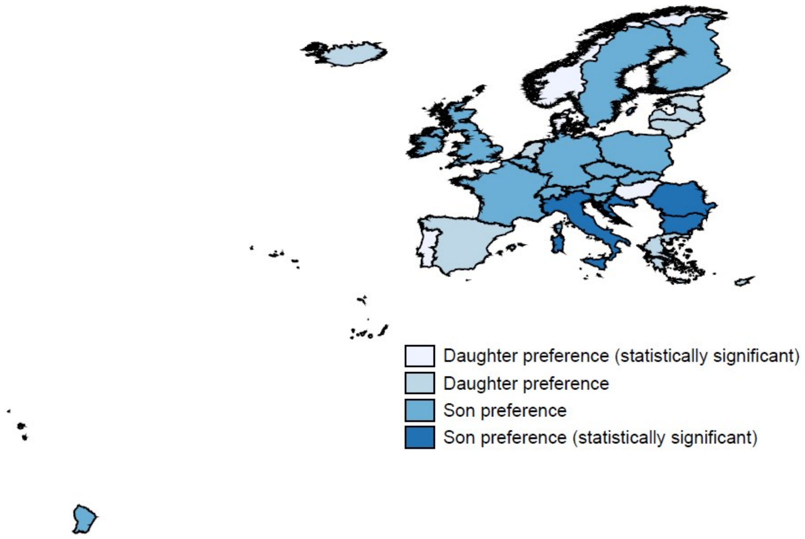


Figure A6: Gender Preferences of Children in 31 EU-SILC countries

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

3.1 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)[51]). A growing body of research examines how a child's gender may be associated with differential treatment by parents from birth, which could to gender 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, Warner, and Acock 1994); and a birth outside marriage is more likely to be followed by marriage if the child is a son (Lundberg and Rose 2003a). 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, Donahue, and Haskins 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 2017). 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). However, a more recent study by Kabátek and Ribar (2017) concluded that strained relationships in families with teenage daughters likely lie behind the higher divorce rate of couples who have daughters. At the same time, the aforementioned mechanism does not explain why single mothers with daughters are less likely to marry, especially if their children are of preteen or younger age. Kabátek and Ribar (2017)

[51] 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.

do not check whether the last fact holds for Dutch households, which are their population of interest. My paper aims to simultaneously estimate the impact of the firstborn gender on family formation and dissolution, conditional on firstborn age.[52] 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 divorce to change with children's ages, but also incorporates the length of cohabitation of spouses and the right-censoring in marriage spells.[53]

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. Therefore, if similar mechanisms to those underlying the results of the 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.

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 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 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 divorce 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 divorce.[54] My results show that the effect on probability of divorce[55] is negative, and the effect on probability of marriage is also negative. Additionally, I find an indication that having teenage daughters increases the probability of divorce in line with

[52] Plausibility of the assumption of the firstborn gender randomness is discussed in the Method section.
[53] 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.
[54] Kabátek and Ribar (2017) argue that parents do not foresee difficulties in relationship with teenage daughters when daughters are of a younger age.
[55] 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.

Kabátek and Ribar (2017).[56] 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 divorce. Investigation of the precise nature of these determinants is currently beyond the scope of this paper.[57]

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 sup- porting 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.

3.2 Data

3.2.1 RLMS-HSE Data

The source of data for my analysis is the Russia Longitudinal Monitoring Survey, RLMS-HSE, for the 1994-2018 period. The dataset covers 23 yearly rounds of a national representative survey on the social, health and economic situation in Russia. Two years are missing, 1997 and

[56] This result is less warranted because I cannot observe many teenage children because of the attrition of households over survey waves.
[57] I suggest five possible causes for the negative effect of first-born daughters on divorce 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 divorce are higher when children are female. Third, higher marriage rates among single mothers of first-born sons reduces the cost of divorce for them as they perceive their remarriage prospects to be more favorable. Fourth, public policy is more oriented towards women, because women constitute the majority of voters demographically and vote more actively. Fifth, gender of children induces changes in parents' (especially fathers') personality, which in turn influence marriage stability.

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 by several organizations including the National Research University Higher School of Economics and the Carolina Population Center, University of North Carolina at Chapel Hill.[58]

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. House- holds participating in the survey were selected through a multi-stage probability sampling procedure in order to guarantee cross-sectional national representative- ness. 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.[59]

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.[60]

The RLMS-HSE is a panel survey with a longitudinal component. A point of concern is that of attrition in the panel.[61] 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

[58] 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/.
[59] The target sample size was set at 4 000 households (Kozyreva, Kosolapov, and Tonis 2016). 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, Kosolapov, and Popkin (2016) and also here: https://www.hse.ru/rlms/sample (in Russian).
[60] 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, Kosolapov, and Popkin 2016) initiated in 1989.
[61] 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.

years is reported in, e.g., Kozyreva, Kosolapov, and Popkin (2016), 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.[62]

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

[62] 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.

3.2.2 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 married women[63] and 1,367 were observed in more than one survey wave. Correspondingly, 357 firstborns were born to single women,[64] of whom 255 were observed in more than one survey wave. 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 3.5 in Appendix B. Descriptive statistics are presented in Table 3.1. 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 3.1: 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

[63] Of these marriages, 1,231 have a known start date. Estimation of models in which the time under risk starts from the year of marriage formation rather than from the year of firstborn arrival yields estimates close to my reported results.
[64] 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.

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

Notes. The results are based on 1,367 observations, 701 boys and 666 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 3.2. 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 marry faster.15 Second, mothers of sons appear to be 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 3.2: 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
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 divorce, which I observe in the data and which implies that single mothers of daughters are less willing to participate in the survey. As for the lower survey response rate from less-educated individuals, this is reported in earlier studies.

Table 3.3: Numbers and shares of first-born children living with married and single mothers by age and gender

	1 wave before birth[a]		The year of birth		1-year-olds	
	Mar	Unmar	Mar	Unmar	Mar	Unmar
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	
Mar	Unmar	Mar	Unmar	Mar	Unmar	Mar	Unmar
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

Year-olds		7-year-olds		8-year-olds		9-year-olds	
Mar	Unmar	Mar	Unmar	Mar	Unmar	Mar	Unmar
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	
Mar	Unmar	Mar	Unmar	Mar	Unmar	Mar	Unmar
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.

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 married mothers.

The numbers of first-born boys and girls of different ages living with married and unmarried mothers can be seen in Table 3.3. The share of boys among children living with single mothers is higher than for children living with married 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 divorce. Second, mothers of daughters more frequently drop out of the survey after divorce,[65] while single mothers of sons more frequently drop out of the survey after marriage. The second conjecture is supported by the tendency for the share of first-born boys remaining in the survey to increase over time.

To take into account the last-mentioned fact, I calculate the differences between shares of boys living with unmarried and married mothers from Table 3.3 for first- borns aged 0-10 and show them on Figure 3.1.

The differences between shares of sons living with single and married mothers tend first to increase and then to decrease. This tendency is compatible with more divorces among mothers of younger first-born sons and fewer divorces among mothers of older first-born sons.

[65] This could happen because mothers with daughters tend to divorce when children are older, as the results in Table 4 show. Moving to a different location is easier with older children.

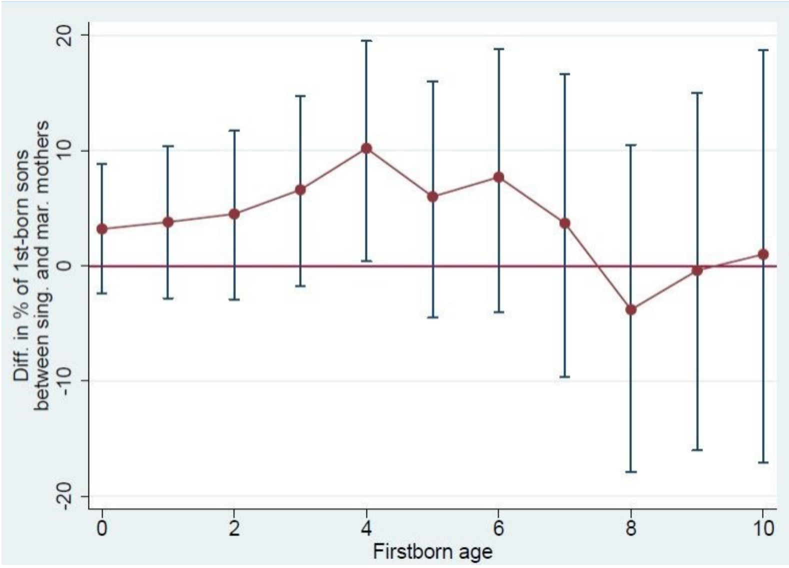


Figure 3.1: Difference between percentages of firstborn sons among single and married mothers conditional on firstborn age

Note: Capped spikes show 90% confidence intervals.

3.3 Method

Following (Kabátek and Ribar 2020), I estimate a complementary log-log (cloglog) discrete-time hazard model of divorce and marriage. This model has a number of advantages. First, it allows for covariates changing with time. Second, the results are straightforward to interpret: exponentiated estimated coefficients can be interpreted as approximate odds ratios. Third, it is widely used in the literature, being a discrete analog of the continuous proportional hazards model. Fourth, the underlying link function more closely approximates the distribution of observed marriage 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 marriage duration data, in which the divorce is relatively rare and hence most marriage durations are large which leads to a left-tailed (or left-skewed) distribution of baseline divorce hazards.[66] Fifth, the assumption of proportional hazards is intuitively plausible in the current setting. The Cox PH model assumes

[66] 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.

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 observed covariates (see, e.g., Hurrell (2015, p. 475) or Wooldridge (2002, p. 690)).

The functional form of the cloglog model is:

$$\Pr[y_{it} = 1|x_{it}] = 1\text{-}\exp(\text{-}\exp(x_{it}^{'}\beta)) \qquad (1a)$$

where the hazard probability of y_{it}, a divorce for a couple *i* or a marriage for a single mother *i* observed in year *t*, is defined as a function of covariates $\mathbf{x}_{it}$ that are specific to the given couple and year. This model includes time-varying covariates (TVC). Statistical software packages, for instance, Stata, can handle the estimation of such models. A more detailed discussion of estimating a Cox PH model with TVC is presented in, e.g., Wooldridge (2002, pp. 711-714) or Box-Steffensmeier and Jones (2004, pp. 95-104), while a discussion of estimating cloglog model with TVC is presented, e.g., in Box-Steffensmeier and Jones (2004, pp. 108-110) or Jenkins (2008, Lesson 6).

The corresponding specification of the predictor (also called *index*) $x_{it}^{'}\beta$ in the model of marriage dissolution with conditioning on firstborn age is:

$$\begin{aligned} x_{it}^{'}\beta = \beta_0 &+ 1\,(\text{FB age}_{\text{range}_{it}} = 0\text{-}5)\,(\beta_{0_{0\text{-}5}} + \beta_{1_{0\text{-}5}}\text{FB daughter}_i) \\ &+ 1\,(\text{FB age}_{\text{range}_{it}} = 6\text{-}18)\,(\beta_{0_{6\text{-}18}} + \beta_{1_{6\text{-}18}}\text{FB daughter}_i) + \sum_{k=1}^{18} \beta_{3j} \\ &\cdot\ 1(\text{ FB years\_obs}_{it} = j)\ +\ z_{it}^{'}\,\beta_4 \qquad (1b) \end{aligned}$$

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 divorce hazard increases by over the divorce hazard in families with firstborns older than 18 in any given year. 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 (1b) each with a dummy included for only one age range, with the other age range being a base category. In all three regression specifications, the constant is not suppressed. The vector **z** includes employment status, age, religious affiliation, living in an urban area, being of Russian ethnicity, the number of children in the family, and educational accomplishment of both spouses (see Table 1).

The functional form for the model of marriage formation is the same as for the model of marriage dissolution:

$$\Pr[y_{it} = 1|x_{it}] = 1\text{-}\exp(\text{-}\exp(l_{it}^{'}\alpha)) \qquad (2a)$$

The specification of the predictor in equation (2.1), which I now denote $l_{it}^{'}\alpha$ with conditioning on firstborn age is:

$$l_{it}^{'}\alpha = \alpha_0 + 1\,(\text{FB age}_{\text{range}_{it}} = 0\text{-}5)\,(\alpha_{0_{0\text{-}5}} + \alpha_{1_{0\text{-}5}}\text{FB daughter}_i) + 1\,(\text{FB age}_{\text{range}_{it}} = 6\text{-}18)\,(\alpha_{0_{6\text{-}18}} + \alpha_{1_{6\text{-}18}}\text{FB daughter}_i) + \sum_{k=1}^{18} \alpha_{3j} \cdot 1(\text{ FB years\_obs}_{it} = j) + w_{it}^{'}\,\alpha_4 \qquad (1b)$$

I also estimate specifications analogous to those of the marriage dissolution model. The vector **w** includes employment status, age, religious affiliation, living in an urban area, being of Russian ethnicity, the number of children in the family, and educational accomplishment for single mothers. Vectors of parameters β and α 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.[67]

For the marital dissolution model, I test two hypotheses: a) first-born teen- and school age daughters cause a different parental divorce rate than their peer first-born sons, i.e. H_0 : $\beta_{1_{6\text{-}18}} = 0$ versus H_A : $\beta_{1_{6\text{-}18}} \neq 0$; b), and first-born daughters of preschool age (0-5 years old) cause a different parental divorce rate than their peer first-born sons, i.e. H_0 : $\beta_{1_{0\text{-}5}} = 0$ versus H_A : $\beta_{1_{0\text{-}5}} \neq 0$. 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

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

marital formation model, I test H_0 : $\alpha_{1_{0\text{-}5}} = 0$ H_A : $\alpha_{1_{0\text{-}5}} \neq 0$. 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.

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 sub- groups 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.[68] 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.[69] Second is construction of the dependent variable - the indicator of family dissolution. While Kabátek and Ribar (2020) focus on formal marriage status,[70] 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

[68] 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, Appleman, and Maksymovych (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).

[69] 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.

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

(2008). Second, women who appear to be divorced 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 only if they are divorced and non-cohabiting in the individual-level data (as reported by a wife).[71] In other words, in the basic marriage dissolution model I consider those women who appear to be divorced based on both individual and household level data as actually. Women who fulfill only one of these conditions or none, are considered to be married.

Similarly, my marriage formation indicator takes a value of 1 when a couple starts cohabiting according to the household-level data file and if they are married in the individual-level data (again as reported by a wife). In other words, in the basic marriage formation model I consider those women who are married both in the individual and household level data to be married; women who fulfil only one condition or none are considered to be single. Finally, the character of association between covariates and the dependent variable, along with estimation results with alternative errors specifications suggest that concerns about non-monotonicity and heteroskedasticity are not justified.[72]

I expect, in line with previous studies, the coefficient $\beta_{1_{6\text{-}18}}$ to have a positive value and the coefficient $\alpha_{1_{0\text{-}5}}$ to have a negative value. At the same time, I expect first-born daughters aged 0-5 years to have a negative impact on family dissolution, i.e. for $\beta_{1_{0\text{-}5}}$ to be negative. I readily see two possible reasons for that. 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; Adushkina 2015; Aivazova 2015; Mkchtrian 2017).[73] The second is the loss in marriage surplus due to loss in the human capital of daughters. In other words, if investment in human capital of children brings higher returns in terms of marriage surplus for

[71] Other living arrangements recorded in the individual-level data file are: non-divorced and non- cohabiting or divorced and cohabiting. I use alternative measures for the family dissolution for robustness checks. Table 7 in Appendix C shows results for the marriage dissolution model with the dependent variable being cohabitation termination according to the household data file.

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

[73] 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 house-hold 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 approachtheir 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 thansons.

daughters than for sons (Abramishvili, Appleman, and Maksymovych 2019), divorce of spouses with a first-born daughter could cause especially high loss in marriage surplus (Currie and Almond 2011; Myck, Oczkowska, and Wowczko 2021; Kim and Shapiro 2021). These two explanations, as well as three more that follow, are based on the related academic literature and narratives in common present socioeconomic discourse. However, they are not likely to exhaust the list of all possible explanations.

3.4 Results and Discussion

3.4.1 Results for Marriage Dissolution

Estimates of the model (1a) with six specifications of predictor (1b) are presented in Table 4.[74]
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.[75] The estimated effect of young (0-5 years old) first-born daughters on the probability of marriage dissolution is negative. It becomes statistically significant after adding covariates. These results confirm the expectations based on conjectures about

[74] Results for cohabitation termination as a dependent variable are presented in Table 7. 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.

[75] 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 3.4: The impact of the firstborn gender on family dissolution from complementary log-log model estimation.

Explanatory var-s:	Marriage dissolution (1)	Marriage dissolution (2)	Marriage dissolution (3)	Marriage dissolution (4)	Marriage dissolution (5)	Marriage dissolution (6)
Firstborn 0-5 years old is daughter	0.97	0.68	0.93	0.50		
	(0.20)	(0.14)*	(0.19)	(0.17)**		
Firstborn 6-18 years old is daughter	2.27	2.16			2.24	2.16
	(0.96)**	(0.94)*			(0.95)**	(0.93)*
Firstborn 0-5 years old	0.031	0.25	1.65	0.83		
	(0.004)***	(0.12)***	(0.58)	(0.92)		
Firstborn 6-18 years old	0.022	0.26			0.42	1.10
	(0.009)***	(0.37)***			(0.18)**	(1.45)
N of marriage spells	1,367	1,367	1,367	1,367	1,367	1,367
N of marriage-year observations	7,163	7,069	7,163	7,069	7,163	7,069
Log-likelihood	-624.55	-586.56	-634.18	-587.98	-632.47	-587.78
Socio-demographic constrols	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 (1) and (2), dummiesfor both age ranges are included and the base category is all firstborns older than 18. In columns(3)-(6) the base category also includes first-born sons and daughters aged either 6-18 (columns (3)-(4)) or 0-5 (columns (5)-(6)). 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).

the relationship with parents of spousesand higher loss of marriage surplus by divorcing parents of daughters. Results ofthe estimation with cohabitation termination as a dependent variable in Table 7 in Appendix C are in line with the results in Table 4. Besides the two explanations already proposed, three further mechanisms might account for this finding.

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

The fourth reason is a cumulative effect of existing policies, especially social policies. The fact that policies cannot be gender neutral is discussed in Washington (2008). For example, Cygan-Rehm, Kuehnle, and Riphahn (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[76], generous public pensions (relative to average salary) that disproportionately benefit women due to a significant gender gap in life expectancy,[77] indexing public sector salaries, which is likely to be more beneficial to women, who constitute a majority of public sector employees (including education and health care), or "gender asymmetry" in family law (Klimashevskaya 2021).[78] 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 when their firstborn is a daughter.

Finally, the fifth 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 A.

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 statistically significant, corresponding to a bell-shaped form of the empirical divorce hazard.

3.4.2 Results for Marriage Formation

Estimates of model (2a) with four specifications of the predictor (2b) are presented in Table 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.

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

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

[78] 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 (Laktiukhina and Antonov 2017).

Table 3.5: The impact of the firstborn gender on family formation from a complementary log-log model estimation.

Explanatory var-s:	Marriage formation (1)	Marriage formation (2)	Marriage formation (3)	Marriage formation (4)
Firstborn is daughter	0.93 (0.19)	0.45 (0.20)*		
Firstborn 0-5 years old is daughter			0.54 (0.16)**	0.45 (0.15)***
Firstborn 0-5 years old			1.3 (0.46)	1.07 (0.46)
N of single mothers observed	255	255	255	255
N of marriage-year observations	1,040	1,037	1,040	1,037
Log-likelihood	-325.64	-302.18	-327.44	-299.69
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 (5), 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).

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 divorce. Thus, the higher marriage rate among single mothers of first-born sons is outweighed to some extent by a higher divorce rate among married mothers of first-born sons.

The possibility that the estimates obtained are driven by family background of mothers is examined in Appnedix C.

3.5 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 divorce. 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 divorce, while having daughters aged 6-18 predicts a higher chance of parental divorce. 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 discordant with 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 divorce use data from countries for which a son preference has been reported.[79] Son preference has not been established for Russia so far, 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 five such features may be pointed out: relationship between spouses and their parents, higher returns to children human capital investment for daughters than for sons, chances of shotgun marriages depending on the gender of children, women constituting the majority of voters and having relatively high electoral activity, and not very strong reliance on tradition and norms in building family relationships among young spouses. Examining the plausibility of these possible mechanisms will be a focus of my further research.

[79] 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, Carvalho, and Lleras- Muney 2014), Australia (Kippen, Evans, and Gray 2006).

3.6 Appendix A: Impact of changes in extraversion and neuroticism of fathers on marriage stability

Neuroticism increased only among fathers of daughters aged 0-5 while extraversion remained 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 scoreson neuroticism of spouses and likelihood of divorce. This is, in turn, attributed to the negative relation between neuroticism and marital satisfaction, which was con- firmed 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).[80] A similar relationship was con- firmed 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 (2017) 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 im- pact on divorce. Hardly any research has been conducted so far to examine this question in the Russian context.

As for extraversion, its possible impact is less clear-cut. On the one hand, extraversion is assumed to be related to positive emotions (Donellan, Conger, and Bryant 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, von Scheve, and Park (2017) and

[80] 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.

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 suchan association in an Australian nationally representative sample. Moreover, Solomonand Jackson (2014) report a positive relationship between extraversion and marital satisfaction. The latter, according to Hirschberger, Srivastava, and Marsh (2009), negatively predicts prospective marriage dissolution for men.[81] Therefore, the di- rection 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, inthe 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.[82] Zelenskaia and Liders (2015) say that extraversion is associated with the presence of "communicative resources" needed for discussing the role structure of the family.[83] Nevertheless, it remains unclear how the role of extraversion at the onset of a 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 im- balances 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. More-

[81] Marital satisfaction around the first child's transition to school is the strongest predictor.
[82] In addition, Shvetsova and Kondrasheva (2015) report that young husbands assign relatively high value to friend networks (compared to wives and older spouses).
[83] This fact resonates with the conclusion by Somville (2019) that having a daughter reduces male violence against a partner.

over, 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.

3.7 Appendix B

Table 3.6: Data Description for Selected Variables

Marriage termination	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
Marriage 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
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 no religion	Dummy for whether a mother reports adherence 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 ethnicity	Dummy for whether a mother reports being of Russian
Father is Russian ethnicity	Dummy for whether a father reports being of Russian
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

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

3.8 Appendix C

3.8.1 Robustness check for marriage dissolution estimate

Table 3.7: The impact of the firstborn gender on cohabitation termination from complementarylog-log model estimation.

Explanatory var-s:	Cohabitation termination (1)	Cohabitation termination (2)	Cohabitation termination (3)	Cohabitation termination (4)	Cohabitation termination (5)	Cohabitation termination (6)
Firstborn 0-5 years	1.05	0.77	1.02	.77		
old is daughter	(0.20)	(0.15)	(0.20)	(0.15)		
Firstborn 6-18 years	1.93	1.85			1.91	1.85
old is daughter	(0.72)*	(0.70)*			(0.71)**	(0.70)*
Firstborn 0-5 years	0.04	0.77	1.19	0.46		
old	(0.004)***	(0.15)	(0.40)	(0.50)		
Firstborn 6-18 years old	0.03	0.45			0.59	1.91
	(0.01)***	(0.62)			(0.24)	(2.41)
N of marriage spells	1,367	1,367	1,367	1,367	1,367	1,367
N of marriage-year observations	7,163	7,069	7,163	7,069	7,163	7,069
Log-likelihood	-712.73	-657.13	-709.02	-658.37	-707.50	-657.65
Socio-demographic controls	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 a wife does not report widowhood. Explanation of base categories and interpretation of estimates are provided in the notes under Table 4.

3.8.2 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.[84] The estimates of the first-born daughters' impact on marriage dissolution and formation after controlling for women's family background are presented in Table 8. Variables characterizing the

[84] 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.

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 0in 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. Not reporting a year of birth of a parent likely correlates with parental divorce 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). Also, among characteristics of the family background measured in the RLMS-HSE survey, the ones chosen have the fewest missing observations. The estimates of the first-born gender impact in Table 8 appear to accord with the results in Tables 4 and 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 divorce hazard, in line with expectations. The level of statistical significance for this effect is somewhat lower than 0.1, but it mightincrease 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 fathers' year of birth notably accelerates the marriage of single mothers. This might be related to a possible correlation between not knowing ornot being willing to report a father's year of birth, and less-demanding expectations of a potential husband.

Table 3.8: 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:	Marriage dissolution (1)	Marriage dissolution (2)	Marriage dissolution (3)	Marriage formation (4)	Marriage formation (5)
Firstborn 0-5 years old is daughter	0.76	0.75			0.56
	(0.18)	(0.18)			(0.21)*
Firstborn 6-18 years old is daughter	2.03		2.04		
	(0.89)*		(0.89)*		
Firstborn 0-5 years old	0.26	0.61			0.94
	(0.14)**	(0.80)			(0.46)
Firstborn 6-18 years old	0.32	1.26			
	(0.48)		(1.71)		
Firstborn daughter	0.53				
				(0.27)	
Father's occupation not known	1.45	1.44	1.49	0.79	0.78
	(0.42)	(0.42)	(0.43)	(0.28)	(0.27)
Mother's occupation not known	0.99	0.99	0.97	1.79	1.77
	(0.44)	(0.45)	(0.43)	(1.24)	(1.21)
Father's birth year not known	0.56	0.55	0.57	2.34	2.33
	(0.27)	(0.26)	(0.27)	(0.84)***	(0.84)**
Mother's birth year not known	1.40	1.49	1.34	0.47	0.61
	(1.22)	(1.31)	(1.16)	(0.46)	(0.49)
N of (marriage) spells	900	900	900	197	197
N of (marriage-)year observations	5,882	5,882	5,882	919	919
Log-likelihood	-477.31	-478.65	-477.85	-241.97	-241.05
Socio-demographiccontrols	Yes	Yes	Yes	Yes	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 4 for columns(1)-(3) and Table 5 for columns (4)-(5).

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