CERGE
Center for Economics Research and Graduate Education
Charles University Prague

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

**Sergii Maksymovych**

Dissertation

Prague, August 2023

**Sergii Maksymovych**

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

Dissertation

**Prague, August 2023**

Dissertation Committee

Štěpán Jurajda (CERGE-EI; chair)

Alena Bičáková (CERGE-EI)

Randall K. Filer (The City University of New York)

Patrick Gaulé (University of Bristol)

Daniel Münich (CERGE-EI)

Mariola Pytliková (CERGE-EI)

Referees

Jan Kabátek (Melbourne Institute of Applied Economic and Social Research)

Ján Palguta (CUNEF Universidad)

# Table of Contents

# Abstract

The first chapter examines how household living conditions are related to alternative allocations of control over decision-making in the household. The degree of control exerted by a particular individual over different decisions and household living conditions are taken from responses to an extensive multi-national household questionnaire. 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. These findings support the notion that effects of social assistance targeted at women might actually not be driven primarily by female empowerment.

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). First, we use recent EU-SILC data from several Balkan and Scandinavian countries to confirm that the gender of the firstborn predicts the likelihood of a given family having three children or more — a common measure of parental gender preference. We confirm son preference in certain Balkan countries and daughter preference in Scandinavian countries. Both having a first child of the preferred gender and of the more costly gender can decrease the probability of having three or more children because parents may already be content or may lack sufficient resources, respectively. Next, we use information on household consumption to differentiate the two explanations. We argue that under the differential cost hypothesis, parents of children of the more

costly gender should spend more on goods for children and less on household public goods, as well as on parental personal consumption. In contrast, having children of the preferred gender should increase spending on household public goods since such families have higher marriage surplus and are more stable. 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. The first two findings are in accord with findings in the literature. The third finding is specific to the Russian context. My analysis is based on the Russia Longitudinal Monitoring Survey (RLMS-HSE) data for the 1994-2018 period. I estimate complementary log-log (cloglog) regressions of divorce and marriage (for single mothers) on firstborn gender, age, and a set of household socio-demographic characteristics. My findings support the conclusion that the impact of children's gender on family living arrangements depends on family socioeconomic conditions and thus has a different character and magnitude in different contexts.

# Acknowledgments


When as a youth I waxed more bold,<br>
Time *strolled*<br>
When I became a full-grown man,<br>
Time RAN<br>
*Henry Twells, Times Paces*

My years at CERGE-EI seem to have passed by in a blink of an eye, filled with eye-opening revelations, rewarding academic experiences, and valuable professional contacts. Completion of my studies would not have been possible without the support of dedicated professionals and supporters to whom I owe a huge debt of gratitude.

First of all, I am deeply indebted to my supervisor Štěpán Jurajda. Štěpán provided me with expertise and patience all along, opening me to tremendous possibilities of an evidence-based data-driven approach to economics and social sciences in general. Whenever I requested feedback, he responded promptly and to the point. He offered constructive and far-sighted advice. As it always turned out, I could not do better than to pay heed to it. I am also extremely grateful to Alena Bičáková. Alena first supervised my research as a chair and then as a committee member. She has been generous with precious guidance and sympathy. As the DDGS, effectively facilitated the submission of my thesis. Further, I express my deepest appreciation to other members of my committee. Randall Filer and Patrick Gaulé contributed insightful, detailed and plentiful comments on different drafts of chapters comprising this thesis. Daniel Münich made helpful remarks about the enhancement of my study and opened the way to obtaining research data. Mariola Pytliková put forth useful propositions on strengthening my argument in the thesis.

I must also thank members of DPW and DW committees, WP editors, participants of conferences and a brown bag seminar. In particular, my thanks go to Nikolas Mittag, Byeongju Jeong, Jan Zápal, Filip Pertold, Sofia Tommlerová, Luděk Kouba, and Ihor Kulyniak. My gratitude also extends to Krešimir Žigić and Levent Çelik. Consultations with them helped me to recognize my comparative advanatages. Scientists who presented their research at CERGE-EI seminars helped me to understand the nuts and bolts of doing research in economics. Alex Armand, Max Weber, Simon Clark, and Maksym Brukhanov helped even more by commenting on my research in person. I also wish to thank specialists at the Academic Skill Center, especially to Deborah Nováková, Andrea Downing, and Robin-Eliece Mercury. They offered highly constructive recommendations on academic matters, academic writing being just one of them. All in all, my teachers have created strong incentives to constantly expand knowledge of methods and the subject matter in any studied area.

I would also like to extend my gratitude to those who created exceptionally favorable conditions for studies and research. I am indebted to specialists at the Jan Kmenta CERGE-EI Library, above all to Tomáš Pavela, Lucie Vasiljevová, and Jana Koudelková. They have ensured access to the superlative library collection and cooperated in finding sources, obtaining data, and building up information literacy deserving utmost appreciation. I must also thank specialists at the SAO, most of all, Tereza Kulhánková, Eva Kellnerová, Iva Havlíčková, and Lenka Miklánková. They provided unbroken counselling and support that have been instrumental in bringing my studies to completion. Next, I appreciate help from Grayson Krueger with final editing and help from Anna Papariga with arranging the defense. Furthermore, I must give a lot of credit for my dissertation to people who maintained superb educational infrastructure at CERGE-EI, at Charles University and in Prague as well as to those who contribute to the existence of the functional markets of academic literature. Needless to say, I am thankful to Charles University and the CERGE-EI Foundation for financially supporting my studies and teaching.

Last but not least, I extend my sincere thanks to my parents, Alla and Mykola, for their unfailing commitment and unfaltering faith. There are others who spared me attention, time, and concern and are not mentioned here. Altogether, I am grateful to everyone who contributed to what makes me tick and who made my time at CERGE-EI precious.

Czech Republic, Prague — Sergii Maksymovych
August 2023

# 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 family characteristics are shaped by constraints on decisions made by household members. I find a certain uniformity in how the collective nature of decision making in households mediates the relationship between constraints and family characteristics. I consider two constraints on the decisions of household members: the assignment of decision control over different spheres in the household and the 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 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 employ these measures simultaneously for husbands and wives, whereas most studies on the subject examine 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, both female and male control of decisions are concurrently correlated with material conditions of the household. Second, predominantly male control over decisions and, even more so, predominantly female control, is associated with poorer material conditions of the household than when control of household decision making is balanced. The first fact, together with results in the related literature, indicates that using income transfers targeted to women as a proxy variable for female control might not be plausible. The second fact indicates that income transfers might not affect household outcomes only through female empowerment, but their effects are actually driven by increased collaboration of household members, i.e., by an increase in balanced control. Further, 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 the validity of not only the unitary model of the household, but also the collective model which represents household decision making as a tug-of-war game. A more realistic model of the household, which could better account for the observed effects of income transfers, should reflect the collective nature of decisions in the household.

In the second chapter, my co-authors and I aim to identify which of two possible causes of gender preference is more prevalent in Balkan and Scandinavian countries: 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 that allows 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 these steps with my 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: Denmark, Norway, and Sweden. Next, we examine which of the two explanations - gender bias or differential costs - is more prevalent 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 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 do so. 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. As in my first chapter, the second chapter shows that considering the household as a single decision-making entity is not sufficient to explain the 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's age. I improve upon earlier studies in several ways. First, I use data from an 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 that uses longitudinal data and controls for the age of children examines only family dissolution (not 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 it has been among the countries with the highest reported divorce rates for decades. I focus on children in two different age ranges, pre-school (0-5) and school-age children (6-18). I group children this way because I expect that, based on the peculiarities of the Russian context, there will be differential effects of preschool sons and daughters on marriage formation and dissolution.

I find that 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 latter effect accords with a previous study, while the former 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, the complexity and diversity of social phenomena might limit the applicability of deductive reasoning and corroborative techniques. As A. Deaton (Deaton 2010, p. 452) put it, "Technique is never a substitute for the business of doing economics." This is why accumulation of useful knowledge and understanding might require "abduction" (a concept introduced in Heckman and Singer (2017)). This means looking for consilience across bodies of evidence and across studies. It also means revising the models, hypotheses, and data analyzed to provide defensible explanations for surprising phenomena. In my work, I have followed this approach to show that the choices of households, elementary units in the economy, are determined to a large extent by continually evolving conditions rather than by particular generic principles.

# Chapter 1

# Decision making in the household and material deprivation

## 1.1 Introduction

The problem of optimal organization of households has been considered since antiquity (Spiegel 2013) and much modern research attempts to establish links between the allocation of control in the family and households' economic situation (Pollak 1985). Specifically, recent work focuses on the nexus between the extent of women's control and household living conditions (Duflo 2012). Most of the evidence suggests that, compared to income or assets in the hands of men, income or assets in the hands of women are associated with improvements in child health and with higher household expenditure shares on nutrition, health, and housing (Bobonis 2009; Duflo 2003; Lee and Pocock 2007; Lundberg, Pollak, and Wales 1997; Lundberg and Ward-Batts 2000; Hoddinott and Haddad 1995) [1] The mechanisms underlying these associations remain an active research area (Lundberg and Pollak 2007). Women's income is often interpreted as a proxy for women's control

[1] There are also several studies that do not find a strong positive association between women's income and household living conditions; e.g., Braido, Olinto, and Perrone (2012), Haushofer and Shapiro (2016), Thomas (1990).

over decision-making, but the nature and extent of measurement error remains unclear (Cameron and Trivedi 2005).

In contrast to other studies, I use a novel approach and obtain additional findings on the subject. Namely, in this paper 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 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 condition of the household compared to balanced control of household decision making. [2]

These findings are important for the interpretation of the effects of windfall income transfers [3] directed to women on household outcomes. These effects are often attributed to increased female control [4] (Bobonis 2009). However, such interpretation would be correct if the windfall income was a suitable proxy variable for the female control. It is true if the windfall income is correlated only with the female control and not with other, potentially omitted, determinants of household outcomes. But, Ashraf (2009) suggests that windfall income transfers also impact men's behavior: men negotiate the use of transfers (more or less inten-

[2] In Section 4, I indirectly test for the possibility that this association is a consequence of reverse causality and reject this notion.

[3] Windfall income 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).

[4] Some studies, however, attribute this finding to the "labelling effect" or "spending inertia" (Lundberg and Pollak 2007).

sively depending on who receives the transfer and on the allocation of control in the household before transfers [5] ) and men may adjust their contributions to the household budget depending on how the transfer is spent. Moreover, Natali et al. (2016) find that a cash transfer increases balanced control over a number of household decisions, but does not increase solely female control [6]. Windfall income may thus be correlated with two unobserved variables influencing household outcomes: female control and male control. Further, the present study implies that both predominantly female control and predominantly male control predict negative household outcomes. This underlies a need to measure both the male and female control separately. Most of the existing work implicitly assumes that changes in female control are necessarily accompanied by equivalent and opposite changes in male control [7]. 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, 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.[8]

In order to interpret my foregoing results more broadly, I address limitations of my research design. Specifically, the EU-SILC data does not allow me to study the association of windfall income and female and male control. Moreover, my OLS

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

[6]Another study (Haushofer and Shapiro 2016) finds that cash transfers do not effect female empowerment in households, but their measure of female empowerment is based on reported instances and attitudes to home domestic violence.

[7]This is the case in the framework of collective models; see, e.g., Almas et al. (2015).

[8]The available evidence on the impact of cash transfers on "household dynamics" is scarce and mixed (Simon 2019). Some studies find that men become more involved in household matters when their wives receive cash transfers (Ambler and Brauw 2017) while other studies find that cash transfers reinforce pre-existing gender roles (Concern Worldwide 2011). This occurred, however, in communities where men traditionally used to make final decisions on household matters while women were delegated the managing of household finances. All in all, the effect ultimately depends on the design of the intervention as well as on peculiarities of the context (Bastagli et al. 2016). Nevertheless, researchers broadly agree that the design of cash-transfer interventions should take into account the responses of men which largely determine the efficiency of interventions (Arnold, Conway, and Greenslade 2011).

results could be affected by measurement error or reverse causality [9]. 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 and the gender-gap in weekly work-hours. The two 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 (in terms of the effect direction) supporting the notion that the mode of decision-making affects the material status of households.

Further, I present evidence on the association between decision-making in the household and income pooling. The concept of household income pooling arises in two strands of empirical literature: the literature that tests the unitary model of the household, and in the socio-economic literature that explores 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, it means that the budget constraint of the household contains the sum of individual incomes 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 person of 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 individual incomes of partners do 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 incomes to a "common pot", "kitty", etc (Pahl

[9]For example, when one of spouses "screws up", another takes full control over household matters. In this case the predominant individual control is the outcome of deteriorating household condition. And the coefficient on balanced decision making will be positively biased. But, the reverse causality can also go in opposite direction. For instance, when a partner having predominant control "screws up", another one steps in the decision making rather than completely retakes it. In this case the balanced control is the outcome of worsening household condition. And the coefficient on balanced decision making will be negatively biased.

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. That is because they can satisfy their needs by taking money from the common pool if the predominantly deciding household member does not care about them. When incomes are 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 situation is to arise, I check if individualized decision making is also accompanied by individualized family finances (i.e. no income pooling).

I find households that pool income are more likely to use a more balanced decision-making mode. This finding is important to both previously mentioned strands of literature. Namely, 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 incomes, thus violating the assumption of the unitary model. Those who pool incomes are also more likely to decide together. Therefore, they resemble households in the unitary model [10]. It may be that the recurrent rejections of income pooling in the literature could be driven by households that do not pool incomes.

## 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 ref-

[10]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.

erence 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 excepts 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 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.

Country samples can arguably be regarded as representative of respective populations. 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, constently high non-response 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 decsions on everyday shopping (25.2%) and decsions on purchases for children (33.4%). Non-response rate for the primary variables is about 1% (see also Eurostat, 2016).

## 1.3 Empirical strategy

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, Guio, and Marlier 2010). I do not use all 11 criteria, but only 7. My measure is equal to the sum of 7 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 [11]. The distribution of households by the sum of dummies is shown in Figure 1.A.1. If the sum of these dummies for a given household exceeds 2, the composite measure of material deprivation takes value 1 [12].

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" [13]. 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 described below. Consistent responses constitute more than 90% of all responses to any considered question. I focus

[11] 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.

[12] The results do not change much qualitatively nor quantitatively if the threshold for the sum of component dummies is equal to 3

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

on responses about the following five decisions [14] : '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 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 1.A.1 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" (Almas 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. Almas 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 labelling 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.' [15].

[14]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.

[15]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 (Almas et al. 2015) is based on the collective household model. This model implies that an

Turning to the empirical specification, let $\mathbf{Dec}_i$ 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 the material deprivation indicator in question for a household $i$, $y_i$. I estimate the following empirical model:

$$y_i = \boldsymbol{\beta}\mathbf{Dec}_i + \boldsymbol{\alpha}\mathbf{X}_i + \epsilon_i \tag{1.1}$$

where $\mathbf{X}_i$ is a vector of respondent, spousal, and household characteristics, and $\epsilon_i$ is the error term. The null-hypothesis is $H_0 : \boldsymbol{\beta} = 0$. The methods used for estimation are OLS and 2SLS. Concerning the covariates $\mathbf{X}_i$, the main ones included are: family income, number of children, number of daughters, length of cohabitation of spouses, living in a rural area, being unemployed, employment status, hours spent on job market work, 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 the use of specification 1.1 is contained in Appendix 1.B.

Equation 1.1 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 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.

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).

I use a specification with instrumental variables because 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 an overly large coefficient in absolute value. If, however, one of the 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 an overly small coefficient in absolute value. Regarding the instrumental variables, I use the rate of enrolment 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, the accessibility of institutional pre-school childcare has been found to have significant consequences for female activity status. Several recent studies mention this (Bičáková 2016; Kališková, Münich, and Pertold 2016; Bičáková and Kališková 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. For this reason, she does not have enough control over family finances, in particular, over decisions on the 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 difficult to see any other channel through which it could influence household material deprivation measures once activity status and the incomes of household members are controlled for [16].

Second, besides the activity status, employment opportunities also matter for the allocation of control in the family. For example, Morrill and Pabilonia (2015) show that increasing national unemployment rates reduces time spent together also in households with both partners working due to the 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) such 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 prospects will be less likely to resort to divorce in the case of household

[16] 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. As for the correlation, the Pearson r between regional involvement of children in pre-school establishments and women's long-term unemployment is 0.001. For comparison, the correlation between the regional involvement of children in pre-school establishments and the household composite material deprivation measure is 0.14. As for women's employment endogeneity, 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.

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.

For tractability of findings and compatibility with other studies, I analyze a specific sub-sample of households that includes only households composed of one couple of cohabiting partners, a man and a woman, with or without children. Most households report balanced decision-making (Table 1.A.2). 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 1.A.2, it is clear that about 90% of responses are consistent. Table 1.A.3 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% [17], the means of selected household characteristics tend to differ very little between the two groups. This is evident in Table 1.1.

**Table 1.1** 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 |

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

| | | |
|---|---|---|
| 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 |
| 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.

There appear to be no notable differences between households giving consistent and inconsistent responses on decision making. 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. This 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. Further, 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 that gave predominantly consistent responses.[18]

[18] 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.

## 1.4 Results

### 1.4.1 Baseline results

The results of estimating the baseline specification of Equation 1.1 are presented in column 1 of Table 1.2 [19]. 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 (Mills and Begall 2010; Voicu, Voicu, and Strapkova 2007; Aguero and Marks 2011)).

The baseline results exhibit three noticeable features. First, the share of women's income is negatively associated with material deprivation[20]. This result is in line with the findings of other studies [21]. 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

[19]Results for seven constituent indicators are presented in Table 1.A.4

[20]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.

[21]This is result is not driven by the presence of unemployed women in the sample. It also holds for the subsample of employed women

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.

**Table 1.2** Mode of Decision-making and Composite Material Deprivation

| Explanatory vars. | Dep. var.: Material Deprivation | | |
|---|---|---|---|
| | (1) OLS | (2) OLS | (3) IV |
| Woman takes control over 4-5[22] decisions | 0.071 | 0.054 | |
| | (0.010)*** | (0.010)*** | |
| Man takes control over 4-5 decisions | 0.024 | 0.023 | |
| | (0.012)** | (0.012)** | |
| Woman takes control over 2-3 decisions | 0.026 | 0.026 | |
| | (0.005)*** | (0.005)*** | |
| Man takes control over 2-3 decisions | 0.006 | 0.005 | |
| | (0.008) | (0.008) | |
| Control balanced between partners | | | -0.31 |
| | | | (0.18)* |
| Woman takes control over 4-5 decisions*Men's unemployment | | 0.21 | |
| | | (0.035)*** | |
| Regional gender gap in unemployment | | | 0.07 |
| | | | (0.024) |
| Regional share of employment in hi-tech industries | | | -0.143 |

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

Table 1.2 (continued)

| Explanatory vars. | Dep. var.: Material Deprivation | | |
|---|---|---|---|
| | (1) OLS | (2) OLS | IV (3) |
| | | | (0.016)*** |
| Share of population having access to broadband internet connection in a region | | | -0.001 |
| | | | (0.000)*** |
| Regional rate of long-term unemployment | | | 0.018 |
| | | | (0.002)*** |
| Share of woman's income in total household income | -0.024 | -0.06 | -0.045 |
| | (0.014)* | (0.01)*** | (0.016)*** |
| Woman responds the questionnaire | 0.030 | 0.030 | 0.02 |
| | (0.003)*** | (0.003)*** | (0.008)*** |
| Number of children | 0.046 | 0.046 | 0.045 |
| | (0.003)*** | (0.003)*** | (0.003)*** |
| Woman's age | -0.002 | -0.002 | -0.001 |
| | (0.000)*** | (0.000)*** | (0.000)*** |
| Man's age | -0.001 | -0.001 | -0.002 |
| | (0.000)*** | (0.000)*** | (0.001)*** |
| Woman has tertiary education | -0.121 | -0.120 | -0.116 |
| | (0.006)*** | (0.006)*** | (0.008)*** |
| Man has tertiary education | -0.129 | -0.129 | -0.122 |
| | (0.006)*** | (0.006)*** | (0.011)*** |
| Woman has secondary education | -0.053 | -0.053 | -0.047 |
| | (0.005)*** | (0.005)*** | (0.006)*** |
| Man has secondary education | -0.061 | -0.061 | -0.051 |
| | (0.005)*** | (0.005)*** | (0.007)*** |

Table 1.2 (continued)

| Explanatory vars. | Dep. var.: Material Deprivation | | |
|---|---|---|---|
| | (1) OLS | (2) OLS | IV (3) |
| Woman is employed full-time | -0.068 | -0.068 | -0.064 |
| | (0.007)*** | (0.007)*** | (0.008)*** |
| Man is employed full-time | -0.099 | -0.097 | -0.096 |
| | (0.007)*** | (0.007)*** | (0.008)*** |
| Woman is employed part-time | -0.056 | -0.056 | -0.039 |
| | (0.009)*** | (0.007)*** | (0.01)*** |
| Man is employed part-time | -0.022 | -0.021 | -0.028 |
| | (0.015) | (0.015) | (0.018) |
| Woman is self-employed | -0.056 | -0.057 | -0.052 |
| | (0.007)*** | (0.007)*** | (0.008)*** |
| Man is self-employed | -0.107 | -0.106 | -0.119 |
| | (0.006)*** | (0.006)*** | (0.007)*** |
| Log woman's earnings | 0.003 | -0.01 | 0.003 |
| | (0.001)*** | (0.001)*** | (0.001)*** |
| Log man's earnings | 0.003 | 0.003 | 0.003 |
| | (0.001)*** | (0.001)*** | (0.001)*** |
| Log household disposable income | -0.19 | -0.18 | -0.18 |
| | (0.003)*** | (0.003)*** | (0.004)*** |
| Own accommodation | -0.12 | -0.12 | -0.123 |
| | (0.005)*** | (0.005)*** | (0.005)*** |
| Densely populated area | 0.004 | 0.004 | -0.002 |
| | (0.004) | (0.004) | (0.004) |
| Partnership lasts 5 years | 0.022 | 0.022 | 0.02 |
| | (0.006)*** | (0.006)*** | (0.007)*** |
| Partnership is not official | 0.016 | 0.016 | 0.019 |

Table 1.2 (continued)

| Explanatory vars. | Dep. var.: Material Deprivation | | |
|---|---|---|---|
| | (1) OLS | (2) OLS | IV (3) |
| | (0.007)** | (0.007)** | (0.009)** |
| Couple separated | 0.07 | 0.065 | 0.06 |
| | (0.03)** | (0.03)** | (0.03)* |
| Couple in divorce | 0.05 | 0.046 | 0.035 |
| | (0.01)** | (0.015)** | (0.017)** |
| Number of daughters | -0.003 | -0.003 | -0.003 |
| | (0.004) | (0.004) | (0.004) |
| Country dummies | Yes | Yes | Yes |
| $R^2$ | 0.25 | 0.25 | 0.21 |
| $N$ | 62,420 | 62,420 | 55,482 |

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

*Note:* 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 1.1. 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 1.4.3. 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 rate of involvement of 4-year-olds in formal childcare and the gender gap in weekly work hours.

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 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 1.1. The results of controlling for potential confounders as well as of 2SLS estimation are reported next.

### 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 1.1. The results of the estimations are shown in column 2 of Table 1.2. 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 1.1 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.

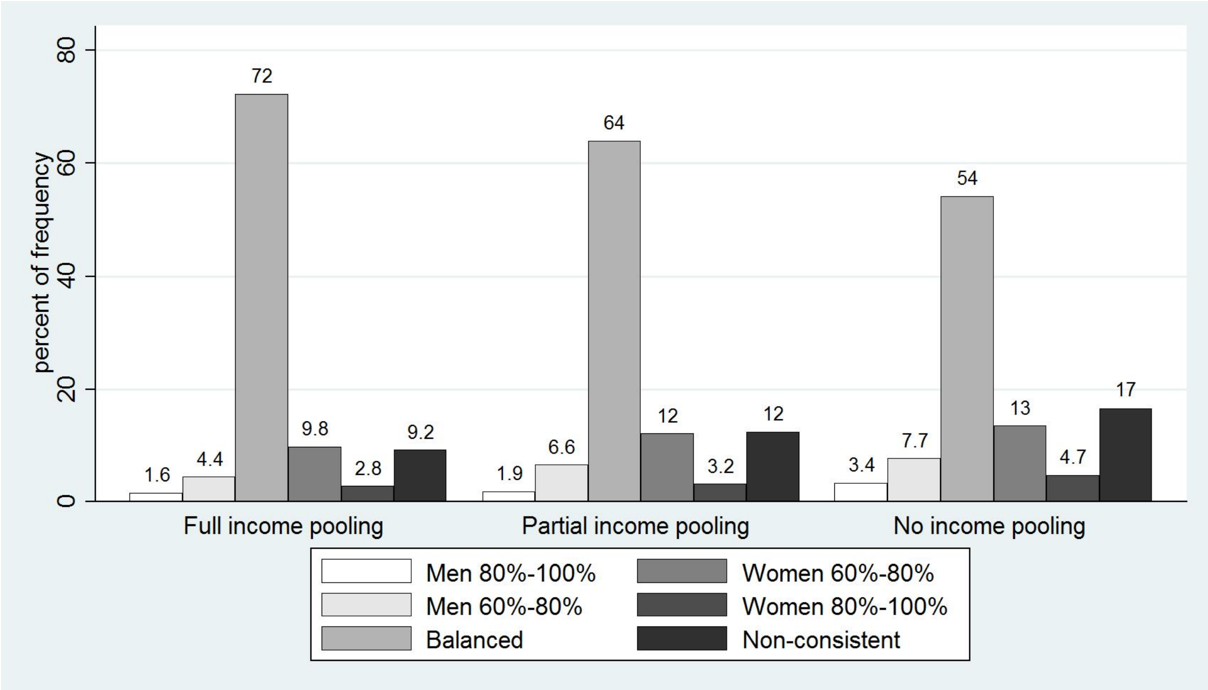



**Figure 1.1:** Shares of families by decision-making mode across the regimes of family finances

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 1.2. 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. Estimated coefficients on decision-making variables and on corresponding interaction terms with different household characteristics are presented in Table 1.A.5

### 1.4.3 Instrumental variable 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 1.1 so that the RHS contains only one dummy for balanced decision-making rather than four dummies for remaining modes.

The results of 2SLS estimation are shown in column 3 of Table 1.2. 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.[23] 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

[23]Moreover, Card (1999) explains the fact that 2SLS estimates of returns to schooling persistently exceed OLS estimates by treatment heterogeneity, i.e. returns to schooling are higher among those groups whose schooling decisions are more likely to be affected by structural innovations in the schooling system. Similarly, in my case those families that are most likely to switch either to a different mode of decision making because of lacking childcare facilities or to intra-household specialization because of different market work intensity for men and women could also be teetering on the brink of material deprivation. Thus, shifts in the distribution of responsibility would more likely plunge them into poverty.

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

This paper establishes 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 Briere and Rawlings 2006), perhaps should not be unambiguously preferred to household-specific targeting.

## 1.A Appendix A

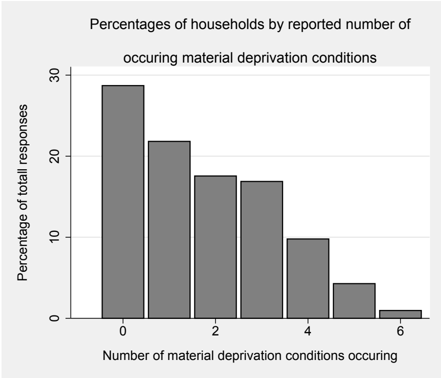



Figure 1.A.1

**Table 1.A.1** 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 |

**Table 1.A.2** 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 1.A.3** 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.

**Table 1.A.4** 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<br>(0.012)*** | 0.046<br>(0.007)*** | 0.053<br>(0.017)*** | 0.066<br>(0.010)*** | 0.034<br>(0.007)*** | 0.070<br>(0.010)*** | 0.049<br>(0.009)*** |
| A man takes control over 4-5 decisions | 0.026<br>(0.014)** | 0.031<br>(0.008)*** | 0.029<br>(0.022) | 0.039<br>(0.012)*** | 0.013<br>(0.008)* | 0.021<br>(0.012)* | -0.002<br>(0.012) |
| A woman takes control over 2-3 decisions | 0.011<br>(0.005)*** | 0.024<br>(0.003)*** | 0.024<br>(0.008)*** | 0.023<br>(0.005)*** | 0.012<br>(0.003)*** | 0.024<br>(0.005)*** | 0.011<br>(0.005)** |
| A man takes control over 2-3 decisions | 0.014<br>(0.007)** | 0.025<br>(0.005)*** | 0.030<br>(0.012)*** | 0.005<br>(0.007) | 0.005<br>(0.005) | -0.007<br>(0.007) | -0.037<br>(0.007)*** |
| A woman responds the household questionnaire | -0.003<br>(0.003) | 0.006<br>(0.002)*** | -0.002<br>(0.005) | 0.041<br>(0.003)*** | 0.002<br>(0.002) | 0.041<br>(0.003)*** | 0.029<br>(0.003)*** |
| Woman's income share | -.004<br>(0.007) | -.006<br>(0.004) | -.015<br>(0.008)* | -.006<br>(0.007) | .000<br>(0.004) | .000<br>(0.007) | .010<br>(0.006)* |
| N of hhs. | 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.

**Table 1.A.5** Household Economic Outcomes and the Mode of Decision Making Interacted with Household Characteristics

| Modes of decision-making relative to balanced | Interaction variables | | | | | | |
|---|---|---|---|---|---|---|---|
| | Tertiary education | Hours of employment | Hours of housework | Man's long-term unemployment | Length of cohabitation | Income pooling | Unregistered partnership |
| A woman takes control over 4-5 decisions | 0.072<br>(0.011)*** | 0.056<br>(0.013)*** | 0.030<br>(0.025) | 0.054<br>(0.010)*** | 0.12<br>(0.03)*** | 0.12<br>(0.02)*** | 0.064<br>(0.011)*** |
| A man takes control over 4-5 decisions | 0.026<br>(0.013)** | -0.01<br>(0.017) | 0.026<br>(0.018) | 0.024<br>(0.012)** | 0.08<br>(0.03)*** | 0.078<br>(0.023)*** | 0.021<br>(0.013)* |
| A woman takes control over 2-3 decisions | 0.028<br>(0.006)*** | 0.024<br>(0.007)*** | -0.020<br>(0.014) | 0.017<br>(0.006)*** | 0.03<br>(0.01)** | 0.040<br>(0.011)*** | 0.027<br>(0.006)*** |
| A man takes control over 2-3 decisions | 0.004<br>(0.009) | -0.007<br>(0.012) | 0.001<br>(0.012) | 0.003<br>(0.008) | 0.003<br>(0.018) | 0.015<br>(0.014) | 0.004<br>(0.008) |
| A woman takes control over 4-5 decisions*(Variable in the column heading) | -0.0002<br>(0.030) | 0.001<br>(0.0005)** | 0.001<br>(0.0006)* | 0.21<br>(0.035)*** | -0.054<br>(0.031)* | -0.059<br>(0.022)*** | 0.080<br>(0.034)** |
| A man takes control over 4-5 decisions*(Variable in the column heading) | -0.008<br>(0.014) | 0.002<br>(0.0005)*** | -0.0002<br>(0.001) | -0.003<br>(0.073) | -0.072<br>(0.034)** | -0.077<br>(0.027)*** | 0.031<br>(0.043) |
| A woman takes control over 2-3 decisions*(Variable in the column heading) | -0.015<br>(0.035) | 0.0001<br>(0.0003) | 0.0009<br>(0.0004)*** | 0.179<br>(0.024)*** | -0.004<br>(0.015) | -0.018<br>(0.012) | -0.011<br>(0.019) |
| A man takes control over 2-3 decisions*(Variable in the column heading) | 0.006<br>(0.018) | 0.0005<br>(0.0003) | 0.001<br>(0.0007) | 0.06<br>(0.04) | 0.003<br>(0.019) | -0.014<br>(0.017) | 0.009<br>(0.026) |
| N of hhs. | 62,420 | 61,719 | 31,447 | 62,420 | 62,420 | 62,212 | 62,420 |
| $R^2 adj.$ | 0.25 | 0.25 | 0.25 | 0.25 | 0.25 | 0.25 | 0.25 |

*Notes:* Robust standard errors are in parentheses. * denotes statistical significance at 10 percent, ** at 5 percent, and *** at 1 percent.

## 1.B Appendix B

This appendix presents a simple stylized model borrowed with few changes from one related study (Andaluz and Molina 2007). The model has two purposes. First, it answers the research question in a simplified stylized framework. Second, it justifies the use of empirical specification 1.1 for data analysis. I shall explain why exactly this model was chosen. Then, I present the model.

Broadly, there are distinguished three general models of the household. Detailed presentation of these models and further references can be found, e.g., in Lundberg and Pollak (2008). The unitary model is the least suitable to answer my research question since it assumes only one decision-maker in a family. In the sample analyzed, observations in which only one spouse takes all the decision control happen very rarely. As for the collective model, it allows for more than one decision-maker, but it imposes the restrictive assumption of Pareto efficiency [24] on equilibrium household allocation. In this way, the collective model makes derivation of testable restrictions on the response of household demands to 'distribution factors' that affect the household allocation possible. Since this study does not aim to derive testable restrictions on household demands, it seems reasonable to place the intended empirical estimation into the framework of the cooperative bargaining model. This model relies on weaker assumptions than the collective model.

This particular specification can produce demand equations to be estimated using the available data under some additional common assumptions. In the cooperative bargaining model, there are two possible equilibrium states of a household.

[24] More precisely, this model requires that the household utility function has the so-called "Pareto property". It means that the function is strictly increasing in individual household utility. That is why to avoid confusion the household utility function is often called "welfare function" (analogous to Bergson-Samuelson social welfare function (SWF)) (Apps and Rees 2008). However, in most unitary models this property necessarily implies Pareto optimality or Pareto efficiency of household resource allocation, i.e. first-best allocation of household resources. This condition appears to be quite restrictive.

The first is a non-cooperative one [25]. Formally it is represented as the equilibrium in the Cournot game. This equilibrium is not necessarily Pareto-efficient. The second possible equilibrium state is a cooperative one which is determined through the bargaining between partners. Formally it represents a solution to the Nash-bargaining problem, characterized by Pareto-efficiency. The bargaining outcome completely determines intra-household allocation including consumption structure, work-leisure choice, and assignment of home tasks (Lundberg and Pollak 2008). The partners are referred to as the husband ($j = h$) and the wife ($j = w$) The objective of a partner $j$ is to maximize a utility function $W^j(U^j, U^{-j})$. The utility function $W$ itself depends on the own utility from consumption $U^j$ and on that of another partner $U^{-j}$. Each partner has a utility function of the form $W^j = U^j + U^{-j}$, with $s \in [0, 1]$ denoting the degree of altruism of the partners, which is the same for both of them. Let $U^j$ be the direct utility of individual consumption of a family member $j$ consuming an amount $c^j$ of the private good and amounts $q^h$ and $q^w$ of the public good provided by the husband and the wife respectively. The utility of individual consumption for each partner takes the following form:

$$U^j(c^j, q^h, q^w) = c^j + \alpha \mathrm{Ln} q^w; (j = h, w; 0 < \alpha < 1) \tag{1.B1}$$

The first is a non-cooperative equilibrium analogous to the one in (Lundberg and Pollak 1993) in which partners play a Cournot game. Each agent decides, given the decisions made by the other player, both consumption of the private good and contribution to the household public good. Formally, each partner solves the

[25] Terms "cooperative" and "non-cooperative" are used here in the sense of cooperative game theory, which assumes that players can make binding, costlessly enforceable commitments.

following optimization problem:

$$
\begin{aligned}
&\max_{c^j,q^j} W^j(U^j, \overline{U^{-j}}), \quad j = h, w \\
&\text{s.t. } c^j + p^j q^j = I^j \\
&\qquad q^{-j} = \overline{q^{-j}}
\end{aligned}
\tag{1.B2}
$$

where $p^j$ is the price of the household public good provided by $j$, $I^j$ is the income of individual $j$, and the price of the private good is normalized to one. The price $p^j$ is the shadow price of a unit of the public good. It is shadow because there is no market for household public goods. Typical examples are childcare, preparing meals, and maintaining accommodation. However, there are also other examples more suitable for the current research, like trying to save energy or putting effort in searching for lower prices or more effective solutions when buying furniture and appliances, or housing renovation. The second optimization constraint in 1.B2 means that the optimizing partner treats the amount of the public good provided by another partner as given. The solution to that problem is characterized by the following amounts of the public good provided by each partner:

$$
{q_{NC}^h}^* = \frac{\alpha(1+s)}{p_h} \quad {q_{NC}^w}^* = \frac{(1-\alpha)(1+s)}{p_w} \tag{1.B3}
$$

where $p_h$ and $p_w$ stand for prices of the public good for the husband and the wife respectively. Assuming that both partners face the same price of the public good [26], it is possible to write an expression for the total amount of the public good provided by the partners:

$$
{q_{NC}}^* = {q_{NC}^h}^* + {q_{NC}^w}^* = \frac{(1+s)}{p} \tag{1.B4}
$$

The second equilibrium considered is a cooperative one. In this case individuals

[26] This assumption comes at no cost because the results do not change, but it helps to simplify the math.

choose the allocation of goods which maximizes the product below:

$$N = [W^h(c^h, q^h, q^w) - W^h_{NC}(Z)][W^w(c^w, q^h, q^w) - W^w_{NC}(Z)] \\ \text{s.t. } p(c^h + c^w) + p^h q^h + p^w q^w = I^w + I^h = I \tag{1.B5}$$

the subscripts $C$ and $NC$ stand for "non-cooperative" and "cooperative" respectively. The non-cooperative equilibrium does not mean dissolution of a household but rather specialization of its members on separate spheres of activity according to conventional social norms with little coordination between them. The allocation which maximizes 1.B5 is called generalized Nash-bargaining solution. It is characterized by the following amounts of the public good:

$$q^{h*}_C = \frac{2\alpha}{p_h} \quad q^{w*}_C = \frac{2(1-\alpha)}{p_w} \tag{1.B6}$$

Since $p_h = p_w$, the total amount of the public good provided by the partners takes the following form:

$$q_C{}^* = q^{h*}_C + q^{w*}_C = \frac{2}{p} \tag{1.B7}$$

It is clear from Equations 1.B4 and 1.B7 that $q_{NC}{}^* < q_C{}^*$ [27] or the amount of the public good provided in non-cooperative equilibrium is lower than the amount of the public good provided in cooperative equilibrium. Using Equations 1.B4 and 1.B7, a total amount of the public good provided can compactly be rewritten in the following way:

$$q_i = \theta_0 + \theta_1 d_i \tag{1.B8}$$

where $d_i = 1$ if a household $i$ is in non-cooperative equilibrium and $d_i = 0$ if a household $i$ is in cooperative equilibrium and $\theta_0 = \frac{2}{p}$, $\theta_1 = -\frac{(1-s)}{p}$.

Then, I can argue that $q$ in the Equation 1.B8 constitutes a latent variable

[27] Again, if prices of public goods were not assumed to be equal, it is clear that each term in Equation 1.B7 is larger than the corresponding term in Equation 1.B4, so the conclusion holds.

underlying the binary indicator of material deprivation. That is because the public good is the total effort of partners to find the best allocation of household resources. The more such effort they contribute, the better the material condition of the household will be. Thus, Equation 1.B8 constitutes an underlying latent-variable formulation for Equation 1.1, while the mode of decision making is a proxy variable for $d$, the dummy for the cooperative bargaining equilibrium in the household.

# Chapter 2

# Parental gender preference in the Balkans and Scandinavia: gender bias or differential costs?

**(co-authored with Zurab Abramishvili and William Appleman)**

## 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 data for the Balkans. Specifically, we find a higher probability of having a third child among families with a first-born girl for a pooled EU-SILC 2004-2015 cross-sectional sample from four Balkan countries: Bulgaria, Croatia, Slovenia, and the Republic of Serbia.[1] We also confirm daughter preference for three Scandinavian countries, i.e. Denmark, Norway, and Sweden, which had been established previously by Andersson et al. (2006) and Hank and 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 and Welch 1976; Lundberg 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

[1]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/

the allocation of household resources.

As for the gender bias, we understand it as a fixed difference between parental utilities derived from having a son and having a daughter. This difference does not depend on family characteristics, is exogenous to decisions of parents, and likely stems from gender-related kinship organization norms in a given context. If gender bias 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, and 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 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 and Rose 2003a).

Regarding the differential costs explanation, it encompasses two cases. The first is the differential fixed costs of raising sons and daughters. The second is differential prices of children "quality" in respect of sons and daughters. The fixed costs of children in our paper are identical to "prices" of sons and daughters in Dahl and Moretti (2008). The difference between fixed costs of raising sons and daughters is similar to gender bias in that it is constant and independent of parental decisions. However, its existence is not related to the existence of gender bias. The size of fixed costs of raising children is based on the prevailing view of normal conditions necessary for proper development of children in the context in question. The differential price of children's "quality" (we prefer using "human capital" instead of "quality") arises in view of two factors in parental choices. The first factor is that parents value characteristics of human capital (academic achievements, sport successes, artistic performances, etc.) of sons and daughters differently (i.e., perceive them to be of different quality and, thus, have different marginal utilities). The second factor is that prices of acquiring particular attributes of human capital differ between sons and daughters. The first factor might look similar to gender bias because it is a part of parental preferences and the second factor might look similar to differential fixed costs because it is a part of parental constraints. However, these two factors jointly determine parental choice of outlays on children's human capital. Parental choice of outlays on children depends in theory on the ratio of the price of children's human capital to the marginal utility of children's human capital rather than on magnitudes of these two factors separately. We label this ratio as a the price of children's human capital [2] and regard it as an element of parental choice constraints. Therefore, gender bias characterizes parental preferences while differential costs characterizes parental choice constraints.

If the differential cost explanation is true, parents who have a child of a gender related to higher outlays should have fewer children thereafter, spend less on themselves (both parents), spend less on adult public goods, and spend more on children (i.e., have more children's items in their household). Moreover, parents having a child of a gender related to higher outlays should report higher sums needed to make ends meet. Conversely, if the gender bias explanation specified above is correct, parents will report lower sums because they should spend more on household public goods that exhibit economies of scale in consumption. Finally, if the differential human capital price explanation is true, parental individual consumption should be characterized by a specific pattern. Namely, fathers of sons in the Balkan countries should have higher personal consumption than fathers of

[2] A more precise label could be "a price of children's quality" (assuming that children's "quality" is measured in units of parental utility derived from having children). Still, we tend to not use this label because it is uncommon in the literature.

daughters. At the same time, mothers of sons in the Balkan countries should have lower personal consumption than mothers of daughters. For the Scandinavian countries, this should be vice versa. Under gender bias, however, the pattern of parental personal consumption should be diametrically opposite.

We test between the two explanations by checking which relationships hold for the household-level data. We find that Balkan households with first-born daughters replace furniture less frequently than households with first-born sons. Moreover, in households with first-born daughters, 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 children's home items in households with first-born daughters (at 5% significance level). Moreover, we do not find a systematic impact of the gender of their children on parental consumption. We argue based on conclusions in Lundberg (2005) and Lundberg and Rose (2003b), 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.

## 2.2 Literature review

The evidence on the impact of parental gender preference pertains to developing economies (Barcellos, Carvalho, and Lleras-Muney 2014; Jiang, Li, and Sanchez-Barricarte 2016; Altindag 2016) and developed economies (Dahl and Moretti 2008;

Andersson et al. 2006; Pollard and Morgan 2002; Brockmann 2001; Kabátek and Ribar 2020). Authors attribute this impact to parental preference for a particular gender of children. In developing economies, parents usually have more children (progress to higher parities) when their firstborn is a daughter (Filmer, Friedman, and 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 and Moretti 2008; Choi and Hwang 2015), but daughter preference in others (Andersson et al. 2006; Brockmann 2001; Kabátek and Ribar 2020).[3] 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 and De Jong 2010). Consequences include shorter breastfeeding period for girls (Jayachandran and Kuziemko 2011), worse health and nutritional status of girls (Arnold 1992), and biased sex ratios (e.g., Jayachandran (2017), Guilmoto and Duthe (2013)). In more developed economies, Kippen, Evans, and Gray (2006) and Dahl and Moretti (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[4] could lead to a female "underclass", because poorer parents would prefer daughters and richer ones prefer sons (Trivers and Willard 1973). Other 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, and Han-

[3] 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.

[4] Such technologies may include infanticide, sex-selective abortion, or poorer health care.

son 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 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 and Welch 1976; Lundberg 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., Jayachandran and Kuziemko (2011), Dahl and 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 and Kuziemko 2011) or explain it by a predilection (Dahl and 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 and 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 (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.[5] 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, and 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 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 and Rose 2003b).

Turning to the difference in costs of raising sons and daughters, the literature

[5]Figures 2.A.4 and 2.A.5 in the Appendix illustrate a more detailed explanation of the difference between the gender bias and the cost difference.

considers two cases.[6] 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 topic (van Praag and Warnaar 1997), which frequently calculate so-called *normative budgets*.[7] 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, and Stevens 2016). Available empirical evidence suggests that parental expectations are important for parental spending (Hao and 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 and 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 and Warnaar 1997), is theoretically plausible (Deaton and 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

[6] 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 and 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.

[7] 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 and Carlson 2010). The constant cost of children is also assumed in, e.g., Dahl and Moretti (2008), Hazan and Zoabi (2015), Leung (1991), Sienaert (2008), Bojer (2002); and Raurich and Seegmuller (2017).

method measures gender difference in costs of children relying upon the subjective scales method (Leyden approach) proposed and substantiated in van Praag and 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 and Doepke (2003); and Hazan and 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 have less siblings due to substitution of quality for quantity (Galor 2011; Aaronson, Lange, and Mazumder 2014).[8]

We use a set of home items as measures of parental investment (Cunha, Heckman, and 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 a gender associated with higher outlays 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 child having gender associated with higher outlays should report higher sums needed to make ends meet. However, if the gender bias explanation specified above is correct, parents

[8]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 2.A.3 in the Appendix shows.

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. We analyze only households in which the oldest child is, at most, 12 years old, which compulsory school age in all European countries.

The two causes considered, the gender bias and the differential costs, might actually be in play simultaneously, but our testing points to which cause salient in driving the estimates. We observe higher parental expenditure on daughters in countries with daughter preference, which is consistent a lower price of child human capital for daughters. Whereas the son bias drives son preference in countries where we observe it, outweighing the effects of a higher cost of daughters (which is, however, not as high as in daughter-preferring countries). Moreover, the cross-country correlation between our estimates of the gender preference and the cost difference is stronger than the correlation between our estimates of the gender preference and the conventional measures of gender equality (GGI, GDI, etc.), which arguably approximate gender bias. All these findings taken together indicate that gender preference across countries is more strongly determined by the cost difference than by gender bias. Therefore, a policy intended to neutralize gender preference effects would subsidize the costs of human capital for sons from families which are less well off.

## 2.3 An overview of selected studies on patriarchal and gender norms in the Balkans and Scandinavia

Recent studies report evidence of parental preference for the gender of children at birth in Albania and Kosovo (Guilmoto and Duthé 2013), Serbia (Battaglia, Chabé-Ferret, and Lebedinski 2021), North Macedonia, and Montenegro (UNFPA

2020). Battaglia, Chabé-Ferret, and Lebedinski (2021) explain their results by the persistence of traditional patriarchal values in Serbian culture. Broadly speaking, Kaser (2008, p. 9) argues that patriarchy has resurfaced on both sides of the Bosporus since the demise of Socialist and Kemalist regimes, tracing its historical roots to the Ottoman imperial past. According Kaser, the optimal framework in which traditional patriarchal structures could best develop and unfold was a tributary political and agrarian order, such as the Ottoman Empire. The Ottoman regime and economy were based upon warfare, military expansion, and tributes, and the Ottoman institutional framework fostered patrilineality, patrilocality, and patriarchally oriented customary law. This, in turn, led to a kinship organization based on main three principles: segmentary distance, genealogical distance, and generational distance. These specific patriarchal relations differentiate the Balkans and Asia Minor from the rest of Europe and the Middle East and, in Kaser's narrative, is extremely stable across time and space (Grandits et al. 2020). [9]

As for Croatia and Slovenia, Kaser (2008) does not include these two countries in his analysis. Available evidence, however, indicates that Croatia and Slovenia might possess common characteristics of patriarchal and gender norms. The two countries share a common socialist Yugoslav legacy with the Balkan countries studied by Kaser, who also argues that the socialist state in the former Yugoslavia reinforced patriarchal norms (despite nominally working to counter them) (Kaser 2008, pp. 146, 164) (see also Guilmoto and Duthé [2013]). Moreover, according to international rankings on the gender gap, Croatia is close to other Balkan countries (Morrica et al. 2019; Forum 2018).[10] Despite Slovenia standing higher on international gender gap rankings than other Balkan countries, the processes of "retraditionalization" and "domestication"[11] taking place in the region since the

[9] Beyond that, scholars working on a global taxonomy of family argue that families in Balkans and Anatolia belong to the same type (Grandits et al. 2020). Altindag (2016) also finds evidence of son preference in Turkey.

[10] Beyond that, publications in periodicals (e.g., [Milekic 2018]) highlight the public debate about patriarchal gender norms.

[11] This term refers to development of social norms prescribing women's confinement to home matters and subordination to a husband (Rogers 1980; Baker 1984; Nagel 2013).

demise of socialism have influenced Slovenia (Antončič, Celpak, and Ule 2018). The increased ratio of male to female births in the Western Balkans might be partially explained by reduced fertility, as is done by Jayachandran (2017) for the case of India. At the same time, Lerch (2018) finds that fertility in the Western Balkans declined to a lesser extent than in other CEE countries during the period of transition from state socialism to market economies.

Although Kaser (2008) studies Romania and Greece along with other Balkan countries, we do not analyze these two countries. With Romania and Greece included, our analyzed group of countries could be too heterogenous compared to the group of Scandinavian countries which are characterized by socio-cultural and linguistic similarities. Thus, we decided to focus on Slavic-speaking countries that share a socialist legacy and, with exception of Bulgaria, a Yugoslav legacy.

As for Scandinavian countries, international rankings have long placed them among the most gender-equal societies in the world. Studies focusing on gender perspectives in economic history shed light on how the current state of gender equality in Scandinavia and Nordic countries may have developed. Pylkkänen (2017) argues that Finnish folklore is highly egalitarian with respect to gender, and that its rural gender code was less differentiated or dichotomous than in other European cultures. Laws protecting women's property rights in marriage among the broad strata of population already existed in early modern times in Denmark (Dübeck 2017), Norway (Johansen 2017; Sandvik 2017), Sweden (Fiebranz 2017), and Iceland (Arnórsdóttir 2017). Schultz (2001) finds that changes in global prices of agricultural products in the second half of the 19th century increased the value of women's time relative to men's. Edvinsson and Edvinsson (2017) claim that "capitalist patriarchy" and, particularly, the breadwinner-homemaker household appeared in Scandinavian countries later, was weaker, and ended earlier than in the UK, the US, and the Netherlands. Studies cited elsewhere report daughter preference arising in Scandinavia during recent decades, and other authors observing daughter preference in Europe (Bereczkei and Dunbar 2002; Brockmann 2001) believe that is it driven by economic factors. Therefore, in Scandinavia, there should be no strong institutional basis for a parental gender preference similar to the one underpinning the son preference in the Balkans.

We formally test whether the explanations for a daughter preference in Scandinavia as being driven by economic factors and of a son preference in the Balkans as being driven by son bias are compatible with data. Concerning our analysis of Balkan countries, the increased ratio of male to female births in the Western Balkans might be partially explained by reduced fertility, in line with Jayachandran (2017) for the case of India. At the same time, Lerch (2018) reports that fertility in the Western Balkans declined to a lesser extent than in other CEE countries during the period of transition from state socialism to market economies. Therefore, the parity-two progression ratio is a suitable method to measure parental gender preference in the Balkan countries because relatively many families progress to a third birth.

## 2.4 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.[12] The Scandinavian countries are Denmark, Norway, and Sweden.[13] A primary goal of EU-SILC is

[12]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.

[13]These groupings of countries have been frequently used in previous studies. For instance, Estrin and 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 and Matysiak (2016) 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

to collect cross-sectional and 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) .[14]

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

countries. In a similar vein, Filmer, Friedman, and Schady (2009) pool HNS data into six subsamples by parts of the world and assume no difference in parameters between countries within groups.

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

available only for all children in the household together and not for each child separately .[15]

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 and Moretti 2008; Karbownik and Myck 2017; Ananat and 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 .[16] Our calculated sex-ratio for firstborns is 1.057, close to the commonly accepted value of 1.06 (Grech, Savona-Ventura, and Vassallo-Agius 2002) .[17] 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 2.1 contains descriptive statistics for selected household socio-demographic characteristics separately for all families and for cohabiting couples. Table 2.2 presents descriptive statistics on variables characterising different aspects of the household material condition. We use variables in Table 2.2 as dependent variables and variables in Table 2.1 as covariates.

[15] 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.

[16] 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 and Moretti 2008; Karbownik and Myck 2017) use the threshold of 12 years. Karbownik and 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 and Moretti (2008) find the 12-year cutoff conservative while Ichino, Lindstŕom, and Viviano (2011) and (Ananat and 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

[17] 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

**Table 2.1:** Descriptive statistics - demographics and labour market information.

| | Balkan countries | | | | Scandinavian countries | | | |
|---|---|---|---|---|---|---|---|---|
| | All families | | Married couples | | All families | | Married 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 | 37.56 | 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 education | (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) | (14.872) | (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 |
| hours of work | - | - | (16.689) | (0.162) | - | - | (12.762) | (0.106) |
| Household disposable | 20,469.770 | 265.421 | 20,982.732 | 214.079 | 64,070.609 | 325.596 | 65,957.259 | 450.271 |
| income (euros) | (15,431.683) | (141.036) | (15,550.905) | (150.036) | (57,680.462) | (447.583) | (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 012.

Amongst adult and household material deprivation characteristics, Table 2.2 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 two-parent 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 separated families .[18] Otherwise, the intact families do not appear to differ systematically from all families. That supports our decision to focus the analysis on intact families.

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

**Table 2.2:** Availability of selected items in the home environment for girls and boys.

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

Table 2.2 (continued)

| | Balkan countries | | | | Scandinavian countries | | | |
|---|---|---|---|---|---|---|---|---|
| | All families | | Married couples | | All families | | Married couples | |
| Dependent variables | Mean | Girl-boy difference | Mean | Girl-boy difference | Mean | Girl-boy difference | Mean | Girl-boy difference |
| pairs of properly fitting shoes (men) | 0.597 | -0.000 | 0.634 | 0.002 | 0.408 | -0.012 | 0.435 | -0.012 |
| | (0.490) | (0.007) | (0.482) | (0.007) | (0.492) | (0.007)* | (0.496) | (0.007)* |
| Replace worn-out 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 at least once a month (women) | 0.552 | 0.004 | 0.565 | 0.005 | 0.429 | 0.018 | 0.405 | 0.017 |
| | (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 a leisure activity (men) | 0.254 | -0.005 | 0.276 | -0.006 | 0.317 | -0.009 | 0.338 | -0.009 |
| | ( 0.435) | (0.006) | (0.447) | (0.006) | (0.465) | (0.007) | ( 0.473) | (0.007) |

Table 2.2 (continued)

| Dependent variables | Balkan countries: All families: Mean | Balkan countries: All families: Girl-boy difference | Balkan countries: Married couples: Mean | Balkan countries: Married couples: Girl-boy difference | Scandinavian countries: All families: Mean | Scandinavian countries: All families: Girl-boy difference | Scandinavian countries: Married couples: Mean | Scandinavian countries: Married couples: Girl-boy difference |
|---|---|---|---|---|---|---|---|---|
| *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 |
| | (0.382) | (0.007) | (0.363) | (0.007) | (0.118) | (0.003) | (0.113) | (0.003) |
| Two pairs of properly fitting shoes | 0.845 | 0.006 | 0.867 | 0.007 | 0.983 | 0.000 | 0.986 | -0.002 |
| | (0.362) | (0.006) | (0.340) | (0.006) | (0.128) | (0.003) | (0.118) | (0.003) |
| Fresh fruits and vegetables once a day | 0.866 | -0.010 | 0.885 | -0.006 | 0.982 | -0.003 | 0.983 | -0.003 |
| | (0.341) | (0.006) | (0.319) | (0.006) | (0.134) | (0.003) | (0.127) | (0.003) |
| One meal with 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 |
| | (0.365) | (0.006) | (0.345) | (0.006) | (0.108) | (0.002) | (0.103) | (0.002) |
| Books at home suitable for children's ages | 0.844 | 0.006 | 0.863 | 0.009 | 0.983 | 0.006 | 0.984 | 0.005 |
| | (0.363) | (0.006) | (0.344) | (0.006) | (0.131) | (0.003) | (0.126) | (0.003) |
| Outdoor leisure 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 ) |

Table 2.2 (continued)

| | Balkan countries | | | | Scandinavian countries | | | |
|---|---|---|---|---|---|---|---|---|
| | All families | | Married couples | | All families | | Married couples | |
| Dependent variables | Mean | Girl-boy difference | Mean | Girl-boy difference | Mean | Girl-boy difference | Mean | Girl-boy difference |
| 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 special occasions | 0.867 | -0.002 | 0.884 | 0.000 | 0.981 | 0.001 | 0.983 | 0.002 |
| | (0.339) | (0.006) | (0.320) | (0.006) | (0.137) | (0.003) | (0.129) | (0.003) |
| Invite friends over to play | 0.790 | 0.002 | 0.807 | 0.005 | 0.959 | 0.002 | 0.959 | 0.002 |
| | (0.408) | (0.007) | (0.395) | (0.007) | (0.198) | (0.004) | (0.198) | (0.004) |

*Note*: The statistics 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.

## 2.5 Empirical strategy

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, and Lleras-Muney 2014) [19] and spend more on household public goods (Lundberg 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., a gender associated with higher parental outlays, resulting in fewer additional or total births) work 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

[19] These authors also mention that in such households, mothers end their maternal leave earlier. Evidence from the US, however, suggests that fathers of sons tend to work less. At the same time, many authors find sons preferred in the US. The descriptive statistics for the pooled EU-SILC sample show that mothers of daughters actually work more when. daughters are the preferred gender. Nevertheless, a comprehensive testing of this implication for the EU-SILC data is beyond the scope of this paper.

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. [20]. 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 and Wu 2011), the number of children in a family (Dahl and Moretti 2008), the occurrence of divorce (Bedard and Deschenes 2005; Ananat and Michaels 2008), and the area of accommodations (Dujardin and Goffette-Nagot 2009) .[21]

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

[20]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.

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

to make ends meet. The same predictions should hold for parents of first-born sons in son-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 and Moretti (2008), Jayachandran and 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 (e.g., Galor (2011), Aaronson, Lange, and Mazumder (2014), de la Croix and 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:

$$y_i = \beta\textit{(First child girl)}_i + \boldsymbol{\alpha}\boldsymbol{X}_i + \epsilon_i \tag{2.1}$$

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 $\boldsymbol{X}_i$ 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, $\epsilon_i$, can be correlated. The specific set of variables that make up $\boldsymbol{X}$ 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 and Moretti 2008). In that case, parents who have bias towards any gender will progress to parity two independently of the gender of their first-born. 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 and Moretti 2008; Hank and Kohler 2000; Haughton and Haughton 1998; Larsen, Chung, and Das Gupta 1998; Clark 2000; Basu and De Jong 2010). We include higher degree polynomials in the mother's age to account for the conclusions reached by Yamaguchi and 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 (e.g., Wooldridge (2002)). Nevertheless, we also run Probit estimations to check whether our estimates are sensitive to an alternative functional form specification.[22] Since we expect observations not to be iid, but correlated within countries, we cluster the standard errors at the country level. We do not analyze all countries in the Balkans and Scandinavia but subsets of countries in each group in view of the availability of some variables and comparability of countries within each group. In such a case, Abadie et al. (2022) recommend using clustering at the country level.

[22]The Probit estimates correspond to OLS estimates in terms of impact direction and statistical significance.

### 2.5.1 Connection between the differential costs mechanism and household outcomes.

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).

As for the first cause (different one-time costs), it operates similarly to gender bias. This is because gender difference in the fixed costs of raising children does not depend on the behavior of parents. In the same way, the difference between utility of sons and utility of daughters under gender bias does not depend on the behavior of parents. In the case of differential fixed costs, predicting the effects of the gender of the first-born child is straightforward. Parents having a child of the gender characterized by higher fixed costs should spend less on household public goods and on personal consumption because they should have less money left after making necessary outlays on their child. Further, a child in such a household should have more children-specific consumption goods because parents make higher outlays on these goods (assuming that the prices of goods for boys and girls are not very different). Moreover, if differential costs are fixed costs, the conceptual distinction between gender bias and differential costs is clear-cut. The gender bias means difference between parental utilities of sons and daughters while fixed costs of sons and daughters are the same. Conversely, differential costs means different fixed costs of sons and daughters while the parental utilities of sons and daughters are the same. Again, if the two mechanisms are in operation simultaneously, our testing will show which one dominates.

The second of the two foregoing causes could be the difference in the price of human capital or, more precisely, in expenditures per unit of parental discounted utility derived from a child's human capital. It is convenient to consider two cases to show the conceptual difference between this cause and the gender bias. In the first case, the marginal parental utility of an additional unit of human capital of sons equals *ceteris paribus* that of daughters. Then, if the price of an additional unit of human capital is lower for daughters, parents of daughters will tend to spend more on their human capital than parents of sons spend on the human capital of sons. The reasons for that are provided in the next paragraph. In the second case, marginal parental utilities of additional units of human capital of sons and daughters are different while the corresponding prices are the same. We argue that the second case is equivalent to the first case with respect to the parental choice of optimal outlays on children. In other words, if a marginal parental utility of a daughter's human capital is two times higher than that of a son's, parents would decide on acquiring one additional unit of human capital exactly as if the price of the daughter's human capital was two times lower. For the case of the linear marginal parental utility function, the formal proof of this claim seems to be straightforward. The key factor at play here is that parents increase outlays on children's human capital until a ratio of a children's human capital price to a marginal parental utility of children's human capital (i.e., expenditure per unit of parental utility) is equal to that for other goods. Thus, in both aforementioned cases it is a ratio of the marginal utility of children's human capital to the price of children's human capital that determines the parental choice but not any of these two magnitudes separately. Therefore, we say that there is a differential price of children's "quality"[23] when the ratios in question differ for sons and for daugh-

[23]"Quality" might appear a suitable term because it incorporates both notions of children's human capital (characteristics such as academic success, sport achievements, or artistic prowess) and parental utility derived from children's human capital (how valuable are children's achievements and successes to parents). This term, however, is not common in the literature. For this reason, w refer to "differential price of children's human capital" assuming that children's human capital is measured not in observable characteristics (e.g., academic achievement) but in units that have equal value for parents of either sons or daughters, e.g. a currency unit of a daughter's

ters.[24] Moreover, the difference between parental marginal utilities derived from particular attributes of the human capital of sons and daughters might change from one attribute to another. For instance, parents might prefer a son's successes in sports to a daughter's and a daughter's academic achievements to a son's. That is why we consider a unified magnitude (i.e., the aforementioned ratio) which we label as the price of children's human capital and regard as an element of parental choice constraints. All in all, we consider three mechanisms - gender bias, differential fixed costs of raising children, and differential prices of children's human capital (or of children's "quality") - each corresponding to different exogenously given determinants of parental behavior. Gender bias is arguably driven by kinship organization norms related to gender, differential fixed costs arise from norms of consumption related to gender and prices of consumption goods, while differential prices of children's human capital could stem from the state of gender equality on the labor market and in the education system. We unite differential fixed costs and the differential price of children's human capital under one umbrella of differential costs because the operation of the two mechanisms primarily relies on the configuration of particular market prices that could, in principle, be influenced by policy measures. Within the differential price of children's human capital we do not distinguish between the case of differential parental utility derived from the specific characteristics of human capital of sons and daughters and the case of differential prices of these characteristics. This is because it is only the ratio of the two that matters for parental choices that we can observe. Testing the implications of the two aforementioned sub-mechanisms would likely require additional data along with a further study since they can hardly be found in the literature. All in all, the three mechanisms stem from different causes and constitute three

---

expected earning capacity.

[24] Lundberg (2005) considers the second case as an instance of gender bias. However, Lundberg's definition does not take into account the price of children's human capital, which is a key determinant of the parental choice of outlays on children. In addition to that, the case in which both prices and parental marginal utilities of children's human capital differ between sons and daughters is identical to the previous two cases because parents take into account only a ratio of the two magnitudes when they choose outlays on children's human capital.

different channels through which the gender of a first-born child influences household outcomes. Next, we discuss how differential prices of children's human capital operate and which household outcomes they are likely to influence.

The channel through which differential prices influence household outcomes is the 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 [25] is the availability of conditions and items at home which are necessary for normal child development (Cunha, Heckman, and Schennach 2010; Todd and Wolpin 2007; Juhn, Rubinstein, and Zuppann 2015) .[26] The expected effects of the first-born daughter are systematically presented in Tables 2.A.5-2.A.8. 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 countries and sons, respectively, in son-preferring countries. Under this assumption, parents having a child of a gender associated with higher parental outlays, 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 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 and Warnaar (1997).[27]

[25]The most common measure is the years of schooling conditional on household income.

[26]These variables are described in more detail in the second Appendix subsection

[27]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. 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. Specifically, in the case

### 2.5.2 Expected effects directions under gender bias

The expected effect directions are different for different dependent variables and for each group of countries. At start we present predictions for household-level variables. For Balkans, which have son preference, having a first daughter will cause parents to save money for later because they prepare to have more children, so we expect $\beta$ to be positive ($\beta > 0$) when the dependent variable $y$ is the ability to deal with unexpected expenditures (because parents of daughters have higher savings). Also, under son bias, having a daughter adds less to a marriage surplus than having a son. Thus, parents who have a daughter expect their marriage to be less stable and of a shorter duration than it would be if they had a son. That is because the opportunity cost of divorce is lower when marriage surplus is lower. Therefore, they invest less in household public goods because the expected period of receiving utility from them, which equals expected duration of the marriage spell, is shorter. In other words, we expect $\beta$ to be negative ($\beta < 0$) when $y$ is replacing worn-out furniture. Moreover, a family spending more on household public goods will have a higher well being than a family with a similar income that spends less on household public goods. The reason is economies of scale in consumption characterizing public goods. This means that households with first-born daughters in the Balkans will be less able to make ends meet (because they spend less on household public goods characterized by economies of scale in consumption). That is $\beta$ will be negative when $y$ is the ability to make ends meet, and, consequently, $\beta$ will be positive ($\beta > 0$) when $y$ is an amount of money needed

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 2.A.6

to make ends meet. For Scandinavian countries (in view of our findings on the parental gender preference), it should be vice versa in any of the four specifications. Turning to the individual parental consumption, we consider only the case when sons directly increase the utility of fathers (the point asserted in, e.g., Lundberg (2005)). Then, a standard bargaining model of the household predicts a shift of household resources from fathers to mothers. This occurs because boys provide a direct utility bonus to fathers, and intra-household bargaining will distribute some portion of that additional marital surplus to mothers. This redistribution could be observable through less consumption of private commodities by mothers of daughters as increased leisure among mothers of sons, or increased consumption of private commodities typically consumed by women (Lundberg and Rose 2003a). Moreover, according to Barcellos, Carvalho, and Lleras-Muney (2014), parents with a child of the more desired gender should work fewer hours, because they are less likely to expect to increase their family.

Regarding predictions for individual-level variables, when there is son bias in Balkan countries, fathers of sons should have lower personal consumption than fathers of daughters. That is because fathers agree to redirect part of their personal consumption to mothers since fathers receive a utility premium from having sons. Conversely, having a daughter will lead to a father to having higher personal consumption than having a son. That is to say, $\beta$ will be positive ($\beta > 0$) for any of the dependent variables measuring a fathers consumption (the six dummy variables) and, consequently, $\beta$ will be negative ($\beta < 0$) for any of the dependent variables measuring mothers consumption. As for weekly hours of work, it should be that $\beta > 0$ for both parents. For Scandinavian countries, it should be vice versa in any of the seven regression specifications if the daughter preference is driven by gender bias. Predictions about the sign of coefficients are systematized in Tables 2.A.7 and 2.A.8.

### 2.5.3 Expected effects directions under differential costs

#### Expected effects directions under differential fixed costs

We assume, in line with other studies, that 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 at a given time (Kornrich and Furstenberg 2007). Under this assumption, parents with a child of a gender that requires bigger parental outlays (because of higher fixed costs), 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. We measure the ability to make ends meet by a binary variable taking value 1 when a household is able to make ends meet. The predictions above follow from the fact that parents will have less financial means left after making outlays on children than will parents with a child of the less resourced gender. Predictions about household-level consumption formulated in the previous paragraph are opposite to those based on the gender bias mechanism. The expectations for the two groups of countries are presented in Table 2.A.7.

In addition to looking at the four household-level dependent variables mentioned above, we also look at a set of ten home items as measures of parental investment (Cunha, Heckman, and Schennach 2010), which are proxy variables for parental outlays on children. We construct a binary variable taking value 1 if all ten items are available in the household and 0 otherwise. If the difference in fixed costs means that sons and daughters have different needs for these items (as parents perceive it), we expect that $\beta < 0$ (i.e., daughters have less of home items) in the Balkan countries and $\beta > 0$ (i.e., daughters have more of home items) in the Scandinavian countries if the fixed cost difference mechanism applies.

### Expected effects directions under differential price of human capital

Now, we discuss the variable component of the difference between the costs of sons and daughters. We assume (as does, e.g., Galor (2011)) that the mechanism behind the different consumption and birth rates of households with first-born sons and daughters is that of substitution of quality for quantity of children. Specifically, in countries with daughter preference at birth, parents spend more on daughter quality and have fewer children after having one or more daughters. They do so because the marginal parental utility of goods and services provided for daughters is on average higher than that of goods and services provided for sons.

In this kind of scenario, daughters in daughter-preferring countries and sons in son-preferring countries should receive more parental resources. That is why predictions about household-level variables under differential prices of human capital are identical to those under differential fixed costs. In other words, we cannot distinguish between differential prices of human capital and differential fixed costs mechanisms using only available household-level data. Using such data, we can only tell whether it is compatible with a gender bias or differential costs broadly (without specifying whether it is fixed costs or the prices of human capital). However, the differential prices of human capital have specific implications for individual-level consumption.

First, we discuss the personal well-being of fathers under the gender difference in the prices of human capital of children. If the wage gap is in favor of men, fathers of children with cheaper human capital should enjoy higher personal consumption. That is because their earnings add more to the marriage surplus in terms of mothers utility[28] and, in the process of intra-household bargaining, mothers will be more willing to share household resources with fathers to the detriment of their own consumption, to motivate fathers to stay in the family. Therefore, fathers of sons in the Balkan countries should have higher personal consumption than fathers of daughters. At the same time, mothers of sons in the Balkan countries should have

[28]Even if spending time with mothers is an indispensable resource for raising daughters, mothers can afford to spend more time with daughters when fathers earnings are higher.

lower personal consumption than mothers of daughters. For the Scandinavian countries, this should be vice versa.

Household-level data on home items for children can also be used to throw light on a specific form that differential price of human capital might take. Namely, we consider two possible forms of the differential price of human capital. These are the technology hypothesis (parents have a comparative advantage in raising children of the same gender) and gender difference in returns to parental investments. When the technology hypothesis is true, sons in son preferring countries receive less of childrens home items than daughters (because they receive more of relatively expensive fathers time). However, when parental investments in home items have different returns for sons and daughters (as in the example of foreign language learning in Appendix 2.A.1), higher parental provision of home items for the child of the gender with higher returns should concur with parental preference for that gender at birth. In other words, if sons receive less of home items in Balkan countries, this is consistent with the technology hypothesis. If sons receive more of household items, this is consistent with differential returns mechanism. For Scandinavian countries, where we confirm the daughter preference, the respective relationships in the data should have the opposite direction.

## 2.6 Results

### 2.6.1 Estimates of parental gender preferences for children

Table 2.3 presents coefficients on the gender of the firstborn for different specifications of the dependent variable in Equation 2.1 estimated on data from Balkan countries. These results resemble those obtained by Dahl and 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 at any conventional level. In line with the expectations discussed above, the impact of the gender of the firstborn on progression to parity two in column (2) has lower statistical significance and magnitude than the impact on progression to 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 and 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 at 1% level 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, and 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 2.4 presents estimates analogous to those in Table 2.3, 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 at 1% level. 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 at any conventional significance level. The parity three progression results in column (3) are in line with those obtained by Andersson et al. (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 2.A.6 and Table 2.A.3, we present gender preferences across EU countries. Our results are broadly consistent with those obtained in previous literature (Hank and Kohler 2000). We also attempt to evaluate how our results would differ if there were no family disruptions caused by child gender, which

**Table 2.3:** The firstborn-child gender and family size in the Balkans.

| | 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 | 0.030 | -0.001 | 0.013 | 0.011 | .003 |
| | (0.010)*** | (0.008) | (0.005)*** | (0.002)*** | (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 childrens ages are in the range 012. 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.

is frequently reported in the literature (see, e.g., Lundberg (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.

**Table 2.4:** 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 2.3

### 2.6.2 Testing the gender bias explanation

We first test predictions implied by the gender bias mechanism for household-level data. The results are presented in Table 2.5.

Column (1) of Table 5 contains estimates of the impact of the firstborn's gender on replacement of worn-out furniture in the household. The negative and statistically significant estimate for Balkan countries confirms the prediction from 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 lower 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 be able to deal with unexpected expenditures: 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 (with deferred consumption and present-discounted utility flow) and respond positively to the arrival of a child of the preferred gender, countering the negative effect of a reduction in the expected number of children. The same conjecture could apply to the ability required to make ends meet and to the minimum amount of money to make ends meet. Therefore, our results for the household-level data corroborate the presence of son bias in the Balkan countries, but do not corroborate the presence of daughter bias in the Scandinavian countries.

Next, we test predictions implied by the gender bias mechanism for individual-level data. The results are presented in Tables 2.6 and 2.7.

The negative impact of the mother's ability to spend on herself in Balkan countries in column (3) of Table 2.6 is in line with the gender bias explanation.[29] In addition, two more facts hold for intrahousehold allocations in Balkan countries which, at first sight, are out of line with expectations. 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). Meanwhile, the time mothers spend on unpaid household work tends to increase (not reported here), which compensates for a withdrawal from market work. 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

[29] This could be one reason there is no effect for the ability to make ends meet: increases in savings cannot be achieved because a substantial part of personal consumption is conceded in intrahousehold bargaining.

**Table 2.5:** 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 childrens ages are in the range 012. 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.

of housework done by daughters.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. Thus, the results obtained for the individual-level data are consistent with the gender bias explanation for Balkan countries.

As for Scandinavian countries, the estimates of the firstborn's gender impact on parental consumption in Table 2.7 do not differ between fathers and mothers, similarly to the unconditional means in Table 3. Estimates of the impacts on the

**Table 2.6:** 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)** |

*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 2.5.
* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

**Table 2.7:** 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) |

*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 2.5.
* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

ability of mothers to meet with friends and family and to have regular leisure activities do not contradict the gender bias explanation *per se*. However, the estimated impacts on the consumption of fathers should be positive according to the gender bias explanation, but 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 (Choi, Joesch, and Lundberg 2008a; Lundberg and Rose 2002), which exhibit son preference (Dahl and Moretti 2008; Hank and Kohler 2000). All in all, the data does not support the gender bias explanation for Scandinavian countries.

### 2.6.3 Testing the fixed costs explanation

We use the estimates in Table 2.5 to test predictions flowing from the differential fixed costs mechanism. Lower likelihood of being able to replace worn-out furniture by parents of first-born daughters in Balkan countries is out of line with higher fixed costs of sons. Conversely, more frequently available home items in households with first-born daughters in Scandinavia is in line with higher fixed costs of daughters. The previous fact, however, is also compatible with lower prices of acquiring human capital for daughters. Next, we examine individual-level data to see which of the two mechanisms (if any) it supports: differential fixed costs or differential prices of human capital.

### 2.6.4 Testing the difference in prices of human capital between sons and daughters

As we explain in Subsection 2.5.3, we cannot distinguish differential fixed costs from differential prices of human capital using the household-level data. However, two mechanisms lead to different predictions for the individual-level data. We use estimates in Tables 2.6 and 2.7 to check expectations about the sizes of first-born daughter effects on parental consumption.

Mothers of daughters in Scandinavia 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 and Rose (2003a) reported a similar effect for fathers from the US, they offered an explanation based on the son bias idea, but they did not formally test it. Our testing, however, does not support the son bias explanation for Scandinavia. Furthermore, Norwegian data indicates that paternal leave has more pronounced positive effects for daughters than sons (Cools, Fiva, and Kirkeboen 2015), which could explain why fathers in Scandinavian countries substitute time spent on work for time spent on children (rather than leisure). All in all, the differential costs explanation of daughter preference in Scandinavian countries is supported by the data.

For the sake of convenience, we highlight in grey all effects listed in Tables 2.A.7 and 2.A.8 that we find to be statistically significant.

Figure 2.A2 and Figure 2.A3 show cross-country relationships between gender preferences, the gender gap in parental investment, and conventional measures of gender equality. These relationships are in line with our previous points.

## 2.7 Conclusion

We find evidence that parental gender preferences in different countries have different explanations. In Balkan countries, the observed son preference is likely driven by gender bias towards sons, which plays the major role. In Scandinavian countries, the observed daughter-preference is likely driven by a lower price of human capital of daughters (i.e., lower price of child "quality" as we mean it when children are daughters). 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 (Roberts and 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.A Appendix

## 2.A.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. For example, Arnold, Choe, and Roy (1998) assert that some Indian parents prefer sons for reasons connected with religious beliefs and kinship descent, whereas Jacobsen, Møller, and 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 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, and 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.A.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, and 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 and Willard 1973). Furthermore, the wealthiest individuals in societies tend to have sons born more frequently (Cameron and 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.[30]

At the same time, the gender of the firstborn might impact marital stability (Lundberg and Rose 2003a; Mammen 2008; Lundberg, McLanahan, and Rose 2007), family size (Hank and Kohler 2000; Angrist and Evans 1998), and parental time allocation (Lundberg and Rose 2002; Lindstrom 2013; Choi, Joesch, and Lundberg 2008b). This makes "exclusion restrictions a priori unpersuasive" (Lundberg 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 (Lundberg and Rose 2002). However, Pabilonia and Ward-Batts (2007) find $\frac{1}{3}$ of the same effect and at no conventional statistical significance level). An even larger effect, almost 5% of mean annual male work hours, was found in West Germany (Choi, Joesch, and Lundberg 2008b). Meanwhile, Ichino, Lindstr'om, and 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 (2013) 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

[30]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.

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 and 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.A.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, and Schennach 2010; Todd and Wolpin 2007; Juhn, Rubinstein, and Zuppann 2015) and PISA-2000 Xu (2016). Instead, the ten questions we consider largely overlap with the resources-spent and time-with-child subcomponents defined by Juhn, Rubinstein, and 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, and 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 and Caldwell 1980; Bradley and Caldwell 1981; Bradley and Caldwell 1984) and strongly influencing it (Cunha, Heckman, and 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 (2016) 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 2.1 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 and 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 and 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. Then, 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 at the 5% level.

### 2.A.4 Cross-country comparison of gender preference and parental investment

Table 2.A.3 displays the results of estimating gender preference by country. The geographical pattern of the gender preference at birth is depicted in Figure 2.A.6. Our results are broadly consistent with those previously obtained in the literature. As did Hank and Kohler (2000), we find son preference in Italy and France and daughter preference in Portugal and Lithuania. Similar to Andersson et al. (2006), we also find daughter preference in Norway, though not in Sweden.[31] 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 2.A.3. Son preference becomes statistically significant in Slovenia and stops being statistically significant in Croatia (in both cases at the 10% level). 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

[31]Still, our estimates are correlated with ($\rho$=0.6) and statistically significantly at 5% level predict comparable estimates to Hank and Kohler (2000)

of the firstborn's gender on the selected household fertility outcomes are presented in Table 2.4. 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 2.A.1). Still, the last two sets of coefficients are strongly correlated with coefficients for the total number of children. To rationalize the estimates obtained, we plot the coefficients against several existing measures of gender inequality. As Figure 2.A.2 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 is why we use third parity progression results in Figure 2.A.6 and beyond.

In addition, the fact that parents tend to invest more in daughters as measured by the presence of the home items[32] hold for the pooled EU-SILC sample. To test for the gender gap in parental investment we estimate Equation 2.1 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 2.A.10 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 2.A.10. 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

[32]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

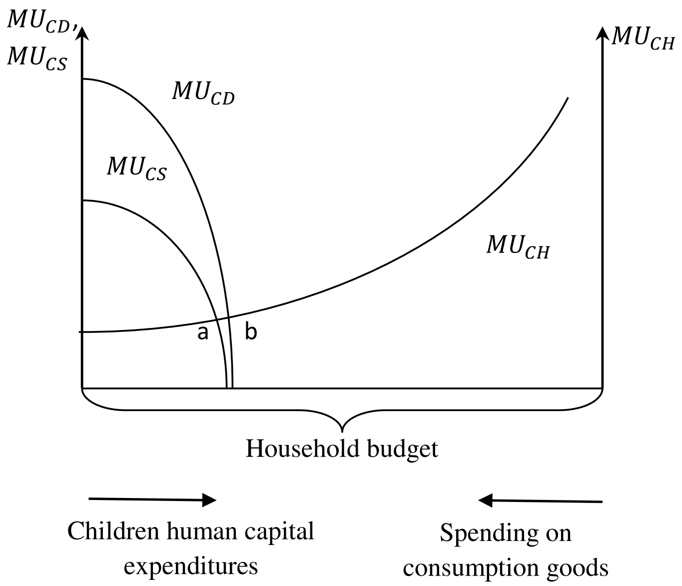
are typical in the literature on gender effects. Meanwhile, the gender preference pattern established before holds for the sub-sample of households from the highest income decile. These results might suggest that society as a whole is attaching increasingly positive significance to female children, an idea that has appeared in previous studies, such as Brockmann (2001) and Andersson et al. (2006). A daughter may assume both the role of a breadwinner and that of a caregiver.[33] As Brockmann (2001) puts it, "in the future, the average girl may well wish to become the mother of a one-daughter family."

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 2.A.3 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 et al. 2008).

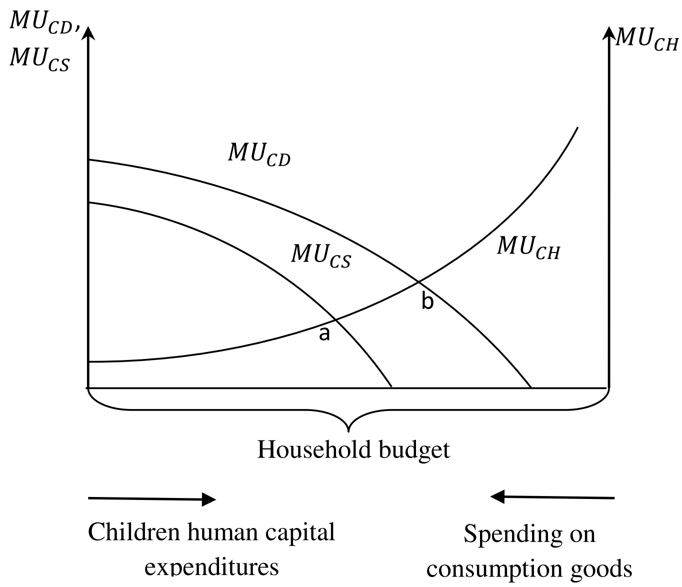
However, Xu (2016) did not find any strong relationship between the gender gap in the home environment measure (similar to ours) and the GGG, though he measured the gender gap by difference in the unconditional mean between genders. Moreover, as explained earlier, our measure is preferable to the one 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.[34] Nevertheless, the latter is commonly used as a measure of parental gender preference.

[33] In this regard some authors speak about the "boy crisis" (Husain and Millimet 2009; Sadowski 2010).

[34] A similar and statistically significant at the 10% level 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).




**(a)** Gender bias




**(b)** Differential cost

Figure 2.A.1: 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 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.

### 2.A.5 Tables and Figures

**Table 2.A.1:** 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 2.A.3 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.

**Table 2.A.2:** 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 childrens ages are in the range 012. 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 2.A.3:** 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 2.3 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 2.A.4 contains names of countries corresponding to the abbreviations.

**Table 2.A.4:** 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 2.A.5:** 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 2.A.6:** 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.

**Table 2.A.7:** Expected directions of the impact of a first-born daughter on household-level consumption in the Balkans and Scandinavia under the gender bias and the differential costs mechanisms

| Dependent variable | Gender bias | Differential cost |
|---|---|---|
| **The Balkans** | | |
| Ability to replace worn-out furniture | $\beta < 0$ | $\beta > 0$ |
| Capacity to deal with unexpected expenditures | $\beta > 0$ | $\beta < 0$ |
| Ability to make ends meet | $\beta < 0$ | $\beta > 0$ |
| Lowest monthly income to make ends meet | $\beta > 0$ | $\beta < 0$ |
| Availability of home items | — | $\beta < 0$ |
| **Scandinavia** | | |
| Ability to replace worn-out furniture | $\beta > 0$ | $\beta < 0$ |
| Capacity to deal with unexpected expenditures | $\beta < 0$ | $\beta > 0$ |
| Ability to make ends meet | $\beta > 0$ | $\beta < 0$ |
| Lowest monthly income to make ends meet | $\beta < 0$ | $\beta > 0$ |
| Availability of home items | — | $\beta > 0$ |

**Table 2.A.8:** Expected directions of the impact of a first-born daughter on individual-level consumption and time-use in the Balkans and Scandinavia under the gender bias and the differential price of human capital mechanisms (assuming that son bias for either gender is exhibited by fathers)

| Dependent variable | Gender bias | Differential cost |
|---|---|---|
| **The Balkans** | | |
| Mother being employed | $\beta > 0$ | — |
| Father being employed | $\beta > 0$ | — |
| Mother's weekly hours of work | $\beta > 0$ | — |
| Father's weekly hours of work | $\beta > 0$ | — |

| | | |
|---|---|---|
| Mother's ability to spend on oneself | $\beta < 0$ | $\beta > 0$ |
| Father's ability to spend on oneself | $\beta > 0$ | $\beta < 0$ |
| Mother having two or more pairs of shoes | $\beta < 0$ | $\beta > 0$ |
| Father having two or more pairs of shoes | $\beta > 0$ | $\beta < 0$ |
| Mother being able to replace worn-out clothes | $\beta < 0$ | $\beta > 0$ |
| Father being able to replace worn-out clothes | $\beta > 0$ | $\beta < 0$ |
| Mother being able to get together with friends | $\beta < 0$ | $\beta > 0$ |
| Father being able to get together with friends | $\beta > 0$ | $\beta < 0$ |
| Mother having regular leisure activity | $\beta < 0$ | $\beta > 0$ |
| Father having regular leisure activity | $\beta > 0$ | $\beta < 0$ |
| **Scandinavia** | | |
| Mother being employed | $\beta < 0$ | — |
| Father being employed | $\beta < 0$ | — |
| Mother's weekly hours of work | $\beta < 0$ | — |
| Father's weekly hours of work | $\beta < 0$ | — |
| Mother's ability to spend on oneself | $\beta > 0$ | $\beta < 0$ |
| Father's ability to spend on oneself | $\beta < 0$ | $\beta > 0$ |
| Mother having two or more pairs of shoes | $\beta > 0$ | $\beta < 0$ |
| Father having two or more pairs of shoes | $\beta < 0$ | $\beta > 0$ |
| Mother being able to replace worn-out clothes | $\beta > 0$ | $\beta < 0$ |
| Father being able to replace worn-out clothes | $\beta < 0$ | $\beta > 0$ |
| Mother being able to get together with friends | $\beta > 0$ | $\beta < 0$ |
| Father being able to get together with friends | $\beta < 0$ | $\beta > 0$ |
| Mother having regular leisure activity | $\beta > 0$ | $\beta < 0$ |
| Father having regular leisure activity | $\beta < 0$ | $\beta > 0$ |

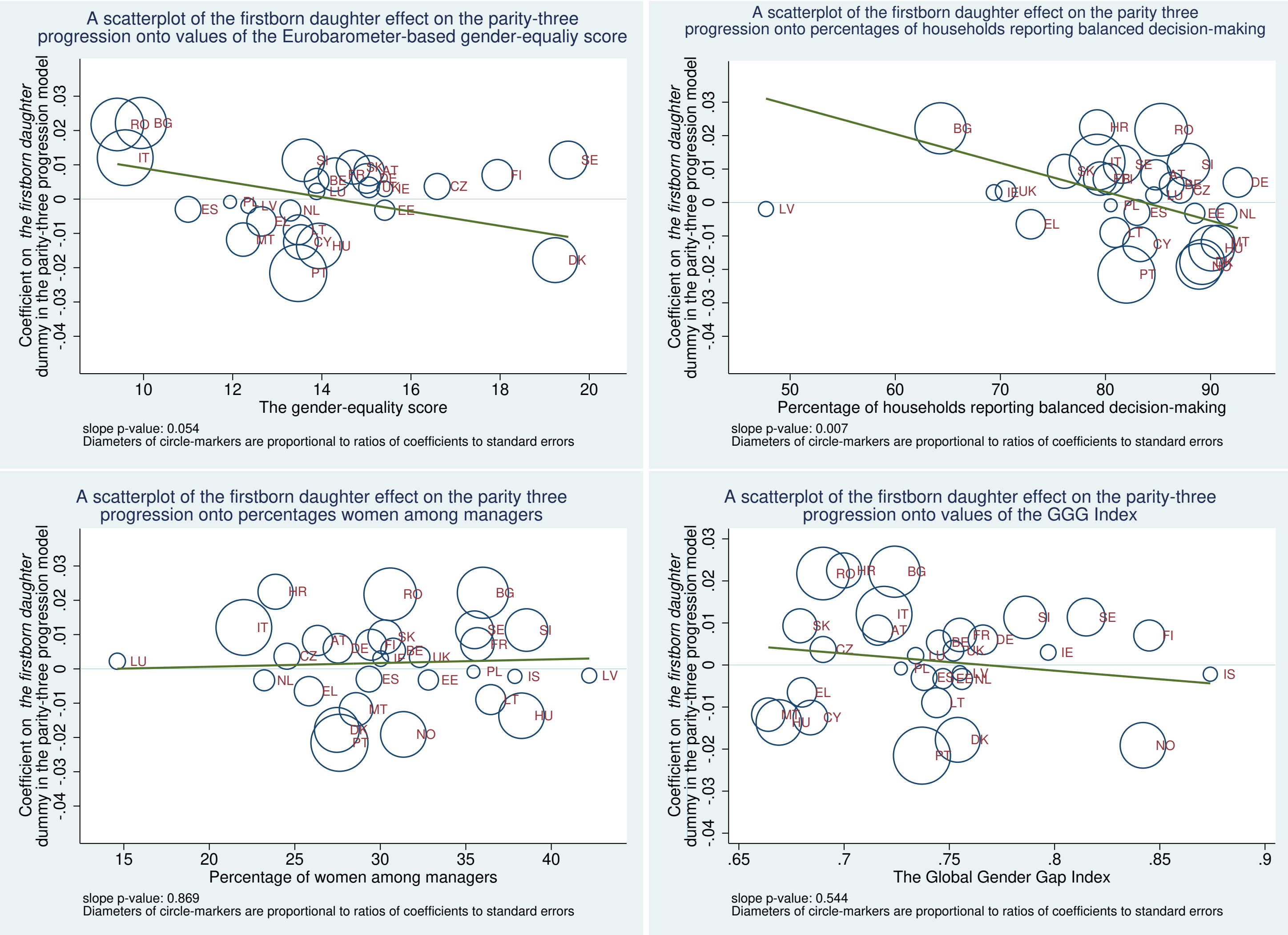



Figure 2.A.2: The relationship between the effect of first-born daughters on third parity progression and specific gender-equality measures across countries

*Notes:* 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 (Eurocommission 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.

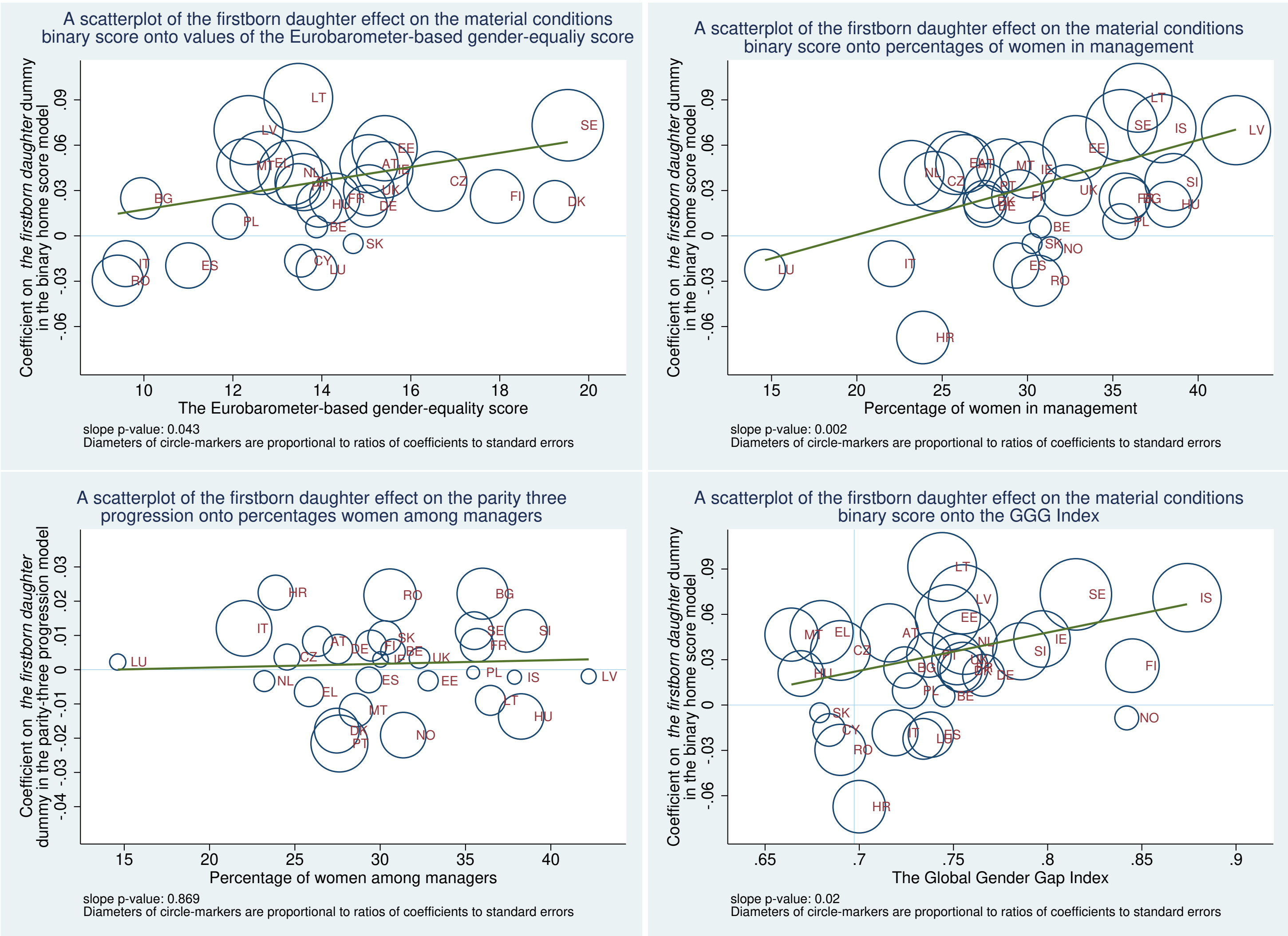
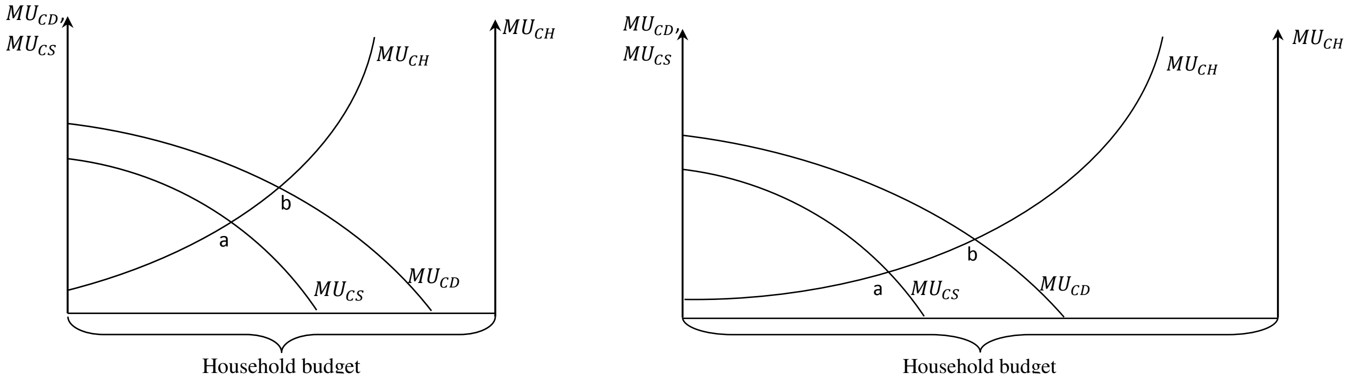



Figure 2.A.3: The relationship between the effect of first-born daughters on child material conditions and specific gender-equality measures across countries.




**(a)** Low-income families

**(b)** High-income families

Figure 2.A.4: Differences in expenditures on children between low-income and high-income households

*Notes:* See the note to Figure 2.A.1 for explantion.

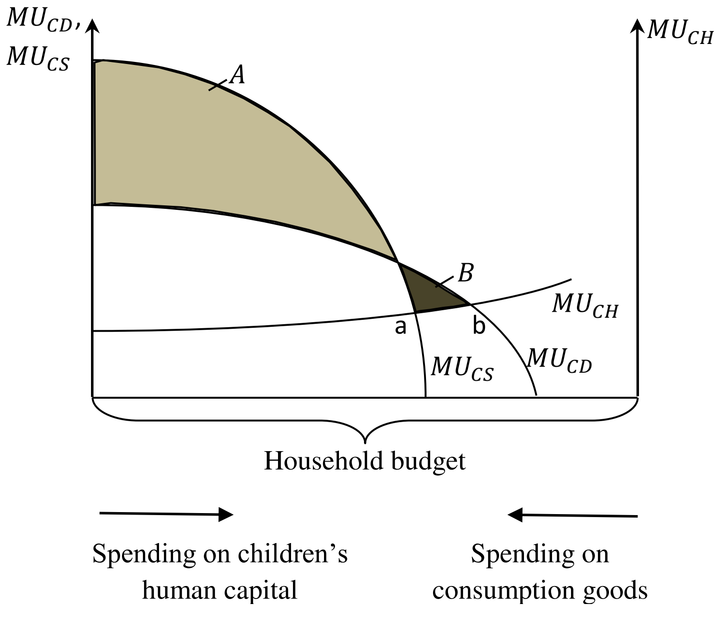



Figure 2.A.5: Coexistence of gender (son) bias and differential cost with the gender bias effect on fertility prevailing.

*Notes:* See the note to Figure 2.A.1 for explanation.

Table 2.A.9: 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 2.A.3.

Table 2.A.10: 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 childrens ages are in the range 012. 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 2.2 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.

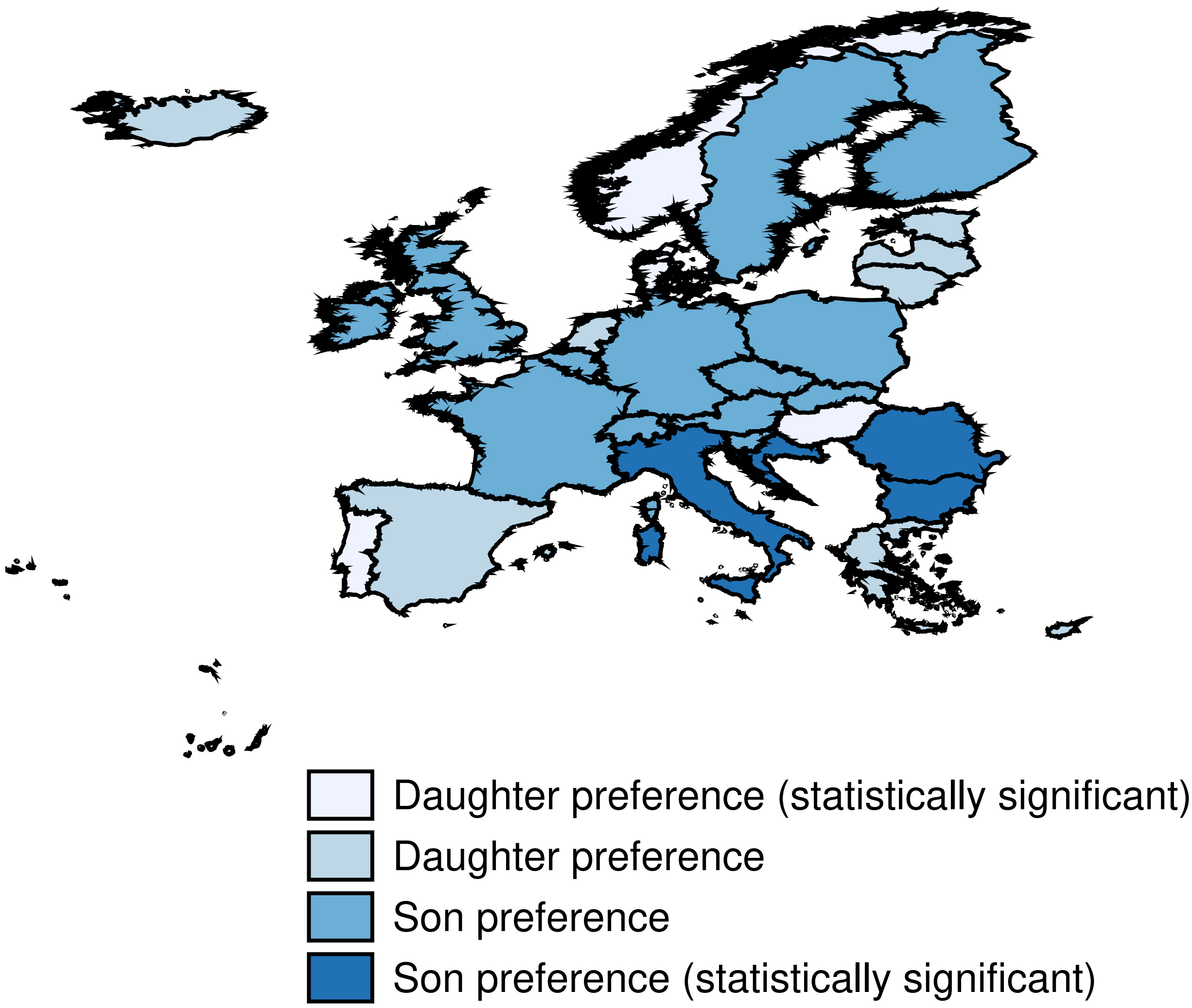



Figure 2.A.6: Gender Preferences of Children in 31 EU-SILC countries

# Chapter 3

# The impact of the first-born child's 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)[1]). A growing body of research examines how a child's gender may be associated with differential treatment by parents from birth, which could contribute to gendered differences in adult market outcomes (e.g., Lundberg (2005) among others). One strand of this literature has found associations between child gender and parental marriage and separation, with implications for the living arrangements of children. Fathers are more likely to be present in the

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

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 2020). Many authors have proposed that parental, and especially fathers' preference for sons may offer an explanation. This explanation has been supported by systematic evidence from several US survey data sets by Dahl and Moretti (2008). However, a more recent study by Kabátek and Ribar (2020) concluded that strained relationships in families with teenage daughters likely lie behind the higher divorce rate of couples who have daughters. This explanation is compatible with the finding of Kabátek and Ribar (2020) that single Dutch mothers of firstborn girls are just as likely to marry as single Dutch mothers of firstborn boys and they take about the same amount of time to do so. However, it would not have accounted for the finding of Dahl and Moretti (2008) that single mothers of first-born daughters in US were less likely to marry.

My paper, like the two aforementioned studies, aims to simultaneously estimate the impact of the firstborn gender on family formation and dissolution, conditional on firstborn age.[2] In contrast with Dahl and Moretti (2008), who also look into family formation and dissolution in their paper, I consider an event-history model that not only allows the risks of separation to change with childrens ages, but

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

also incorporates the length of cohabitation of spouses and the right-censoring in marriage spells. [3] Compared to Kabátek and Ribar (2020), my work employs data on multiple socioeconomic characteristics of households, which allows me to check whether the effect in question is confined to a particular group of the population. Moreover, I look at actual cohabitation and not only at the officially registered relationships, as Kabátek and Ribar (2020) do.

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

I focus on children of different ages, more specifically on pre-school children (0-5 years) and school-age children (6-18). I group children's ages this way because I expect differential effects of preschool children. First, events before five years old can have large long term impacts on adult outcomes (see e.g., Currie and Almond (2011)). Second, having daughters of this age might have an impact on such personal traits of fathers as neuroticism and extraversion (van Lent 2020), which in turn are related to a higher chance of divorce (Diederik and Mortelmans 2018; Fani and Kheirabadi 2011; Zare et al. 2013). Third, single mothers in my data set most often have young children.

Each of the two proposed mechanisms - the son bias and tensions between teen-age daughters and their fathers - implies different predictions for my results. If the overarching son preference holds, lower marriage rates of single mothers of daughters should hold simultaneously with higher separation rates of married mothers of daughters. If the explanation through strained relations with teenage children holds, there should be no effect for either marriage or separation. [4] My results show that the effect on probability of separation [5] is negative, and the effect

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

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

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

on probability of partnership formation is also negative. Additionally, I find an indication that having teenage daughters increases the probability of separation in line with Kabátek and Ribar (2020).[6] My results, based on the Russia Longitudinal Monitoring Survey (RLMS-HSE) data, suggest that while son preference might play a notable role in marriage decisions, it is outweighed by other determinants of the family process, leading to a negative observed effect of preschool firstborn daughters on separation. Investigation of the precise nature of these determinants is currently beyond the scope of this paper. [7]

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

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

[7] I suggest six possible causes for the negative effect of first-born daughters on separation in Russia. First, parents of spouses might be more supportive of a marriage in which daughters are born. Second, losses of marriage surplus due to separation are higher when children are female. Third, higher marriage rates among single mothers of first-born sons reduces the estimated cost of separation for them as they perceive their remarriage prospects to be more favorable. Fourth, mothers of younger sons should be able to remarry faster than mothers of older sons because sons, compared to daughters, are more likely to come to terms with stepfathers when they are younger. Fifth, public policy is more oriented towards women, because women constitute the majority of voters demographically and vote more actively. Sixth, gender of children induces changes in parents' (especially fathers') personality, which in turn influence marriage stability.

## 3.2 Data

### RLMS-HSE data

The source of data for my analysis is the Russia Longitudinal Monitoring Survey, RLMS-HSE, for the 1994-2018 period. The data set covers 23 yearly rounds of a national representative survey on the social, health and economic situation in Russia. Two years are missing, 1997 and 1999. The survey is scheduled annually during fall and winter months exact dates varying from year to year (i.e., one household could be surveyed twice during the same calendar year). The RLMS-HSE is conducted by several organizations including the National Research University Higher School of Economics and the Carolina Population Center, University of North Carolina at Chapel Hill. [8]

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

The RLMS-HSE was established to create a nationally representative survey to monitor the economic and health impact of the reforms in the Russian Federation. Throughout the entire set of surveys, detailed basic household and in-

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

[9]The target sample size was set at 4 000 households (Kozyreva, 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).

dividual data have been collected. The major data components are: economic (detailed income, assets, expenditures and labor force behavior data, including type of employment, earnings, hours and ownership form, i.e., public, private or joint), demographic/sociological (household structure and age-gender composition, background, education and school behavior); and health (24-hour dietary recall, nutrient intake levels, smoking, drinking activity, BMI direct measurement). All in all, there are more than 3,000 variables. [10]

The RLMS-HSE is a panel survey with a longitudinal component. A point of concern is that of attrition in the panel. [11] Some households are inevitably lost from the panel as a result of moving house, splitting up, or other common causes of attrition. The size of attrition across years is reported in, e.g., Kozyreva, 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

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

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

to have higher median incomes and expenditures. Gerry and Papadopoulos (2015) investigate patterns of the RLMS attrition and how it is related to demographic, health and other socio-economic characteristics of participants. The authors confirm the presence of non-random attrition for the RLMS. At the same time, they also conclude that the non-random attrition does not significantly distort estimates of statistical models. [12]

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

## Selected variables and descriptive statistics

I identify 1,788 firstborn children whose mothers participated in the RLMS-HSE survey in the year of their birth, i.e. before they turned 1. Of these 1,788 firstborn children, 1,431 were born to partnered women (either in a married couple or in a consensual union) [13] and 1,367 were observed in more than one survey wave.[14]

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

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

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

Therefore, the firstborns from the analyzed sample are born to already partnered couples and their gender cannot influence the matching or selection of their parents into partnership. Correspondingly, 357 firstborns were born to single women, [15] of whom 255 were observed in more than one survey wave (of them 73 report being married in the individual file, 102 are observed starting partnership of which 55 start marriages and 47 start cohabiting). Therefore, I have two main samples for analysis: a sample of 1,367 firstborns with two parents present and a sample of 255 firstborns born to single mothers.

The variables used for analysis are described in Table 3.B.1 in Appendix 3.B. Descriptive statistics are presented in Table 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 | 1.4% | 1% | 0.4% | 0.36 |

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

| | | | | |
|---|---|---|---|---|
| religious affiliation | | | | |
| Mother reports no religious affiliation | 4% | 3% | 1% | 0.22 |
| Urban area | 74.32% | 76.73% | -2.40% | 0.30 |
| Father is Russian | 89.24% | 87.98% | 1.27% | 0.46 |
| Mother is Russian | 91.61% | 89.26% | 2.35% | 0.14 |
| Father has vocational or tertiary education | 52.00% | 48.05% | 3.95% | 0.14 |
| Mother has vocational or tertiary education | 62.34% | 61.26% | 1.08% | 0.68 |
| Number of family members | 3.96 | 4.01 | -0.05 | 0.54 |
| Mother reporting satisfactory life | 64.25% | 62.44% | 1.82% | 0.49 |
| Father reporting satisfactory life | 60.54% | 58.13% | 2.41% | 0.37 |
| Reporting satisfactory job | 33.5% | 33.5% | 0.04% | 0.99 |
| Having subordinates at work | 9.1% | 10.5% | -1.4% | 0.39 |
| Concerned about losing job | 26.7% | 28.4% | -1.7% | 0.48 |
| Well-being improved in the last year | 22.7% | 22.4% | 0.3% | 0.89 |
| Expecting economic improvement in 12 months | 40.4% | 41.9% | -1.5% | 0.57 |
| Relative economic standing | 15.3% | 14% | 1.3% | 0.50 |
| Feeling empowered | 13.7% | 12.3% | 1.4% | 0.45 |
| Feeling respected | 61.9% | 62.8% | -0.9% | 0.75 |
| Satisfactory material condition | 23.3% | 20.1% | 3.1% | 0.16 |
| Understanding between generations is possible | 28.8% | 28.1% | 0.7% | 0.76 |

| | | | | |
|---|---|---|---|---|
| Knowing foreign language | 22.3% | 23.3% | -1.0% | 0.65 |
| Having disability | 0.6% | 0.5% | 0.1% | 0.76 |
| Smoking now | 12.3% | 11.0% | 1.7% | 0.32 |
| Ever having smoked | 18.8% | 20.7% | -1.9% | 0.38 |
| Drinking alcohol recently | 25.4% | 27.2% | -1.8% | 0.45 |
| Drinking alcohol sometimes | 35.7% | 37.5% | -1.9% | 0.47 |
| Ability to have three meals daily | 33.0% | 34.5% | -1.6% | 0.54 |
| [a] Living in regional center | 47.6% | 44.4% | 3.2% | 0.24 |
| Family size | 3.6 | 3.7 | -0.1 | 0.32 |
| Owning accommodation | 74.8% | 77.3% | % | |
| Number of rooms | 2.13 | 2.10 | 0.03 | 0.48 |
| Area of accommodation ($m^2$) | 33.31 | 33.71 | -0.40 | 0.67 |
| Running water | 89.4% | 88.4% | 1% | 0.56 |
| Owning country house | 14.6% | 17.4% | -2.8% | 0.17 |
| Saving money recently | 17.6% | 13.1% | 4.5% | 0.02 |
| Receiving economic support | 49.7% | 47.1% | 2.6% | 0.33 |
| Having credit debt | 29.1% | 28.0% | 1% | 0.67 |
| Owing money to individuals | 4.3% | 6.1% | -1.8% | 0.13 |

*Notes.* The results are based on 1,367 observations, 701 boys and 666 girls.
[a] The following variables are taken from the household file and descriptive statistics based on 1,354 observations, 698 boys and 656 girls.

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

Descriptive statistics for single mothers are presented in Table 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 form partnerships faster.[16] Second, mothers of sons appear to have lower educational attainment that mothers of daughters. This could be partially explained by their younger average age. Another possible reason is a lower response rate from mothers of first-born daughters, which could be even lower for those with lower educational attainment.

**Table 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 |

[16]More precisely, this would be the case for children born out of wedlock on January 1, the boy's mother is already partnered by the time of the survey (October-December), while the girl's mother is still single. This statement could be supported by the results of a back-of-an-envelope numerical simulation with the probability of partnership formation following a Bernoulli distribution. The difference in the age of mothers appear to be noticeable even during the first year after a firstborn birth (the period for which the numbers in Table 3.2 are calculated). This result is even stronger when younger mothers are more likely to marry (which again makes sense because the observed mean age difference accumulates only during the first year after a birth).

| | | | | |
|---|---|---|---|---|
| Satisfactory life | 40% | 49% | -9% | 0.08 |

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

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

The numbers of first-born boys and girls of different ages living with partnered and single mothers can be seen in Table 3.3. The share of boys among children living with single mothers is higher than for children living with partnered mothers when the children are younger. This could be explained by two simultaneous tendencies. First, married mothers of first-born sons are more likely to separate. Second, mothers of daughters more frequently drop out of the survey after separation, [17] while single mothers of sons more frequently drop out of the survey after partnership formation.[18] The latter fact could explain why the share of first-born boys remaining in the survey increases over time.

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

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

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

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

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

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

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

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

* The difference is statistically significant at a 0.1 level.

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

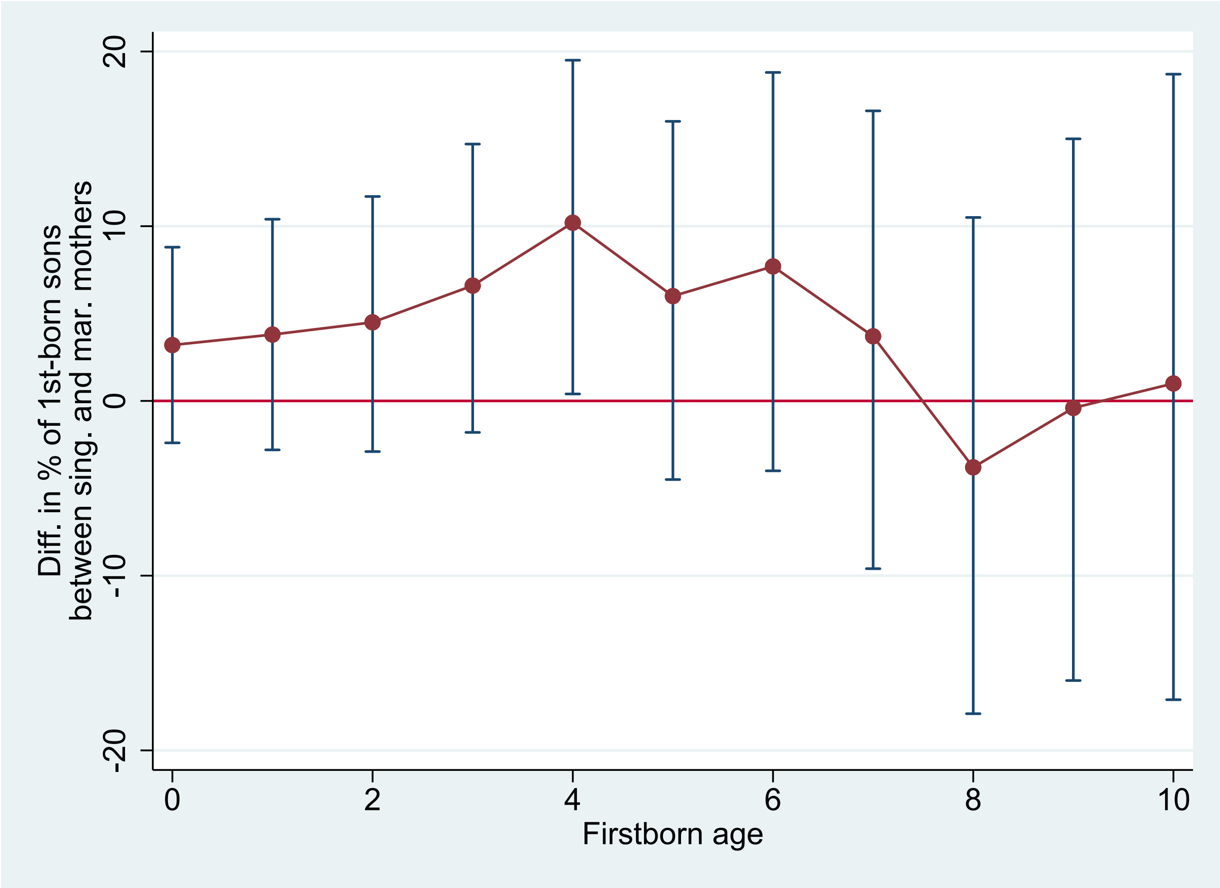



**Figure 3.1:** Difference between percentages of firstborn sons among single and partnered mothers conditional on firstborn age
*Note:* Capped spikes show 90% confidence intervals.

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

## 3.3 Method

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

The functional form of the cloglog model is:

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

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

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

$$\begin{aligned}\mathbf{x}_i'\boldsymbol{\beta} = \beta_0 &+ \mathbf{1}(FB\ age\_range_{it} = 0-5)(\beta_{0_{0-5}} + \beta_{1_{0-5}} \cdot FB\ daughter_i) \\ &+\mathbf{1}(FB\ age\_range_{it} = 6-18)(\beta_{0_{6-18}} + \beta_{1_{6-18}} \cdot FB\ daughter_i)\end{aligned}$$

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

$$+\sum_{k=1}^{18}\beta_{3j}\cdot\mathbf{1}(FB\ years\_obs_{it}=j)+\mathbf{z}_i^{'}\boldsymbol{\beta}_4 \qquad \text{(1b)}$$

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

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

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

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

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

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

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

For identification, I rely upon the assumption that the firstborn's gender is random. This assumption would be violated if there were sex-selective abortions. At the level of the entire sample this assumption appears to be warranted because the sex ratio does not notably differ from that in the population, and the average age of women who give birth to their first child does not notably differ by the gender of the first child. This fact, however, does not rule out a possibility that sex-selective abortions could be biased either towards sons or towards daughters in different subgroups of the population. In this case, the effects of sex-selective abortions across different subgroups could cancel out, so the sex-ratio at birth at the level of the entire population would be close to the natural one. To the

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

best of my knowledge, no evidence, however, has been reported on sex selective abortions biased towards different sexes and confined to particular subgroups of the population in Russia. [21]

While there are few reasons for concern about reverse causality and unobserved heterogeneity, there are several issues regarding the empirical framework that should be pointed out. First is the measurement error in age ranges of firstborns. Specifically, I measure firstborn ages as differences between the year of observation and the year of birth. Therefore, when two consecutive survey waves are less then one year apart, some children have the same age at these two consecutive waves. In such cases, I add one year to their age in the second wave. [22] Second is construction of the dependent variable, the indicator of family dissolution. [23] While Kabátek and Ribar (2020) focus on formal marriage status,[24] I also take into account information about the actual cohabitation of spouses. I do this for two reasons. First, the related literature tends to focus on the actual absence of fathers rather than on official marriage status, as in the paper by Dahl and Moretti (2008). Second, women who appear to single according to the individual-level data quite often have a husband according to the household-level data. That is why my dependent variable takes a value of 1 when a couple stops cohabiting according to the household-level data file and if this couple is not officially married in the individual-level data (as reported by a wife).[25] In other words, in the basic family

[21] As for specific groups of population in other countries, some studies indicate this assumption might not hold. These studies, however, have not been frequently replicated so far, the effect they found is small, and it is not clear how characteristics causing a shift in sex ratio of children are related to marriage stability. More detailed discussion of the firstborn gender randomness assumption plausibility is presented in Abramishvili, 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`).

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

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

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

[25] Other living arrangements recorded in the individual-level data file, besides being officially

dissolution model I consider those women who appear to be not married based on individual data and separated based on household-level data as actually separated. Women who fulfill only one of these conditions or none, are regarded as being in partnership. At the same time, my marriage formation indicator takes a value of 1 when a couple starts cohabiting according to the household-level data file. Finally, the character of association between covariates and the dependant variable, along with estimation results with alternative errors specifications suggest that concerns about non-monotonicity and heteroskedasticity are not justified.[26]

## 3.4 Results and discussion

### 3.4.1 Results for partnership dissolution

Estimates of the model (1.a) with eight specifications of predictor (1.b) are presented in Table 3.4. [27] The presence of older first-born daughters predicts a substantially higher likelihood of divorce than the presence of sons in the same age range. This finding is in line with the result in Kabátek and Ribar (2020), but the effect size is much higher than they find. Nevertheless, their estimated values lie within the 95% confidence interval of my estimated effect. [28] Moreover, it is

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

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

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

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

necessary to take into account the fact that the divorce rate in Russia has been high on the global scale. Specifically, if similar mechanisms to those underlying the results of earlier research are at work in the Russian setting, my estimate of the impact of firstborn's gender on marriage dissolution will be higher than in previous studies.

**Table 3.4:** The impact of the firstborn gender on family dissolution from complementary log-log model estimation.

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

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

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

numbers in Table 3, which include all firstborns that remain in the survey.

**Table 3.5:** The impact of preschool first-born daughters on separation after conditioning on interaction with socio-demographic characteristics.

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

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

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

The estimated effect of young (0-5 years old) first-born daughters on the probability of marriage dissolution is negative. It, however, is not statistically significant. Still, when sons aged 6-18 years become a reference category, the impact of first-born daughters aged 0-5 becomes statistically significant as can bee seen from columns (1) and (2) of Table 3.6.

Table 3.6 also contains results for separation regressions with interaction terms between the dummy for a first-born daughter aged 0-5 and moderator variables, observed socio-economic characteristics of mothers (being Orthodox, Russian, employed; having secondary education; living in an urban area; having above average well-being). Adding these interaction terms makes the coefficient on the dummy for a first-born daughter aged 0-5 comparable in magnitude and statistical significance to that on the dummy for a first-born daughter aged 6-18 in Table 3.4. Therefore, the lower separation among parents of first-born daughters aged 0-5 is driven by belonging to groups of the population with the foregoing characteristics (most notably, by being an Orthodox).

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

As for the effect direction for the preschool daughters, I can put forth six explanations for $\beta_{1_{0-5}}$ to be negative. The first is the relationship with parents of spouses (esp. mothers of husbands), who may be more supportive of the marriage when a child is a daughter (Duflo 2003; Aduškina 2015; Aivazova 2015; Mkchtrian 2017).[29] The second is the loss in marriage surplus due to loss in the human capital of daughters. In other words, if investment in the human capital of children

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

brings higher returns in terms of marriage surplus for daughters than for sons (Abramishvili, Appleman, and Maksymovych 2019), the divorce of spouses with a first-born daughter could cause an especially high loss in the marriage surplus (Currie and Almond 2011; Myck, Oczkowska, and Wowczko 2021; Kim and Shapiro 2021).

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

The fourth reason is that older children and boys fare worse in "stepfamilies" than younger children and girls (Brown 2004). Thus, mothers of younger boys should try not to delay separation, once they opt for it, because remarriage is more challenging when sons get older. This conjecture is also supported by the comparison of baseline results with results of the robustness check. Namely, the fact that the observed effect for younger daughters becomes more statistically significant when separations with the preservation of the official marriage are not counted means that mothers who separate while remaining in official marriage are more likely to be mothers of daughters. [30] This might suggest that mothers of young sons are less willing to preserve a formal marriage because, for instance, the remarriage becomes more problematic when sons get older.

The fifth reason is a cumulative effect of existing policies, especially social policies. The fact that policies can be not neutral with respect to gender is discussed in Washington (2008). For example, Cygan-Rehm, 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 consti-

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

tute a larger share of voters demographically and, at the same time, vote more actively (Goncharenko 2018). Examples of such policies might include mother's exclusive entitlement to maternity capital introduced in 2007, [31] generous public pensions (relative to average salary) that disproportionately benefit women due to a significant gender gap in life expectancy, [32] indexing public sector salaries, which is likely to be more beneficial to women, who constitute a majority of public sector employees (including those in education and health care), or "gender asymmetry" in family law (Klimashevskaya 2021). [33] These circumstances enhance the chances of daughters to be employed, financially secure, and, ultimately, capable of supporting their parents in their dotage. Thus, fathers are likely to have fewer reasons for leaving a family when their firstborn is a daughter.

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

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

As for the baseline hazard, the exponentiated coefficient on the log(time) is significantly less than 1 in all specifications. That is, the estimated baseline hazard decreases with elapsed survival time. This result seems plausible because the divorce hazard is falling during most of a typical marriage after a relatively short period of rising following the start of a marriage. When the quadratic polynomial of time is used in the regression instead of the log(time), the coefficient on the

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

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

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

squared year is negative and mostly statistically significant, corresponding to a bell-shaped form of the empirical divorce hazard.

### 3.4.2 Results for partnership formation

Estimates of model (2.a) with four specifications of the predictor (2.b) are presented in Table 3.5. The first-born daughter delays marriage, even without conditioning on age, as can be seen in column (2). Conditioning on age does not change the result substantially. This could also be partially explained by the fact

**Table 3.6:** The impact of the firstborn gender on family formation from a complementary log-log model estimation.

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

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

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

that, in most observations, children living with single mothers are younger. The impact of first-born daughters on marriage of single mothers is close in magnitude (but opposite in direction) to the impact of 0-5-year-old first-born daughters on separation. Thus, the higher marriage rate among single mothers of first-born sons is outweighed to some extent by a higher separation rate among married mothers of first-born sons.

The possibility that the estimates obtained are driven by family background of mothers is examined in Appnedix 3.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 separation. At the same time, I find that the effect of daughters on parental living arrangements depends on a daughter's age. In particular, having daughters aged 0-5 is related to a lower chance of parental separation, while having daughters aged 6-18 predicts a higher chance of parental separation. The two effects seem to cancel each other in the pooled sample of firstborns aged 0 to 18. The latter effect accords with Kabátek and Ribar (2020) who report the negative impact of teenage daughters (aged 13-18) on the duration of the parental marriage. The former fact is a novel contribution to the extant literature.

Therefore, my findings give a reason to believe that the impact of the children's gender on family living arrangements likely depends on family socioeconomic conditions and thus has a different character and magnitude in different contexts. In this regard, it is worth mentioning that most studies finding that daughters cause separation use data from countries for which a son preference has been reported.[34] Son preference has not been established for Russia so far, [35] and, hence, in the

[34] Despite it not being found for the Netherlands, the country studied by Kabátek and Ribar (2020), it was reported for US (Dahl and Moretti 2008), India (Barcellos, Carvalho, and Lleras-Muney 2014), Australia (Kippen, Evans, and Gray 2006)

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

Russian context it might have a lower impact than in other contexts. Moreover, some features peculiar to the Russian socioeconomic landscape are likely to mediate the relationship between the gender of children and parental marriage stability. At least six such features may be pointed out: relationship between spouses and their parents, higher returns to children human capital investment for daughters than for sons, lower subjective cost of separation among mothers of sons, women constituting the majority of voters and having relatively high electoral activity, older sons having harder time in "stepfamilies", and changes in parental personalities induced by having children of a particular gender. Examining the plausibility of these possible mechanisms will be a focus of my further research.

## 3.A Appendix A

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

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

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

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

As for extraversion, its possible impact is less clear-cut. On the one hand, extraversion is assumed to be related to positive emotions (Donellan, 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 Boertien and Mortelmans (2018), using German and UK data, report a positive association between extraversion and divorce risk; Solomon and Jackson (2014) do not find such an association in an Australian nationally representative sample. Moreover, Solomon and Jackson (2014) report a positive relationship between extraversion and marital satisfaction. The latter, according to Hirschberger, Srivastava, and Marsh (2009), negatively predicts prospective marriage dissolution for men.[37] Therefore, the direction and magnitude of relation between extraversion and divorce likely depends on the relative strengths of influential factors in a specific socio-economic context.

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

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

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

[39]This fact resonates with the conclusion by Somville (2019) that having a daughter reduces male violence against a partner.

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 imbalances caused by World War II had a lasting effect on family structure in Russia, including lower rates of marriage and fertility, higher non-marital births and reduced bargaining power within marriage for women most affected by war deaths. Moreover, the author argues that the mentioned effects were likely magnified by family policies in the former USSR. This is why the effect of increased extraversion of young husbands is likely to be more pronounced than in other socio-cultural contexts.

## 3.B Appendix B

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

| | |
|---|---|
| Separation | Dummy for whether a person who cohabited with a partner in a previous wave is not cohabiting with the same partner and reports being divorced or cohabits with a different partner in a current wave |
| Partnership formation | Dummy for whether a person who did not cohabit with a partner and reported being single in a previous wave is cohabiting with a partner in a current wave |
| 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 | Dummy for whether a mother reports being of Russian ethnicity |
| Father is Russian | Dummy for whether a father reports being of Russian ethnicity |
| Mother has vocational or tertiary education | Dummy for whether a mother reports attaining vocational or tertiary education |
| Father has vocational or tertiary education | Dummy for whether a father reports attaining vocational or tertiary education |
| Number of family members | How many people live in the household |
| Mother reporting satisfactory life | Dummy for whether a mother is fully satisfied or rather satisfied with life at the current moment |
| Father reporting satisfactory life | Dummy for whether a father is fully satisfied or rather satisfied with life at the current moment |
| Having three meals per day | Dummy for being able regularly to have three meals per day |
| Consuming alcohol | Dummy for consuming alcoholic drinks at least sometimes |
| Drinking alcohol last month | Dummy for consuming alcoholic drinks during previous 30 days |
| Ever smoked | Dummy for being a smoker at some time |
| Smoking now | Dummy for being a smoker at the present moment |
| Foreign language | Dummy for knowing other language than languages of former USSR republics |
| Understanding between generations | Dummy for agreeing that understanding between generations is possible |
| Satisfied with material condition | Dummy for being currently satisfied with own material condition |
| Respectability | Dummy for feeling oneself sufficiently respected |

| | |
|---|---|
| Empowerment | Dummy for feeling oneself sufficiently empowered |
| Relative economic standing | Dummy for feeling oneself better off than others |
| Expecting economic improvement | Dummy for expecting improvement in economic situation of a family in next 12 month |
| Well-being improved in the last year | Dummy for improvement of well-being during previous 12 months |
| Unemployment concern | Dummy for being concerned about loosing a job |
| Having subordinates | Dummy for having subordinates at work |
| Satisfied with job | Dummy for being satisfied with one's own job |
| [a] Living in a regional center | Dummy for living in a regional center |
| Dwelling area | Area of an occupied dwelling in square meters |
| Population | Population of a municipality of residence |
| Family size | Number of people residing in a household |
| Rooms | Number of rooms in an occupied dwelling |
| Running water | Dummy for availability of running water in a dwelling |
| Running sewerage | Dummy for availability of running sewerage in a dwelling |
| Country house | Dummy for a family having a country house |
| Saved money last month | Dummy for a family saving money during previous 30 days |
| Help from others | Dummy for a family receiving financial or in kind help from others |
| Credit debt | Dummy for a family having credit debt |
| Owing money to others | Dummy for a family being in debt to other individuals |

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

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

# 3.C Appendix C

## Robustness check for marriage dissolution estimate

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

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

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

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

## Controlling for family background of mothers

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

[40] The author focuses her analysis on the situation of Soviet women in the wake of WWII. Thus the author's claim is limited to women. It might apply to men as well. Nevertheless, it appears intuitively plausible that women who grew up with single mothers might be more confident about

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

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

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.

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

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

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

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

# Bibliography


Aaronson, Daniel, Fabian Lange, and Bhashkar Mazumder. 2014. "Fertility Transitions Along the Extensive and Intensive Margins." *American Economic Review* 104 (11): 3701–3724.

Abadie, Alberto, Susan Athey, Guido W Imbens, and Jeffrey M Wooldridge. 2022. "When Should You Adjust Standard Errors for Clustering?*." *The Quarterly Journal of Economics* 138 (1): 1–35 (10).

Abramishvili, Zurab, William Appleman, and Sergii Maksymovych. 2019. "Parental Gender Preference in the Balkans and Scandinavia: Gender Bias or Differential Costs?" Cerge-ei working paper series 623, Center for Economic Research and Ggraduate Education (CERGE-EI).

Adda, Jerome, Christian Dustmann, and Katrien Stevens. 2016, October. "The Career Costs of Children." Working paper 6158, CESifo.

Aduškina, K. 2015. "Determinants of Marriage Satisfaction in Young Families [Determinanty udovletvorennosti brakom v molodyh sem'jah]." *Proceedings of the 6th International Scientific Conference "Psychological Problems of Modern Family" [VI Meždunarodnaja Naučnaja Konferencija "Psihologičeskie Problemy Sovremennoj Sem'i." Tezisy.]*. Lomonosov Moscow State University, 443–449.

Aguero, Jorge M, and Mindy S Marks. 2011. "Motherhood and Female Labor Supply in the Developing World." *Journal of Human Resources*, pp. 800–826.

Aivazova, D. 2015. "Image of Marriage among Modern Megalopolis Residents [Obraz supružestva u sovremennyh žiteliei megapolisa]." *Proceedings of the 6th International Scientific Conference "Psychological Problems of Modern Family" [VI Meždunarodnaja Naučnaja Konferencija "Psihologičeskie Problemy Sovremennoj Sem'i." Tezisy.]*. Lomonosov Moscow State University, 456–461.

Almas, Ingvild, Alex Armand, Orazio Attanasio, and Pedro Carneiro. 2015. "Measuring and Changing Control: Women's Empowerment and Targeted Transfers." *NBER Working Papers*.

Altindag, Onur. 2016. "Son Preference, Fertility Decline, and the Nonmissing Girls of Turkey." *Demography* 53 (2): 541–66.

Ambler, Kate, and Alan De Brauw. 2017, 3. "The impacts of cash transfers on womens empowerment : learning from Pakistans BISP program." Technical Report, World Bank.

Ananat, Elizabeth O., and Guy Michaels. 2008. "The Effect of Marital Breakup on the Income Distribution of Women with Children." *Journal of Human Resources* 43 (3): 611 – 629.

Andaluz, Joaqin, and Jose A Molina. 2007. "On the sustainability of bargaining solutions in family: sustainability in family decision models." *Review of Economics of the Household*, pp. 405–418.

Andersson, Gunnar, Karsten Hank, Marit Ronsen, and Andres Vikat. 2006. "Gender Family Composition: Sex Preferences for Children and Childbearing Behavior in the Nordic Countries." *Demography* 43 (2): 255–267.

Angrist, Joshua D., and William N. Evans. 1998. "Children and Their Parents' Labor Supply: Evidence from Exogenous Variation in Family Size." *American Economic Review* 88 (3): 449–477.

Antonov, Sergej, and Aleksej Smagin. 2021. "Why Russians Divorce. The Most Frequent Causes. [Počemu rossijane razvodjatsja. Samyje chastyje pričiny.]." *Tinkoff Journal*.

Antončič, Emica, Menka Mencin Celpak, and Mirjana Ule. 2018, 10. Struggles for equality: Feminism in Slovenia | Eurozine.

Apps, Patricia, and Ray Rees. 2008. "Testing Pareto Efficiency of Household Resource Allocations." *CESifo Working Papers*, pp. 1–11.

Arnold, Catherine, Tim Conway, and Matthew Greenslade. 2011. "DFID Cash Transfers Literature Review - The CALP Network." Technical Report, Department for International Development.

Arnold, Fred. 1992. "Sex Preference and Its Demographic and Health Implications." *International Family Planning Perspectives* 18, no. 3.

Arnold, Fred, Minja Kim Choe, and T. K. Roy. 1998. "Son preference, the family-building process and child mortality in India." *Population Studies* 52 (3): 301 – 315.

Arnórsdóttir, Agnes S. 2017. “Death and donation: Different channels of property transfer in late medieval Iceland.” *The Marital Economy in Scandinavia and Britain 1400-1900*, 1, 207–220.

Asendorpf, J. B., and S. Wilpers. 1998. “Personality effects on social relationships.” *Journal of Personality and Social Psychology,* 74:1531–1544.

Ashraf, Nava. 2009. “Spousal Control and Intra-household Decision Making: An Experimental Study in the Philippines.” *American Economic Review* 99 (4): 1245 – 1277.

Attanasio, Orazio, and Valerie Lechene. 2010. “Conditional cash transfers, women and the demand for food.” *Institute for Fiscal Studies Working Papers.*

Autor, David, David Dorn, and Gordon Hanson. 2017, February. “When Work Disappears: Manufacturing Decline and the Falling Marriage-Market Value of Young Men.” Working paper 21173, National Bureau of Economic Research.

Baker, Paula. 1984. “The Domestication of Politics: Women and American Political Society, 1780-1920.” *The American Historical Review* 89 (6): 620.

Baranowska-Rataj, A, and A Matysiak. 2016. “The Casual Effects of the Number of CHildren on Female Employment—Do European Institutional and Gender Conditions Matter?” *Journal of Labor Research.*

Barcellos, Silvia Helena, Leandro S Carvalho, and Adriana Lleras-Muney. 2014. “Child Gender and Parental Investments in India: Are Boys and Girls Treated Differently?” *American Economic Journal: Applied Economics* 6 (1): 157–189.

Bastagli, Francesca, Jessica Hagen-Zanker, Luke Harman, Valentina Barca, Georgina Sturge, Tanja Schmidt, and Luca Pellerano. 2016, 7. “Cash transfers: what does the evidence say?” Technical Report, Overseas Development Institute.

Basu, Deepankar, and Robert De Jong. 2010. “Son Targeting Fertility Behavior: Some Consequences and Determinants.” *Demography* 47 (2): 521–536.

Basu, Kaushik. 2006. “Gender and Say: a Model of Household Behaviour with Endogenously Determined Balance of Power.” *Economic Journal* 116 (511): 558–580 (04).

Battaglia, Marianna, Bastien Chabé-Ferret, and Lara Lebedinski. 2021. “Segregation, fertility, and son preference: the case of the Roma in Serbia.” *Journal of Demographic Economics* 87 (6): 233–260.

Bedard, K, and O Deschenes. 2005. “Sex Preferences, Marital Dissolution and the Economic Status of Women.” *Journal of Human Resources* 40 (2): 411–434.

Ben-Porath, Yoram, and Finis Welch. 1976. "Do Sex Preferences Really Matter?" *Quarterly Journal of Economics* 90 (2): 285 – 307.

Bereczkei, Tamas, and R. I.M. Dunbar. 2002. "HelpingattheNest and SexBiased Parental Investment in a Hungarian Gypsy Population." *https://doi.org/10.1086/344374* 43:804–809.

Bičáková, Alena. 2016. "Gender unemployment gaps in the EU: blame the family." *IZA Journal of European Labor Studies* 22 (5): 1–31.

Bičáková, Alena, and Klara Kališková. 2016. "Career Breaks after Childbirth: The Impact of Family Leave Reforms in the Czech Republic." Working paper 568, CERGE-EI.

Blacklow, Paul. 2002, May. "Intertemporal Equivalence Scales: Measuring the Life-Cycle Costs of Children." Memorandum 8, University of Oslo Department of Economics.

Blau, Francine D. 1997, October. "Trends in the Well-Being of American Women, 1970-1995." Working paper 6206, National Bureau of Economic Research.

Blau, Francine D., and Lawrence M. Kahn. 2017. "The Gender Wage Gap: Extent, Trends, and Explanations." *Journal of Economic Literature* 55 (3): 789–865 (September).

Bobonis, Gustavo J. 2009. "Is the Allocation of Resources within the Household Efficient? New Evidence from a Randomized Experiment." *Journal of Political Economy*, pp. 453–503.

Boertien, Diederik, and Dimitri Mortelmans. 2018. "Does the relationship between personality and divorce change over time? A cross-country comparison of marriage cohorts." *Acta Sociologica* 61 (3): 300–316.

Boertien, Diederik, Christian von Scheve, and Mona Park. 2017. "Can Personality Explain the Educational Gradient in Divorce? Evidence From a Nationally Representative Panel Survey." *Journal of Family Issues* 38 (10): 1339–1362.

Bogan, Vicki L. 2013. "Household investment decisions and offspring gender: parental accounting." *Applied Economics* 31 (45): 4429–4442.

Bojer, Hilde. 2002, May. "The Time Cost of Children and Equivalent Full Incomes." Memorandum 8, University of Oslo Department of Economics.

Bond, Timothy N., and Kevin Lang. 2013. "The Evolution of the Black-White Test Score Gap in Grades K-3: the Fragility of Results." *The Review of Economics and Statistics* 95 (5): 1468–1479.

Bradbury, Bruce. 2004, January. "The Price, Cost, Consumption and Value of Children." Supporting children: English-speaking countries in international context, workshop presentation paper, Princeton University.

Bradley, Robert H, and Bettye M Caldwell. 1980. "The Relation of Home Environment, Cognitive Competence, and IQ among Males and Females." *Child Development* 51:1140–1148.

———. 1981. "The HOME Inventory: A Validation of the Preschool Scale for Black Children." *Child Development* 52:708–710.

———. 1984. "The Relation of Infants' Home Environments to Achievement Test Performance in First Grade: A Follow-up Study." *Child Development* 55:803–809.

Braido, Luis H.B., Pedro Olinto, and Helena Perrone. 2012. "Gender Bias in Intrahousehold Allocation: Evidence from and Unintentional Experiment." *Review of Economics and Statistics*, pp. 552–565.

Brainerd, Elizabeth. 2008. Uncounted Costs of World War II: The Effect of Changing Sex Ratios on Marriage and Fertility of Russian Women, unpublished manuscript.

———. 2016. "The Lasting Effect of Sex Ratio Imbalance on Marriage and Family: Evidence from World War II in Russia." Iza discussion papers 10130, Institute for the Study of Labor (IZA), Bonn.

Brockmann, Hilke. 2001. "Girls Preferred? Changing Patterns of Sex Preferences in the Two German States." *European Sociological Review* 17 (2): 189–202.

Brown, Susan L. 2004. "Family Structure and Child Well-Being: The Significance of Parental Cohabitation." *Journal of Marriage and Family* 66 (2): 351–367.

Burman, Douglas D., Tali Bitan, and James R. Booth. 2008. "Sex Differences in Neural Processing of Language Among Children." *Neuropsychologia* 46 (5): 1349–1362.

Cameron, Adrian Colin, and Pravin K. Trivedi. 2005. *Microeconometrics: Methods and Applications*. Cambridge University Press.

Cameron, Elissa Z., and Fredrik Dalerum. 2009. "A TriversWillard effect in contemporary humans: male-biased sex-ratios among billionaires." *PLoS ONE* 4, no. 1.

Card, David. 1999. "Chapter 30 - The Causal Effect of Education on Earnings." In , edited by Orley C. Ashenfelter and David Card, Volume 3 of *Handbook of Labor Economics*, 1801–1863. Elsevier.

Choi, Eleanor Jawon, and Jisoo Hwang. 2015. "Child Gender and Parental Inputs: No More Son Preference in Korea?" *American Economic Review: Papers and Proceedings* 105 (5): 638–643.

Choi, Hyung-Jai, Jutta M. Joesch, and Shelly Lundberg. 2008a. "Sons, daughters, wives, and the labour market outcomes of West German men." *Labour Economics* 15:795–811.

Choi, Hyung-Jai, Jutta M Joesch, and Shelly Lundberg. 2008b. "Sons, daughters, wives, and the labor market outcomes of West German men." *Labour Economics* 15:795–811.

Clark, Shelley. 2000. "Son Preference and Sex Composition of Children: Evidence from India." *Demography* 37 (1): 95–108.

Claxton, A, A O'Rourke, JZ Smith, and et al. 2012. "Personality traits and marital satisfaction within enduring relationships." *Journal of Social and Personal Relationships* 29 (3): 375–396.

Concern Worldwide, Oxfam GB. 2011. Walking the Talk: Cash Transfers and Gender Dynamics - The CALP Network.

Cools, Angela, and Eleonora Patacchini. 2017, September. "Sibling Gender Composition and Women's Wages." Discussion paper 11001, IZA.

Cools, Sara, Jon H. Fiva, and Lars J. Kirkeboen. 2015. "Causal Effects of Paternity Leave on Children and Parents." *Scandinavian Journal of Economics* 117 (3): 801 – 828.

Cunha, Flavio, James J Heckman, and Susanne M Schennach. 2010. "Estimating the Technology of Cognitive and Noncognitive Skill Formation." *Econometrica* 78 (3): 883–931.

Currie, Janet, and Douglas Almond. 2011. "Human capital development before age five." In *Handbook of Labor Economics*, edited by David Card and Orley Ashenfelter, Volume 4, 1315–1486. Elsevier.

Cygan-Rehm, Kamila, Daniel Kuehnle, and Regina K Riphahn. 2018. "Paid parental leave and families' living arrangements." *Labour Economics* 53:182–197.

Dahl, Gordon, and Enrico Moretti. 2008. "The Demand for Sons." *The Review of Economic Studies* 75:1085–1120.

Davis, Devra Lee, Michelle B. Gottlieb, and Julie R. Stampnitzky. 1998. "Reduced Ratio of Male to Female Births in Several Industrial Countries: A Sentinel Health Indicator?" *Journal of American Medical Association* 279 (13): 1018–1023.

Deaton, Angus. 2010. "Instruments, Randomization, and Learning about Development." *Journal of Economic Literature* 48 (2): 424–55.

Deaton, Angus, and John Muellbauer. 1986. "On measuring child costs: with applications to poor countries." *Journal of Political Economy* 94:720–745.

De la Briere, Benedicte, and Laura B Rawlings. 2006. "Examining Conditional Cash Transfer Programs: A Role for Increased Social Inclusion?" *The World Bank Social Safety Net Primer Series*, no. 0603.

de la Croix, David, and Matthias Doepke. 2003. "Inequality and Growth: Why Differential Fertility Matters." *American Economic Review*, pp. 1091–1113.

Diederik, Boertien, and Dmitri Mortelmans. 2018. "Does the Relationship between Personality and Divorce Change over Time?" *Acta Sociologica*, pp. 300–316.

Donellan, MB, RD Conger, and CM Bryant. 2004. "The Big Five and enduring marriages." *Journal of Research in Personality* 38 (5): 481–504.

Duflo, Esther. 2003. "Grandmothers and Granddaughters: Old-Age Pensions and Intrahousehold Allocation in South Africa." *The World Bank Economic Review.*

———. 2012. "Women Empowerment and Economic Development." *Journal of Economic Literature*, pp. 1051 – 1079.

Dujardin, Claire, and Florence Goffette-Nagot. 2009. "Does public housing occupancy increase unemployment?" *Journal of Economic Geography* 9:823–851.

Dübeck, Inger. 2017. "Property and authority in Danish marital law." *The Marital Economy in Scandinavia and Britain 1400-1900*, 1, 127–140.

Economist. 2018. Why polygamy breeds civil war. `https://www.economist.com/blogs/economist-explains/2018/03/economist-explains-16`. Accessed: 2019-07-04.

Edlund, Lena. 1999. "Son Preference, Sex Ratios, and Marriage Patterns." *Journal of Political Economy* 107, no. 6.

Edvinsson, Rodney, and Therese Nordlund Edvinsson. 2017. "Explaining the Swedish housewife era of 19301970: joint utility maximisation or renewed patriarchy?" *Scandinavian Economic History Review* 65 (5): 169–188. doi: 10.1080/03585522.2017.1323671.

Elizabeth, Vivienne. 2001. "Managing money, managing coupledom: a critical investigation of cohabitants' money management practices." *Sociological Review*, pp. 389–411.

Estrin, Saul, and M Uvalic. 2014. "FDI into transition economies. Are the Balkans different?" *Economics of Transition*, pp. 281–312.

Eurocommission. 2010, February. "Gender equality in the EU in 2009." Special eurobarometer 326, European Commission.

Eurostat. 2015. *ESS handbook for quality reports.* Luxembourg: Publications Office of the European Union.

———. 2017. *European Union Statistics on Income and Living Conditions.* http://ec.europa.eu/eurostat/web/microdata/european-union-statistics-on-income-and-living-conditions [Accessed: 18.11.17].

Evertsson, Lars, and Charlotte Nyman. 2014. "Perceptions and Practices in Independent Management: Blurring the Boundaries Between "Mine", "Yours" and "Ours"." *Journal of Family and Economic Issues*, pp. 65–80.

Exley, Christine L, and Judd B Kessler. 2022. "The Gender Gap in Self-Promotion*." *The Quarterly Journal of Economics*, 01. qjac003.

Fani, Seddighe, and Aghile N Kheirabadi. 2011. "Personality Traits and Mental Divorce." *Procedia - Social and Behavioral Sciences*, pp. 671–675.

Fiebranz, Rosemarie. 2017. "Marital conflict over the gender division of labour in agrarian households, Sweden 17501850." *The Marital Economy in Scandinavia and Britain 1400-1900*, 1, 141–156.

Filmer, Deon, Jed Friedman, and Norbert Schady. 2009. "Development, Modernization, and Childbearing: The Role of Family Sex Composition." *World Bank Economic Review* 23 (3): 371–398.

Forum, World Economic. 2018. The Global Gender Gap Report 2018.

Fusco, Alessio, Anne-Catherine Guio, and Eric Marlier. 2010. "Income and Living Conditions in Europe." Technical Report.

Galor, Oded. 2011. *Unified growth theory.* Princeton : Princeton University Press.

Gerry, Christopher, and Georgios Papadopoulos. 2015. "Sample attrition in the RLMS, 200110." *The Economics of Transition* 23 (2): 425–468.

Goncharenko, Roman. 2018. "Russia votes: Vladimir Putin is counting on female turnout." *Deutsche Welle.*

Grandits, Hannes, Xavier Bougarel, Nathalie Clayer, and Fabio Giomi. 2020. "Patriarchal and heroic re- and deconstructions: a tribute to and critical reflections on four books of Karl Kaser."

Grech, Victor, Charles Savona-Ventura, and P Vassallo-Agius. 2002. "Unexplained differences in sex ratios at birth in Europe and North America." *BMJ* 324 (7344): 1010–1011.

Guilmoto, Christophe Z., and Geraldine Duthe. 2013, December. "Masculinization of births in Eastern Europe." Population & society: Monthly bulletin 506, The French National Institute for Demographic Studies.

Guilmoto, Christophe Z, and Géraldine Duthé. 2013. "Masculinization of births in Eastern Europe." Population and societies, monthly bulletin 506, French National Institute for Demographic Studies.

Guiso, Luigi, Ferdinando Monte, Paola Sapienza, and Luigi Zingales. 2008. "Culture, Gender, and Math." *Science* 320:1164–1165.

Gupta, Bishnupriya. 2014. "Where Have All the Brides Gone? Son Preference and Marriage in India over the Twentieth Century." *Economic History Review* 67 (1): 1 – 24.

Hank, Karsten. 2007. "Parental Gender Preferences and Reproductive Behavior: A Review of the Recent Literature." *Journal of Biosocial Sciences* 39:759–767.

Hank, Karsten, and Hans-Peter Kohler. 2000. "Gender Preferences for Children in Europe: Empirical Results from 17 FFS Countries." *Demographic Research* 2, no. 1.

Hao, Lingxin, and Wei-Jun Jean Yeung. 2015. "Parental Spending on School-Age Children: Structural Stratification and Parental Expectation." *Demography*, no. 52:835–860.

Haughton, Jonathan, and Dominique Haughton. 1998. "Are simple tests of son preference useful? An evaluation using data from Vietnam." *Journal of Population Economics* 11:495–516.

Haushofer, Johannes, and Jeremy Shapiro. 2016. "The Short-Term Impact of Unconditional Cash-Transfers to the Poor: Experimental Evidence from Kenya." *Quarterly Journal of Economics*.

Hazan, Moshe, and Hosny Zoabi. 2015. "Sons or Daughters? Sex Preferences and the Reversal of the Gender Educational Gap." *Journal of Demographic Economics* 81:179–201.

Heaven, PCL, L Smith, SM Prabhakar, and et al. 2006. "Personality and conflict communication patterns in cohabiting couples." *Journal of Research in Personality* 40 (5): 829–840.

Heckman, James J., and Burton Singer. 2017. "Abducting Economics." *The American Economic Review* 107 (5): 298.

Heeringa, Steven G. 1997, March. "Russia Longitudinal Monitoring Survey (RLMS) Sample Attrition, Replenishment, and Weighting in Rounds V-VII ." Technical Report, Survey Design and Analysis Unit, Institute for Social Research, University of Michigan Ann Arbor, MI 48106-1248.

Hirschberger, Gilad, Sanjay Srivastava, and Penny Marsh. 2009. "Attachment, Marital Satisfaction, and Divorce During the First Fifteen Years of Parenthood." *Personal Relationships* 16 (3): 401–420.

Hoddinott, John, and Lawrence Haddad. 1995. “Does Female Income Share Influence Household Expenditures?” *Oxford Bulletin of Economics and Statistics*, pp. 77–96.

Hurrell, Frank E Jr. 2015. *Regression Modeling Strategies. With Applications to Linear Models, Logistic and Ordinal Regression, and Survival Analysis.* Springer International Publishing Switzerland.

Husain, Muna, and Daniel L. Millimet. 2009. “The mythical boy crisis?” *Economics of Education Review* 28 (1): 38 – 48.

Ichino, Andrea, Ally-Ann Lindstr´om, and Eliana Viviano. 2011, April. “Hidden Consequences of a First-Born Boy for Mothers.” Discussion papers 5649, IZA.

Jacobsen, Rune, Henrik Møller, and Gerda Engholm. 1999. “Fertility rates in Denmark in relation to the sexes of preceding children in the family.” *Human Reproduction* 14 (4): 1127–1130.

Jayachandran, Seema. 2017. “Fertility Decline and Missing Women.” *American Economic Journal: Applied Economics* 9 (1): 118–139.

Jayachandran, Seema, and Ilyana Kuziemko. 2011. “Why Do Mothers Breastfeed Girls Less Than Boys? Evidence and Implications for Child Health in India.” *The Quarterly Journal of Economics* 126:1485–1538.

Jensen, Robert T. 2003. “Equal Treatment, Unequal Outcomes? Generating Sex Inequality Through Fertility Behavior.” Working paper, Harvard University John F. Kennedy School of Government.

Jiang, Quanbao, Ying Li, and Jesus J. Sanchez-Barricarte. 2016. “Fertility Intention, Son Preference, and Second Childbirth: Survey Findings from Shaanxi Province of China.” *Social Indicators Research* 125:935–953.

Johansen, Hanne Marie. 2017. “Marriage or money? Legal actions for enforcement of marriage contracts in Norway.” *The Marital Economy in Scandinavia and Britain 1400-1900*, 1, 23–38.

Juhn, Chinhui, Yona Rubinstein, and C. Andrew Zuppann. 2015, December. “The Quantity-Quality Trade-off and the Formation of Cognitive and Non-cognitive Skills.” Working paper 21824, National Bureau of Economic Research.

Jurajda, Štěpán. 2021. *Econometrics III. Lecture Notes. URL: http://home.cerge-ei.cz/jurajda/ecmiv.pdf.*

Kabátek, Jan, and David C Ribar. 2020. “Daughters and Divorce.” *The Economic Journal* 131 (637): 2144–2170 (12).

Kališková, Klara, Daniel Münich, and Filip Pertold. 2016. “Public financing for pre-school places pays off: a cost and benefit analysis.” *IDEA studies*, no. 3.

Kanazawa, Satoshi. 2007. "Beautiful parents have more daughters: A further implication of the generalized TriversWillard hypothesis (gTWH)." *Journal of Theoretical Biology* 244 (1): 133 – 140.

Karbownik, Krzysztof, and Michal Myck. 2017. "Who gets to look nice and who gets to play? Effects of child gender on household expenditures." *Review of Economics of the Household* 15:925–944.

Kaser, Karl. 2008. *Patriarchy after patriarchy : gender relations in Turkey and in the Balkans, 1500-2000.* Berlin : Lit ; London : Distributed in UK by Global Book Marketing.

Katzev, Aphra R, Rebecca L Warner, and Alan C Acock. 1994. "Girls or Boys? Relationship of Child Gender to Marital Instability." *Journal of Marriage and Family* 56 (1): 89–100.

Kim, Minsuk, and Judith Shapiro. 2021. "(Revealed) Preference for Sons and Daughters: Estimates from the Russian Longitudinal Monitoring Survey (RLMS-HSE), 1994-2019." *5th International RLMS-HSE User Conference. Session 6.* Higher School of Economics.

Kippen, Rebecca, Ann Evans, and Edith Gray. 2006. "Parental Preference for Sons and Daughters in a Western Industrial Setting: Evidence and Implications." *Journal of Biosocial Science* 39:583–597.

Klimashevskaya, Olga. 2021. "The Problem of Gender Asymmetry in the Family Legislation of the Russian Federation [Problema Gendernoj Asimmetrii v Semejnom Zakonodatel'stve Rossijskoj Federacii]." *Vlast'*, no. 1:53–58.

Kornienko, Dmitrii S., and Yelena A. Silina. 2020. "Characteristics of Personality and Assessment of Family Environment by Spouses with Different Marriage Duration [Ličnostnye čerty i ocenka semejnoj sredy suprugami s različnym stažem braka]." *Vestnik udmurtskogo universiteta. Seria: filosofia, psichologia, pedagogika.* 2 (1): 48–54.

Kornrich, Sabino, and Frank Furstenberg. 2007. "Investing in Children: Changes in Parental Spending on Children, 1972 to 2007." Unpublished paper.

Kozyreva, Polina, Mikhail Kosolapov, and Barry M Popkin. 2016. "Data Resource Profile: The Russia Longitudinal Monitoring SurveyHigher School of Economics (RLMS-HSE) Phase II: Monitoring the Economic and Health Situation in Russia, 19942013." *International Journal of Epidemiology* 45 (2): 395–401 (02).

Kozyreva, Polina, Mikhail Kosolapov, and Elena Tonis. 2016. "The Russia Longitudinal Monitoring Survey - Higher School of Economics (RLMS-HSE): Phase II, 19942015." *5th Panel Survey Methods Workshop, 20th-21st June 2016, Berlin.*

Larsen, Ulla, Woojin Chung, and Monica Das Gupta. 1998. “Fertility and son preference in Korea.” *Population Studies* 52:317–325.

Lee, Jungmin, and Mark L. Pocock. 2007. “Intrahousehold allocation of financial resources: evidence from South Korean individual bank accounts.” *Review of Economics of the Household*, pp. 41–58.

Lerch, Mathias. 2018. “Fertility and union formation during crisis and societal consolidation in the Western Balkans.” *Population Studies* 72 (5): 217–234. doi: 10.1080/00324728.2017.1412492.

Leung, Siu Fai. 1991. “A Stochastic Dynamic Analysis of Parental Sex Preferences and Fertility.” *Quarterly Journal of Economics*, pp. 1063–1088.

Li, Lixing, and Xiaoyu Wu. 2011. “Gender of Children, Bargaining Power, and Intrahousehold Resource Allocation in China.” *Journal of Human Resources*, pp. 295–316.

Lindstrom, Elly-Ann. 2013. “Gender Bias in Parental Leave: Evidence from Sweden.” *Journal of Family and Economic Issues* 34:235–248.

Lino, Mark, and Andrea Carlson. 2010. “Estimating Housing Expenses on Children: Comparison of Methodologies.” *Journal of Legal Economics* 16 (2): 61–79.

Lundberg, Shelly. 2005. “Sons, Daughters, and Parental Behavior.” *Oxford Review of Economic Policy* 21, no. 3.

———. 2012. “Personality and marital surplus.” *IZA Journal of Labor Economics* 3, no. 1.

Lundberg, Shelly, Sara McLanahan, and Eliana Rose. 2007. “Child Gender and Father Involvement in Fragile Families.” *Demography* 44 (1): 79–92.

Lundberg, Shelly, and Eliana Rose. 2002. “The Effects of Sons and Daughters on Men’s Labor Supply and Wages.” *The Review of Economics and Statistics* 84, no. 2.

———. 2003a. “Child Gender and the Transition to Marriage.” *Demography* 40 (2): 333–349.

———. 2003b, February. “Investments in Sons and Daughters: Evidence from the Consumer Expenditure Survey.” Unpublished manuscript, Joint Center for Poverty Research.

Lundberg, Shelly, and Jennifer Ward-Batts. 2000. “Saving for Retirement: Household Bargaining and Household Net Worth.” Working paper. michigan retirement research center, University of Michigan.

Lundberg, Shelly J, and Robert A Pollak. 1993. "Separate Spheres Bargaining and the Marriage Market." *Journal of Political Economy*, pp. 988–1010.

———. 2007. "The American Family and Family Economics." *Journal of Economic Perspectives* 21 (2): 3–26.

———. 2008. "Family decision making." In *The New Palgrave: A Dictionary of Economics.*

Lundberg, Shelly J, Robert A Pollak, and Terrence J Wales. 1997. "Do Husbands and Wives Poool Their Resources? Evidence from the United Kingdom Child Benefit." *The Journal of Human Resources*, pp. 463–480.

Mammen, Kristin. 2008. "The Effect of Children's Gender on Living Arrangements and Child Support." *American Economic Review: Papers and Proceedings* 98, no. 2.

Mansour, Hani, and Terra McKinnish. 2013, December. "Couples' Time Together: Complementarities in Production versus Complementarities in Consumption." Iza discussion papers 7848, Institute for the Study of Labor (IZA).

McKenzie, David J. 2005. "Measuring inequality with asset indicators." *Journal of Population Economics* 18:229–260.

McLanahan, Sara, Elisabeth Donahue, and Ron Haskins. 2005. "Introducing the Issue." *The Future of Children* 15 (2): 3–12.

Milekic, Sven. 2018, 5. Croatia Needs More Dads  and Fewer Fathers.

Mills, Melinda, and Katia Begall. 2010. "Preferences for the sex-composition of children in Europe: A multi-level examination of its effect on progression to a third child." *Population Studies*, pp. 77–95.

Mkchtrian, Gayane. 2017. "What is better, to have a son or a daughter? [Lučše imet' syna ili doč'?]." *Ziarul de Garda.*

Moffit, Robert. 2005. "Remarks on the Analysis of Causal Relationships in Population Research." *Demography* 42:91–108.

Morrica, Valerie, Tara Sharafudheen, Paul Andres Corral Rodas, Ursula Casabonne, and Zuzana Boehmova. 2019, 12. Investing in Opportunities for All Croatia Country Gender Assessment.

Morrill, Melinda, and Sabrina Pabilonia. 2015. "What effects do macroeconomic conditions have on the time couples with children spend together?" *Review of Economics of the Household* 13 (4): 791–814 (December).

Mott, Frank L. 1994. "Sons, Daughters and Fathers' Absence:: Differentials in Father-Leaving Probabilities and in Home Environments." *Journal of Family Issues* 15 (1): 97–128.

Myck, Michal, Monika Oczkowska, and Izabela Wowczko. 2021. "Gender Preferences in Central and Eastern Europe as Reflected in Partnership and Fertility Outcomes." Iza discussion papers 14244, Institute for the Study of Labor (IZA).

Nagel, Mechthild. 2013. "PATRIARCHAL IDEOLOGIES AND WOMENS DOMESTICATION." *The End of Prisons*, 1, 147–167.

Natali, Luisa, Sudhanshu Handa, Amber Peterman, David Seidenfeld, and Gelson Tembo. 2016. "Making Money Work. Unconditional cash transfers allow women to save and re-invest in rural Zambia." Innocenti working paper, UNICEF Office of Research.

Nikolajeva, L. A. 2018. "Psychological Pecularities of Different Types of Modern Family [Psihologičeskje Osobennosti Različnyh Vidov Sovremennoj Sem'i]." *Čelovečeskii factor: Social'nyj Psiholog* 36 (2): 197–206.

Norberg, Karen. 2004. "." *Proceedings of the Royal Society* 271 (15555): 2403–2410.

Outram, Quentin. 2015, January. "Son Preference Aming the Edwardian Middle Class." Unpublished manuscript.

Pabilonia, Sabrina W, and Jennifer Ward-Batts. 2007. "The Effect of Child Gender on Parents' Labor Supply." *American Economic Review: Papers and Proceedings* 97 (2): 402–406.

Pahl, Jan. 2005. "Individualisation in Couple Finances: Who Pays for the Children?" *Social Policy and Society*, pp. 381–391.

———. 2008. "Family finances, individualisation, spending patterns and access to credit." *The Journal of Socio-Economics*, pp. 577–591.

Pollak, Robert. 1985. "A Transaction Cost Approach to Families and Households." *Journal of Economic Literature*, pp. 581 – 608.

Pollard, Michael S., and S. Philip Morgan. 2002. "EMERGING PARENTAL GENDER INDIFFERENCE? SEX COMPOSITION OF CHILDREN AND THE THIRD BIRTH." *American Sociological Review* 67 (4): 600 – 613.

Pylkkänen, Anu. 2017. "Forming the marital economy in the early modern Finnish countryside." *The Marital Economy in Scandinavia and Britain 1400-1900*, 1, 75–88.

Ragan, K. 2013. "Taxes and Time Use: Fiscal Policy in a Household Production Model." *American Economic Journal: Macroeconomics*, pp. 168–192.

Raurich, Xavier, and Thomas Seegmuller. 2017, July. "Growth and bubbles: the interplay between productive investment and the coost of rearing children." Working paper 26, AMSE.

Roberts, Russel, and Roy Baumeister. 2011, November. Baumeister on Gender Differences and Culture.

Rogers, Barbara. 1980. *The domestication of women : discrimination in developing societies.* Kogan Page.

Sadowski, Michael. 2010. “Putting the ”Boy Crisis” in Context.” *Education Digest: Essential Readings Condensed for Quick Review* 76 (3): 10–13.

Sandstrom, Glenn, and Lotta Vikstrom. 2015. “Sex preference for children in German villages during the fertility transition.” *Population Studies* 69 (1): 57–71.

Sandvik, Hilde. 2017. “Decision-making on marital property in Norway, 15001800.” *The Marital Economy in Scandinavia and Britain 1400-1900*, 1, 111–126.

Schady, Norbert, and Jose Rosero. 2007. “Are Cash Transfers Made to Women Spent Like other Sources of Income.” *The World Bank Policy Research Working Papers*, no. 4282.

Schneebaum, Alyssa, and Katharina Mader. 2013, October. “The gendered nature of intra-household decision making in and across Europe.” Department of economics working paper series 157, WU Vienna University of Economics and Business, Vienna.

Schultz, T. Paul. 2001. “Chapter 8 Women’s roles in the agricultural household: Bargaining and human capital investments.” In *Agricultural Production*, Volume 1, Part A of *Handbook of Agricultural Economics*, 383 – 456. Elsevier.

Seidl, Christian. 1995. “The Desire for a Son Is the Father of Many Daughters: A Sex Ratio Paradox.” *Journal of Population Economics* 8, no. 2.

Shvetsova, M.N., and E.A. Kondrasheva. 2015. “Gender-Specific Representations of the Family Values of Spouses with Different Experience of Marriage.” *Modern problems of science and education* 1 (1): 1575–1575.

Sienaert, Alex. 2008, May. “Some Child Cost Estimates for South Africa.” Working paper series 15, CSAE.

Simon, Claire E. 2019. “The effect of cash-based interventions on gender outcomes in development and humanitarian settings | UN Women Headquarters.” Technical Report, United Nations Entity for Gender Equality and the Empowerment of Women (UN Women).

Simonoff, J.S. 2003. *Analyzing Categorical Data.* Springer Texts in Statistics. Springer New York.

Solomon, Brittany C., and Joshua J. Jackson. 2014. "Why do personality traits predict divorce? Multiple pathways through satisfaction." *Journal of Personality and Social Psychology* 106 (6): 978–996.

Somville, Vincent. 2019, December. "Having a Daughter Reduces Male Violence Against a Partner." Discussion paper series in economics 24/2019, Norwegian School of Economics, Department of Economics.

Spiegel, Henry W. 2013. "Xenophon (c430 bc - c 355 bc)." In *The New Palgrave: A Dictionary of Economics*.

Thomas, Duncan. 1990. "Intra-Household Resource Allocation: An Inferential Approach." *Journal of Human Resources*, pp. 635–664.

Todd, Petra E, and Kenneth I Wolpin. 2007. "The Production of Cognitive Achievement in Children: Home, School, and Racial Test Score Gaps." *Journal of Human Capital* 1 (1): 91–136.

Trivers, Robert L., and Dan E. Willard. 1973. "Natural Selection of Parental Ability to Vary the Sex Ratio of Offspring." *Science* 179 (4068): 90–92.

Udry, Christopher. 1996. "Gender, Agricultural Production and the Theory of the Household." *Journal of Political Economy*, pp. 1010–1046.

UNFPA. 2020. Preventing Son Preference and Undervaluing of Girls in Eastern Europe and Central Asia.

van Lent, Max. 2020. "Fathering Daughters and Personality." Technical Report, Leiden University.

van Praag, Bernard M.S., and Marcel F. Warnaar. 1997. "Chapter 6 The cost of children and the use of demographic variables in consumer demand." In *Handbook of Pupolation and Family Economics*, edited by M.R. Rosenzweig and O. Stark, Volume 1, 241 – 273. Elsevier.

Vogl, Tom S. 2013. "Marriage Institutions and Sibling Competition: Evidence from South Asia." *Quarterly Journal of Economics* 128 (3): 1017 – 1072.

Vogler, Carolyn, and Jan Pahl. 1994. "Money, power and inequality within marriage." *Sociological Review*, pp. 263–288.

Voicu, Bogdan, Malina Voicu, and Katarina Strapkova. 2007. "Engendered housework. A cross-european analysis." *IRISS Working Papers*.

Voroncova, Julija, and Viktor Ermolaev. 2016. "Gender Perceptions of Family among Generation Z: Cross-Cultural Analysis [Gendernyje predstavlenija o sem'je u pokolenija Z: kross-kul'turnyj analiz]." *Psiholog* - (3): 79–95.

Washington, Ebonya L. 2008. "Female Socialization: How Daughters Affect Their Legislator Fathers Voting on Womens Issues." *American Economic Review* 98 (1): 311332.

Wooldridge, Jeffrey M. 2002. *Econometric analysis of cross section and panel data.* MIT Press.

Xu, Jun. 2016. “Patriarchy, Gendered Spheres, or Evolutionary Adaptation? A Cross-National Examination of Adolescent Boys and Girls Access to Home Resources.” *Chinese Sociological Review* 48 (3): 209–247.

Yamaguchi, Kazuo, and Linda R. Ferguson. 1995. “The Stopping and Spacing of Childbirths and their Birth-History Predictors: Rational-Choice Theory and Event-History Analysis.” *American Sociological Review* 60:272–298.

Yoon, Yong J. 2006. “Gender Imbalance: The Male/Female Sex Ratio Determination.” *Journal of Bioeconomics* 8:253–268.

Yusof, Selamah Abdullah, and Jarita Duasa. 2010. “Household Decision-Making and Expenditure Patterns of Married Men and Women in Malaysia.” *Journal of Family and Economic Issues* 31 (3): 371–381 (September).

Zare, Bahare, Rohany Nasir, Khariul A Mastor, and W S Wan Shahrazad. 2013. “Personality Traits, The Risk of Divorce and Marital Satisfaction: An Intrapersonal Model.” *The Social Sciences*, pp. 466–472.

Zelenskaja, M., and A. Liders. 2015. “The Role of Empathy in Spousal Relations [Rol’ empatii v suprižeskih otnošenijah].” *Proceedings of the 6th International Scientific Conference "Psychological Problems of Modern Family" [VI Meždunarodnaja Naučnaja Konferencija "Psihologičeskie Problemy Sovremennoj Sem’i." Tezisy.].* Lomonosov Moscow State University, 497–501.