Working Paper Series 643
(ISSN 1211-3298)

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

Zurab Abramishvili
William Appleman
Sergii Maksymovych

CERGE-EI
Prague, August 2019

ISBN 978-80-7343-450-2 (Univerzita Karlova, Centrum pro ekonomický výzkum a doktorské studium)
ISBN 978-80-7344-503-4 (Národohospodářský ústav AV ČR, v. v. i.)

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

Zurab Abramishvili †, William Appleman, Sergii Maksymovych ‡

July 28, 2019

Abstract

There is much 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.

JEL codes: J13, J16, O15

Keywords: gender preference, gender differences, parental influence, household expenditure

*We would like to thank Štěpán Jurajda, Randall Filer, Alena Bičáková, and Patrick Gaulé for their insightful, detailed and generous comments on several drafts of this paper and to Daniel Münich for useful comments and access to the research data. Moreover, our paper benefited substantially from discussions with Nikolas Mittag, Simon Clark, and Maxym Brukhanov and from questions and remarks by participants in the CES 2016 and MiC 2018 conferences and the research seminar at Mendel University in Brno. Deborah Nováková did scrupulous and helpful proofreading. The authors are responsible for all errors. The research was supported by the CERGE-EI Foundation.

†International School of Economics at Tbilisi State University, 16 Zandukeli Street, Tbilisi 0108, Georgia

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

1 Introduction

The impact of the gender of the first-born child on the number of children in a family has been repeatedly observed in many countries. We confirm son preference using the parity-three progression method applied to a pooled EU-SILC 2004-2015 cross-sectional sample from four Balkan countries: Bulgaria, Croatia, Slovenia, and the Republic of Serbia.[1] We also confirm daughter preference for three Scandinavian countries, i.e. Denmark, Norway, and Sweden, which had been identified 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 the allocation of household resources. We test between the two explanations by checking which relationships hold for the household-level data.

We find that Balkan households with more female children replace furniture less frequently than households with fewer female children. Moreover, in households with more female children, mothers report a lower ability to spend on themselves. Additionally, for Balkan countries we find no difference in parental investment in male and female children and no impact of the gender composition of children on the ability to make ends meet or the minimum amount of money needed to make ends meet. We argue, based on earlier studies, that these findings are consistent with the gender bias explanation and not with the differential expenses explanation. For Scandinavian countries we find no impact of the gender composition of children

[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/

on replacing furniture or on consumption of other household public goods, and we find significantly larger parental investment in households with more female children. Moreover, we do not find a systematic impact of the gender of their children on parental consumption. We argue based on conclusions in Lundberg (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 Literature review

The evidence on the impact of parental gender preference pertains to developing economies (Barcellos et al., 2014; Jiang et al., 2016; Altindag, 2016) and developed economies (Dahl and Moretti, 2008; Andersson et al., 2006; Pollard and Morgan, 2002; Brockmann, 2001). 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 et al., 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).[2] Consequences of parental gender preference have mostly been researched for developing economies. The main consequence is that girls, on aver-

[2]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.

age, 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 et al. (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[3] could lead to a female "under-class", 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 et al., 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

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

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.[4] 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 et al., 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 considers two cases.[5] 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),

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

[5]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.

which frequently calculate so-called *normative budgets*.[6] 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 et al., 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 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

[6]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).

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 et al., 2014).[7]

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

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

driving the estimates. We expect to find support for gender bias and no impact of differential costs, because cost difference should play a lesser role in European economies compared to developing economies characterized by pronounced son bias (Brockmann, 2001). However, 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.

3 Data and sample statistics

We use a data set from the European Union Survey of Income and Living Conditions (EU-SILC) for 2004 - 2015. The data set is collected annually by national statistical offices in cooperation with Eurostat from nationally representative samples, which covered the EU-28 and several non-EU countries in 2015. In 2004, only 15 countries were covered by the survey. Our analysis is based on data from four Balkan countries and three Scandinavian countries. The Balkan countries are Bulgaria, Croatia, Slovenia, and the Republic of Serbia.[8] The Scandinavian countries are Denmark,

[8]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

Norway, and Sweden.[9] A primary goal of EU-SILC is 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) .[10]

change qualitatively.

[9]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 countries. In a similar vein, Filmer et al. (2009) pool HNS data into six sub-samples by parts of the world and assume no difference in parameters between countries within groups.

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

Table 1: Descriptive statistics - demographics and 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 father	0.114	-0.005	-	-	0.106	0.003	-	-
	(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 at first birth [a]	26.44	0.035	27.06	0.03	28.91	0.04	29.09	0.07
	(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 tertiary education	0.178	-0.005	0.195	-0.007	0.363	0.002	0.402	0.004
	(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 hours of work	28.100	-0.106	28.738	-0.159	27.985	0.341	28.001	0.340
	(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 hours of work	-	-	37.156	0.082	-	-	37.810	-0.165
	-	-	(16.689)	(0.162)	-	-	(12.762)	(0.106)
Household disposable income (euros)	20,469.770	265.421	20,982.732	214.079	64,070.609	325.596	65,957.259	450.271
	(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 accomodation	0.767	-0.003	0.763	-0.004	0.920	-0.005	0.929	-0.004
	(0.423)	(0.004)	(0.425)	(0.004)	(0.271)	(0.002)**	(0.257)	(0.002)**
N of hhds	24,951		22,027		28,352		25,294	

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

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

[a] These statistics were calculated only for families in which the mother is younger than 41 and older than 17 and had her first child at the age of 16 or older and child ages are in the range 0–12.

Two main advantages of this data set are important for our analysis. First, it contains information on the age and gender of all adults and their children living in the household. Second, the ad-hoc modules from 2009, 2010 and 2013-2015 contain detailed information on material deprivation of adults and children in the household. There are also two significant drawbacks. First, not all children might be present in the household at the time of the survey for some reason (e.g., because they study or work elsewhere). We cannot be sure that the firstborn child lives in the household. Second, the information on material deprivation of children is available only for all children in the household together and not for each child separately .[11] 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 .[12] Our calculated sex-ratio for firstborns is 1.057, close to the commonly accepted value of 1.06 (Grech et al., 2002) .[13] 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

[11] 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.

[12] 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 et al. (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

[13] 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

gender of children influences household composition, we limit our analysis, for the most part, to married and cohabiting couples. Table 1 contains descriptive statistics for selected household socio-demographic characteristics separately for all families and for cohabiting couples. Table 2 presents descriptive statistics on variables characterising different aspects of the household material condition. We use variables in Table 2 as dependent variables and variables in Table 1 as covariates. Amongst adult and household material deprivation characteristics, Table 2 also presents the average frequency of the ten home environment items for children along with girl-boy differences. One can readily see that girls are more likely to have books, have an opportunity to invite friends, and to host celebrations. These differences are small, however, and hover around one percent of the standard deviation of the corresponding items. This is less than reported by Xu (2016). The largest differences between all families and intact, i.e. married and cohabiting, families appear to be in food and clothing. Specifically, the girl-boy difference is significant for all families, but disappears for intact families. This could be explained by more limited resources of non-intact families .[14] Otherwise, the intact families do not appear to differ systematically from all families along the considered characteristics. That supports our decision to focus the analysis on intact families.

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

Table 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 to make ends meet	1,486.629	11.577	1,507.179	8.119	4,725.007	44.569	4,823.201	84.074
	(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 furniture	0.278	-0.008	0.290	-0.005	0.888	-0.006	0.905	-0.004
	(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 on oneself (men)	0.540	0.003	0.573	0.005	0.383	-0.013	0.408	-0.014
	(0.498)	(0.007)	(0.495)	(0.007)	(0.486)	(0.007)*	(0.492)	(0.007)**

Table 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
Availability of two pairs of properly fitting shoes (women)	0.615	-0.003	0.627	-0.001	0.437	0.017	0.411	0.015
	(0.487)	(0.007)	(0.484)	(0.007)	(0.496)	(0.007)**	(0.492)	(0.007)**
Availability of two pairs of properly fitting shoes (men)	0.597	-0.000	0.634	0.002	0.408	-0.012	0.435	-0.012
	(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								

Table 2 (continued)

Dependent variables	Balkan countries				Scandinavian countries			
	All families		Married couples		All families		Married couples	
	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference
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)
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)

Table 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
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)
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)**

Table 2 (continued)

Dependent variables	Balkan countries				Scandinavian countries			
	All families		Married couples		All families		Married couples	
	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference
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.

4 Empirical Analysis

Our paper tests between two alternative explanations for parental gender preference. Each has different implications for household economic behavior. The gender bias hypothesis implies that households with a first-born child of the desired gender save less (Barcellos et al., 2014) [15] 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., of the more expensive gender, 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 ex-

[15] 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.

plores the relationship between household financial assets market participation and the gender of children. Specifically, Bogan estimates a regression in which the dependent variable is stock or bond ownership while the explanatory variables are dummies for only female and only male children or a proportion of female children in the household. However, since the explanatory variable in both specifications (the dummies for same-gender children and share of daughters) might be decided by households and, thus, may be endogenous, therefore, such estimates cannot be taken as evidence of a causal relationship between the variables in question. [16]. 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) .[17]

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

[16]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.

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

daughter should be less capable of dealing with unexpected financial expenditures (because they save less), spend less on themselves, be more likely to replace worn-out furniture, be more able to make ends meet, and need less money to make ends meet. The same predictions should hold for parents of first-born sons in son-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 et al., 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:

$$y_i = \beta(\textit{First child girl})_i + \boldsymbol{\alpha}\boldsymbol{X}_i + \epsilon_i \tag{1}$$

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, ϵ_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 et al., 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 (Wooldridge, 2002, p.456). Nevertheless, we also run Probit estimations to check for consistency with the OLS-based results .[18] Since we expect observations not to

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

be iid, but correlated within countries, we cluster the standard errors at the country level.

In regard to testing for differential costs of sons and daughters, we assume that the cost of children consists of two components, constant (one-time cost) and the variable (outlays on human capital). Researchers commonly use this assumption in models featuring parental investment in children. The fixed component of child cost primarily represents the time cost of rearing children during infancy, whereas the variable component represents parental expenditures on child human capital. Thus, if our analysis finds that parental outlays on children of one gender are larger, there could be two causes: larger one-time costs or lower price of human capital (parental discounted utility derived from child human capital). The mechanism behind the second cause is that of substitution of quality for quantity of children. For example, parents may spend more on daughter "quality" and have fewer children after daughters. If this explanation is true, daughters in daughter-preferring countries should receive more parental investments. One measure of parental investments used in the literature [19] is the availability of conditions and items at home which are necessary for normal child development (Cunha et al., 2010; Todd and Wolpin, 2007; Juhn et al., 2015) .[20] The expected effects of the first-born daughter are systematically presented in Table A7. We use the 2009/2010 and 2013-2015 EU-SILC data on availability of such items in households to test if daughters tend to have better material conditions in daughter-preferring countries and sons, respectively, in son-preferring countries. Under this assumption, parents having a child of the more expensive gender, in addition to having a lower progression ratio, should also have lower expenditures on private consumption and household public goods, be less able to deal with unexpected financial expenditures, be less able to make ends meet, and need more money to make ends meet. The ability to make ends meet is measured by

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

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

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). [21]

5 Results

5.1 Estimates of parental gender preferences for children

Table 3 presents coefficients on the gender of the firstborn for different specifications of the dependent variable in Equation 1 estimated on data from Balkan countries. These results resemble those obtained by Dahl 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. In line with the expectations discussed above, the impact of the gender of the firstborn on progression to parity two in column (2) is much less statistically significant and lower than the impact on progression to parity three and has much lower percent effect. The numbers in column (3) show the probability

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

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 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 et al. (2009) in Central Asia, South Asia, Middle East, and North Africa. It is this result which is most commonly interpreted in the literature as a manifestation of son preference.

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

	Breakdown by number of children				
	(1) Total number of children	(2) Two or more children	(3) Three or more children	(4) Four or more children	(5) Five or more children
First-born child being a girl	0.030	-0.001	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 children's ages are in the range 0–12. The estimation method used is weighted OLS with probability weights reflecting non-random sampling within and between countries. The table presents estimated effects of the firstborn being a daughter compared with the baseline case of the firstborn being a son. The effect is a ratio of the estimated OLS coefficient on the firstborn's gender dummy to the baseline value of the dependent variable. The dependent variables are the total number of children and a set of binary indicators for specific numbers of children. The control variables, besides the gender of the firstborn, are: the dummy for a first-born daughter, gender of the first two children, cubic polynomial of mother's age, squared polynomial of mother's age at first birth, length of cohabitation of spouses, mother's education, father's education, mother's employment, father's employment, household disposable income, living in urban area, tenure status, year and country dummies.

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

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

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

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

Table 4 presents estimates analogous to those in Table 3, but for Scandinavian countries. These results are notably different from the results for Balkan countries. First, the impact of a first-born daughter on progression to parity three in column (3) is negative and statistically significant. Despite having a similar absolute value, the effect is half of the Balkan effect, because a larger share of Scandinavian families progresses to parity three. Second, impacts of a first-born daughter on the total number of children and on progression to other parities have small absolute magnitudes and are not statistically significant. The parity three progression results in column (3) are in line with those obtained by Andersson 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 A6 and Table A3, 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 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.

5.2 Testing between the gender bias and differential cost explanations

The gender bias explanation implies two patterns in household-level allocations. [22] First, expenditures on household public goods should be higher when the firstborn is of the preferred gender (Lundberg, 2005). Specifically, if a son increases marital surplus more than a daughter, then the birth of a son reduces the probability of divorce and increases the incentive of partners to invest further in the marriage, i.e. the family as a whole (Lundberg and Rose, 2003b). Second, saving should be less, because parents anticipate fewer births in the future (Barcellos et al., 2014). To test the first implication, we estimate the impact of a first-born daughter on the frequency of replacing furniture in the household. Lundberg and Rose (2003b) consider furniture an important household public good along with automobiles and housing conditions as proxies for housing expenditures. Spending on automobiles and housing, however, can be directly influenced by child gender composition. As Lundberg and Rose (2003b) note, observed differences in housing spending could be influenced the need for grater space to accommodate the size and activity of sons or the desire for a higher quality neighborhood to reduce the probability of risky behavior by boys or probability of crimes against girls. Concerning automobiles, having

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

one might make more sense when a couple has sons, who are possibly expected to be more handy with cars and, for whose socialization, access to an automobile can often be more important. Meanwhile, expenditures on furniture do not appear to be directly influenced by the gender of children. Column (1) of Table 5 contains estimates of the firstborn's gender impact on replacement of worn-out furniture in the household. The negative and statistically significant estimate for Balkan countries confirms the prediction from the son bias explanation of the observed gender preference. To support the daughter bias explanation for Scandinavian countries, the estimate would need to be positive, which is not the case. Regarding the prediction that savings should be less in families with a firstborn of the preferred gender, we test that by estimating the impact of the firstborn's gender on the ability to deal with unexpected expenditures, assuming that households with higher savings are more likely to respond positively to this question, the estimate should be positive in Balkan countries and negative in Scandinavian countries. The estimates obtained in column (2), however, are small in magnitude and not statistically significant. For Balkan countries, this result could be reconciled with son preference by the fact that common savings are also a household public good and respond positively to the arrival of a child of the preferred gender, countering the negative effect of reduction in expected number of children.

Higher expenditure on household public goods may also be consistent with the comparative advantage a father has in raising sons, i.e. the so called "technology" explanation, according to Dahl and Moretti (2008). The gender bias and technology explanations have different implications for consumption patterns of fathers and mothers. The gender bias explanation suggests lower consumption of mothers of daughters while the technology explanation implies it to be higher. Specifically, if sons directly increase the utility of fathers, then a standard bargaining model of the household predicts a shift of household resources from fathers to mothers. This

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

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

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

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

Table 6: Impact of a first-born girl on employment 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 5.

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

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

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

Notes: The standard errors of estimates on the sub-sample for Scandinavian countries are clustered at the country level. For details on sampling and estimation see the note under Table 5.
* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

redistribution could be observable through lower consumption of private commodities by mothers of daughters. The negative impact of the mother's ability to spend on herself in Balkan countries in column (3) of Table 6 is in line with the gender bias explanation. In addition, two more facts hold for intrahousehold allocations in Balkan countries. First, mothers of daughters are less likely to be employed. Second, fathers of daughters report more time spent on leisure. The first could be explained by self-selection into unemployment of mothers whose comparative advantage in raising daughters results in an even greater opportunity cost than for similar mothers of sons (otherwise, first-born daughters would also negatively impact the intensive margin of mother's employment). Still, such self-selection of mothers into employment would not undermine our results, because the "technology" explanation implies lower progression to parity three when fathers have a sufficiently high comparative advantage in raising sons and a sufficiently wide wage gap in favor of men (Gugl and Welling, 2012). Despite the existence of a wide gender wage gap, our estimates do not support the existence of a sizable comparative advantage of fathers in raising sons in the Balkans, which would be evident from fewer hours of work and higher personal consumption reported by fathers with first-born sons, as explained earlier. Finally, the fact that fathers have more leisure could be explained by longer hours of housework done by daughters .[23] Thus, the obtained results are consistent with the gender bias explanation for Balkan countries.

For Scandinavian countries, there is no firstborn gender effect on either furniture replacement nor the ability to deal with unexpected expenditures (the first two columns of Table 5). Moreover, the estimates of the firstborn's gender impact on parental consumption in Table 7 do not differ between fathers and mothers, similar to the unconditional means in Table 3, which would be in line with parental comparative advantage. In other words, the parental consumption difference between fathers and mothers, as the difference between fathers and mothers of the unconditional means in Table 3, both point to the parental comparative advantage explanation.[24] That is because mothers of sons should redirect household resources to fathers to keep them in the family due to their important role in raising sons (Lundberg, 2005). At the same time, estimates of the impacts on ability of mothers to meet with friends and family and to have regular leisure activity do not contradict the gender bias explanation *per se*. However, the estimated impacts on father's consumption should be positive according to the gender bias explanation and it is not. Fathers of daughters work fewer hours but they do not redirect that time to leisure. Moreover, the fewer hours worked by fathers of daughters is not likely to drive the observed daughter preference because similar effects were found in the US and West Germany (Lundberg and Rose, 2002; Choi et al., 2008), 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.

[23] 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.

[24] It cannot be the main driving cause for the observed gender preference in Scandinavia, because the gender wage gap should be in favor of women (Gugl and Welling, 2012) and that is not the case. Still, this result is consistent with a comparative advantage of intact families with daughters in producing "child quality".

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

The Scandinavian results do show the expected higher outlays on daughters consistent with the differential cost explanation. Households with first-born daughters are more likely to have the entire set of ten important children consumption items. However, neither ability to make ends meet nor the minimum amount of money to make ends meet depend on the gender of the firstborn. Nevertheless, for the top income decile, the minimum amount of money needed to make ends meet is larger for families with a first-born daughter.[25] Mothers of daughters appear to more frequently forgo regular leisure activity and substitute it with apparently less costly socialization through meeting with friends and family. Moreover, more hours worked by mothers of daughters suggest that they are willing to substitute leisure for outlays on daughters. At the same time, fathers of daughters tend to work less than fathers of sons. When Lundberg and Rose (2002) reported a similar effect for fathers from the US, they offered an explanation based on the son bias idea but did not formally test it. Our testing, however, does not support the son bias explanation. Furthermore, Norwegian data indicates that paternal leave has more pronounced positive effects for daughters than sons (Cools et al., 2015) . That could be a reason fathers in Scandinavian countries substitute time spent on work for time spent on children (rather than leisure).[26] All in all, the differential expenses explanation of

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

[26]Examining data from detailed time-use surveys could shed more light on this issue.

daughter preference in Scandinavian countries is supported by the data.

Figure A2 and Figure A3 show cross-country relationships between the gender preference, the gender gap in parental investment, and conventional measures of gender equality. These relationships are in line with our previous points. [27]

6 Conclusion

We find evidence that parental gender preferences in different countries are caused by different reasons. In Balkan countries, the observed son preference is likely driven by gender bias towards sons, that plays the major role. In Scandinavian countries, the observed daughter-preference is likely driven by a lower cost of daughter quality (which incorporates gender-specific personal characteristics and their usefulness for parents). To measure the effect of the gender difference in the cost of children precisely we would need to observe its random variation. Evidence of a lower price for female human capital is most pronounced in more gender-equal societies in line with trends of institutional change in modern societies in favor of women (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.

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

References

(2010). Gender equality in the EU in 2009. Special Eurobarometer 326, European Commission.

Aaronson, D., Lange, F., and Mazumder, B. (2014). Fertility Transitions Along the Extensive and Intensive Margins. *American Economic Review*, 104(11):3701–3724.

Adda, J., Dustmann, C., and Stevens, K. (2016). The Career Costs of Children. Working Paper 6158, CESifo.

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

Ananat, E. O. and Michaels, G. (2008). The effect of marital breakup on the income distribution of women with children. *Journal of Human Resources*, 43(3):611 – 629.

Andersson, G., Hank, K., Ronsen, M., and Vikat, A. (2006). Gender Family Composition: Sex Preferences for Children and Childbearing Behavior in the Nordic Countries. *Demography*, 43(2):255–267.

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

Arnold, F. (1992). Sex Preference and Its Demographic and Health Implications. *International Family Planning Perspectives*, 18(3).

Arnold, F., Choe, M. K., and Roy, T. K. (1998). Son preference, the family-building process and child mortality in india. *Population Studies*, 52(3):301 – 315.

Autor, D., Dorn, D., and Hanson, G. (2017). When work disappears: Manufacturing decline and the falling marriage-market value of young men. Working Paper 21173, National Bureau of Economic Research.

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

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

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

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

Ben-Porath, Y. and Welch, F. (1976). Do sex preferences really matter? *Quarterly Journal of Economics*, 90(2):285 – 307.

Blacklow, P. (2002). Intertemporal Equivalence Scales: Measuring the Life-Cycle Costs of Children. Memorandum 8, University of Oslo Department of Economics.

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

Bojer, H. (2002). The Time Cost of Children and Equivalent Full Incomes. Memorandum 8, University of Oslo Department of Economics.

Bond, T. N. and Lang, K. (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, B. (2004). The Price, Cost, Consumption and Value of Children. Supporting children: English-speaking countries in international context, workshop presentation paper, Princeton University.

Bradley, R. H. and Caldwell, B. M. (1980). The Relation of Home Environment, Cognitive Competence, and IQ among Males and Females. *Child Development*, 51:1140–1148.

Bradley, R. H. and Caldwell, B. M. (1981). The HOME Inventory: A Validation of the Preschool Scale for Black Children. *Child Development*, 52:708–710.

Bradley, R. H. and Caldwell, B. M. (1984). The Relation of Infants' Home Environments to Achievement Test Performance in First Grade: A Follow-up Study. *Child Development*, 55:803–809.

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

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

Cameron, E. Z. and Dalerum, F. (2009). A Trivers–Willard effect in contemporary humans: male-biased sex-ratios among billionaires. *PLoS ONE*, 4(1).

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

Choi, H.-J., Joesch, J. M., and Lundberg, S. (2008). Sons, daughters, wives, and the labor market outcomes of West German men. *Labour Economics*, 15:795–811.

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

Cools, A. and Patacchini, E. (2017). Sibling Gender Composition and Women's Wages. Discussion Paper 11001, IZA.

Cools, S., Fiva, J. H., and Kirkeboen, L. J. (2015). Causal effects of paternity leave on children and parents. *Scandinavian Journal of Economics*, 117(3):801 – 828.

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

Dahl, G. and Moretti, E. (2008). The Demand for Sons. *The Review of Economic Studies*, 75:1085–1120.

Davis, D. L., Gottlieb, M. B., and Stampnitzky, J. R. (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.

de la Croix, D. and Doepke, M. (2003). Inequality and Growth: Why Differential Fertility Matters. *American Economic Review*, pages 1091–1113.

Deaton, A. and Muellbauer, J. (1986). On measuring child costs: with applications to poor countries. *Journal of Political Economy*, 94:720–745.

Dujardin, C. and Goffette-Nagot, F. (2009). Does public housing occupancy increase unemployment? *Journal of Economic Geography*, 9:823–851.

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

Edlund, L. (1999). Son Preference, Sex Ratios, and Marriage Patterns. *Journal of Political Economy*, 107(6).

Estrin, S. and Uvalic, M. (2014). FDI into transition economies. Are the Balkans different? *Economics of Transition*, pages 281–312.

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

Eurostat (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].

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

Galor, O. (2011). *Unified growth theory*. Princeton : Princeton University Press.

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

1011.

Gugl, E. and Welling, L. (2012). Time with sons and daughters. *Review of Economics of the Household*, 10:277–298.

Guilmoto, C. Z. and Duthe, G. (2013). Masculinization of births in Eastern Europe. Population & Society: Monthly bulletin 506, The French National Institute for Demographic Studies.

Guiso, L., Monte, F., Sapienza, P., and Zingales, L. (2008). Culture, Gender, and Math. *Science*, 320:1164–1165.

Gupta, B. (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, K. (2007). Parental Gender Preferences and Reproductive Behavior: A Review of the Recent Literature. *Journal of Biosocial Sciences*, 39:759–767.

Hank, K. and Kohler, H.-P. (2000). Gender Preferences for Children in Europe: Empirical Results from 17 FFS Countries. *Demographic Research*, 2(1).

Hao, L. and Yeung, W.-J. J. (2015). Parental Spending on School-Age Children: Structural Stratification and Parental Expectation. *Demography*, (52):835–860.

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

Hazan, M. and Zoabi, H. (2015). Sons or Daughters? Sex Preferences and the Reversal of the Gender Educational Gap. *Journal of Demographic Economics*, 81:179–201.

Husain, M. and Millimet, D. L. (2009). The mythical 'boy crisis'? *Economics of Education Review*, 28(1):38 – 48.

Ichino, A., Lindstr'́om, A.-A., and Viviano, E. (2011). Hidden Consequences of a First-Born Boy for Mothers. Discussion Papers 5649, IZA.

Jacobsen, R., Møller, H., and Engholm, G. (1999). Fertility rates in denmark in relation to the sexes of preceding children in the family. *Human Reproduction*, 14(4):1127–1130.

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

Jayachandran, S. and Kuziemko, I. (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, R. T. (2003). Equal Treatment, Unequal Outcomes? Generating Sex Inequality Through Fertility Behavior. Working paper, Harvard University John F. Kennedy School of Government.

Jiang, Q., Li, Y., and Sanchez-Barricarte, J. J. (2016). Fertility Intention, Son Preference, and Second Childbirth: Survey Findings from Shaanxi Province of China. *Social Indicators Research*, 125:935–953.

Juhn, C., Rubinstein, Y., and Zuppann, C. A. (2015). The quantity-quality tradeoff and the formation of cognitive and non-cognitive skills. Working Paper 21824, National Bureau of Economic Research.

Kanazawa, S. (2007). Beautiful parents have more daughters: A further implication of the generalized trivers–willard hypothesis (gtwh). *Journal of Theoretical Biology*, 244(1):133 – 140.

Karbownik, K. and Myck, M. (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.

Kippen, R., Evans, A., and Gray, E. (2006). Parental Preference for Sons and Daughters in a Western Industrial Setting: Evidence and Implications. *Journal of Biosocial Science*, 39:583–597.

Kornrich, S. and Furstenberg, F. (2007). Investing in Children: Changes in Parental Spending on Children, 1972 to 2007. Unpublished paper.

Larsen, U., Chung, W., and Das Gupta, M. (1998). Fertility and son preference in Korea. *Population Studies*, 52:317–325.

Leung, S. F. (1991). A Stochastic Dynamic Analysis of Parental Sex Preferences and Fertility. *Quarterly Journal of Economics*, pages 1063–1088.

Li, L. and Wu, X. (2011). Gender of Children, Bargaining Power, and Intrahousehold Resource Allocation in China. *Journal of Human Resources*, pages 295–316.

Lindström, E.-A. (2013). Gender Bias in Parental Leave: Evidence from Sweden. *Journal of Family and Economic Issues*, 34:235–248.

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

Lundberg, S. (2005). Sons, Daughters, and Parental Behavior. *Oxford Review of Economic Policy*, 21(3).

Lundberg, S., McLanahan, S., and Rose, E. (2007). Child Gender and Father Involvement in Fragile Families. *Demography*, 44(1):79–92.

Lundberg, S. and Rose, E. (2002). The Effects of Sons and Daughters on Men's Labor Supply and Wages. *The Review of Economics and Statistics*, 84(2).

Lundberg, S. and Rose, E. (2003a). Child Gender and the Transition to Marriage. *Demography*, 40(2):333–349.

Lundberg, S. and Rose, E. (2003b). Investments in Sons and Daughters: Evidence from the Consumer Expenditure Survey. Unpublished manuscript, Joint Center for Poverty Research.

Mammen, K. (2008). The Effect of Children's Gender on Living Arrangements and Child Support. *American Economic Review: Papers and Proceedings*, 98(2).

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

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

Outram, Q. (2015). Son Preference Aming the Edwardian Middle Class. Unpublished manuscript.

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

Pollard, M. S. and Morgan, S. P. (2002). Emerging parental gender indifference? sex composition of children and the third birth. *American Sociological Review*, 67(4):600 – 613.

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

Raurich, X. and Seegmuller, T. (2017). Growth and bubbles: the interplay between productive investment and the coost of rearing children. Workin Paper 26, AMSE.

Roberts, R. and Baumeister, R. (2011). Baumeister on gender differences and culture.

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

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

Seidl, C. (1995). The Desire for a Son Is the Father of Many Daughters: A Sex Ratio Paradox. *Journal of Population Economics*, 8(2).

Sienaert, A. (2008). Some Child Cost Estimates for South Africa. Working Paper Series 15, CSAE.

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

Trivers, R. L. and Willard, D. E. (1973). Natural selection of parental ability to vary the sex ratio of offspring. *Science*, 179(4068):90–92.

van Praag, B. M. and Warnaar, M. F. (1997). Chapter 6 the cost of children and the use of demographic variables in consumer demand. In Rosenzweig, M. and Stark, O., editors, *Handbook of Pupolation and Family Economics*, volume 1, pages 241 – 273. Elsevier.

Vogl, T. S. (2013). Marriage institutions and sibling competition: Evidence from south asia. *Quarterly Journal of Economics*, 128(3):1017 – 1072.

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

Xu, J. (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, K. and Ferguson, L. R. (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, Y. J. (2006). Gender Imbalance: The Male/Female Sex Ratio Determination. *Journal of Bioeconomics*, 8:253–268.

Appendix

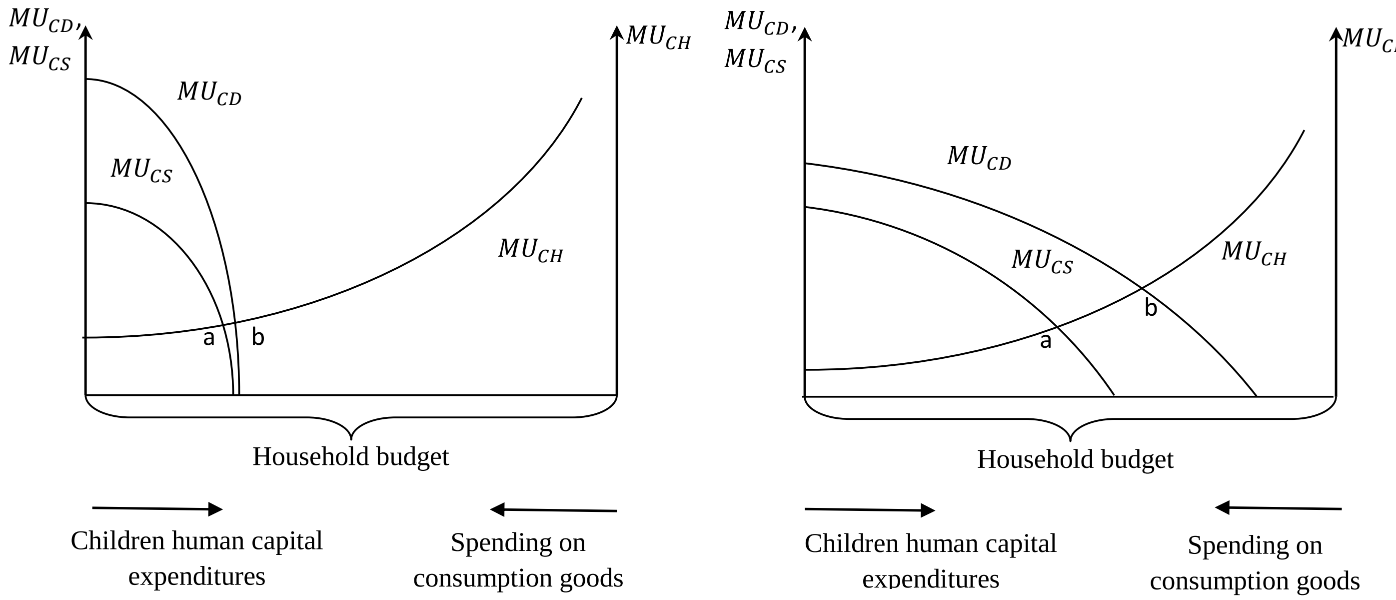
The distinction between the gender bias and differential costs concepts

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 et al. (1998) assert that some Indian parents prefer sons for reasons connected with religious beliefs and kinship descent, whereas Jacobsen et al. (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 et al., 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.

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 et al., 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 coun-



(a) Gender bias

(b) Differential cost

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

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

try fixed-effects, and repeat the analysis after dropping the top 1% of wealthiest households in each country from the sample.[28]

At the same time, the gender of the firstborn might impact marital stability (Lundberg and Rose, 2003a; Mammen, 2008; Lundberg et al., 2007), family size (Hank and Kohler, 2000; Angrist and Evans, 1998), and parental time allocation (Lundberg and Rose, 2002; Lindström, 2013; Choi et al., 2008). This makes "exclusion restrictions a priori unpersuasive" (Lundberg, 2005). To solve this problem,

[28]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 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 not at a statistically significant level). An even larger effect, almost 5% of mean annual male work hours, was found in West Germany (Choi et al., 2008). Meanwhile, Ichino et al. (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. Lindström (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 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.

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 et al., 2010; Todd and Wolpin, 2007; Juhn et al., 2015) and PISA-2000 Xu (2016). Those surveys are more extensive. Instead, the ten questions we consider largely overlap with the resources-spent and time-with-child subcomponents defined by Juhn et al. (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 et al. (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, 1981,

1984) and strongly influencing it (Cunha et al., 2010), the raw score of the EU-SILC ad-hoc modules should also be correlated with and have an impact upon child development. Furthermore, the responses from the PISA-2000 survey analyzed by Xu (2016) contain more detailed information, but correspond directly with the EU-SILC questions on participating in regular leisure activity, availability of a suitable place to study, and having books at home. Xu argues that precisely those items are important for a child's adult outcomes and supports the point by referring to multiple related studies.

To test for a gender-gap in children's material conditions at home, we use five alternative dependent variables in equation 1 for measuring material condition. The first is a pure sum of the binary indicators of the presence of the first ten material conditions listed in the previous paragraph. This sum corresponds to the so-called HOME index used in the literature. One problem with this variable is susceptibility to monotonous transformations, also known as the scaling problem (Bond 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. The probit and the Poisson regressions measure the probabilities of acquiring the next most necessary items. Finally, the frequency histogram of the raw sum of indicators (Figure A1) shows that around one-half of the households possess all ten items. Therefore, we introduce one more binary alternative dependent variable. It takes a value of 1 for households which possess all specified items and a value of 0 for the other households. This specification of the dependent variable is the most intuitively appealing to us and we rely upon it in the main analysis. Nevertheless, under all specifications of the dependent variable the results of the analysis are qualitatively similar and the estimated coefficients of primary interest are statistically significant.

Cross-country comparison of gender preference and parental investment

Table A3 displays the results of estimating gender preference by country. The geographical pattern of the gender preference at birth is depicted in Figure A6. Our results are broadly consistent with those previously obtained in the literature. As did Hank 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.[29] We also attempt to evaluate how our results would differ if there were no family disruptions caused by gender of children, which is frequently reported in the literature (see, e.g., Lundberg, 2005, for a review). The results are presented in Table A3. Son preference becomes statistically significant in Slovenia and stops being statistically significant in Croatia. However, the estimates obtained after including Slovenia and excluding Croatia from son preferring countries do not differ qualitatively and do not differ

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

much quantitatively from those reported here. The rank correlations between the country-level impacts of the firstborn's gender on the selected household fertility outcomes are presented in Table 4. The absence of a strong correlation between estimated impacts on progression to parity two and parity three suggests different factors driving these impacts as we expected above.

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

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

In addition, the fact that parents tend to invest more in daughters as measured by the presence of the home items[30] hold for the pooled EU-SILC sample. To test for the gender gap in parental investment we estimate Equation 1 with several alternative measures of child material conditions on the LHS. We primarily focus on the specification with the binary home indicator (the dummy variable for all 10 items)

[30] 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

on the LHS. Table A8 displays estimates for this specification on a pooled sample. The results suggest that daughters, on average, receive more parental investments in terms of home items. For example, the number in column 1 means that families with first-born girls are 1.5% more likely to have all 10 items. This estimate is robust to the alternative sets of covariates, as can be seen in the rest of Table A8. Still, this effect is not large, remaining between 1,7% and 2% of the standard deviation of the binary home indicator. Results of this scale are typical in the literature on gender effects. Meanwhile, the gender preference pattern established before holds for the sub-sample of households from the highest income decile. These results might suggest that society as a whole is attaching increasingly positive significance to female children, an idea that has appeared in previous studies, such as Brockmann (2001) and Andersson et al. (2006). A daughter may assume both the role of a breadwinner and that of a caregiver.[31] 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 A3 displays the three strongest relationships. Most importantly, there is a strong relationship with the Global Gender Gap (GGG) score, calculated by the World Economic Forum (we used the most recent 2016 data). This index is also strongly related to the gender gap in PISA math achievement (Guiso 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.

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

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.[32] Nevertheless, the latter is commonly used as a measure of parental gender preference.

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

Table A1: Coefficients corrected for selection bias

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

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

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

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

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

*

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

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

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

Countries[a]	(1) Total number of children	(2) Two or more children	(3) Three or more children	(4) Four or more children	(5) Five or more children	Obs
AT	-0.0181	-0.0245*	0.0083	-0.0050	0.0015	6,574
BE	-0.0074	-0.0139	0.0054	0.0007	0.0004	7,694
BG	0.0206	-0.0112	0.0222**	0.0096*	0.0011	3,509
CH	0.0353	0.0364**	0.0013	0.0013	-0.0017	4,461
CY	-0.0422*	-0.0330**	-0.0125	0.0032	0.0002	5,675
CZ	-0.0123	-0.0167*	0.0037	-0.0002	0.0001	10,329
DE	-0.0141	-0.0179*	0.0060	-0.0012	-0.0010	9,790
DK	-0.0183	-0.0023	-0.0178*	0.0012	0.0007	7,889
EE	-0.0147	-0.0091	-0.0032	0.0027	-0.0017	6,594
EL	-0.0040	-0.0075	-0.0065	0.0045	0.0041***	8,147
ES	-0.0292**	-0.0277***	-0.0030	0.0003	0.0008	16,054
FI	-0.0027	-0.0031	0.0070	-0.0000	-0.0011	13,145
FR	0.0209*	0.0102	0.0072	0.0005	0.0029**	14,496
HR	0.0878**	0.0507*	0.026**	0.0127	0.0031	1,742
HU	-0.0082	0.0057	-0.0137**	-0.0027	0.0015	11,281
IE	0.0002	0.0094	0.0030	-0.0074	-0.0007	5,636
IS	-0.0059	0.0009	-0.0022	-0.0028	-0.0014	5,711
IT	0.0091	-0.0032	0.0121***	-0.0004	0.0002	21,486
LT	-0.0352	-0.0096	-0.0090	-0.0098**	-0.0040*	3,742
LU	-0.0068	-0.0069	0.0022	0.0020	-0.0029*	8,084
LV	-0.0172	-0.0204	-0.0020	0.0028	0.0008	5,102
MT	-0.0170	-0.0013	-0.0118	-0.0019	-0.0013	2,872
NL	0.0021	0.0039	-0.0033	-0.0001	0.0001	11,942
NO	-0.0385**	-0.0210*	-0.0191*	0.0006	0.0007	8,108
PL	0.0049	-0.0037	-0.0008	0.0023	0.0035**	18,374
PT	-0.0794	-0.0486***	-0.0216***	-0.0074**	-0.0008	6,044
RO	0.0293	0.0028	0.0218**	0.0075*	-0.0027	4,948
RS	0.0619	0.0378	0.0214	0.0044	-0.0017	1,221
SE	0.0240	0.0112	0.0114*	0.0019	-0.0006	9,228
SI	0.0140	-0.0147	0.0113	0.0093***	0.0036***	10,544
SK	0.0191	-0.0025	0.0093	0.0072*	0.0018	5,802
UK	-0.0155	-0.0104	0.0034	-0.0085*	-0.0012	9,288

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

Notes: See notes for Table 3 for data samples, variable definitions, and included control variables. The columns contain estimated country-level effects of firstborn daughters on the corresponding variables in the column headings.

[a] Table A4 contains names of countries corresponding to the abbreviations.

Table A4: Abbreviations for countries

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

Source: Eurostat

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

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

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

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

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

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

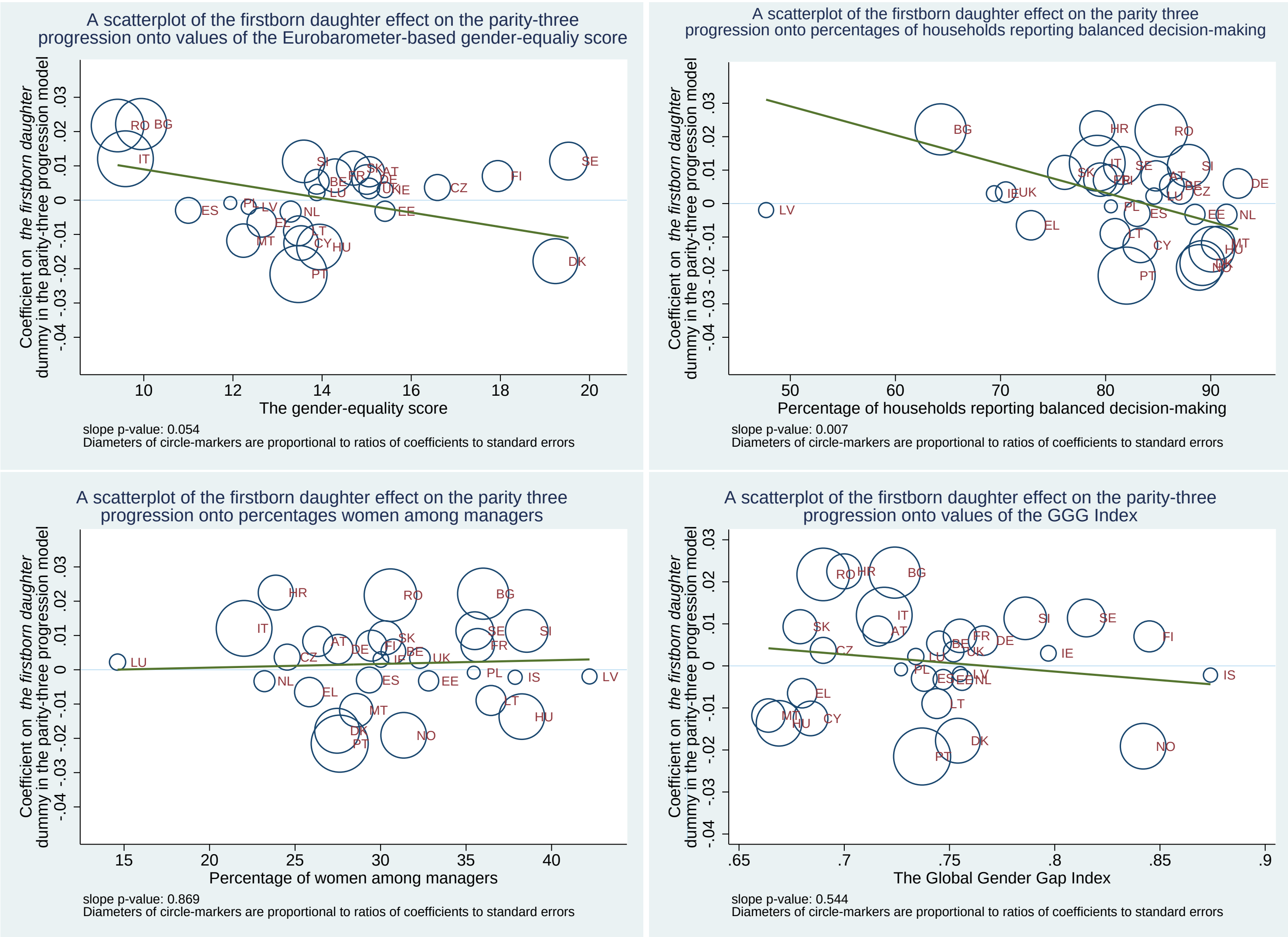


Figure A2: The relationship between the effect of first-born daughters on third parity progression and specific gender-equality measures across countries. We calculate the Eurobarometer-based gender equality score for a particular country as a sum of the country's ranks in responses to questions about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

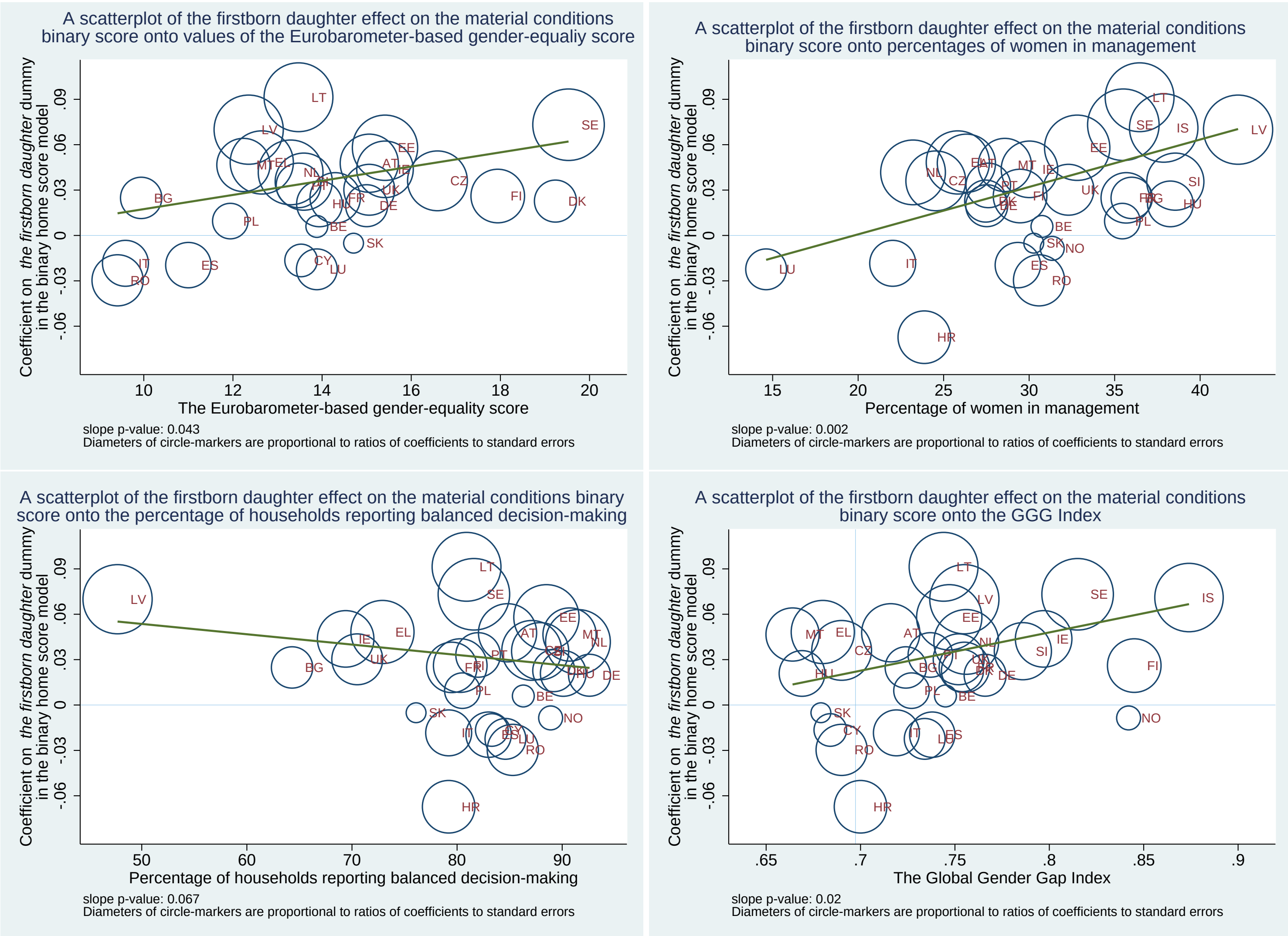
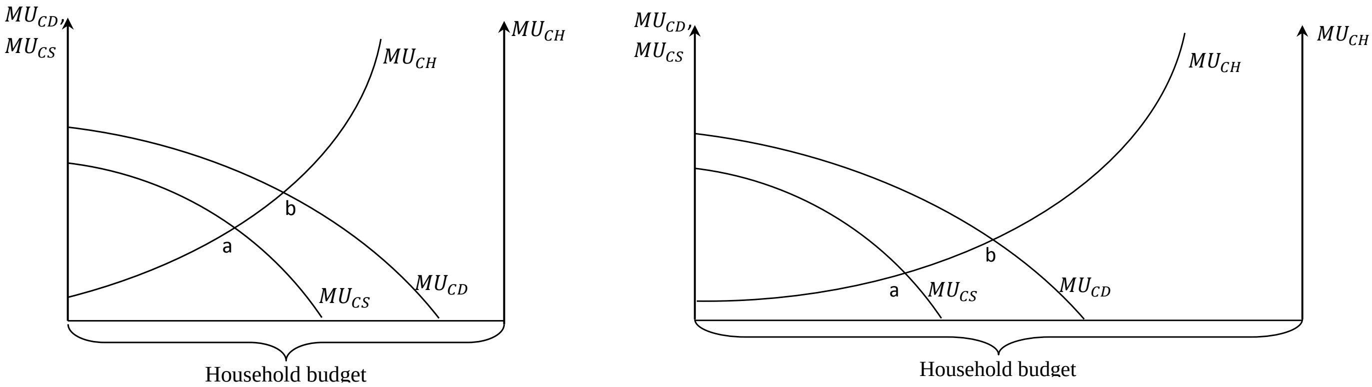


Figure A3: The relationship between the effect of first-born daughters on child material conditions and specific gender-equality measures across countries.



(a) Low-income families (b) High-income families

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

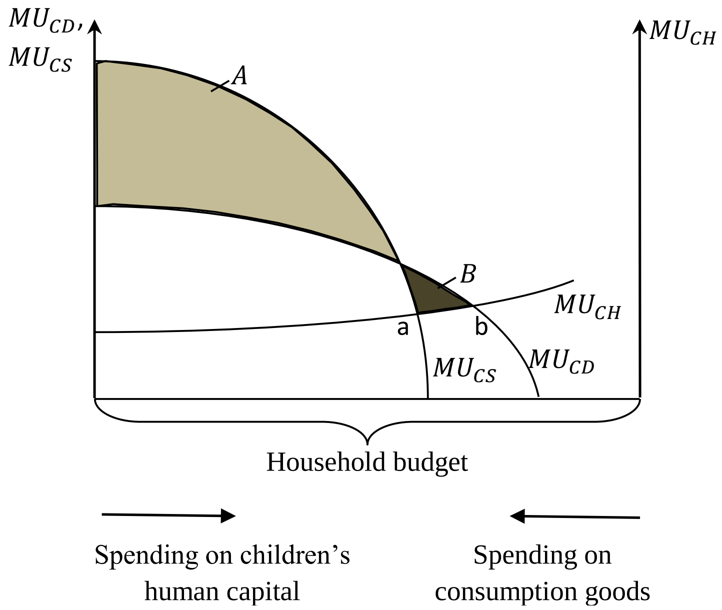


Figure A5: Coexistence of gender (son) bias and differential cost with the gender bias effect on fertility prevailing.
Notes: See the note to Figure A1 for explanation.

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

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

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

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

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

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

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

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

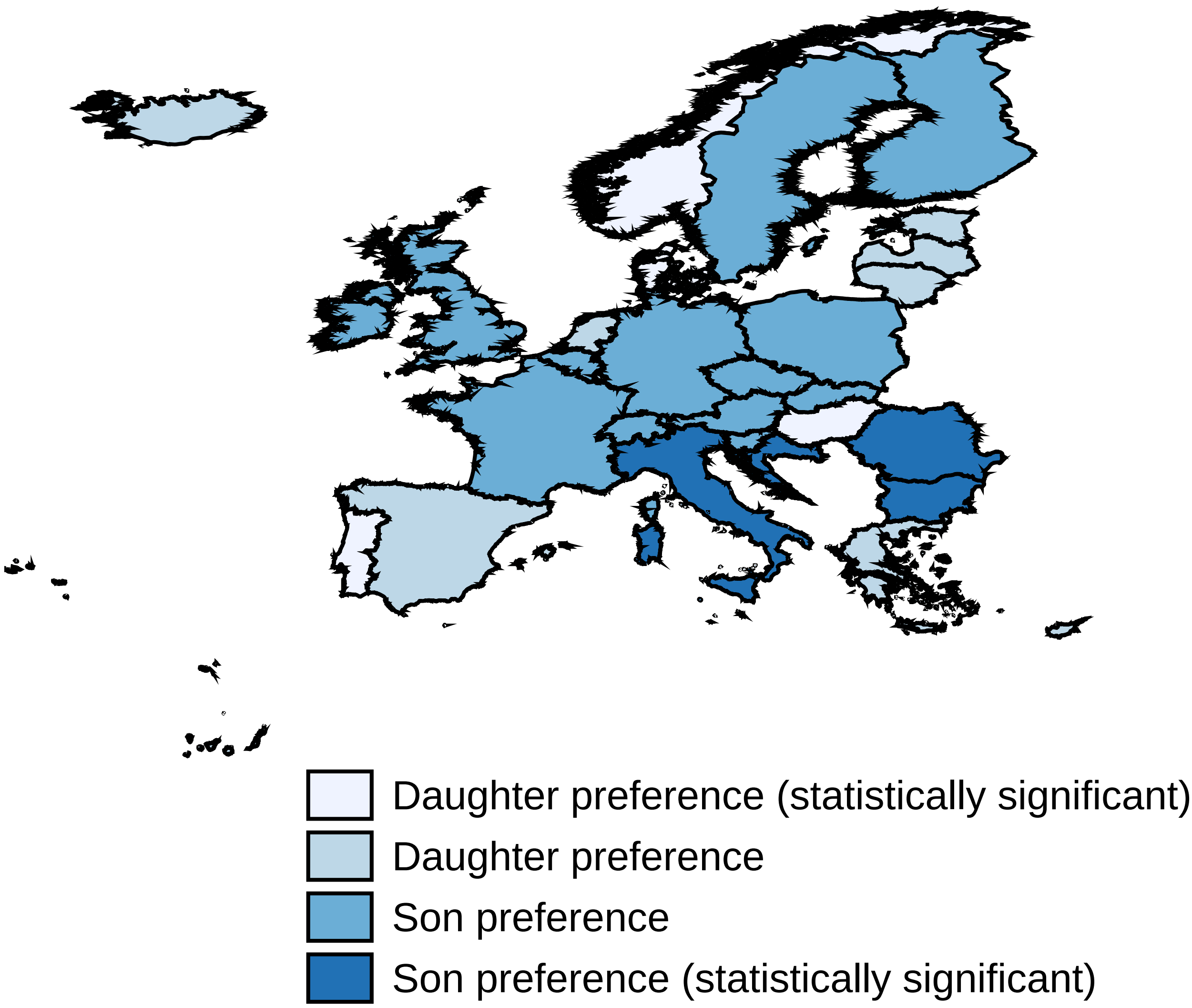


Figure A6: Gender Preferences of Children in 31 EU-SILC countries

Abstrakt

Existuje mnoho studií, které naznačují přítomnost preference rodičů pro konkrétní pohlaví dítěte. Hlavním cílem článku je otestovat dvě hlavní vysvětlení existence těchto preferencí. Konkrétně se jedná o rozdílné náklady na výchovu synů a dcer na straně jedné a upřednostňování jednoho pohlaví na straně druhé (dané rozdílným užitkem rodičů z určitého pohlaví dítěte nebo charakteristik výhradně spojených s jedním z pohlaví). Nejdříve využíváme nedávná EU-SILC data z několika balkánských a skandinávských zemí k potvrzení hypotézy, že pohlaví prvorozeného dítěte předpovídá pravděpodobnost dané rodiny mít tři a více dětí. Ukazatel počtu tří a více dětí je běžná míra genderové preference rodičů. Potvrzujeme preferenci synů v některých balkánských zemích a preferenci dcer ve skandinávských zemích. Prvorozené dítě s preferovaným pohlavím i vyšší náklady na výchovu dětí určitého pohlaví mohou snížit pravděpodobnost rodiny mít tři nebo více dětí. Efekt lze odůvodnit spokojeností rodičů v prvním případě a nedostatkem zdrojů v případě druhém. Dále využíváme informace o spotřebě domácností k rozlišení vlivu uvedených vysvětlení. Tvrdíme, že za předpokladu platnosti hypotézy o rozdílnosti nákladů na výchovu by rodiče dětí s více nákladným pohlavím měli utratit více za zboží pro děti a méně za společné zboží pro domácnost a osobní spotřebu rodičů. Naopak mít děti upřednostňovaného pohlaví by mělo zvýšit výdaje na společné zboží domácnosti, protože takové rodiny mají vyšší přebytečný užitek z manželství a jsou stabilnější. Naše důkazy podporují vysvětlení rozdílnými náklady na výchovu v zemích s preferencí dcer.

Working Paper Series
ISSN 1211-3298
Registration No. (Ministry of Culture): E 19443

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

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

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

Phone: + 420 224 005 153
Email: office@cerge-ei.cz
Web: http://www.cerge-ei.cz

Editor: Byeongju Jeong

The paper is available online at http://www.cerge-ei.cz/publications/working_papers/.

ISBN 978-80-7343-450-2 (Univerzita Karlova, Centrum pro ekonomický výzkum a doktorské studium)
ISBN 978-80-7344-503-4 (Národohospodářský ústav AV ČR, v. v. i.)