

# Hidden consequences of a first-born boy for mothers

New evidence and a critical assessment of the literature \*

Andrea Ichino  
University of Bologna

Elly-Ann Lindström  
IFAU, Uppsala

Eliana Viviano  
Bank of Italy

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## Abstract

We show that in the US, UK, Sweden and Italy women whose first child is a boy are less likely to work in a typical week and work fewer hours than women with first-born girls. The puzzle is why women in these countries react in this way to the sex of their first child, which is chosen randomly by nature. We consider two explanations. As Dahl and Moretti (2008) we show that first-born boys positively affect the probability that a marriage survives, but differently from them and from the literature on developing countries, we show that after a first-born boy the probability that women have more children increases. The negative impact on fertility deriving from the fact that fewer pregnancies are needed to get a boy is more than compensated by the positive effect on fertility deriving from the greater stability of marriages, which Dahl and Moretti (2008) do not consider because they focus on married women only. Second, we show that the gender of the first-born child is likely to affect labor supply also directly, i.e. independently of fertility and marriage. The existence of this direct channel suggests that the sex of the first child may not be a good instrument for fertility and marital stability in equations aimed at identifying the causal effects of these variables on female labor supply.

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# 1 Introduction

Women whose first child is a boy are less likely to work in a typical week and, if they work, they do so for fewer hours than women with first-born girls. We observe this fact in the US, the UK, Italy and Sweden, using representative samples of women aged between 18 and 55 who had their first child between 18 and 40. The estimates are statistically significant and translate into quantitatively relevant labor income losses over the lifetime. The puzzle is why women in these countries react in this way and by so much to the sex of their first child, which to a large extent is chosen randomly by nature.<sup>1</sup>

It is a puzzle because a large body of evidence suggests that, if the first-born child is a girl, parents continue to procreate until a son arrives, while they tend to stop otherwise. For example, Jayachandran and Kuziemko (2009) show that the “desire for a son” of Indian parents induces them to wean their first-born girls faster, because breastfeeding suppresses post-natal fertility preventing the possibility to conceive again in order to have a boy.<sup>2</sup> By contrast, having a first-born boy should give mothers more possibilities to work also because it reduces the length of fertile periods. This is indeed the finding of Chun and Oh (2002) for Korea who use the sex of the first child as an instrument for fertility in a labor supply equation for women, finding that mothers who have an additional child because their first-born child is a girl reduce the probability of labor force participation by 27.5%.

We show that the puzzle has several possible explanations, some related to fertility and marriage and others directly depending on the sex of the first child. The quality of the available data does not allow us to nail down precisely the relative importance of these different explanations. While we are aware of this important shortcoming of our paper, we believe that the value of our contribution consists in the joint analysis of this set of explanations, from the specific viewpoint of understanding the link between the sex of the first child and female labor supply. The literature has so far considered these explanations in a disjoint manner and typically from different viewpoints or for other purposes.

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<sup>1</sup>The gender of the first child is random at conception but not exactly random at birth because the success of pregnancies is correlated slightly with socio-economic characteristics of parents for biological and evolutionary reasons. For example, Catalano et al., (2005a,b) shows that the birth ratio increased after the September 11 attack or in times of high unemployment. Moreover, the sex of a born-alive child (also the first-born) is (weakly) correlated with socio-economic characteristics of the mother, making mothers in good condition more likely to have sons (Cox, 2007, Trivers and Willard, 1973). The definition of “good condition” may vary but for example Almond and Edlund (2007) find that better educated, married and younger women bear more sons.

<sup>2</sup>The biological and behavioral literatures investigate the differences in breastfeeding duration by gender of the child. See, for example, Margulis, Altmann and Ober (1993).

We show that in advanced economies the sex of the first child affects fertility in two opposite ways, of which only one is likely to be at work in less developed economies like the ones considered by the previously cited studies. As in these economies, also in advanced countries a first-born girl induces more fertility *within married couples*, because parents continue to procreate until they get a son (the “desire for a son” effect). On the other hand, a first-born girl reduces substantially the stability of a marriage (the “divorce” effect). This is the second effect that we study and that has been first shown by Bedard and Dechénes (2004) who find that the rate of marital dissolution is 4% higher for women whose first-born child is a girl. Since women in unstable marriages have fewer children over their lifetime, the sex of the first-born has ambiguous effects on fertility in countries where divorces are more likely.

Interestingly, also Dahl and Moretti (2008), in their suggestive collage of evidence that American families prefer boys over girls, show both effects, i.e. that a first-born son increases the expected duration of marriages and reduces fertility *in a sample of married couples*. But they do not consider the full implications of their findings for the effect on fertility of a first-born son *on all women*, independently of their marital status. We use their data and show that in the US, when all mothers are considered, a first-born son increases the number of children over the lifetime of a mother. And the same effect prevails in the UK, Italy and in Sweden. In other words, the negative impact on fertility deriving from the fact that fewer pregnancies are needed to get a boy, is more than compensated by the positive effect deriving from the greater stability of marriages.

Therefore, the two effects of a first-born son act in opposite ways on fertility, but in advanced countries the “divorce” effect dominates the “desire for a son” effect, and, via the prevailing increase in fertility, a first-born boy reduces labour supply.

This combination of effects may explain why we find that in the US, UK, Italy and Sweden a first-born son decreases on average the labor force participation of women via fertility and marital stability. But we further consider the possibility that there might be also direct effects of the first-born sex on female labor supply, even when the risks of new pregnancies or divorce are not an issue.

To explore this possibility we concentrate on two specific subsamples of our populations. First we look at a sample of US widows in which, for arguably exogenous reasons, marriage has been interrupted. Second, we consider only mothers with a first child less than two years old, thus looking at the probability of returning to work in the immediate aftermath of the first child birth. In both these cases, we observe that in the US, the UK and Italy a first-born

boy reduces female labor market participation, even if fertility and divorce are unlikely to be relevant for the observed pattern. More difficult is to nail down the exact nature of these direct effects and why we do not find significant direct effects in Sweden. The evidence that we review suggests three possible direct channels: differences in morbidity of boys and girls, differences in the cost of raising boys and girls and, residually, pure parental differences in the utility of spending time with boys or girls. Of these three channels, we show that the last one gets relatively more credit.

Other papers in this literature (like Lundberg, 2005) are primarily concerned with the effect of children's sex on fertility, on marital stability or on time and resources allocation. Only a few studies are focused explicitly on parental labor market behavior. For example, Lundberg and Rose (2002) and Wulff, Pablonia and Ward-Batts (2007) investigate the labor supply effect of the number of children by sex, but restrict the analysis to married couples and disregard the fact that since fertility and marital stability are affected by the first child's sex, the total number of children by sex is not randomly assigned.

Ananat and Michaels (2008) use the sex of the first child as an instrumental variable to estimate, in quantile regressions, the causal effect of divorces on women's income. Their findings suggest that at the top of the income distribution women who experience a divorce are induced to work more because of the decreased possibility to share income with their husbands and because they typically have the skills and the contacts to do it. At the bottom of the income distribution, instead, women experience income losses that are more disruptive in relative terms. However, since we find that the sex of the first-born child affects mother's labour force participation also directly and independently of marital stability, our result suggests that the sex of the first child may not be a valid instrument for analyzing the effect of divorce. The same problem may exist for other papers like the already cited Chun and Oh (2002).

The rest of the paper is organized as follows. In Section 2 we estimate the overall effect of a first-born son on women's labor supply in the US, UK, Italy and Sweden. In Section 3 we discuss the indirect explanations, i.e. those that link the sex of the first child to mothers' labor supply via fertility and marital stability. The set of direct links are instead analyzed in Section 4. Section 5 concludes.

## 2 First child gender and mothers' labor supply.

The facts that motivate this paper are described in Table 1. In the top panel we look at the probability that a woman is employed (the dependent variable is a dummy equal to one if the person declares to be employed). In the bottom we look at the mean weekly hours worked (except in Sweden where we observe only annual labor income). Non-employed people are included with zero hours to avoid possible bias determined by the fact that women self-select into employment.<sup>3</sup> We focus on all women aged between 18 and 55 at the time of the interview, who had their first child between 18 and 40 years of age and whose first child is no more than 15. We limit the sample at age 15 of the first-born child for two main reasons. First, because children older than 15 may potentially enter the labour market and their decision may affect maternal labour supply, for instance because of income sharing within the household. Second, in most datasets we can identify only mothers with cohabiting children. Since in all the countries considered children tend to leave the household not before 18 and the time spacing between the first and the second child in our sample is around 3 years, the 15 year cut-off minimizes the possibility that we are measuring the sex of the second child instead of the first (just because the first has already left the household). Our estimates, in any case, are robust with respect to this cut-off.<sup>4</sup>

In the first column we use the 1960–2000 waves of the US Census. Consider first the probability of being employed. On average, during the period 1960–2000, 53.6% of the women whose first-born child is a girl are working. Since the sex of the first child is random, women whose first born child is a boy are statistically identical to those who had a girl, but the probability to be employed of the former is 0.4% lower and the difference is statically significant. For the last three available Census waves (1980, 1990 and 2000), we can use hours worked per week at the time of the interview as a measure of labour supply. The bottom panel of the first column shows that if the first born child is a girl, mothers work on average 20.3 hours, but in the case of a boy the working time is reduced by 0.5%. All these estimates are obtained controlling for a quadratic function of age and for interview year dummies.

In column 2 and 3 of the same table we find that these results are confirmed using two

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<sup>3</sup>We have also carried out tobit estimates where the dependent variable is hours worked. They give even stronger results for all the countries.

<sup>4</sup>In particular, we have carried out estimates using a sample of mothers with children aged no more than 12 as in Dahl and Moretti (2008). This lower cut-off allows us to control for the fact that during the sixties in the US a not negligible share of children were used to leave the household around age 16. Our results are fully confirmed and available upon request.

other independent sources of data for the US: the Current Population Survey (CPS, for the years 1990-2008) and the National Health Interview Survey (NHIS, for the years 2005-2008). Actually, using these alternative data sources, with the same sample selection and the same regression specification, the estimated effects of a first-born boy are quantitatively larger and at least equally significant despite the smaller sample sizes. The probability of working decreases by 0.5% in the CPS and by 1.3% in the NHIS, while the correspondent percent effects on hours worked per week are respectively -1.3% and -2.7%.

Columns 4 and 5 of the table further show that these effects are observed not only in the US but also in the UK and in Italy, focusing on similarly defined samples of women and using the same regression specification. In the UK Census of 1991 (column 4), the probability of working in case of a first born girl is 52.0% and decreases by -1.4% in case of a boy. The hours worked per week in case of a girl are 12.4 and decrease by 1.9% in case of a boy. Similar results are reported in the fifth column for Italy. In this case the data come from the Labor Force Survey (LFS, for the years 2004-2008). As expected, Italy is the country where fewer women work (49% if the first born child is a girl) and their employment rate is even lower, by 1.2%, if the first born child is a boy. In terms of hours worked per week the reduction due to a first-born boy is 1.3%. Finally in Sweden, the probability of working is still negatively affected by the sex of the first child (even if the sign is not significant), but the effect of a first-born boy on labour income, used as a proxy for hours worked (as for instance in Gerber and Mitchell, 2009), is again negative and the size is comparable with what found for the other countries.

Therefore, in these countries and datasets mothers whose first-born child is a boy (because of a random choice of nature) work less than those who instead have a first-born girl. Over a lifetime these effects are quantitatively important and translate into large labor income losses. If we calculate this loss using the average hourly pay of a woman aged between 18 and 55 and we assume 35 years of lifetime work, these losses (using 2007 as the base year) amount to roughly 27,000 dollars in the US, to 16,000 euros for Italian women, 25,000 pounds for UK women and 35,000 krona in Sweden (slightly more than 5,000 US dollars).<sup>5</sup> From a different

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<sup>5</sup>These calculations are based on the data on earnings of women published by the US Census Bureau and the National Statistical Office, for US and UK respectively. Data for Italy are drawn from EU-SILC 2007, which report annual earnings. All data refer to gross earnings. In order to calculate the lifetime value of the loss we have used an annual return of 3%. By the use of the MORG-CPS we have also carried out a regression where the dependent variable is labour income and the RHS variable is a dummy equal to 1 if the first child is a boy and 0 otherwise. As before these estimates are obtained using a sample of women who had their first-born child between age 18 and age 40 and whose child is not older than 15 in the reference period. According to these results, having a boy as first child costs 21 dollars per week more than having a girl, at

perspective, our estimates also imply that in the US each year more than 50,000 women aged between 18 and 55 with at least 1 child do not work, simply because their first child is a boy. These figures are equal to 20,000 in Italy and 24,000 in the UK.

We carried out a similar exercise on the Italian Time Use Survey, which allows us to identify the sex of the first-born child. This additional evidence, not reported to save space but available upon request, confirms the results presented in this section. We find that women whose first child is a boy tend to spend more time at home with children, for care and surveillance and, consequently, to work less.

While this collage of evidence and its implications for lifetime earnings and employment rates are undoubtedly solid and hard to dispute, the interesting puzzle is what explains them. We identify two main set of channels for which we find evidence in the available data. One is indirect in the sense that the effect of the first child sex on mothers' labor supply involves intermediate effects on fertility and marital stability. This indirect channel is discussed in Section 3. The second set involves instead direct channels and is discussed in Section 4.

### **3 Fertility and marital stability as indirect channels for the effects of a first-born son.**

A large body of evidence for developing countries<sup>6</sup> suggests that if the first-born child is a girl, parents continue to procreate until a son arrives, while they tend to stop otherwise. A similar evidence is presented by Dahl and Moretti (2008) for the US. Because of this *desire for a son* effect, which implies more pregnancies in the case of a first-born girl, we should see less labor market participation among mothers of first-born girls. This effect should be greater in countries where economic, cultural and institutional factors do not allow pregnant women to do (some) work during pregnancy. But given the evidence described in the previous section, this cannot be the only effect at work in developed economies like the US, the UK, Sweden or Italy, where we see that it is a first-born son that reduces female labor supply.

Indeed in advanced economies the sex of the first child affects fertility in a second way that works in an opposite direction. As shown by Bedard and Dechénes (2004) the rate of marital dissolution is 4% higher for women whose first-born child is a girl. We refer to this second channel as to the *divorce* effect. Since women in unstable marriages have fewer children over

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least for the first 15 years of life of the first child. If we consider a worker with a stable employment working around 48 weeks per year, this loss is around 1,000 dollars per year.

<sup>6</sup>See for example Jayachandran and Kuziemko (2009) and Chun and Oh (2002).

their lifetime, the gender of the first-born child has ambiguous effects on fertility in countries where divorces are more likely. On the one hand, a first-born boy increases the probability of marital stability (the “divorce effect”) and, as marital stability implies more births, it may also increase fertility. On the other hand, having a first-born boy reduces the need of other pregnancies (the *desire for a son* effect).

If both these effects are at work in developed economies we should see different results depending on whether the analysis is restricted to women in married couples or is instead extended to all women independently of their marital status. In Table 2 we compare the estimates of Dahl and Moretti (2008) for the US Census, which are based on a sample of only married women, with estimates obtained using all women in the same dataset. To replicate the sample selection of Dahl and Moretti (2008) we focus on women aged between 18 and 40, who must have had their first child in the same age range and whose first child must be not older than 12. Moreover, as in their specifications, all regressions include a quadratic function of age, educational attainment, race and year dummies. We cannot replicate exactly their estimates as we do not have access to the same US Census sub-samples they have (our data derives from a smaller share of the population), but our results closely resemble what they get.

Panel (a) reports estimates in which the dependent variable is the total number of children. If we consider only married women (column 1) a first-born boy reduces this measure of fertility but the estimated coefficient is statistically not significant, exactly as in Dahl and Moretti (2008). Columns 2, 3 and 4 are based instead on all women and the interesting finding is in the last column, which includes the gender of the first child, marital status and the interaction between these two variables. In this specification we see that in general a first-born boy has a small positive and significant effect on fertility, but if the women is married the effect changes sign while remaining statistically significant. In other words, as in developing countries, a first-born boy reduces fertility among married couples but has the opposite effect on women in general.

The other panels of the table break the evidence by the number of children replicating the results obtained by Dahl and Moretti (2008): as in their paper, we find that for married women the probability of having two or more children increases with a first-born boy but the other probabilities (three or more; four or more) decreases significantly when the first child is male. Based on qualitatively similar evidence on married women, Dahl and Moretti (2008) conclude that having a first-born boy reduces fertility. We suggest the possibility that

their result is affected by selection bias, because the probability that a woman has 3 or more children is strongly correlated with her marital status, as married women have on average more children than unmarried women. At the same time, the probability of being married is in turn influenced by the sex of the first-born child.

We therefore argue that the effect of a first-born boy should be estimated using the sample criteria and the specification of Table 1, which includes all women independently on their marital status. Our results are reported in Table 3. Here we consider all women aged between 18 and 55 who had their first child between 18 and 40 and whose first child is no more than 15 at the time of the interview. The dependent variable is equal to 1 if the woman has at least two children and zero otherwise. All specifications include a quadratic function of the age of the mother and year dummies when several years are pooled.

For all countries and datasets the probability of having more than 1 child increases when the first-born child is a boy. All the estimates are statistically significant and the effect ranges between 0.5% and 2% in the US, 1.1% in the UK, 0.9% in Italy and 0.6% in Sweden. Thus, differently than in those developing countries for which some evidence exists, in the US, UK, Italy and Sweden mothers whose first child is a boy have higher fertility and tend to work less. But we still need to show that this occurs via the channel of greater marital stability induced by a first-born boy. This is a finding of Dahl and Moretti (2008) which we confirm in our datasets. The results are presented in Table 4 which reports regressions in which the dependent variable is equal to 1 if the women is married at the time of the interview, and 0 if never married, separated or divorced (widows are excluded). The sample selection is the same as in Table 1: women aged between 18 and 55 who had their first child between 18 and 40 and whose first child is aged no more than 15. Using the US Census, in column 1, the probability of marriage is 86.4% if the first-born child is a girl and increases by 0.05% in the case of a boy. The percent effects are considerably larger in the CPS (0.2%). Again consistently with the findings of Table 1 and 3, the bigger effect is observed in the NHIS (0.7%). In the UK, Italy and Sweden the results are in the same ball park as the US estimates.<sup>7</sup>

These results are therefore consistent with the hypothesis that in these advanced economies, which differ substantially from developing countries, mothers whose first child is a boy tend to work less because their marriage is more stable and its stability increases fertility.

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<sup>7</sup>We have carried out these estimates also for a sample of women aged at least 42 in order to control also for completed fertility and we obtained very similar results.

## 4 Direct channels.

Is it possible that the sex of the first child affects labor supply also in a direct way, i.e. independently of fertility and marital stability? The question is relevant because the random nature of this variable makes it an attractive instrumental variable for the estimation of the effect of fertility or marital stability on female labor force participation or incomes, as for example in Chun and Oh (2002) or Ananat and Michaels (2008) respectively.<sup>8</sup>

As an attempt to detect the existence of direct channels we restrict the analysis to two specific but informative sub-samples of our populations. First, we select a sample of US widows, that is of women whose marriage is broken by an arguably exogenous shock. Table 5 reports the effect of the sex of the first child in this sample. Like all women in Table 1, widows whose first child is a boy have a lower probability to work. The effect on hours worked is negative as well, although not significant. These results cannot be due to the fact that widows with a first-born son are more likely than those with a first-born girl to share the income of a partner who is more stable precisely because of the sex of the first child. An alternative possible explanation of these effects for widows, is that also for them, like for non-widows, a first-born son reduces labor supply because it induces long run expectations of a more stable marriage and of a future life with many children and not much work. But column 3 of the Table shows that widows with a first-born boy do not have a higher number of children than widows whose first child is a girl. This means that even in the presence of such expectations, the husband's death has occurred early enough to shut down the fertility effect, which is, therefore, practically not a work even if it might have been at work in expectations after the birth of the first child.<sup>9</sup> Finally, the effect that we find for widows could be due to the fact that the estimates in Table 5 are biased because women whose first child is a boy have a higher probability to be married (a necessary condition to be widow), as shown by Dahl and Moretti (2008). So this sample may not be randomly selected. Nevertheless, even keeping in mind this possibility, the negative effect of a first-born son on the labor supply of widows can hardly be explained just in terms of marital stability and fertility, and thus suggests the possibility of other direct channels.

The second exercise that we perform to detect the existence of other direct channels is

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<sup>8</sup>Also Angrist, Lavy and Schlosser (2005) use the sex of the first child as an instrument for fertility in some ethnic groups in Israel, although the bulk of their paper uses instruments based on sibling composition.

<sup>9</sup>Ideally, one would need to know the time of the death of the spouse and test whether a fertility effect of a first-born son is present for women who become widows only late in their life, and disappears for those who lose their husband early in life. Unfortunately this information is not available.

described in Table 6 in which we replicate the regressions presented in Table 1 restricting the sample to mothers whose first child is aged no more than 2. The reason for this restriction is that in such a sample of women, the effects of further fertility should be (at least temporarily) not at work.<sup>10</sup> In the US, the UK and Italy, but not in Sweden, mothers whose newly first-born is a boy have a lower probability of working than mothers who have a newly born girl (as in the total sample of women with children). Also the effect on hours worked is negative but the coefficient is statistically significant only in the UK and Italy.

In columns 6 we present a different but similarly informative test for Italy, again based only on women whose first child is still very young.<sup>11</sup> In this country working women who have a child and who were already employed before or during their pregnancy, must, by law, be on maternity leave 2 months before and 3 months after the birth of their child (with no wage reduction). Moreover, on a fully voluntary basis, they can have 6 additional months of leave, which is paid at a rate that is 30% of the original wage. The Italian Labor Force Survey asks women to report if they are on voluntary maternity leave. We have selected women who have a job before the birth of the first child and have a child aged no more than 1 at the time of the interview. We have then created a dummy variable equal to 1 if the woman is on voluntary leave and 0 otherwise. The results show that if the first born child is a boy the probability that the mother is on voluntary maternity leave is 1.6 percentage points higher than if the first-born child is a girl (a 5% increase of the probability of being on leave). Also in this case, as for widows, unless expectations of a more stable marriage and of greater future fertility induced by a male first child are the reason for such immediate effect, it is reasonable to suspect the existence of direct channels independently of fertility and marital stability.<sup>12</sup>

Further evidence of a direct effect is offered by the Italian Health Survey conducted by Istat in 2004-2005.<sup>13</sup> The survey interviews a representative sample of more than 24,000 households and asks questions on breastfeeding concerning the last child of all women with at least one child. Our sample consists of 1,400 women who had their last child when their

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<sup>10</sup>In all countries the probability that a woman has a second birth just within the first year of life of the first child is lower than 4 percent.

<sup>11</sup>The restriction in this case is that the first child must not be older than 23 months.

<sup>12</sup>Perhaps cultural reasons explain why Johansson (2010) finds instead different results in Sweden. She investigates parental leave during the first 23 months of life of a first-born child and finds that women have higher probability of being in parental leave if the first child is a girl and men have higher probability if the first child is a boy. In Table 1 we have seen that also in Sweden, like in Italy, mothers of first born sons tend work less and receive lower labor incomes. Differently than Italy, though, the effect of a male first child on parental leave of mothers does not seem to be coherent with the effect on labor supply, which might be due to the indirect effect only.

<sup>13</sup>*Condizioni di salute e ricorso ai servizi sanitari.*

age was between 18 and 40. Controlling for a quadratic function of the child's birth-weight, we find that on average women breastfeed their daughters for 31 weeks after birth and sons for 33 weeks. This difference is significant at standard levels. The duration of breastfeeding is negatively related to mothers' labour supply, as shown by Baker and Milligan (2006) for Canada. As already mentioned, Jayachandran and Kuziemko (2009) interpret instead the propensity of Indian mothers to breastfeed their first children for a shorter period if they are girls, as evidence in favor of the "desire for a son" effect that leads them to wean daughters faster in order to have the possibility to become pregnant again in search of a male child. But in the case of Italy the result refers to the last child and thus suggests the possibility of a more general and pervasive differential treatment of boys and girls by mothers that is independent of fertility and marital stability.

The existence of direct channels seems therefore to be supported by a collage of different pieces of evidence. But the question that follows immediately from this observation is what are these possible direct channels through which the sex of the first child can affect labor supply independently of fertility and marital stability. We can think of at least three potential channels: differences in morbidity between boys and girls, differences in the costs of raising boys and girls and pure preferences for spending more time with boys than with girls.

The first of these three possibilities appears, *prima facie* as the most obvious. Medical statistics show that boys have higher incidence of many congenital and non-congenital (including behavioural) diseases than girls at birth and especially during childhood. For instance, according to the US Center for Disease Control and Prevention, which publishes reports on the incidence of various health conditions distinct by sex and age, during 2007 the share of boys with mental health conditions, which includes emotional and behavioral areas of health like relationships and learning, was twice that of girls in the US.<sup>14</sup> Our own direct evidence confirms that differences in child morbidity by gender are sizeable. Using the NHIS and limiting the attention to first-born children aged no more than 1 year old, we find that 21% of boys have some health condition, against 16% of girls.

Child morbidity can affect labor supply of mothers through several channels. On the one hand, mothers might increase labor supply to ensure that their children have access to health insurance coverage, especially in the US where, until recently, there was no public health insurance system. On the other hand, especially when the child is affected by emotional

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<sup>14</sup>See for instance NCHS Data Brief, No. 8, September 2008 available at [www.cdc.gov/nchs/data/databriefs/](http://www.cdc.gov/nchs/data/databriefs/).

and behavioural problems, mothers might prefer to spend more time with their children and reduce labor supply.<sup>15</sup> The net effect is ultimately an empirical question, but one that cannot be answered easily because of potential reverse causality and simultaneity: children whose mothers work more might be more frequently in bad health, while at the same time mothers might adjust their labour supply in response to their children health status. Frijters et al. (2009) try to disentangle these effects using child handedness as an instrumental variable with Australian data. Mothers whose children have developmental problems are shown to be more likely to reduce their labor supply. Incidentally, this paper also shows that boys are more at risk of poor development.

We cannot replicate the strategy of Frijters et al. (2009), but we can exploit the fact that the NHIS contains information on congenital diseases, which are unlikely to depend on the labour supply of mothers.<sup>16</sup> In our sample the share of first-born children with congenital diseases is 8% among boys and 6% among girls (the difference is significant at the 5% level). We use congenital diseases to verify whether they can explain at least part of the direct effect found in our data. The evidence, that we omit to save space but is available from authors, does not support this possibility. When we augment the NHIS regression described in column 3 of Table 6 including a dummy for congenital diseases, the negative effect of having a boy as a first child remains negative and essentially unchanged in size and significance. This means that the higher congenital morbidity of boys does not explain why mothers who have boys as a first child are less likely to participate in the labor market. When we use working hours as the dependent variables results do not change indicating that this conclusion is valid for both the intensive and extensive margin of labor supply.<sup>17</sup>

Equally not convincing is the hypothesis that the direct effect is driven by differences in the cost of raising boys and girls. On the one hand, the higher male morbidity that we have just mentioned might be one reasons of these differential costs, suggesting that boys may be more costly. But other sources of difference point in the opposite direction. One is

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<sup>15</sup>On a similar ground Belfied and Kelly (2010), using the Early Childhood Longitudinal Survey - Birth Cohort, show that the intensity and the duration of breastfeeding, instead of formula-feeding, have positive effects on health status of children. Thus, as long as the duration and the intensity of breastfeeding are negatively related to labour supply, mothers might reduce their labour supply as a consequence of children bad health.

<sup>16</sup>Congenital diseases recorded by the NHIS include: cell anemia, mental retardation, Down syndrome, muscular dystrophy, autism, cystic fibrosis, cerebral palsy, congenital heart disease, troubles in seeing, troubles in hearing.

<sup>17</sup>Also Dahl and Moretti, using the NHIS, show that morbidity does not affect marriage survival of US couples.

education. In US, UK, Italy and Sweden girls have higher probability of attending high school and university. This should imply that girls cost more than boys. Official estimates of the cost of raising boys and girls in Sweden (The Swedish Consumer Agency, 2001) show that the overall costs of raising children are either equal for both sexes, or that boys cost marginally more to raise (which again works in the opposite direction of our findings). Another possible story is that girls, as adults, have less job opportunities than boys and lower salaries. Mothers would then prefer to work more for bequest motives if they have a girl. However, a survey by Taubman (1991) gives no support to this hypothesis, as there seems to be no effect of child gender on bequests. On the basis of this mixed results it seems difficult to conclude that differences by gender in the cost of raising children are a plausible explanation of the direct effect consistently highlighted in Section 4 for so many different countries and datasets.

This leaves us with the possibility that mothers value direct child care and, because they prefer sons to daughters, work less if they have sons. In other words, preferences for sons would not only imply, as already mentioned, that a first-born girl induces greater fertility because parents want a boy. It also implies that time spent with sons enters the utility of mothers with a greater weight than time spent with daughters. This is essentially the conclusion of Dahl and Moretti (2008) who cite the Gallup poll on gender preferences, which has been conducted six times since 1941 asking the following question: “Suppose you could have only one child. Would you prefer it to be a boy or a girl?”.<sup>18</sup> Since the first wave of the poll, US respondents have shown a gender preference for boys. In 2007, 37% of the 1,000 adult participants stated to prefer a boy, 28% a girl and the others have no preferences. Differences are statistically significant at standard levels. Moreover, men are more likely than women to prefer boys, while women have more equal preferences. This is shown for example in the polls conducted from 2000 to 2003 in which more than 45% of men answered that they would prefer to have a boy, while only 20% reported they would like to have a girl. Women instead have nearly equal preferences (35 % would prefer a girl, 30% a boy, 35% either). The International Gallup Poll conducted in 1997 confirms the existence of a relatively strong preference for boys in many Asian countries, but also in France and the UK. In the 2007 poll respondents who had a preference for boys were also asked to motivate their choice. Among men, 23% answered that boys can relate to other men (especially the father) better, 20% said that they desire a boy to carry on the family name, only 2% say that girls are more expensive to raise. According to this evidence, mothers would spend more time with boys because fathers value direct child

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<sup>18</sup>See e.g. [www.gallup.com](http://www.gallup.com).

care and mothers have lower market salaries and lower bargaining power within the couple.

In the light of the evidence provided by these Gallup polls, it is not clear how to interpret the apparently contrasting evidence of Baccara et al. (2010) who show that potential adopting parents are significantly more attracted by girls than by boys (11.5% probability against 7.9%). In any case, all these pieces of evidence support the hypothesis that the utility of at least some parents directly depends on the gender of children.

To conclude there is no clear cut evidence concerning the exact nature and relative importance of the direct channels linking the sex of the first-born to the labor supply behaviour of mothers. But it seems hard to dispute that some direct channel exist, and this raises an identification problem in those papers like Chun and Oh (2002), or Ananat and Michaels (2008) that use the sex of the first child as an instrumental variable to estimate the causal effect of fertility or of marital stability on the labor supply or incomes of mothers.

## 5 Conclusions

We have shown that in the US, the UK, Italy and Sweden women whose first child is a boy are less likely to work in a typical week and they do so for fewer hours than women with first-born girls. The estimates are statistically significant and translate into quantitatively relevant labor income losses over the lifetime. These effects of the first child's sex are the combined result of at least three important sets of channels. To begin with, a first-born son reduces fertility because fewer pregnancies are needed to have a son (the *desire for a son* effect). Because of lower fertility, mothers of first-born sons should work less, and this is typically the evidence found in developing countries. But the sex of the first child affects fertility also in an opposite way by making the marriage more stable in case of a first-born boy (the *divorce* effect). We show that in advanced economies this effect dominates and fertility decreases when the first child is a male. In relation to this result, a contribution of our paper is to highlight the importance of using data on *all* women not only *married* women to study the effects of the first child gender on labor market outcomes.

Finally, there is also evidence of third and *direct* negative set of effects of a first-born boy on mothers' labor supply. We uncover these effects by focusing on subsamples of our populations in which fertility and a marital stability are unlikely to explain the observed labor market outcomes (i.e. women with one child less than two years old and widows). The presence of these direct effects suggest that the sex of the first child may not be a valid instrument for

fertility and marital stability, as instead has been done by some studies in equations aimed at estimating the causal effect of family size and/or divorce on different measures of women's well being.

Our study considers, instead, the overall "Intention To Treat" effect of this instrument on the lifetime labor supply of mothers. While the intermediate causal effects through which the sex of the first child displays its consequences are certainly interesting and important per se for a positive analysis, estimates of the overall "Intention To Treat" effect are less demanding in terms of identification assumption and not less relevant in order to evaluate whether it would be desirable to neutralize the overall undesirable consequences of this random event chosen by nature, independently of the intermediate channels through which it displays its effects. In order to understand whether the overall effect of a first-born son corresponds to social preferences and should be accepted as such or instead neutralized, its size must be first estimated precisely. Providing this knowledge for an informed decision has been the goal of our study.

## 6 Data Appendix

Census data for US refer to years 1960–2000 are collected within the IPUMS International project and are available at <https://international.ipums.org/international/>. They are roughly a 2% random sample of the US population in the Census years and contain both personal and household identifiers and socio-demographic characteristics. Not all information is available for all years. For instance, the number of marriages is available only for years 1960 and 1980; the total number of children of women is available only for year 1990.

Census data for UK refer to years 1991 and are collected within the IPUMS International project and are available at <https://international.ipums.org/international/>. They are roughly a 1% random sample of the UK population in the Census years and contain both personal and household identifiers and socio-demographic characteristics.

Current Population Survey data are drawn from the NBER site <http://www.nber.org/cps/>. The universe is the civilian noninstitutional population of the United States living in housing units and members of the Armed Forces living in civilian housing units on a military base or in a household not on a military base. In this paper we use the March supplement, which includes detailed information not only on labor supply, but also on socio-demographic characteristics of individuals and households. About 57,000 households are currently interviewed, containing approximately 112,000 persons 15 years old and approximately 31,000 children 0-14 years. We use data form 1990 to 2008.

National Health Interview Survey (NHIS) is conducted since 1957 and it is aimed at collecting information on a broad range of health topics. Data are collected yearly through cross-sectional household interview and are available for free at <http://www.cdc.gov/nchs/about/major/nhis/>. The sample over-weights both Black person and Hispanic persons and it covers roughly 35,000 households and 87,000 individuals. Basic demographic information is available for all household members. Instead, information on health conditions is collected only for one adult and one child selected randomly in the household.

Not all the individuals in the sample respond to all sections of the questionnaire and only a sub-sample of children reports their health status. In this paper we focus on the health status of the first-born child and we select the sub-sample in which the first-born child has been interviewed. We also focus only on years 2005-2008 as in these years information on illness is more detailed than previous years and fully comparable.

The Italian labor Force Survey is conducted by the Italian Statistical Office, Istat, and includes around 80,000 households and 200,000 individuals, which are interviewed in different weeks of the year. The sample units are “de facto” households, composed of people living together even if with no formal arrangement. Each year 4 releases are available, in January, March, July and October. Because of a break in the sample design and questionnaire in 2004, data comparability before 2004 is not ensured. The public-use files contain a household identifier and detailed socio-demographic characteristics also for individuals aged less than 16. Public-use files are released by Istat, but they are not free of charge. As the CPS no retrospective information on fertility and marital status is included.

The Swedish data, provided by Statistics Sweden, is a population-wide panel data set (LISA) based on administrative records. Detailed socio-demographic variables are observed on a yearly basis from 1990 and onwards for all individuals at least 16 years old. Moreover, households are identified and all children (regardless of age) are linked to their biological parents by the Multigenerational registry (Flergenerationsregistret). No information on hours worked is available.

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Table 1: First child gender and labor supply: US, UK, Italy and Sweden.

	US		UK		Italy		Sweden	
	Census 1960-2000	CPS 1990-2008	NHIS 2005-2008	Census 1991	LFS 2004-2008	LISA 2004		
<b>(a) - Probability of working</b>								
The first child is a boy	-0.002	*** -0.003	* -0.008	-0.007	*	-0.006	***	-0.001
St. err.	0.000	0.002	0.006	0.005		0.002		0.001
Baseline: The first child is a girl	0.536	*** 0.635	*** 0.647	*** 0.520	***	0.493	***	0.802
St. err.	0.000	0.001	0.003	0.021		0.002		0.001
Percent effect (%)	-0.366	-0.502	-1.269	-1.406		-1.160		-0.153
<b>(b) - Hours worked per week (Annual labour income in Sweden)</b>								
The first child is a boy	-0.092	*** -0.200	*** -0.464	*** -0.232	*	-0.195	**	-5.790
St. err.	0.021	0.071	0.087	0.139		0.078		2.864
Baseline: The first child is a girl	20.294	*** 15.491	*** 17.460	*** 12.358	***	15.272	***	1130.594
St. err.	0.014	0.049	1.122	0.657		0.058		2.834
Percent effect (%)	-0.452	-1.294	-2.657	-1.881		-1.278		-0.512
Sample size	3,422,119	302,925	25,339	45,068		198,801		699,805

Source and notes: Authors' calculations. Women aged between 18 and 55 who had their first child between 18 and 40 years and whose first child is aged no more than 15. In the top panel the dependent variable is a dummy equal to 1 if the person is employed and 0 otherwise, except for Sweden, in which case it is equal to 1 if the person has positive labour income, 0 otherwise. In the bottom panel, for the US Census the dependent variable is equal to the number of hours worked per week in all jobs during the previous year; for the CPS, the NHIS and LFS, it is equal to the number of hours worked in the week preceding the interview. For the UK Census, it is equal to the usual working time in all jobs. For Sweden, it is equal to annual labor income (in hundreds SEK). All models include a quadratic in age of the mother and year dummies. Data on hours worked are available only for the period 1980-2000. Significance levels: \*\*\* = 1%; \*\* = 5%; \* = 10%.

Table 2: Re-assessing the effect of first child gender on fertility. US.

	US Census 1960–2000				
	Married (1)	(2)	All (3)	(4)	
<b>(a) - Total number of children</b>					
First-born boy	-0.0012 0.0012	0.0008 0.0011	0.0000 0.0011	0.0053 0.0027	**
Married			0.2073 *** 0.0015	0.2105 *** 0.0021	***
First-born boy*Marr.				-0.0063 0.0029	**
<b>(b) - Probability of having two or more children</b>					
First-born boy	0.0024 *** 0.0006	0.0034 *** 0.0006	0.0029 *** 0.0006	0.0051 *** 0.0015	***
Married			0.1381 *** 0.0008	0.1394 *** 0.0012	***
First-born boy*Marr.				-0.0026 0.0016	*
<b>(c) - Probability of having three or more children</b>					
First-born boy	-0.0022 *** 0.0005	-0.0015 *** 0.0005	-0.0017 *** 0.0005	0.0004 0.0012	
Married			0.0530 *** 0.0007	0.0543 *** 0.0010	***
First-born boy*Marr.				-0.0025 0.0013	*
<b>(d) - Probability of having four or more children</b>					
First-born boy	-0.0013 *** 0.0003	-0.0011 *** 0.0003	-0.0011 *** 0.0003	-0.0002 0.0007	
Married			0.0116 *** 0.0004	0.0122 *** 0.0005	***
First-born boy*Marr.				-0.0011 0.0007	
Obs.	2,029,913	2,437,284	2,437,284	2,437,284	

Source and notes: Authors' calculations. Women aged between 18 and 40 who had their first child between 18 and 40 years and whose first child is aged no more than 12, as in Dahl and Moretti (2008). To replicate Dahl and Moretti (2008) all models include a quadratic in age, educational attainment, race and year dummies. In panel (a) the dependent variable is the total number of children. In panels (b), (c) and (d) the dependent variables are dummies equal to 1 if the woman has two or more (three or more, four or more) children and 0 otherwise. Significance levels: \*\*\* = 1%; \*\* = 5%; \* = 10%.

Table 3: First child gender and fertility. US, UK, Italy and Sweden.

	US		UK		Italy		Sweden	
	Census 1960-2000	CPS 1990-2008	NHIS 2005-2008	Census 1991	LFS 2004-2008	LISA 2004		
<b>Probability of having more than 1 child</b>								
The first child is a boy	0.003	***	0.005	***	0.012	**	0.007	*
St. err.	0.000	0.002	0.005	0.004	0.005	0.002	0.005	**
Baseline: The first child is a girl	0.641	***	0.625	***	0.614	***	0.656	***
St. err.	0.000	0.001	0.041	0.020	0.041	0.002	0.020	0.001
Percent effect (%)	0.489	0.845	1.993	1.127	0.867	0.549	0.867	0.552
Sample size	3,422,119	302,925	25,339	44,813	198,801	699,805		

Source and notes: Authors' calculations. Women who had their first child between 18 and 40 years and whose first child is aged no more than 15. The dependent variables is dummies equal to 1 if the woman has at least 2 children and 0 otherwise. All models include a quadratic in age of the mother and year dummies. Significance levels: \*\*\* = 1%; \*\* = 5%; \* = 10%.

Table 4: First child gender and marital status of the mother. US, UK, Italy and Sweden.

	US		UK		Italy		Sweden	
	Census 1960-2000	CPS 1990-2008	NHIS 2005-2008	Census 1991	LFS 2004-2008	LISA 2004		
The first child is a boy	0.007	***	0.009	***	0.013	***	0.003	**
St. err.	0.000	0.001	0.005	0.003	0.001	0.001	0.003	0.001
	<b>Probability of being married</b>							
Baseline: The first child is a girl	0.864	***	0.800	***	0.780	***	0.846	***
St. err.	0.000	0.001	0.004	0.017	0.001	0.001	0.017	0.001
Percent effect (%)	0.047	0.181	0.695	0.397	0.127	0.127	0.925	0.670
Sample size	3,392,600	300,535	23,121	44,813	196,445	697,873		

Source and notes: Authors' calculations. Women who had their first child between 18 and 40 years and whose first child is aged no more than 16. The dependent variables are dummies equal to 1 if the woman is married and 0 otherwise. Widows are excluded. All models include a quadratic in age of the mother and year dummies. Significance levels: \*\*\* = 1%, \*\* = 5%, \* = 10%.

Table 5: The direct effect of the first child. US, Women who are widows at the time of the interview.

	US Census 1960-2000		
	Probability of working	Hours worked per week	Probability of having more than 1 child
The first child is a boy	-0.010	-0.145	-0.005
St. err.	0.006	0.243	0.006
Baseline: The first child is a girl	0.635	23.280	0.563
St. err.	0.000	0.014	0.011
Percent effect (%)	-1.563	-0.621	-0.888
	29,519	29,519	29,519

Source and notes: Authors' calculations. Women who had their first child between 18 and 40 years of age. US Census years 1960-2000 for the estimates on the probability of working, years 1980-2000 for hours worked (no data on hours worked in years 1960 and 1970). The dependent variable for the "Probability of working" specification is a dummy equal to 1 if the person is employed and zero otherwise. The number of hours worked per week refer to hours in all jobs in the previous year. Non-employed women are assigned zero hours worked. All models include a quadratic in age of the mother and year dummies. Significance levels: \*\*\* = 1%; \*\* = 5%; \* = 10%.

Table 6: The direct effect of the first child. US, UK and Italy. Women whose first child is aged less than 2.

	US		UK	Italy	Sweden	
	Census 1960-2000	CPS 1990-2008	NHIS 2005-2008	Census 1991	LISA 2004	
First child boy	-0.003	** -0.014	** -0.033	* -0.020	** 0.016	0.000
St. err.	0.001	0.006	0.017	0.013	0.008	0.003
Baseline						
First child girl	0.528	*** 0.559	*** 0.590	*** 0.395	*** 0.283	*** 0.754
St. err.	0.000	0.001	0.012	0.009	0.006	0.004
Percent effect (%)	-0.518	-2.475	-5.569	-5.019	-3.543	0.023
		<b>Hours worked</b>		<b>Hours worked</b>	<b>Prob. work</b>	<b>Prob. maternal leave</b>
First child boy	-0.035	-0.344	-1.041	-0.666	* -0.485	** 0.734
St. err.	0.063	0.240	0.665	0.398	0.222	5.910
Baseline						
First child girl	26.387	*** 18.877	*** 21.246	*** 10.257	*** 12.386	*** 836.382
St. err.	0.014	0.049	0.474	0.290	0.163	9.175
Percent effect (%)	-0.134	-1.820	-4.902	-6.493	-3.912	0.088
Sample size	356,750	24,911	3,537	5,319	22,415	12,159
						88,061

Source and notes: Authors' calculations. Women who had their first child between 18 and 40 years of age and whose first child is aged less than 2. US Census years 1960-2000 for the estimates on the probability of working, years 1980-2000 for hours worked (no data on hours worked in years 1960 and 1970). The dependent variable for the "Probability of working" specification is a dummy equal to 1 if the person is employed and zero otherwise, except for Sweden, in which case it is equal to 1 if the person has positive labour income, 0 otherwise. The number of hours worked per week refer to hours in all jobs in the previous year in the US Census, to the usual working time in the UK Census, to the hours worked the previous week in the Italian LFS. Non-employed women are assigned zero hours worked. The probability of maternal leave is estimated on a sample of employed women. The dependent variable is a dummy equal to 1 if the woman is absent from work during the week preceding the interview because of voluntary maternal leave. For Sweden, labour income in expressed in hundreds SEK. All models include a quadratic in age of the mother and year dummies. Significance levels: \*\*\* = 1%; \*\* = 5%; \* = 10%.