Hidden consequences of a first-born boy for mothers

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In the US, the UK, Italy and Sweden women whose first child is a boy work less than women with first-born girls. As Dahl and Moretti (2008) we show that first-born boys positively affect the probability that a marriage survives. Differently from them 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.
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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. 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. But the real puzzle is why women in these countries react

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in this way and by so much to the sex of their first child, which to a large extent is chosen randomly by nature.

It is a puzzle because a large body of evidence for developing countries 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 (2011) 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.[1] Therefore, having a first-born boy should give mothers more possibilities to work, not only because mothers with a lower number of children have less need to spend time on childcare, but also because during pregnancies mothers typically reduce labor supply. Such a positive relationship between a first-born boy and labor market participation for developing countries has been found, for example, by Chun and Oh (2002) in Korea. They 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 is a girl reduce the probability of labor force participation by 27.5%.

We show that the puzzle is solved if we consider that in advanced economies the sex of the first child affects fertility in opposite ways, of which only one is likely to be at work in less developed economies. 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 found e.g. by Dahl and Moretti, 2008). On the other hand, a first-born girl reduces substantially the stability of a marriage (the "divorce" effect). This second effect has been shown by Bedard and Dechéens (2004) but they do not extend its implications to female labor supply. They find that the rate of marital dissolution is 4% higher for women whose first-born child is a girl. The

[1] The biological and behavioral literatures investigate the differences in breastfeeding duration by gender of the child. See, for example, Margulis, Altmann and Ober (1993).

"divorce" effect has been further studied also by Ananat and Michaels (2008) who use the sex of the first child as an instrumental variable to estimate, in quantile regressions, the causal effect of separations on women's income at different points of the income distribution.

The "divorce effect" has an impact on the labor supply of women that goes through two channels . First, the higher marital stability following the birth of a boy generates less need to work for mothers, because they can expect income support from fathers. This is a rather obvious consequence of the "divorce effect". But in addition, we prove the existence of a second and less obvious channel. Marital stability following a first-born boy generates higher fertility and fertility reduces mothers' labor supply.

As a result of the combination of all these effects, the sex of the first-born child has ambiguous effects on fertility in countries where divorces are more likely and therefore the ambiguity extends to the effect on mothers' labor supply.

Interestingly, also Dahl and Moretti (2008), in their suggestive collage of evidence that American families prefer boys over girls, show 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 (and on labor supply) 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, in Italy and in Sweden. In other words, the effects of a first-born son act in opposite ways on fertility *when all women* (with young first child) are considered, but in advanced countries the "divorce" effect dominates the "desire for a son" effect, and, specifically via the prevailing increase in fertility, a first-born boy reduces labor supply.[2]

[2]One may argue that after a first-born boy, the "desire for a son" effect reduces fertility for all women independently of marriage, while the "divorce effect" increases the stability of marriages and thus fertility for married women only. If this were the case, our results should be the opposite of what we find, because

These results imply that the overall pattern of pregnancies, after the first one, may be affected by the sex of the first born. While the sex of the first child is randomly assigned, the sex and number of subsequent children are endogenous with respect to parental preferences and behavior. For this reason Lundberg and Rose (2002), in their study of the relationship between labor supply and parenthood, fail to find evidence that the labor supply of women varies according to the sex of their *full set* of children. Although they focus correctly on all men and women, not just on married parents, they do not realize the endogeneity of the full parity. Only when the gender of the first born is considered without conditioning on marriage, the effect a child's sex on fertility and female labor supply emerges clearly, as shown in this paper.

We do not rule out the possibility that other channels may exist, through which a first-born son may affect the labor supply of women. We do not have data to study convincigly these complementary channels as well, but our analysis does not need to exclude them. First, there is some evidence that sons have more health and behavioral problems than daughters at the early stages of their life, while, possibly, the opposite happens at puberty, when, for example, psychological morbidity (e. g. anorexia) tends to affect girls more than boys (see e.g. World Health Organization, 2009). Mothers may then work less in the case of first-born sons because, on average, physical or psychological morbidity is higher for them (Frijter et al., 2009). Second, mothers may value direct child care and, because they prefer sons to daughters, work less if they have sons. In other words, time spent with sons might enter the utility of mothers with a greater weight than time spend with daughters.[3] Third, and opposite with respect to the previous effects, it is sometimes suggested that the costs of raising children depend on gender. Our evidence may be seen as more compatible with higher costs of raising girls, instead

unmarried women are subject only to the first effect while an ambiguity could emerge only for married women who would be subject to both effects. This issue is discussed extensively in section IV.

[3] However, according to the Gallup polls on gender preferences men are more likely than women to prefer boys, while women have more equal preferences. See e.g. www.gallup.com.

of boys, but the existing literature on these differential costs is inconclusive.[4]

We want to emphasize that the contribution of this short note is just to bring together and reconcile different already well known pieces of a puzzle that have never been jointly considered but that are all necessary to fully understand the hidden consequences of a first-born son on maternal labor supply. We believe that this recomposition of the puzzle, although limited in scope, must be brought to the attention of the profession.

In Section I we describe the basic facts that motivate our paper. Section II discusses the randomness of the first child sex. Its effects on marital stability and fertility among all women are described in Section III. Section IV brings the pieces of the puzzle together and Section V concludes.

I. 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 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.[5] In the bottom 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). 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, children older than 15 may potentially enter the labor market and their decision may affect maternal labor supply, for instance because of income

[4]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 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.

[5]We have also carried out tobit estimates where the dependent variable is hours worked. They give even stronger results for all the countries.

sharing within the household.

Second, in some of the the datasets we can identify only mothers with cohabiting children. This is the case of all the US datasets, the UK Census and the Italian LFS,[6] but not the case of the BHPS and the Swedish data.

In the Appendix we provide a short description of the datasets and of the motivations for their use.

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. 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 non-negligible share of children used to leave the household around age 16. Our results are confirmed in this different sample and available upon request.[7]

In the first column we use the 1980–2000 waves of the US Census. Consider first hours worked per week at the time of the interview as a measure of labor supply. The top 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. The bottom panel of the first column reports the probability of being employed, available for the Census waves from 1960 to 2000. On average, during the period, 53.6% of the women whose first-born child is a girl are working. Since the sex of the first child is random,

[6] This potential source of bias, however, is negligible for the Italian LFS as in Italy in the years under consideration the share of children leaving parents' home before age 18 is lower than 1%.

[7] In principle the PSID might help us to recover all the offsprings of the householdhead and the wife and minimize measurement error, at least for data collected after 1985. (Non-cohabiting children are not measured before 1985). However, as it tipycally happens in panel data, non-response rates are correlated with household structure and marital instability. In the next sections we will show that the probability of these events may depend of the sex composition of children. This prevents the use of the 1985-2009 waves of the PSID.

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 statistically significant.

In column 2 and 3 of the same table we find that these results are confirmed using two other independent sources of data for the US: the Current Population Survey (CPS, for the years 1994-2011, March supplement) and the National Health Interview Survey (NHIS, for the years 2005-2010). 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 relatively larger and at least equally significant despite the smaller sample sizes. Hours worked decrease by -0.8% in the CPS and by -1.5% in the NHIS, while the correspondent percent effects on the probability of working decreases by 0.4% in the CPS and by 1.4% in the NHIS.

Columns 4 and 5 of the table further show that these effects are observed not only in the US but also in the UK, focusing on similarly defined samples of women and using the same regression specification. In the UK Census of 1991 (column 4) the hours worked per week in case of a girl are 12.4 and decrease by 1.9% in case of a boy. 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 results based on the BHPS are even stronger. Hours worked decrease by 12.5% in case of a first born girl, while the probability to work decreases by 9.5% (column 5).

Results for Italy are reported in the sixth column. In this case the data come from the Labor Force Survey (LFS, for the years 2004-2011). 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 effect of a first-born boy on labor income used as a broader measure of labor supply, as for instance suggested by Gerber and Mitchell

(2011)[8], is again negative and the size is comparable with what found for the other countries. The probability of working is still negatively affected by the sex of the first child even if the sign is not significant.

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. Our estimates 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.[9] From an earnings' perspective, these effects translate into large labor income losses. Using the average hourly pay of a woman aged between 18 and 55[10], we calculate that over the 15 years following the birth of the first child these losses (using 2007 as the base year) amount to roughly 8,300 dollars in the US, to 5,000 euros for Italian women, 8,000 pounds for UK women and 11,000 krona in Sweden (slightly more than 1,600 US dollars).[11]

[8]Labor income incorporates increases of labor supply at both the extensive (hours) and the intensive (effort per hour) margins, inasmuch as effort per hour is reflected in the wage.

[9]We carried out a similar exercise on the American and Italian Time Use surveys. 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.

[10]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, hours worked per week and months worked per year. Hourly wages are then estimated by assuming that on average women do not vary their working time during the year. All data refer to gross earnings. In order to calculate the lifetime value of the loss we have used a discount factor equal to 3%. By the use of the MORG-CPS we have also carried out a regression where the dependent variable is labor 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 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.

[11]Unfortunately we cannot calculate lifetime effects because we cannot estimate the effect of the sex of the first child after 15 years from the her/his birth. This is because at age 16 children can enter the labor market and affect labor mothers' supply. If we assume a permanent effect and calculate this loss over 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).

II. Is the sex of the first child exogenous?

The estimates presented in the previous section are valid only if the sex of the oldest child living with his/her mother is exogenous. Lack of exogeneity may arise because of two channels. First, the gender of the first child is random at conception but may not be random at birth because the success of pregnancies could be correlated with socio-economic characteristics of parents for biological and evolutionary reasons. For example, Catalano et al., (2005a,b) show that the birth ratio increased after the September 11 attack and, more generally, 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. In general, this evidence is rather inconclusive, but this hypothesis cannot be *a priori* excluded.[12]

Endogeneity might arise also if child custody after divorce is affected by the sex of the first child. Dahl and Moretti (2008) argue that in case of divorce fathers get more frequently the custody of boys instead of girls. So, identifying the sex of the first child looking at the sex of the oldest cohabiting child might lead to exclude from our labor supply estimates all mothers who divorced and did not get child custody. However, although the US and the UK Census, the CPS, the NHIS, and the Italian LFS may be affected by this potential source of bias, results based on the BHPS and the Swedish data are not. In these last two datasets we observe all children, independently of cohabitation.[13] Moreover, Dahl and Moretti (2008) select a sample of children who live with a divorced parent in their analysis on child custody. They define maternal custody as the situation in

[12] Oster (2005) describes the effect of hepatitis B as one of the causes of the birth sex ratio in favor of boys in many Asian countries, but her evidence, at least for China, has been widely criticized.

[13] In the BHPS we have information on the sex of all the children of women in the sample, even in the case of child death.

which they observe a child living with the mother at the time of the Census and zero if the child is in the custody of the father. They find that girls have a higher probability to live with a divorced mother. We argue that their evidence may be affected by selection bias. This is because within the sample of children with divorced parents the sex of the first-born child is not random, since, as shown by themselves, the probability of divorce is higher for mothers with girls.[14] This effect will be discussed more deeply in Section III.[15]

To provide evidence in favor of our assumption that the sex of the oldest child living with his/her mother is exogenous, we regress the sex of the first child on a set of mothers' socio-demographic characteristics, like age, educational attainment, and for the US and the UK, also race and interaction terms between education and race. Estimates for the US Census are distinct by year as the issue of custody is more likely to be relevant for recent years than for the Sixties. For simplicity we report estimates for 1960 and 2000, but the results for the other years are very similar. Table 2 reports the F statistics for the significance of all covariates. For all the countries and datasets the F statistics are always below 2 and the hypothesis that none of the covariates help to explain the sex of the first child cannot be rejected.

III. The effects of a first-born son on fertility and marital stability

The collage of evidence presented in Section I and its implications for earnings and employment rates are undoubtedly solid and cannot be disregarded. They are likely to depend directly and indirectly on a wide set of factors that will need to be explored. But what we find most interesting, and we study in this short note,

[14] Dahl and Moretti, 2008, would like to estimate the probability of mother/father custody after divorce. Ideally, to assume that custody arrangements are exogenously determined, they would need a sample of divorced women, assigning a dummy equal to 1 to mothers who get custody of the child and 0 otherwise. Census data do not allow them to carry out this exercise because they cannot identify mothers who cannot live with their children. Therefore they select a sample of children in the custody of one of the two parents and set a dummy equal to 1 if the parent is the mother and 0 otherwise. This is a case of selection on outcomes.

[15] Another source of bias may arise if boys and girls were affected by different infant mortality rates. To our knowledge, there is no evidence supporting this hypothesis in the developed countries considered in this paper.

is how they relate to the effects of a first-born child on fertility and marriage.

A large body of evidence for developing countries[16] 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 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).[17]

We therefore argue that the effects of a first-born boy on fertility and marital stability 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

[16] See for example Jayachandran and Kuziemko (2011) and Chun and Oh (2002).
[17] According to the US Census 1960-2000 married women have on average 1.96 children, unmarried have 1.73. In the UK Census these values 1.95 and 1.73 respectively. In Italy 1.68 and 1.34, in Sweden 2.02 and 1.72.

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.

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, between 1.1% and 7.3% in the UK, 0.7% in Italy and 0.6% in Sweden. Thus, differently than in those developing countries for which some evidence exists, in the US, the UK, Italy and Sweden mothers whose first child is a boy have higher fertility and tend to work less.

Our claim is that this happens 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 (0.6%). In the UK, Italy and Sweden the results are in the same ball park of the US estimates.[18]

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. Of course the greater marital stability after a

[18]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.

first-born son may also reduce maternal labor supply independently on fertility because mothers can count on the income support of their husbands (see e.g. Ananat and Michaels, 2008). Nevertheless, the evidence presented in this Section implies that the higher fertility induced by marital stability cannot be disregarded as an explanation of the observed evidence, and this is the novel finding of our study.

IV. A missing result in the literature

If both the "divorce" and the "desire for a son" 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, and this explains why our results differ from Dahl and Moretti (2008). In Table 5 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) we can conclude that a first-born boy reduces this measure of fertility. Columns 2 and 3 are based instead on all women and the interesting finding is in column 3, which includes the gender of the first child, marital status and the interaction between these two variables. In this decomposition we see that in general a first-born boy has a small positive and significant coefficient, but if the woman is married it changes sign while remaining statistically significant[19].

This evidence may seem puzzling given that after a first-born boy, the "desire for a son" effect should reduce fertility for all women independently of marriage, while the "divorce effect" should increase the stability of marriages and thus fertility for married women only. If this were the case, in Column 3 of Table 5 we should find that the main effect of a first-born boy on fertility for all women is negative, while it becomes positive for married women. However, Dahl and Moretti (2008) have shown that a first born boy increases the probability of shotgun marriages, that end up being more stable precisely because they originate from a first-born boy. So, what we call the "divorce effect" applies to all women, not only to married women. Second the "desire for a son" effect is probably stronger among married women than among unmarried women, since, as shown again by Dahl and Moretti (2008), it mostly originates from fathers. For at least these reasons it is not puzzling that our results based on *all women* differ from those of Dahl and Moretti's (2008) which are based on *married women* only.

Columns 4-6 of panel (a) break the evidence by the number of children replicating the results obtained by Dahl and Moretti (2008) for the probability that a woman has at least two children after a first-born boy. Once again the interaction term between being married and having a first-born boy is positive.

Dahl and Moretti (2008) replicate their exercise on the sample of women who are at their first marriage and find even stronger results. Panel (b) uses the same sample for years 1960 and 1980 and the same specifications of panel (a). The effect of a first-born child decreases by one half if we consider the sample of all women and becomes positive and not statistically different from zero when the model includes also the interaction term between marital status and the sex of the first child. The same holds for the probability of having at least one more child after the first. For all estimates presented in Table 5 formal tests do not

[19] The formal test reported in Table 5 supports the hypothesis that the first-born boy coefficient estimated on the sample of married women differs from the one estimated on the sample of all women.

support the hypothesis of equality of the "first-born boy" coefficient of models 1 and 4 (sample of married women) and the ones corresponding to other models (sample of all women), at least at the 10 percent level.

Based on the evidence similar to the one reported in columns 1 and 4, 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 more than one child and the total number of children are strongly correlated with her marital status, as married women have on average more children than unmarried women.[20] At the same time, the probability of being married is in turn influenced by the sex of the first-born child.[21]

There are few other studies that focus explicitly on the relationship between the sex of children and parental labor market behavior. Most of them operate the same sample selection of Dahl and Moretti (2008), based on the parents being married, that we consider inappropriate for the question that this literature wants to address. For example, Wulff Pabilonia and Ward-Batt (2007) investigates the effect of the first child's gender on parental labor supply in the US. Their results are typically mixed and statistically insignificant, but they do find that Asian men work fewer hours compared to white men if they have a son. No effect is instead found for mothers of any race and ethnicity. Lundberg and Rose (2002), using PSID data, look at a sample of fathers and mothers and find some effect of the first child gender only on fathers. Differently than other studies they correctly base their estimates on all men independently of marital status. However, their estimates are affected by other sources of bias. First, their sample includes also men with no children and this heterogeneity is unlikely to be random. Second, they regress labor supply on the total number and sex composition of the boys and girls that men have, which, as we have argued above, is not randomly assigned.

[20]According to the US Census, on average married women have 2 children and unmarried 1.7. A similar difference can be found in the other countries.

[21]For robustness check, we have replicated these estimates also for all the other countries here considered, finding similar results. Estimates are available upon request.

V. Conclusions

This short note brings together and reconciles some already well known pieces of a puzzle that have never been jointly considered to emphasize the consequences of a first born son on maternal labor supply. 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. Our estimates are statistically significant and translate into quantitatively relevant labor income losses over the lifetime. The effect of the first child sex is the combined result of at least two 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 more, 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 increases when the first child is a male. As a result, in the countries that we consider, a first-born boy decreases maternal labor supply. Our study emphasizes the importance of using data on *all* women not only on *married* women to study the effects of the first-child gender on mothers' labor market outcomes.

VI. References

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VII. Appendix: The data

Census data for US refer to years 1960–2000 are collected within the IPUMS International project and are available at www.international.ipums.org. 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. Differently from other versions of the Census data, the information about the total number of marriages is available only for years 1960 and 1980.

Current Population Survey data are drawn from the NBER site www.nber.org/cps/. The universe is the civilian non-institutional 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 old. We use data from 1994 to 2011, because of a large redesign of the survey occurred in 1994.

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 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. In 2005 the NHIS was subject to major chganges in the sample structure and questionnaire. For this reason we focus on years 2005-10 (last available). In this paper we have used the NHIS also to checxk whether there is strong evidevce of correlation between the health status of the first child and mother's behavior. We have not found any clear evidence and we have choosen to do not report this additional set of results.

Census data for UK refer to year 1991 and are collected within the IPUMS International project and are available at `www.international.ipums.org`. They are roughly a 1% random sample of the UK population and contain both personal and household identifiers and socio-demographic characteristics. Data on the Census 2011, as reported by IPUMS, do not contain a household identifier and cannot be used in this paper.

The British Household Panel Survey (BHPS) is conducted yearly since 1991 and data are available through the UK data archive, after a free-of-charge registration. The last available wave is wave 18 (year 2009). The 1991 sample includes more than 8,000 households and 23,000 individuals. The BHPS collects also detailed retrospective information on fertility and marital status of individuals before 1991. It also collects gender and age of children not cohabiting with the parents. The BHPS is subject to some non-negligible panel attrition, i.e. increasing levels of non-response with each successive wave of the panel. Since marital dissolution is endogenous to the sex of the first child and panel attrition is typically strongly correlated to marital dissolution, we prefer to avoid such potential source of bias and in this paper we use only data on wave 1.

The Italian Labor Force Survey is conducted by the Italian Statical 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, April, July and October. Because of a break in the sample design and questionnaire in 2004, data comparability before 2004 is not ensured. Istat produces two types of files: one for public-use and a full version for public Institutions, like for instance Italian Universities. The public-use files contain a household identifier for around 95% saample households (for privacy concerns) and detailed socio-demographic characteristics also for individuals aged less than 16. In this paper we use the full-version files. Public-use files are released by Istat, free of charge. As the CPS no retrospective information on fertility and marital status is included. The estimates presented in this paper go from January 2004 to July 2011, the last file available.

The Swedish data, provided by Statistics Sweden, is a population-wide panel data set (LISA) based on administrative records. Detailed demographic variables are observed on a yearly basis 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.

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

	US			**UK**		**Italy**	**Sweden**
	Census	CPS	NHIS	Census	BHPS	LFS	LISA
	1960-2000	1994-2011	2005-10	1991	1991	2004-11	2004
	Hours worked per week (Annual labour income in Sweden)						
First-born boy	-0.092	-0.169	-0.340	-0.232	-1.750	-0.193	-5.790
St. err.	0.021	0.074	0.200	0.139	0.898	0.059	2.864
Baseline: girl	20.294	22.597	22.021	12.358	14.005	15.280	1130.594
St. err.	0.014	0.054	0.154	0.657	0.660	0.043	2.834
Percent effect	-0.452	-0.750	-1.543	-1.881	-12.495	-1.260	-0.512
	Probability of working						
First-born boy	-0.002	-0.003	-0.009	-0.007	-0.051	-0.006	-0.001
St. err.	0.000	0.001	0.005	0.005	0.028	0.002	0.001
Baseline: girl	0.536	0.645	0.606	0.520	0.540	0.494	0.652
St. err.	0.000	0.001	0.004	0.021	0.021	0.001	0.001
Percent effect	-0.366	-0.419	-1.422	-1.406	-9.514	-1.263	-0.153
No. obs.	3,422,119	272,664	35,192	45,068	1,186	336,564	699,805

Notes: 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, for the US Census the dependent variable is equal to the number of hours worked per week in all jobs during the previous year. Data on hours worked are available only for the period 1980-2000. 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). In the bottom 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 labor income, 0 otherwise. All models include a quadratic in age of the mother and year dummies.

Table 2—First child gender and mothers' socio-demographic characteristics.

	US				UK		Italy	Sweden
	Census	Census	CPS	NHIS	Census	BHPS	LFS	LISA
	1960	2000	1994-2011	2005-10	1991	1991	2004-11	2004
F stat.	1.35	1.22	1.19	0.99	1.08	1.19	0.97	0.36
No. obs.	169414	1,108,779	272,664	35,192	44,813	1,186	336,564	699,805

Notes: F-statistics for the null hypothesis that all regressors except the constant are not statistically sinificant. 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. All models include the age of the mother and dummies for educational attainments and year dummies when referred to pools of different years. Estimates based on US Census and the CPS include also race and the interactions between race and educational attainments. Estimates based on the CPS include also time dummies. Estimates on UK Census include a dummy for the native-non native status and the interactions between native and educational attainments. Estimates on BHPS include age of leaving school and race and the interaction term between the two.

Table 3—First child gender and fertility in the US, the UK, Italy and Sweden.

	US			UK		Italy	Sweden
	Census	CPS	NHIS	Census	BHPS	LFS	LISA
	1960-2000	1994-2011	2005-10	1991	1991	2004-11	2004
First-born boy	0.0031	0.0045	0.0135	0.0074	0.0476	0.0040	0.0030
St. err.	0.0004	0.0018	0.0050	0.0044	0.0265	0.0017	0.0010
Baseline: girl	0.641	0.627	0.632	0.656	0.648	0.548	0.543
St. err.	0.000	0.001	0.004	0.020	0.020	0.001	0.001
Percent effect	0.489	0.720	2.139	1.127	7.343	0.726	0.552
No. obs.	3,422,119	272,664	35,192	44,813	1,186	336,564	699,805

Notes: Women who had their first child between 18 and 40 years and whose first child is aged no more than 15. The dependent variables are 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.

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

	US			**UK**		**Italy**	**Sweden**
	Census	CPS	NHIS	Census	BHPS	LFS	LISA
	1960-2000	1994-2011	2005-2010	1991	1991	2004-2011	2004
First-born boy	0.0067	0.0076	0.0104	0.0039	0.0135	0.0022	0.0030
St. err.	0.0004	0.0015	0.0044	0.0034	0.0058	0.0009	0.0010
Baseline: girl	0.864	0.794	0.764	0.846	0.792	0.919	0.670
St. err.	0.000	0.001	0.003	0.017	0.017	0.001	0.001
Percent effect	0.047	0.192	0.580	0.397	0.731	0.102	0.149
No. obs.	3,392,600	272,664	23,121	44,813	1,174	336,564	699,873

Notes: 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.

Table 5—Re-assessing the effect of first child gender on fertility in the US.

	Total Number of children			Two or more children		
	Panel a					
	Married	**All women**		**Married**	**All women**	
	(1)	(2)	(3)	(4)	(5)	(6)
First-born boy	-0.0028	-0.0002	0.0053	0.0015	0.0029	0.0049
St. err.	0.0012	0.0011	0.0024	0.0006	0.0006	0.0013
Married			0.2208			0.1420
St. err.			0.0020			0.0011
Boy*Married			-0.0079			-0.0032
St. err.			0.0027			0.0014
Test: H_0: Coeff. Boy = to column 1(4)						
p-values		0.0147	0.0008		0.0166	0.0082
N. obs	2,223,685	2,765,483	2,765,483	2,223,685	2,765,483	2,765,483
	Panel b					
	In first marriage	**All women**		**In first In marriage**	**All women**	
	(1)	(2)	(3)	(1)	(2)	(3)
First-born boy	-0.0110	-0.0063	0.0024	-0.0018	0.0004	0.0030
St. err.	0.0022	0.0019	0.0038	0.0011	0.0010	0.0020
Married			0.1920			0.1200
St. err.			0.0032			0.0016
Boy*Married			-0.0131			-0.0046
St. err.			0.0044			0.0023
Test: H_0: Coeff. Boy = to column 1(4)						
p-values		0.0661	0.0005		0.0297	0.015
N. obs	702,865	926,939	926,939	702,865	926,939	926,939
Memo:						
% married women	80.4					
% first marriage	75.8					

Notes: 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 columns 1-3 the dependent variable is the total number of children. In columns 3-6 the dependent variables are dummies equal to 1 if the woman has two or more children and 0 otherwise. In panel b - column 1 only women in first marriage are included. In columns 2,3,5,6 the p-value corresponds to a test of equality of the first-born boy coefficient (reported in that column) with the one obtained using married women only.