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ABSTRACT

In this paper, we use 2008-2013 American Community Survey data to update and further probe evidence on son preference in the United States. In light of the substantial increase in immigration, we examine this question separately for natives and immigrants. Dahl and Moretti (2008) found earlier evidence consistent with son preference in that having a female first child raised fertility and increased the probability that the family was living without a father. We find that for our more recent period, having a female first child still raises the likelihood of living without a father, but is instead associated with lower fertility, particularly for natives. Thus, by the 2008-2013 period, any apparent son preference in fertility decisions appears to have been outweighed by factors such as cost concerns in raising girls or increased female bargaining power. In contrast, some evidence for son preference in fertility persists among immigrants. Immigrant families that have a female first child have significantly higher fertility and are more likely to be living without a father (though not significantly so). Further, gender inequity in source countries is associated with son preference in fertility among immigrants. For both first and second generation immigrants, the impact of a female first-born on fertility is more pronounced for immigrants from source countries with less gender equity. Finally, we find no evidence of sex selection for the general population of natives and immigrants, suggesting that it does not provide an alternative mechanism to account for the disappearance of a positive fertility effect for natives.

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I. Introduction

Economists have long been interested in studying the unequal treatment of women and men in families. These inequities may even manifest early in the life cycle as a son preference. Specifically, a child's gender has been shown to impact family structure and future fertility, sex-selective abortion, and sex differences in parental time inputs, as well as access to health care and nutrition.[1] Some of the strongest evidence of son preference comes from developing countries. Sen (1990), for example, inferred that there were millions of "missing women" in China and India, due largely to neglect in health care and nutrition. With the advancement and increased availability of "safe, effective, inexpensive and accessible technologies to determine the sex of a fetus and to abort unwanted pregnancies," sex selective abortion came to play a major role in unbalanced gender ratios, with (male/female) sex ratios at birth rising in a number of countries, mostly in Asia (Bongaarts 2013, p. 185).[2] Moreover, as Anderson and Ray (2010) note, unequal survival rates from specific diseases such as AIDS can also create sex imbalances in the population. Beyond leading to lower female birth or survival rates, the unequal treatment of women and men also potentially affects gender inequality in the family and in society.

Studies by sociologists and psychologists, but more recently by economists as well, have also found evidence of differences in the behavior of parents of sons and daughters even in developed countries. In a comprehensive review, Lundberg (2005) points to two fairly robust findings: sons increase family stability and, overall, fathers tend to spend more time with sons than daughters; however, recent research by Baker and Milligan (2016) finds that parents of

[1] See, for example, Dahl and Moretti (2008), Anderson and Ray (2010), Almond and Edlund (2008), Abrevaya (2009), Almond, Edlund and Milligan (2013), and Lundberg (2005).
[2] See, for example, Sen (2003) on India and Ebenstein (2010) on China. More recently, (male/female) sex ratios at birth in Korea, which used to also be extremely high, have declined to natural biological levels along with other indicators of son preference (Choi and Hwang, forthcoming).

preschoolers invest more time in girls than boys in teaching activities (e.g., reading to children). While differences may be identified, as Lundberg (2005) notes, it is unclear whether they reflect son preference or constraints, like differences in the productivity of fathers and mothers in parenting sons vs. daughters or differences in costs of boys vs. girls. In a landmark study for the United States, Dahl and Moretti (2008) found evidence consistent with son preference. They estimate that having a female first child increased the probability of living without a father and also raised fertility. However, the evidence on fertility in the United States and other developed countries is somewhat mixed, with several studies finding evidence of a negative effect of girls on fertility for some countries including the U.S. (Abrevaya 2009; Ichino, Lindström, and Viviano 2014; Andersson, Hank, Ronsen and Vikat 2011).

In this paper, we revisit the question of son preference, adding to the literature in several ways. First, we use 2008-2013 American Community Survey (ACS) data to update and further probe Dahl and Moretti's (2008) son preference results for heterogeneity by immigrant status. Updating is important because their analysis of family structure used data from 1960-2000, and, more significantly, their results for fertility were for the 1960-80 period. Like Dahl and Moretti (2008), we exploit the apparent randomness of the gender of the first child to address the endogeneity of family structure and fertility to child gender.[3] Moreover, we address the mixed evidence from other studies by more closely following the Dahl and Moretti (2008) research design and explicitly exploring the impact of any departures. Further, in light of the increase in immigration and research showing that more recent immigrant waves tend to come from countries with a more traditional gender division of labor than in the United States (Blau, Kahn and Papps 2011), it is desirable to analyze immigrants and natives separately.

[3] Below, we discuss in detail whether one can in fact make such an assumption about the sex of the first child.

Among the population in the aggregate, as well as among the native-born separately, we find that having a female first child continues to raise the likelihood of living without a father. However, in contrast to Dahl and Moretti's (2008) earlier findings, we find that for the overall population, as well as among natives separately, having a female first child is now associated with *lower* fertility, significantly so for natives. Thus, for the U.S., by the 2008-2013 period, son preference among natives in their fertility decisions appears to have been reversed or outweighed by factors such as a higher cost of raising girls or increased female bargaining power.

Considering immigrants separately, we find some evidence that having a female first child contributes to the incidence of living without a father, although the impact is statistically insignificant. However, in contrast to our findings for natives, we do find a positive effect of a female first child on fertility, suggesting son preference in fertility among immigrants. This interpretation is further supported by evidence that, for both first and second generation immigrants—immigrants and their native-born children—having a girl has a more positive effect on fertility for those from source countries with less gender equity, as measured by the World Economic Forum's (WEF) Global Gender Gap Index and other indicators. (The second generation was examined using the 1995-2014 Current Population Surveys, which have information on the birthplaces of respondents' parents.)

We also examine another indicator of son preference, sex selection, by estimating the impact of the sex composition of previous children on the probability that a given birth is a boy. We find no evidence that sex selection characterizes the aggregate native and immigrant populations, although previous work indicates it occurs for some groups (e.g., Almond and Edlund 2008; Abrevaya 2006). This suggests that an increase in sex selection among natives as an alternative manifestation of son preference is not driving our fertility findings.

II. Literature Review and Our Contribution

As noted above, Dahl and Moretti (2008) made a major contribution in finding evidence of son preference in the United States. Specifically, using data for the 1960-2000 period, they found that first-born girls were less likely to be living with their father than first-born boys. While such a result is consistent with fathers' preference for sons, there are other possible explanations. One is that raising girls is more expensive than raising boys, making fathers more reluctant to shoulder this burden (e.g. Lundberg 2005; Dahl and Moretti 2008). We note that this possibility is especially plausible in more recent decades given girls' greater propensity to attend college than boys' beginning in the 1980s (Goldin, Katz and Kuziemko 2006), and we will present some direct evidence on the relative expense of raising girls that is consistent with the cost argument. Another possibility is that parents believe the lack of a male role model is more harmful for boys than girls or that fathers have a comparative advantage in raising sons (e.g. Lundberg 2005; Dahl and Moretti 2008). This is consistent with recent empirical evidence suggesting the negative effects of growing up with economic disadvantage, and particularly in a single-mother family, are more harmful for boys than girls. For example, Autor, Figlio, Karbownik, Roth, and Wasserman (2019) find larger negative effects for boys than for girls on a number of education-related outcomes of being born to low-educated, unmarried mothers and raised in disadvantaged neighborhoods.[4] And, as another example, Bertrand and Pan (2013) find that being raised in a single-mother household has major negative consequences for boys' noncognitive development but much less so for girls.

[4] These outcomes include being kindergarten-ready, incidence of truancy and behavioral problems in elementary and middle school, performance on standardized tests, and high school graduation.

In order to distinguish between a preference for sons from these other explanations for the effect of girls on family structure, Dahl and Moretti (2008) examine the impact of a female first child on a couple's subsequent fertility. While son preference implies that the probability of having additional children should be *higher* for all-girl than for all-boy families, the alternatives discussed above suggest, if anything, it should be lower. If girls are more costly (e.g., due to education expenses), the probability of having an additional child for all girl families should be lower than for all boy families. Similarly, if fathers have a comparative advantage in raising boys, this would also make girls more expensive to bring up and lower the probability of having additional children.

Using Census data for 1960-80, Dahl and Moretti (2008) find that the effect on fertility of having a female first child is positive, supporting an interpretation that their finding for family structure reflects son preference. We are able to substantially update Dahl and Moretti's (2008) fertility analysis because our 2008-2013 ACS data contain crucial information on marital history that they argue is needed for the fertility analysis and has been otherwise unavailable in Census data since 1980: whether the respondent had been married more than once. As explained below, it may be argued that using a sample of women in first marriages provides the cleanest test of son preference in fertility.

As we mentioned, evidence on the impact of girls on fertility is mixed, with a number of studies finding a negative effect for the United States and other developed countries, unlike Dahl and Moretti (2008). Abrevaya (2009) found, for whites in the U.S., that families whose first child is a boy are significantly more likely to have a second child than a family whose first child is a girl. Moreover, Ichino, Lindström, and Viviano (2014) found for the U.S., UK, Italy and Sweden that a first child boy increased the probability that a woman would have more children.

In addition, Andersson, Hank, Ronsen and Vikat (2011) found for Denmark, Norway and Sweden (but not Finland), fertility effects consistent with daughter preference for third births, with no effect of child gender on second births. These results are not directly comparable to Dahl and Moretti's (2008), however, due to differences in the empirical designs. Specifically, unlike Dahl and Moretti, the samples used were not restricted to married women. Indeed, Ichino et al (2014) show they are able to replicate Dahl and Moretti's finding of a negative effect of first child boy (implying a positive effect of first child girl) on fertility when they restrict the sample to married couples.[5]

Ichino, Lindström, and Viviano (2014) add an additional outcome variable to consider when examining son preference. Specifically, using a sample of all women, they study the impact of a male first child on the mother's labor supply, finding a negative effect. They offer two reasons for such an impact. First, a boy first child raises the probability of being married, which lowers the mother's labor supply; second, within marriages, a boy first child raises fertility, which also lowers the mother's labor supply.

Another strand of research on son preference examines North Americans with a heritage from countries that have been known to practice sex-selective abortion, an extreme form of son preference: China, Korea, and India. For example, Almond and Edlund (2008), using 2000 U.S. Census data, found evidence consistent with sex selection increasing the male/female sex ratio at the birth of the third child for Chinese, Asian Indian, or Korean women. Similar findings for Chinese and Asian Indian mothers have been found by Abrevaya (2009), who used California birth record data. Almond, Edlund and Milligan (2013), using Canadian Census data, found a similar result for first and second generation South and East Asian immigrants. Moreover,

[5] In an additional difference in design, Andersson, Hank, Ronsen and Vikat (2006) use the gender composition of the first two children as explanatory variables for the third birth, rather than sex of the first child.

Abrevaya (2009) finds a positive relationship between having girls and subsequent fertility for U.S. immigrants from these areas, as did Almond, Edlund and Milligan (2013) for Canada. The finding of son preference in influencing family gender composition has been recently questioned by Persaud, Kalantry, Citro and Nandi (2015), using more recent U.S. data, who found some evidence in favor of a preference for diversity rather than for son preference among these groups.

In this paper, we add important new findings to the literature on son preference that cast doubt on its continued prevalence among the native born in the United States. In particular, our finding that there is no longer a positive effect on the future fertility of a first child girl is significant because it raises a question about whether the findings on living without a father are due to son preference or one of the alternative interpretations discussed above. In obtaining our results we either adhere to the Dahl and Moretti (2008) research design or track the impact of any departures, ensuring that we are able to reach conclusions about trends over time in the impact of child gender on family structure and fertility in the U.S. based on comparable data. Further, while maternal labor force participation is not a major focus of our paper, in light of the findings of Ichino, Lindström, and Viviano (2014), we examine it briefly.

Given the growth in the immigrant share of the population and the tendency of immigrants to come from countries with a more traditional gender division of labor (Blau, Kahn and Papps 2011), we present results analyzing immigrants and natives separately. This investigation is also motivated by recent research highlighting the role of culture in affecting gender-related outcomes such as fertility and labor supply.[6] Specifically, we compare results for living without a father and fertility among first or second generation immigrants whose source

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[6] See, for example, Fernandez and Fogli (2006, 2009), Blau (1992), Antecol (2000), Blau, Kahn and Papps (2011), Blau, Kahn, Liu and Papps (2013), and Blau and Kahn (2015) for studies of the impact of culture on female labor supply and fertility behavior among first and second generation immigrants. See also Nollenberger, Rodríguez-Planas, and Sevilla (2016) for an impact on the gender math gap.

countries differ with respect to the WEF's Global Gender Gap Index or alternative indicators of female status including female labor force participation rates and (boy-to-girl) sex ratios among births. These comparisons in effect provide an estimate of the impact of different "doses" of son preference as indicated by measures of gender equity differences across source countries.

Finally, while alternative manifestations of son preference may be viewed as complementary pieces of information in establishing son preference, it is also possible that they are to some extent substitutes. For example, if sex selection has become more prevalent throughout the general population, it might help to account for the decline we observe in son preference in fertility. Thus, we briefly examine the impact of the sex composition of previous children on the probability that a given birth is a boy. As noted previously, we find no evidence that sex selection has come to characterize the general population of natives and immigrants. This rules out the possibility that an increase in sex selection among natives could account for our fertility findings for this group.

III. Data and Research Design

Our central goal is to study the incidence of living without a father and fertility in order to make inferences about son preference in the contemporary United States. Following Dahl and Moretti (2008), our key explanatory variable for studying son preference is the gender of the first child, which, in the absence of sex selective abortion, is expected to be roughly random.[7] In contrast, if child gender influences family structure (Dahl and Moretti 2008) or if there are son-biased fertility stopping rules (Choi and Hwang forthcoming), the gender composition of

[7] Others who employ this specification for similar reasons include, for example, Ichino, Lindstrom, and Viviano (2014) and Choi and Hwang (forthcoming). We present evidence on randomness below where we consider the possibility that maternal condition could affect the gender of the first child.

subsequent births will be endogenous, potentially biasing the estimate of child gender on family structure and fertility. To see this assume that, given son preference, a first born girl is expected to increase the probability that the mother is unpartnered (due to an unwed birth or divorce) and that single parenthood is likely associated with lower fertility. If this is the case, if we condition on, say, the sex of the first two children in studying the probability of living without a father, we would be missing some of the effect of the first child sex on family breakup, which would have lowered the probability of having a second child. Moreover, for both living without a father and for fertility (even if it is restricted to married couples), conditioning on the sex of say the first two children would produce a self-selected sample with respect to the strength of parents' preferences for boys and for family size; that is, sex-biased stopping rules would lead to biased comparisons between, say, families with two boys vs. two girls (Choi and Hwang forthcoming).

We use American Community Survey (ACS) data for 2008-2013 to study the behavior of the full population and of natives and immigrants separately. And, since information on parental birthplace is not available in the ACS, we use the March Current Population Survey (CPS) data for 1995-2014 to study the immigrant second generation (native-born individuals with at least one foreign born parent). We use a wider time window for the CPS to increase sample size.[8]

We begin by estimating equation (1):

$$(1) \qquad y_{it} = \beta F_{it} + \gamma' X_{it} + \phi_t + \theta_{r(i)} + \varepsilon_{it}$$

where for each woman *i* in year *t*, y_{it} is an outcome variable including a binary for living without a father and indicators of fertility (the number of children and, in some specifications, a binary for having n or more children, n = 2, 3, or 4); F_{it} is a binary equal to one if the firstborn

[8] Information on parental birthplace became available in the CPS starting in 1994. We begin our analysis of the CPS with the March 1995 wave because the 1994 survey had insufficient detail on parents' birthplaces.

child is female, X_{it} is a vector of controls (including an intercept), ϕ_t and $\theta_{r(i)}$ are respectively year and region (based on 9 Census categories) fixed effects, and ε_{it} is a disturbance term.

The vector X includes a cubic in respondents' age, respondents' education (based on < HS, HS only, Some College and College Degree), and race/ethnicity (based on non-Hispanic White, non-Hispanic Black, non-Hispanic Asian, non-Hispanic Other, and Hispanic). For analyses estimated on married samples, analogous spouse education, race/ethnicity and age variables are also included. Standard errors are heteroskedasticity robust and regressions are weighted by adjusting the ACS sampling weights so that each sample year of data contributes equally to the estimation.

When we study fertility, we focus on married couples. As mentioned above, having a first child girl raises the probability of single parenthood or living without a father, which likely lowers fertility. Thus, using a sample that includes nonmarried as well as married individuals would combine the direct effect of son preference on desired fertility and the negative indirect effect of son preference on fertility via its impact on single parenthood. Even if couples do have a preference for sons, if the first child is a girl, the couple is less likely to form or sustain a marriage, reducing the opportunity for the effect of son preference on fertility to be observed in realized fertility. Thus, studying fertility among married individuals is likely to give a more accurate assessment of desired fertility than including nonmarried women. The inclusion of the indirect effect of first child girl on fertility through its effect on single parenthood is also of concern because we seek evidence on the impact of son preference on fertility to aid in the interpretation of the single parenthood effect as a manifestation of son preference or as due to other factors. Moreover, since women who have been previously married and divorced would have spent some of their time unpartnered, following Dahl and Moretti (2008), in some additional specifications, we capitalize on the marital history information in the ACS to restrict the sample to women in their first marriage who are married to men also in their first marriage.[9] However, given that even a first marriage may be endogenous to the birth or expected birth of a girl, in our main specifications, we also estimate our fertility models on a sample of all women. Using a sample of all women likely results in an underestimate of the effect of a female first child on fertility preferences due to son preference. This is the case because, again, this estimate combines the direct effect of first child girl on fertility preferences with its negative indirect effect on fertility via single parenthood. Nonetheless, it may be of interest as a parameter summarizing the total effect (direct plus indirect effect) of a first child girl on a woman's fertility.

Finally, we note that we do not control for the age of the mother at first birth in our fertility regressions. This is because we view mother's age at first birth as endogenous to the fertility decision itself and thus an inappropriate control. A particular concern is that, for immigrants, maternal age at first birth may be an indicator of the broader cultural factors that we are trying to capture with our source country variables.[10]

To further probe our findings from estimating equation (1) for immigrants in the ACS and for second generation respondents in the CPS, we explore the impact of source country characteristics on the response to a first child girl on the probability of living without a father and fertility by estimating equation (2):

$$y_{it} = \beta_0 F_{it} + \beta_1{}' S_{it} + \beta_2{}' F_{it} \times S_{it} + \beta_3{}' Z_{it} + \beta_4{}' I_{it} + \alpha' X_{it} + \zeta_t + \eta_{r(i)} + \mathrm{u}_{it} \quad (2)$$

[9] While Dahl and Moretti (2008) only apply this restriction to the wife, it seems reasonable to apply it to the husband as well since a child or children born during a previous marriage could affect his preferences for the number and sex of children in the current marriage.

[10] Nonetheless, when we include such a control in our analyses, the fertility results (available on request) are similar to those reported here.

where, the additional variables in (2) are S—an indicator or set of indicators of gender equity in the (own or parental) source country; Z—a set of (own or parental) source country characteristics used as further controls; I—a set of variables referring to individual characteristics specific to immigrants; the disturbance term u; and we again include year and region fixed effects (ζ and η). Regressions are weighted by adjusting the ACS or CPS sampling weights so that each sample year of data contributes equally to the estimation and standard errors are clustered at the (own or parental) source country level.

The variables measuring source country gender equity in S include both main effects and interactions with first child girl.[11] For immigrants, these variables relate to conditions in the country from which the individual migrated; for the second generation, these variables relate to parental source country. Our key variable is the WEF's Global Gender Gap Index (which we term the "Equity Index"). This is an annual index computed for each country that is based on the treatment of women on four dimensions: a) Economic Participation and Opportunity; b) Educational Attainment; c) Health and Survival; d) Political Empowerment.[12] The index is calibrated so that higher values signify *more favorable* outcomes for women. We use the average of the 2006 and 2007 values of the Equity Index for each source country; the variable is therefore measured before the data we use from the ACS, which as noted covers the years 2008-13. The Index is not available prior to 2006.

While the Equity Index provides an overall indicator of the favorableness of a country's environment to women, it imposes a specific weighting on its components. To investigate the separate impact of some important gender-related source country characteristics, in additional

[11] For descriptions of and sources for the source country variables, see the Data Appendix.
[12] This Index has been used as an indicator of gender equality in a number of other studies. See, for example, Guiso, Monte, Sapienza, and Zingales (2008); Zentner and Mitura (2012); Fryer and Levitt (2010); and Nollenberger, Rodríguez-Planas, and Sevilla (2016).

specifications, we replace the Equity Index with either (1) the female labor force participation rate relative to that of men or (2) the relative labor force participation variable and the country's (boys/girls) sex ratio at birth. Note that the measure of source country female labor supply we employ is women's labor force participation relative to men's (female LFP/male LFP). This relative measure is appropriate in that it captures the gender division of labor explicitly. A further advantage is that it implicitly adjusts for problems in measuring the labor force, particularly at different levels of economic development, at least to the extent that such problems affect men's and women's measured participation rates similarly. We follow the WEF in left-censoring the sex ratio at birth at 1.059 to identify son preference rather than natural biological variation. The labor force participation and sex ratio variables are averaged over the 2000-2007 period, thus measuring source country conditions for a reasonable period prior to our ACS data.

The regressions with source country variables also include controls for the main effects of basic influences on family structure and fertility, Z, including total fertility and the log of GDP per capita in the source country, which are also averaged for the 2000-7 period. By including a measure of total fertility in the source country, we are interpreting the impact of having a female first child, controlling for overall tastes for family size in the source country.

The immigrant-related variables I in equation (2) for the immigrant sample include the woman's years since migration and years since migration squared, and, in the fertility regressions, include those variables for her spouse as well as an indicator for whether her spouse was an immigrant. For the second-generation sample, the immigrant-related variables include indicators for whether her spouse was an immigrant and whether her spouse was second generation.[13]

Finally, we examine sex selection. Our goals in doing so are two-fold. First, we seek to confirm that it is reasonable to view the sex of the first child as exogenous—given the fundamental importance of this issue for our research design, we turn to it shortly below. Second, we wish to examine the extent of sex selection for the broad population (beyond the subgroup of individuals of Chinese, Korean, and Indian origin who have been previously examined) to see whether sex selection now constitutes a significant complement to or substitute for son preference in family structure and fertility decisions for broad swaths of the population. In particular, sex selection could provide an alternate channel for influencing the sex composition of children and thus help to account for our failure to find evidence consistent with son preference in the fertility decisions of natives. We consider the latter exercise a robustness check and explore this question after our examination of family structure and fertility. We study whether sex ratios are outside what Anderson and Ray (2010) have identified as a normal biological range of 1.03 to 1.07 (boys to girls), as well as whether the sex composition of prior children influences the sex composition of subsequently born children.

The databases used in this and previous studies do not provide a fertility history for the woman who is the respondent. Thus, following the existing literature (e.g., Dahl and Moretti 2008; Almond and Edlund 2008), we infer birth order, number, and sex composition of children from the children present in the household. We thus impose some sample restrictions in order to increase the likelihood that we are observing all the children born to the respondent. We

[13] Note that apart from broad region controls, we do not control for an immigrant's residence in an "enclave" with others from the same source country. This is because location is endogenous and part of the cultural or attitudinal effect we seek to capture with the source country variables.

construct two samples: the core sample places stronger restrictions on sample composition, but allows us to better identify important variables, such as the gender of the first-born child. The extended sample is more representative of the population, but as explained below we are less confident that important variables are accurately measured.

The core sample includes women, ages 18-40, who are the household head or spouse of the household head, with one or more children, where the oldest child is twelve years old or younger and all children are born in the U.S. These age restrictions on mothers and the oldest child present are made to reduce the probability that there is an older child who has left the household. Households with adopted, step or foster children are dropped in the ACS. This restriction is not implemented in the CPS because we are unable to identify adopted and step children in the CPS in all sample years, but we are able to drop CPS households with foster children. Same-sex couples, respondents living in group quarters, respondents born abroad to American parents, widows, as well as mothers with multiple children born in the same year and quarter (ACS) or same year (CPS) are excluded. When the dependent variable is fertility, the sample is additionally limited to married women with a spouse present, and, in some specifications, to women in their first marriage who are married to men also in their first marriage. Regressions that use the core sample are weighted by household weights that are normalized to provide equal weighting for each sample year.

The extended sample expands the core sample by including father-only families and parents who are not the household head or spouse of the head (i.e., in subfamilies). Men are included in the sample only if their children do not have a mother in the household and if they are ever married (i.e., never married men are excluded).[14] We also expand the sample to include

[14] We follow Dahl and Moretti (2008) in excluding never married fathers because they rarely have custody of their children.

step and adopted children since we are not able to identify these categories of children for subfamilies. We continue to exclude foster children, but, in the spirit of inclusiveness, not their households. Additionally, regressions that use the extended sample are weighted by person weights (as suggested by IPUMS when one is analyzing members of subfamilies: https://usa.ipums.org/usa/volii/subfamilies.shtml , accessed 9/10/19) that are normalized to provide equal weighting for each sample year.

We choose to limit the core sample to women who are the household head or spouse of the household head in order to properly identify family relationships. The ACS and CPS ask each individual for their relationship to the household head. This provides us with the necessary information to match each child in the primary family to his/her parent(s). However, in both the ACS and the pre-2007 CPS, in the case of subfamilies (family units that live in someone else's household), parent and child links "are not based on explicit survey items about how one is related to others in one's household. Rather, they are educated guesses based on other variables" (https://usa.ipums.org/usa/volii/subfamilies.shtml , accessed 9/10/19).[15] We are concerned that, in the absence of precise information, including subfamiles will add measurement error to our data. Furthermore, due to data limitations the identification of step or adopted children is not possible for subfamilies. As the gender of these children is potentially endogenous their inclusion can bias our results, which depend on the gender of the first child being as good as random. It is also unclear when they entered the household and thus they may contribute to errors in identifying the gender of the first child at the time marital and fertility decisions were

[15] See also Schroeder (undated). To see how ambiguity in parent-child links may arise, consider a household in which a child who is the grandchild of the household head is present and the household also includes, e.g., (i) two unmarried daughters of the head of childbearing age or (ii) an unmarried son and daughter of the head of childbearing age; or (iii) a married and a single daughter of the head of childbearing age, etc. In instances like these, IPUMS uses information on age, sex, marital status and other variables with sequential assignment rules to link children to (possible) parents.

made. Consequently, we drop subfamilies in our core sample in order to reduce measurement error and bias. Further, we do not include fathers because we can be less confident than for mothers that the children in the household represent all the respondent's children and thus less confident we have correctly identified the sex of the first born child. This is the case because mothers generally receive custody of children in the case of unmarried births or divorce. Below we present data suggesting that our sample restrictions in the core sample do correctly identify women's number of children ever born in the overwhelming majority of cases.

While the core sample is desirable due to its accuracy, it does exclude two categories—subfamilies and single father families—that may be of interest, particularly in exploring the determinants of living without a father. Moreover, the extended sample more closely corresponds to that used by Dahl and Moretti (2008) thus facilitating a comparison to their results and enabling us to chart trends over time in the impact of a first-child girl on living without a father and fertility. Thus, while we focus on the core sample, we initially present some results for the extended sample in our main tables, and also ascertain the robustness of all our findings to using the extended sample. We do indeed find that our results are robust to estimation on the extended sample.

Finally, we exclude respondents who were born abroad to American parents because it is difficult to categorize such individuals as either natives or immigrants (foreign born). In models that use country characteristics, we additionally exclude respondents who report being born in US territories or country aggregates. We also exclude respondents born in countries with low frequency and a high number of missing values in the data or countries with missing data on labor force participation in the source country.[16]

For selected waves of the June CPS, we know the total number of children ever born to each female respondent during the ACS sample period (2008, 2010 and 2012).[17] While these samples are of course much smaller than the ACS, they allow us to determine, for our core sample, the degree to which the number of children living in the household (our measure of fertility) accords with the number of children ever born to the female respondent. The data suggest that our sample restrictions lead to a sample for which these numbers are well matched. This may be seen in Appendix Table A-1, which shows the extent to which the number of children we assign to each woman using our sample restrictions matches the number of children ever born to that woman (based on the June CPS). The first thing to note is that our other sample restrictions do substantially contribute to the accuracy of the match, above and beyond restrictions on the age of the woman. In our married sample we correctly match 92% of the cases for both immigrants and natives, while in the sample of all women, we match 91% of the cases for both immigrants and natives. Subgroups such as Asian immigrants, Hispanic immigrants, and second generation immigrants are also well-matched, with match rates ranging from 90 to 94%.

Recall that, because fertility is endogenous, we follow Dahl and Moretti (2008) in focusing on the sex of the first child, rather than the sex composition of all children present. However, if couples practice sex selection on first births, the gender of the first birth would not be random. This is unlikely, given that, even for subgroups in which there is evidence of sex

[16] These countries include Antigua and Barbuda, Grenada, Bermuda, Micronesia, St. Kitts & Nevis, Marshall Islands, and Dominica. For CPS analyses, we also drop respondents born in countries not included in the 1995 list of countries. This restriction drops respondents born in Ivory Coast and Mongolia.
[17] This information is also available for 1995, 1998, 2000, 2002, 2004, and 2006.

selection, it has not been found to be present on the first birth (Almond and Edlund 2008), but it is important to confirm this. Moreover, several studies have found that the environment can affect the sex ratio at birth, results that may also call into question the assumption that the sex of the first child is indeed exogenous. For example, some researchers have found that conditions of stress lower the sex ratio (boys to girls) at birth. Examples of such conditions include alcoholism (Barreca and Page 2015), pollution (Sanders and Stoecker 2015), terrorist attacks (Catalano, Bruckner, Marks and Eskanazi 2006), earthquakes (Fukuda, Fukuda, Shimizu, and Møller 1998), and maternal anxiety disorders (Subbaraman et. al 2010). While these conditions may be extreme, there is other evidence of the impact of maternal circumstances on the gender of children even beyond this. Trivers-Willard (1973) hypothesized that natural selection would result in a relationship between parental resource status and the sex ratio at birth, with mothers in good condition having a higher ratio of sons. Norberg (2004) found that women who were living with a spouse or a partner before the child's conception or birth had a higher ratio of sons than those who were not. And, Almond and Edlund (2007) found that married women, younger women and more highly educated women were more likely to bear sons. These findings regarding marriage and living with a partner raise a possible concern about reverse causality, and we consider this issue below.[18]

In light of the possibility that sex selective abortion is a feasible option,[19] as well as research on the impact of maternal condition on the sex ratio at birth, we next examine data on

[18] There is also some mixed evidence on the impact of economic prosperity (a perhaps negative indicator of stress) on the sex ratio at birth, with Catalano and Bruckner (2005) finding that prosperity raised the incidence of boys in Sweden, while Fernández, et al (2011) found that a recession also raised the incidence of boys in Cuba.

[19] Although in principle, Assisted Reproductive Technologies (ART) can be used for nonmedical sex selection purposes, this potential method of sex selection is extremely rare. According to the Society for Assisted Reproductive Technology, a member organization that registers 95% of in vitro fertilization cycles in the United States (www.sart.org), 63,286 babies were born in the United States in 2013 using in vitro fertilization (http://www.sart.org/news/article.aspx?id=14570. In such cases, Pre-implantation Genetic Diagnosis (PGD)—the technology that in principle can be used for sex selection-- is used 4-6% of the time, and of these, 9% involved

the sex of the first child and examine the impact of our explanatory variables on the probability of a first child girl.

Appendix Table A2 presents the sex ratio (male to female) of the first birth for natives and immigrants for (i) all married women; (ii) married women in their first marriage; and (iii) all women regardless of marital status. (All of the other sample restrictions are retained.) For all three samples, the first child sex ratio (boys to girls) is well within the range suggested by Anderson and Ray (2010) as indicating a biologically-normal ratio, i.e., 1.03 to 1.07. The table shows the boy-to-girl ratio, and the 95% confidence interval endpoints.[20] In particular, the sex ratio ranges from 1.043 (for immigrants among all women) to 1.057 (for natives among women in their first marriage). The ratio is slightly lower for all women than for the married samples, although the differences across samples are not statistically significant. This slightly lower ratio of boys to girls among all women than among married women is consistent with the birth of a girl increasing the probability of a woman being a single parent (Lundberg 2005; Dahl and Moretti 2008).

While the overall incidence of male births is within biological norms, it is possible that variations in this incidence are related to maternal condition. Therefore, we examine the impact of our explanatory variables on the probability that the first child is a girl. The results are shown in Appendix Table A3. First, among natives, the only significant effect is that black women are more likely to bear a son, an effect that has been noted in previous literature and is believed to reflect biological differences (Anderson and Ray 2010). Second, among immigrants, the less

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nonmedical sex selection in 2005 (Baruch, Kaufman and Hudson 2008). These percentages imply that an upper bound of (0.06)·(0.09)·(63286)=342 babies born in the United States using nonmedical sex selection. This represents a miniscule fraction (0.00009) of the 3,912,181 births registered in the United States in 2013 (Martin et al. 2015).

[20] We calculate these endpoints based on the endpoints of the 95% confidence intervals for the means of the fraction of first children who are boys.

educated are more likely to bear daughters, an effect consistent with Almond and Edlund's (2007) findings, although we do not find a significant effect for natives. Moreover, among immigrants, there is regional variation in the probability of having a first-born daughter, and women migrating from high-fertility countries are less likely to bear daughters, an effect that is marginally significant. Importantly, however, the key gender equity index variable is not significantly related to the sex of the first child.

Appendix Table A3 shows that in most cases, our explanatory variables are not significantly related to the sex of the first child. Moreover, our explanatory variables control for some of the factors that previous research has found to influence the sex ratio at birth, including education, race, age, economic conditions (at least as proxied by the individual characteristics, region and year). Thus, even if the sex of the first child is endogenous, this effect is at least partially absorbed by the control variables. Further, while some individual coefficients are found to be significant, based on the F statistics reported in Table A3, we reject the hypothesis that the variables are jointly insignificant for only one of the three specifications (the regression for immigrants that includes source country characteristics). Nonetheless, we acknowledge that, if the sex of the first child is affected by unmeasured maternal conditions, then regressions that condition on this variable could yield biased results. Later in the paper, we perform a bounding exercise to ascertain the likely maximal impact of maternal conditions on the estimated effect of first child girl on living without a father and also consider whether the impact of maternal condition could account for the pattern of results reported in this paper. We conclude that this is very unlikely.

IV. Living Without a Father and Fertility: Aggregate Results

We begin our data analysis of son preference by presenting regression results for the determinants of living without a father and fertility at the aggregate level in order to characterize the United States as a whole, including both pooled results for the full population and for natives and immigrants separately. In the next sections, we consider educational differences in these outcomes for both natives and immigrants and cultural (source country) differences for immigrants and the second generation.

Table 1 shows native, immigrant and pooled results for the determinants of living without a father for the core and the extended samples. We find for both samples that having a first child girl significantly raises the probability of living without a father both overall and for natives separately, with effects ranging from 0.0031-0.0032 (1.1-1.2% of mean) for the core sample, and 0.0050-0.0056 (1.6-1.7% of the mean) for the extended sample. The higher point estimates that we obtain for the extended sample are quite similar to that reported by Dahl and Moretti (2008) who used a similar sample definition. For immigrants the effect is similar across the two samples (0.0023-0.0027 or 1.1-1.6% of the mean) but the point estimates are not significant.

As discussed above, there is a concern that maternal condition could result in reverse causality from marriage to sex of the first child, biasing our results. However, using existing evidence on the impact of marital status on the sex of a newborn, we can place an upper bound on any such bias. For example, Almond and Edlund (2007) find that being married increases the likelihood that a newborn is male by 0.001. Using the theoretical result in Basu (2015) and the coefficients on first child girl for the pooled samples in Table 1, we estimate that these coefficients (0.0032 for the core sample and 0.0050 for the extended sample) are biased upward by at most 0.0006 and 0.0007 respectively, or only 14-19% of the reported coefficient estimates.[21] We believe that these are overestimates of the bias because some of the apparent effect of marriage on child gender at birth found by Almond and Edlund (2007) is likely due to the greater likelihood of "shotgun marriages" if the child is a boy (Dahl and Moretti 2008), rather than the biological conditions of pregnancy.

Table 2 shows companion results for the impact of a female first child on total fertility among married women for the core and extended samples, again estimated for natives and immigrants separately and for both groups pooled. In Panels A-D, we show estimates for the all marriages and first marriages samples, while in Panels E and F, we show results for all women (including those who are not married spouse present).

Among married women (Panels A-D), the results are very similar across the core and extended samples and are not sensitive to the restriction to first marriages. First, among natives, having a first child girl has a small but significantly negative effect on fertility ("Total # of Children"), with the effects ranging from -0.0070 to -0.0078 or -0.4 percent of the mean fertility level. This significant negative effect shows up for both the probability of having two or more and three or more children. We also find negative point estimates for immigrants and natives pooled ("Both"), although the coefficient estimates are smaller in magnitude and are not statistically significant. These results strongly contrast with Dahl and Moretti's (2008) finding from the 1960-1980 Censuses that, among first marriages, having a female first child raises fertility by 0.3% of the mean.

[21] Basu (2015) shows that the OLS bias = $(Var(z)(1-r^2_{zx})^{-1}(a_2/(1-a_1a_2))\sigma^2$, where z is the first child girl dummy variable, r^2_{zx} is the squared correlation coefficient between z and the inner product of coefficients and variables other than first child girl from the equation for living without a father, a_2 is the effect of living without a father on first child girl, a_1 is the coefficient of interest, and σ^2 is the variance of the regression error. The figures in the text use the OLS estimates for a_1 and Almond and Edlund's (2007) estimate of 0.001 for a_2. Because the OLS bias on a_1 is positive, using the OLS a_1 in Basu's (2015) formula produces a slight overestimate of the OLS bias in this case.

Second, Table 2 (Panels A-D) also shows that, in contrast to the findings for natives, there does appear to be a fertility effect consistent with son preference among married immigrant women. The impact of having a female first child among immigrants is very similar across the core and extended samples and for all marriages and first marriages. It ranges from 0.015 (0.8% of the mean) to 0.018 (1.0% of the mean) and is highly significant in each case. These effects are larger than those obtained for the full population for the 1960-80 period examined by Dahl and Moretti (2008). As in Dahl and Moretti's estimates, this effect does not show up until beyond the margin of having two or more children, probably because having at least two children is so prevalent. To the extent that having a preference for boys characterizes values emphasizing traditional gender roles, the contrast between immigrants and natives shown in Table 2 suggests that, overall, immigrants have more traditional values than natives.[22] Below, we probe this possibility, explicitly examining the impact of source country characteristics on the effect of a female first child on living without a father and fertility among immigrant and second generation women.[23]

The results for all women shown in Table 2 (Panels E and F) are very similar to those for married women for both the core and extended samples, with first child girl having a significantly negative effect on native fertility and a significantly positive effect on immigrant fertility. Moreover, the estimated effects of first child girl for immigrants and natives pooled for all women are significantly negative, in contrast to the insignificant negative effects estimated

[22] Blau, Kahn and Papps (2011) found that immigrants had a more traditional division labor in the home than natives, as indicated by women's labor supply behavior, which reflected the lower female- to-male labor supply ratios in immigrant source countries compared to the United States.

[23] Not surprisingly, the impact of a first child girl tends to affect child spacing in the opposite direction as fertility. We found that a first child girl generally had a positive effect on child spacing for natives and a negative effect for immigrants in both all and first marriages (with the exception of a negative effect obtained for natives in first marriages in the extended sample). However, only the effect for natives in all marriages in the core sample was significant: 0.015 (se 0.008).

for married women. Also of interest, the estimated effects of first child girl for all women tend to be slightly smaller (more negative or less positive) than those estimated for the married samples. These smaller coefficient effects are as expected if first child girl increases the probability of a mother being unpartnered (single or divorced) and unpartnered women have lower fertility. This would result in a more negative effect of first child girl for natives and a less positive effect of first child girl for immigrants for all women than for their married counterparts. This is what we observe with the exception of the effect for native women in the extended sample, which is now slightly less negative than for married women, but still significant.

As noted earlier, first child gender may affect the mother's labor supply through its effect on fertility (Ichino, Lindström and Viviano 2014). In Appendix A4, we examine this for the core sample, showing results for the impact of first child girl on the mother's labor force participation for married women and for all women (regardless of marital status). The table shows that for native women, a female first child has positive effects on labor supply, while for immigrant women, the effects are negative. Although these effects are significant in only one case (natives among all women), they are fairly large in absolute value (1.1 to 1.7 times) relative to their standard errors. The estimated effects are consistent with the fertility patterns shown in Table 2. Specifically, a female first child lowers fertility and raises maternal labor supply for natives, while raising fertility and lowering maternal labor supply for immigrants. The effects are slightly smaller in absolute value for all women than for married women, although not significantly so. These results are as expected for immigrants but not for natives, given the positive effect of first child girl on living without a father for both groups and the consequent reduction in fertility. For immigrants, this works to reduce the negative effect of first child girl on women's labor supply for all women compared to married women. However, for natives, the positive effect of first child girl on women's labor force participation is a bit smaller for all women than for married women, which is counter to what we would have expected.

It is also of interest to know how the estimated effects of the impact of first child girl on family structure and fertility have changed over time so we can ascertain whether these possible manifestations of son preference have been increasing or decreasing. Fortunately, we are able to examine these trends using Dahl and Moretti's (2008) data for 1960-2000 combined with our own data for 2008-13. For the Dahl and Moretti data we use some results reported in their paper as well as own regression estimates based on their data (available at: http://econweb.ucsd.edu/~gdahl/sons-code.html) to obtain separate effects for each decade in their period. For comparability with Dahl and Moretti for the most recent period, we use our extended sample for immigrants and natives pooled and adopt their specification.[24] Our results are shown in Figures 1-3.

In Figure 1, we chart the trends in the impact of first child girl on living without a father. In addition, following Dahl and Moretti (2008), in Figure 2, we decompose the effect of first child girl on living without a father into its component channels. Specifically, a child can be living without a father if i) the mother has never married; or ii) there has been a divorce; or iii) the mother has custody of the children in the event of divorce.[25] Two overall conclusions arise from the Figures. First, while the impact of first child girl on the probability of living without a father rose steadily from about 0.2 to 0.8 percentage points between 1960 and 1990, it then fell sharply to about 0.5 percentage points by 2000 and has remained at roughly that level since then.

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[24] Dahl and Moretti control for decade of birth fixed effects, three (rather than four) categories of education (<HS. HS and College), and do not have a separate indicator for Hispanics. A further difference for the fertility regressions is that Dahl and Moretti do not control for spouse characteristics and include women who are married spouse absent; they also determine first marriage based solely on the wife's marital history (rather than both the husband's and wife's as we do). Our results were virtually unaffected by the slight differences in specification.

[25] For details on this decomposition, see the discussion in Dahl and Moretti (2008), p. 1088.

Second, the relative importance of the never married channel to the impact of first child girl on living without a father has grown, increasing particularly sharply between 2000 and 2008-2013. There is also, as noted by Dahl and Moretti (2008) a noticeable growth in the importance of the custody channel relative to the divorce channel compared to the earliest years (1960 and 1970).

The results for fertility are shown in Figure 3 for all married women and, where available, women in their first marriages. Figure 3 indicates that, between 1960 and 1980, the effect of first child girl was positive for both first marriages and all marriages. These effects were significant for both samples for the 1960-80 period pooled and for 1980 separately. For 1990 and 2000, there is no information on first marriages; however, the effect of a female first child on fertility for all marriages falls to virtually zero (-0.0002) by 1990 and further to -0.0034 by 2000, with both effects insignificant. In our data, the effects for first and all marriages are both about -0.003 and insignificant, or roughly the same magnitude as for 2000's sample of all marriages. These results strongly suggest that the disappearance of a positive effect of first-child girl on fertility reflects a fairly permanent shift. And, as we have seen, for natives, this effect is now significantly negative. One additional piece of data lends support to the conclusion that these are relatively long-term shifts. Below, we examine similar specifications for the 1995-2014 CPS, in order to see results separately by immigrant generation. For natives with native parents (3rd+ generation)—who comprise the majority of natives—our results are similar to those obtained for natives in Table 2: a female first child leads to lower fertility.

V. Living Without a Father and Fertility: Heterogeneity by Education

We now explore whether the aggregate findings for living without a father and fertility vary by education. Figures 4 and 5 show the results for the impact of a first child girl on living without a father (Figure 4) and fertility (Figure 5) disaggregated by education group. For natives, with the exception of high school graduates, results are broadly consistent with the aggregate results presented in Tables 1 and 2, although the magnitude and significance of the estimated coefficients varies. Except for those with exactly a high school degree (high school graduates), having a female first child is found to have a positive effect on the probability of living without a father and a negative effect on fertility. The results for high school graduates, however, show a pattern more consistent with son preference on both dimensions. The impact of a female first child on living without a father is larger for high school graduates than for the other education groups, although not significantly so, while the effect on fertility is large, positive and significant. The more traditional patterns that we find for the high school educated compared to the more highly educated are consistent with findings in the literature that the more highly educated are less likely to have traditional gender role attitudes.[26]

For immigrants, the results for living without a father are mixed. The effect of a female first child on living without a father is positive for high school and below and significant or larger than its standard error. However, the impact for those with more than high school is negative in both cases, although not significant. These mixed results by educational attainment contribute to the weaker results for the impact of first child female on living without a father for immigrants than natives in the aggregate, which were obtained in Table 1. For fertility, the impact of a female first child is more consistent with the effect being positive for each education group. Interestingly, the estimated effect is especially large (and significant) for those with a high school degree.

[26] See, for example, Campbell and Horowitz (2016); and Davis and Greenstein (2009).

VI. Source Country Characteristics and First Child Girl Effects for Immigrants and the Second Generation

Overall, we do not find direct evidence of son preference for fertility for natives or in the aggregate, although the results for living without a father are possibly consistent with this phenomenon. However, immigrants, on average, do exhibit son preference in their fertility behavior. This potentially provides supporting evidence for the (weaker) findings of female headship among immigrants and might indicate son preference along this dimension as well. If the immigrant results for living without a father and fertility reflect son preference, they would likely be tied to source country variables reflecting gender equity in these countries. We thus examine these relationships for both immigrants and second generation natives—evidence of such an association would provide further support for interpreting the estimated effects as indicators of son preference.

Table 3 presents results for living without a father and fertility for immigrants.[27] The key results concern the interactions between first child girl and indicators of gender equity in the source country; as noted above, all regressions additionally control for total fertility and the log of GDP per capita in the source country. In Columns 1 and 4, we summarize women's status using an Equity Index (based on the World Economic Forum's Global Gender Gap Index). In additional specifications, we investigate the separate impact of some important gender-related source country characteristics by replacing the Equity Index with either (i) the female labor force participation ratio (the rate relative to that of men) (Columns 2 and 5) or (ii) the female labor force participation ratio and the country's sex ratio at birth (boys/girls) (Columns 3 and 6). We

[27] In the fertility regressions, we show results for all married couples but results were similar when immigrants were restricted to their first marriage.

find strong evidence of an effect of source country gender equity on fertility but relatively little evidence of such an effect for living without a father.

Looking first at the results for the probability of living without a father shown in Columns 1-3, we see that the effect of first child girl is weaker for women migrating from countries with higher Equity Index scores; the interaction effect is not significant but is larger than its standard error in absolute value. However, we do not obtain significant results for the labor force participation ratio in either specification or for the sex ratio (the latter is also "wrong signed"); and these estimated effects are small relative to their standard errors.

In contrast, the results for fertility shown in Columns 4-6 provide strong evidence of a link between these source country characteristics and immigrant fertility behavior in the United States that is robust across all of our alternative measures of gender equity in the source country. Column 4 shows a significant negative interaction between first child girl and the Equity Index; Column 5 shows a significant negative interaction between first child girl and the female relative LFP rate; and Column 6 shows a significant negative interaction between first child girl and the female relative LFP rate and a significant positive interaction between first child girl and the sex ratio at birth. Thus, in all cases we find that the apparent preference for boys is stronger among immigrants coming from societies with lower gender equity.

To illustrate the magnitude of the source country effects, we evaluate the interaction effect between first child girl and the Equity Index shown in Column 4 using the sample distribution of the Equity Index. Specifically, we contrast the effect of first child girl on the fertility of women migrating from a country at the 75th percentile of the Equity Index in our immigrant sample (0.6796) with that of women migrating from a country at the 25th percentile

(0.6459).[28] As examples, Thailand has an index near the 75th percentile, and Mexico's index is near the 25th percentile. In contrast, the Equity Index for the United States is 0.7022, or above the 75th percentile for immigrants, providing further evidence that gender roles are more traditional on average in immigrant source countries than in the United States. We find that the impact of first child girl on fertility is relatively large and statistically significant for women coming from a country at the 25th percentile of the Equity Index (0.0183, se 0.0050) and smaller and not statistically significant for women coming from a country at the 75th percentile (0.0092, se 0.0077). This simulation shows that the impact of first child girl on fertility of women coming from a country where women have lower gender equity is large and highly statistically significant, while, for women coming from countries where women have higher gender equity, it is small and not statistically significant.[29]

The results for immigrant fertility are consistent with son preference in the aggregate for this group, as well as the importance of source country characteristics or culture in affecting the degree of son preference. An interesting question is whether immigrant preferences for sons and the cultural differences implied in the impact of source country characteristics persist into future generations. To address this question, we take advantage of information on parental birthplace for respondents in the CPS, which, in conjunction with information on where they themselves were born, allows us to study son preference separately for the foreign born (the 1st generation), natives with either parent foreign born (the 2nd generation) and natives with native-born parents (the 3rd+ generation). Our results are shown in Tables 4 and 5.

Table 4 shows the overall effects of first child girl on living without a father and fertility for each immigrant generation. These results can give an indication of whether the CPS is

[28] These percentiles are implicitly weighted by the (weighted) frequency of immigrants from each source country.
[29] A similar exercise using the results of Columns 5 and 6 yields a similar conclusion.

yielding broadly similar overall results to the ACS and also provide our first look at the 2nd generation before we consider finer distinctions among them based on parental source country. In terms of the signs of the effects, the CPS results are broadly similar to those obtained from the ACS. For the probability of living without a father, we find a positive effect of first child girl for both the 3rd + generation—the majority of natives—and the 1st generation, although the estimated effects are not statistically significant. For fertility, as in the ACS, a female first child leads to significantly lower fertility for the 3rd + generation and significantly higher fertility for the 1st generation.[30] The results for the 2nd generation are a bit anomalous. On the one hand, we find that first child girl raises fertility, although the effect just misses statistical significance at conventional levels, suggesting that the 2nd generation is a traditional group on average. On the other hand, in contrast to the other groups, first child girl has a negative effect on the probability of living without a father, although this effect is not statistically significant.

Regardless of overall average effects for the second generation, the more interesting question is how their behavior tracks with parental source country gender equity. This is shown in Table 5. Looking first at the results for fertility, where we had stronger results for the immigrant generation, we see that our findings for the second generation are qualitatively quite similar to those we obtained for immigrants. Specifically, those whose parent(s) came from a country with a higher Equity Index exhibit significantly less son preference in fertility than those whose parent(s) came from a country with a lower Equity Index. Further, the interaction effect for Girl*LFP Ratio is significantly negative (Columns 2 and 3), while the interaction with Sex

[30] Although the fertility results for immigrants and natives in the CPS are similar in sign to the results we obtained with the ACS, the magnitudes of the estimated effects are larger in the CPS than in the ACS. This was not due to our wider time window in the CPS analysis (1995-2014): when we restricted the CPS analysis to the same years as the ACS—2008-2013, the effects for immigrants and 3rd+ generation natives remained larger than in the ACS.

Ratio at Birth is positive, though not statistically significant (Column 3).[31] Overall, the results suggest that cultural transmission of son preference in fertility from source country to immigrants continues into the second generation.

Looking next at the impact of parental source country variables on living without a father, perhaps not surprisingly given the weak evidence we found for the immigrant generation, the second generation results offer little evidence in the expected direction. We do find a positive and significant interaction between first child girl and sex ratio at birth; however, the Equity Index and Labor Force Participation Ratios interactions with First Child Girl are "wrong signed," and insignificant.

VII. Sex Selection

Our findings for fertility do not suggest son preference overall on the part of natives, while we do find some evidence for immigrants that having a girl raises fertility, particularly for immigrants from countries with lower gender equity, and second generation individuals whose parents came from such countries. However, sex selection can serve as another, perhaps substitute option for exercising son preference, and, nearly a decade ago, Dahl and Moretti (2008) warned that the United States might see increases in such behavior due to technological advances in sex selection technology (p. 1087). We therefore have examined this issue by studying the impact of first child sex on the sex of the second child among those with at least two children, and the impact of the sex composition of the first two children on the sex of the third child among those with at least three children.

[31] As was the case with the overall effects for immigrants and natives in Table 4, the interactions for second generation women shown in Table 5 are larger in magnitude for the Equity Index and LFP Ratio than for immigrants shown in Table 3; however, the effect for Girl*Sex Ratio at Birth is smaller for second generation women than for immigrants.

Table 6 shows these results. For both natives and immigrants, we find no evidence of son preference through sex selection. For natives, the boy/girl ratios are all between 1.034 and 1.058, within the normal biological range of 1.03-1.07 suggested by Anderson and Ray (2010). Moreover, the probability that the second or third child is a boy is never higher for all girl families than for all boy families. What associations there are between the sex composition of past and future births for natives appear to be consistent with a biological tendency of future children to be of the same sex as previous children. [32] Among immigrants, in all cases, the 95% confidence interval for the sex ratio is within the biologically normal 1.03-1.07 range. Moreover, as was the case for natives, the probability that the second or third child is a boy is never higher for all girl than for all boy families. These results for immigrants, like the ones for natives, do not suggest sex selection.

While some researchers have found evidence of sex selection for specific groups such as women of Chinese, Indian or Korean heritage (Almond and Edlund 2008; Abrevaya 2009; Almond, Edlund and Milligan 2013; Persaud, Kalantry, Citro, and Nandi 2015),[33] the results of Table 6 suggest that these instances are not numerous enough to characterize the full population of immigrants or natives. Therefore, the disappearance of the effect of a first girl on fertility for the general population and the reversal of the effect of a first girl on fertility that we found for natives likely indicate a general weakening of son preference in the United States rather than a shift to sex selective abortion as an alternative mechanism.

VIII. Discussion

[32] Some researchers have found such a correlation between the sex composition of previous births and the sex of future children (Ben-Porath and Welch 1976; Gellatly 2009), although some find no such pattern (Rodgers and Doughty 2001; Jacobsen, Møller and Mouritsen 1999).

[33] Interestingly, while most authors found evidence of son preference, the recent paper by Persaud, Kalantry, Citro, and Nandi (2015) reports results consistent with a demand for diversity.

We find that effect of first child girl now has a negative effect on fertility in the aggregate and is significantly negative for natives. Our findings for fertility contrast with those of Dahl and Moretti (2008) for an earlier period but are consistent with those reported in Ichino, et. al (2014), Andersson, et. al (2011), and Abrevaya (2009). Our results remove a crucial piece of supporting evidence for the current period that the positive effect of a first girl on the probability of living without a father is due to son preference as opposed to other factors.

Looking at trends in the effects of first child girl, we found that the positive impact of first child girl on living without a father has declined since 1990, while, since 1980, the impact of first child girl on fertility has fallen from a positive effect in the earlier period to the insignificantly negative effect in the aggregate and the significantly negative effect for natives that we find for the current period. This combination of results suggests that son preference has declined in the aggregate and particularly for natives.

Do parents in the aggregate and natives in particular now prefer daughters? This might be the case but we note that there are alternative explanations for our findings than daughter preference. One factor that could contribute would be a rise in the costs of raising girls. This is plausible in that, today, the majority of college students are female, in contrast to the higher male representation during the 1960-80 period (Goldin, Katz and Kuziemko 2006). Thus, it is possible that families who have girls today are more likely to anticipate higher college costs for their children than in the past. Kornrich and Furstenberg (2013) provide some direct evidence on the relative cost of raising girls and boys. They find that, in 1972-3, households with all boys spent significantly more on their children than households with all girls; this gap was largely accounted for by educational expenses. However, by 2006-7, this pattern had reversed, with all girl households spending significantly more than all-boy households, particularly, again, on education (p. 16).[34]

The cost data are consistent with the pattern of our results for the current period in the aggregate and especially for natives that a female first child both lowers fertility and raises the probability of living without a father. They are also broadly consistent with Dahl and Morreti's (2008) findings suggesting son preference, at least for the pre-1980 period (see Figures 1-3). Since girls were less expensive than boys, costs cannot account for the positive effect of female first child on living without a father, although the lower costs of girls do provide an alternative interpretation for the positive effect of first child girl on fertility. Moreover, the increase in the cost of girls is consistent with the decline in the effect of female first child on fertility since 1980 (Figure 1). However, despite these rising costs, the effect of first child girl on living without a father has also decreased since 1990, falling to below the 1980 level and only slightly higher than the 1970 level. This change could not have been driven by costs and thus suggests a decline in son preference likely played a role. Moreover, the spending data could themselves reflect a reduction in son preference, since they indicate that families are now willing to spend more on girls, particularly their education.

The increased expenditure on girls and the less positive effect of girls on fertility than in the past could also reflect an increase in bargaining power of wives as their labor force participation and relative wages have increased. Some data suggest aggregate preferences for the sex of a child have not changed and that son preference appears to be a male phenomenon—that is, that while men have a preference for boys over girls, women show no preference either way

[34] In their online appendix (available at: https://link.springer.com/article/10.1007%2Fs13524-012-0146-4), the authors show that these patterns hold up in a regression context and for mixed gender families.

(Newport 2011).[35] Thus, the disappearance of a positive effect of a first child girl in the aggregate and for natives could indicate that women have a greater say in this decision than previously. This is plausible in light of rising female labor force participation and relative wages (Blau and Kahn 2017). Another possible factor potentially influencing fertility and family structure (single female parenthood) relates to the economic returns to boys relative to girls. While in principle these returns could have been increasing with the rise in female labor force participation and the decline in the gender pay gap, it seems unlikely that material returns play a major role in a developed country like the U.S., where children do not have an important role in supporting their parents economically.[36] However, increases in life expectancy might result in a greater value being placed on the future caretaker role of daughters than in the past, since women shoulder a disproportionate share of elder care (Grigoryeva 2017).

The findings for 1960-80 suggest that preferences for boys outweighed such considerations (if any) during that period. Our findings imply either a reduction among natives in preferences for boys and/or an increase in the impact of these other factors sufficient to outweigh any preference for boys. The possible impact of these various factors makes us reluctant to interpret our finding of a negative effect of first female child on native fertility as indicating a shift from son to daughter preference, although of course it might.

Maternal condition is another factor potentially influencing single female parenthood (living without a father). As we have seen, a reasonable bounding exercise suggests that the impact of this factor, if any, is not large. Here we consider whether the impact of unmeasured

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[35] Specifically, in 2011, when asked about sex preference supposing that one could have only one child, men preferred a boy to a girl by a margin of 49 to 22 percent, whereas women were split roughly equally with 31 percent preferring a boy and 33 percent preferring a girl. (A higher proportion of women (36 percent) than men (28 percent) also said responded "Doesn't matter," "Not sure," or "No opinion." See Newport (2011).

[36] Lundberg (2005) notes that differences in material returns are not expected to play a large role in developed ("wealthy") countries, although, this factor receives considerable emphasis in analyses of son preference in developing countries.

maternal condition could be consistent with the pattern of our findings. In considering this possibility, it is useful to specify the possible effects. For living without a father, the concern would be that an unmeasured factor like stress on the mother raises both the probability of a female first child and the probability of her being unpartnered, since stress is likely to contribute to the breakup of the couple or the failure of the parents to form a family in the first place. For fertility, the expectation would be that unmeasured maternal condition would raise the probability of a female first child but lower fertility. This is based on evidence that maternal stress lowers one's own fertility (Louis, et. al 2011) and the fertility of one's daughters (Plana-Ripoli, et al (2016). Moreover, health problems associated with being over- or underweight, as well as the excessive intake of caffeine, tobacco and alcohol have been found to also reduce fertility (Ruder, Hartman and Goldman 2009).

The pattern of our findings does not fit what would be expected based on the impact of unmeasured maternal condition. A potential correlation between having a first-born girl and poor maternal condition could help explain the positive effect of first-born girls on living without a father and its negative effect on fertility that we find for natives. However, it cannot explain our finding for immigrants that a first-born daughter *raises* fertility as well as the probability of living without a father. Moreover, when we look at the trends in the impact of a first-born girl on living without a father and fertility for the full population, changes in maternal condition do not appear to be a plausible explanation for the pattern. Specifically, the positive impact of first child girl on living without a father has declined since 1990. For changing maternal conditions to explain this development would suggest an improvement in maternal conditions. However, since 1980, the impact of first child girl on fertility has decreased. For this change to be due to maternal conditions, would suggest maternal stress and maternal conditions have worsened.

Further, our findings for source country variables reflecting gender equity for immigrants also cast doubt on the idea that maternal conditions could explain our results. Specifically, we might expect that women migrating from countries with lower status of women would find adapting to US society more stressful than those migrating from societies with more equal status of women. Nonetheless, a female first child has more pronounced positive effects on future fertility for immigrant women from source countries with less gender equality, again, seemingly in contrast to what one might expect from the medical literature.

IX. Conclusion

In this paper, we have used 2008-2013 ACS and 1995-2014 CPS data to generate new findings on the extent of son preference in the United States. In light of the large increase in immigration and the changes in immigrant source countries towards countries with a more traditional status for women than in the United States (Blau, Kahn, and Papps 2011), we introduce a new dimension into this literature by analyzing natives and immigrants separately. Perhaps most importantly, we find that, among native women, as well as among the aggregate population (immigrants and natives pooled), having a female first child reduces future fertility (significant effects for natives and insignificant in the aggregate). This result stands in sharp contrast to earlier research by Dahl and Moretti (2008) which found for the 1960-80 period that having a girl led to higher fertility levels among the aggregate population. As with their earlier work, we do continue to find that first child girl increased the likelihood of living without a father. However, our fertility findings cast doubt on son preference as the explanation for this relationship.

Our findings for fertility and living without a father are more consistent with the hypothesis that girls are more expensive to raise, or that boys especially benefit from having a father living with them than with the hypothesis of son preference. We presented data from Kornrich and Furstenberg (2013) that showed that, indeed, raising girls, which was cheaper in the 1970s than raising boys, had by the 2000s become more expensive than raising boys. Other factors may have further contributed to the disappearance of son preference for natives. Some evidence suggests that women do not share men's preferences for sons (Newport 2011). Thus, the fertility changes may reflect an increase in women's bargaining power in the family, perhaps due to rising female labor force participation and relative wages. Another possibility is that, with rising life expectancy, parents have come to more highly value the caretaker roles daughters disproportionately shoulder. These various shifts, in conjunction with the rising costs of girls, may have reversed or outweighed son preference in fertility for this group.

For immigrants, we also find a positive effect of a female first birth on living without a father, although it is not statistically significant. In contrast to natives, however, we find that, for immigrants, having a first child girl significantly raises future fertility, providing direct evidence consistent with son preference for fertility for this group. Moreover, such fertility preferences are stronger for immigrants coming from countries with lower gender equity and also appear to carry over into the second generation. In contrast, we found little evidence that the impact of having a girl on living without a father was stronger among first and second generation immigrants from source countries with lower gender equity. Thus, in contrast to the fertility results, we do not provide support that the relationship between first child girl and living without a father for immigrants or the second generation is tied to son preference.

We also studied the issue of sex selection, a perhaps extreme manifestation of son preference. Overall, despite warnings that sex selection could spread among the wider population (Dahl and Moretti 2008), we found no evidence of such behavior in the aggregate native and immigrant populations for our relatively recent period of analysis. The findings for sex selection reinforce our overall conclusion that preference for sons appears to have diminished among US natives in that sex selection does not provide an alternative mechanism to account for the disappearance of a positive effect of first child girl on fertility.

Data Appendix

Variable Definitions

Variables from the ACS and CPS

Race and Ethnicity

- We control for race and ethnicity using a set of indicator variables for five mutually-exclusive categories: White non-Hispanic, Black non-Hispanic, Hispanic, Asian non-Hispanic, and other non-Hispanic.
- Respondent is classified as Hispanic if the respondent reports being Hispanic or reports race as Spanish, Portuguese, Mexican, Puerto Rican, Latin American Indian, South American Indian, or Mexican American Indian.
- Respondent is classified as black non-Hispanic if the respondent reports being any detailed race that includes black (except for Black and Chinese, Black and Asian Indian, or Black and Korean) and is not classified as Hispanic.
- Respondent is classified as Asian non-Hispanic if the respondent is not classified as Hispanic or black non-Hispanic and reports race as Asian or any mixed race including Asian.
- Respondent is classified as white non-Hispanic if the respondent is not classified as Hispanic, black non-Hispanic, or Asian non-Hispanic and reports race as white.
- Respondent is classified as other non-Hispanic if none of the above classifications apply.

Immigrant Status and Years Since Migration

- Respondents are classified as natives if their birthplace is one of the fifty states or the District of Columbia.
- For foreign-born persons and persons born in outlying U.S. areas, we define years since migration as the lesser of age or reported years in the United States.

First, Second, Third + Generation (CPS only)

- Respondents are classified as 1st generation if they report their birthplace as outside the fifty states or the District of Columbia.
- Respondents are classified as 3rd + generation if they report that they and both of their parents were born in were born in the fifty states or the District of Columbia.
- Respondents are classified as 2nd generation if they were born in the fifty states or the District of Columbia and they report that either of their parents was born outside the United States. Parental source country characteristics are allocated based on mother's birthplace, if she is foreign born, and father's birthplace otherwise.

Living Without a Father

- We classify a respondent as living without a father if the respondent is female, unmarried (where married, spouse absent is considered married), and has at least one child. (The oldest child must be 12 years of age or younger for sample inclusion.) In the core sample, we include only women who meet these requirements and are listed as head of household. Note that if the respondent has an unmarried partner present, she can still be classified as a female head of household.

- Single fathers are included in the extended sample. Single fathers are included and given a 0 for living without a fatheronly if their children do not have a mother in the household and if they are ever married (i.e., never married men are excluded).

First Marriage

- Respondents are classified as being in a first marriage if both the respondent and her spouse report that their current marriage is their first marriage.

Country Characteristics Variables

Total Fertility

Total fertility data comes from the World Bank, available at http://data.worldbank.org/indicator/SP.DYN.TFRT.IN. In the regressions with country characteristics, we include 2000-2007 country averages of total fertility.

GDP Per Capita

Most GDP per capita data comes from the World Bank, available at http://data.worldbank.org/indicator/NY.GDP.PCAP.PP.KD. For Taiwan, data comes from the Chinese Statistical Yearbook 2013, available at http://ebook.dgbas.gov.tw/public/Data/3117141132EDNZ45LR.pdf. GDP for Argentina, Burma and Syria is constructed from UN Stats data on GDP by Type of Expenditure at current prices and at constant 2005 prices in national currency units, available at http://data.un.org/Data.aspx?d=SNAAMA&f=grID%3A101%3BcurrID%3ANCU%3BpcFlag%3A0 and http://data.un.org/Data.aspx?q=gdp&d=SNAAMA&f=grID%3a102%3bcurrID%3aNCU%3bpcFlag%3a0 , respectively. PPP conversion rates come from http://icp.worldbank.org/icp/QueryResults.aspx?r=-1&ds=0&y=3&ws=1. We use the World Bank methodology to convert to GDP per capita, PPP. In the regressions with country characteristics, we include the natural log of 2000-2007 country averages of GDP per capita.

Ratio of Female to Male Labor Force Participation

Data on male and female labor force participation come from the International Labor Organization's Key Indicators of the Labor Market. We use labor force participation for the population 15 years of age and older. In the regressions with country characteristics, we include 2000-2007 country averages of the ratio of female to male labor force participation.

Sex Ratio at Birth

Sex ratio at birth comes from UN Data, available at http://data.un.org/Data.aspx?q=sex+ratio+at+birth&d=PopDiv&f=variableID%3a52. We follow the WEF in censoring the sex ratio at birth at 1.059 to identify son preference. In the regressions with country characteristics, we include 2000-2007 country averages of sex ratio at birth.

Equity Index

The equity index is based on the World Economic Forum's Global Gender Gap Index from "The Global Gender Gap Report, 2012," available at http://www3.weforum.org/docs/WEF_GenderGap_Report_2012.pdf. In the regressions with country characteristics, we include 2006-2007 country averages of the index, unless a 2006 value is not available, in which case we use the earliest value available up until 2012. (Note that the index first became available in 2006.)

Sample Selection and Weighting

Unless otherwise noted, analyses with the American Community Survey (ACS) use data from the 2008-2013 waves and analyses with the Current Population Survey (CPS) use data from the 1995-2014 March CPS. Regressions for the core sample are weighted by household weights that are normalized to provide equal weighting for each sample year; regressions for the extended sample are weighted by person weights (as suggested by IPUMS when one is analyzing members of subfamilies: https://usa.ipums.org/usa/volii/subfamilies.shtml , accessed 9/10/19) that are normalized to provide equal weighting for each sample year.

Our core sample includes women between the ages of 18 and 40, who are the head of household or spouse of the household head, with one or more children, where the oldest child is twelve years old or younger and all children are born in the U.S. Households with adopted, step or foster children are dropped in the ACS. We unable to identify step or adopted children in the CPS, but we are able to drop CPS households with foster children. Same-sex couples, respondents living in group quarters, respondents born abroad to American parents, widows, as well as mothers with multiple children born in the same year and quarter (ACS) or same year (CPS) are excluded. When the dependent variable is fertility, the sample is additionally limited to married women with a spouse present, and, in some specifications, to women in their first marriage who are married to men also in their first marriage.

The extended sample expands the core sample by including father-only families and parents who are not the household head or spouse of the head (i.e., in subfamilies). Men are included in the sample only if their children do not have a mother in the household and if they are ever married (i.e., never married men are excluded). We also expand the sample to include step and adopted children since we are not able to identify these categories of children for subfamilies. We continue to exclude foster children, but, in the spirit of inclusiveness, not their households.

In analyses that include country characteristics, we exclude respondents who report being born in US territories or country aggregates. We also exclude respondents born in countries with low frequency and a high number of missing values in the data or countries with missing data on labor force participation. These countries include Antigua and Barbuda, Grenada, Bermuda, Micronesia, St. Kitts & Nevis, Marshall Islands, and Dominica. For CPS analyses, we also drop respondents born in countries not included in the 1995 list of countries. This restriction drops respondents born in Ivory Coast and Mongolia.

Conflict of Interest: The authors declare that they have no conflict of interest.

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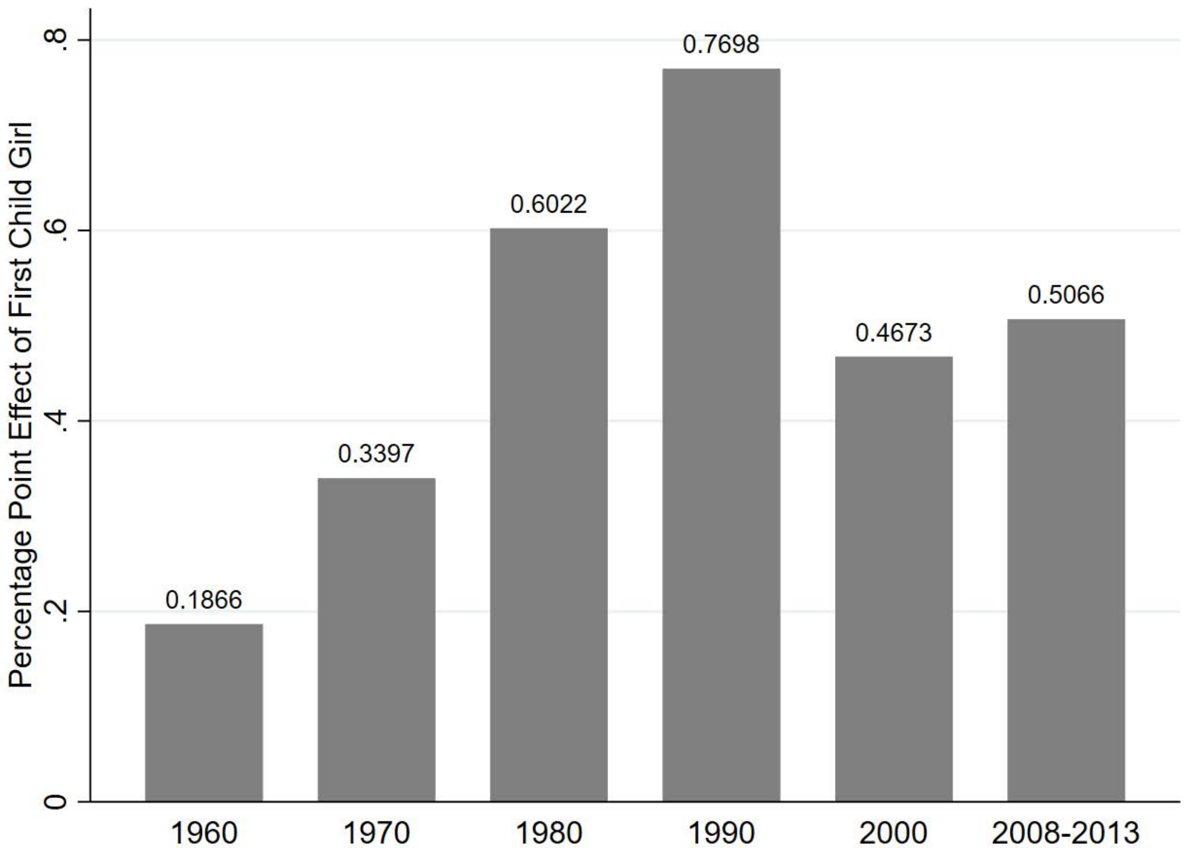
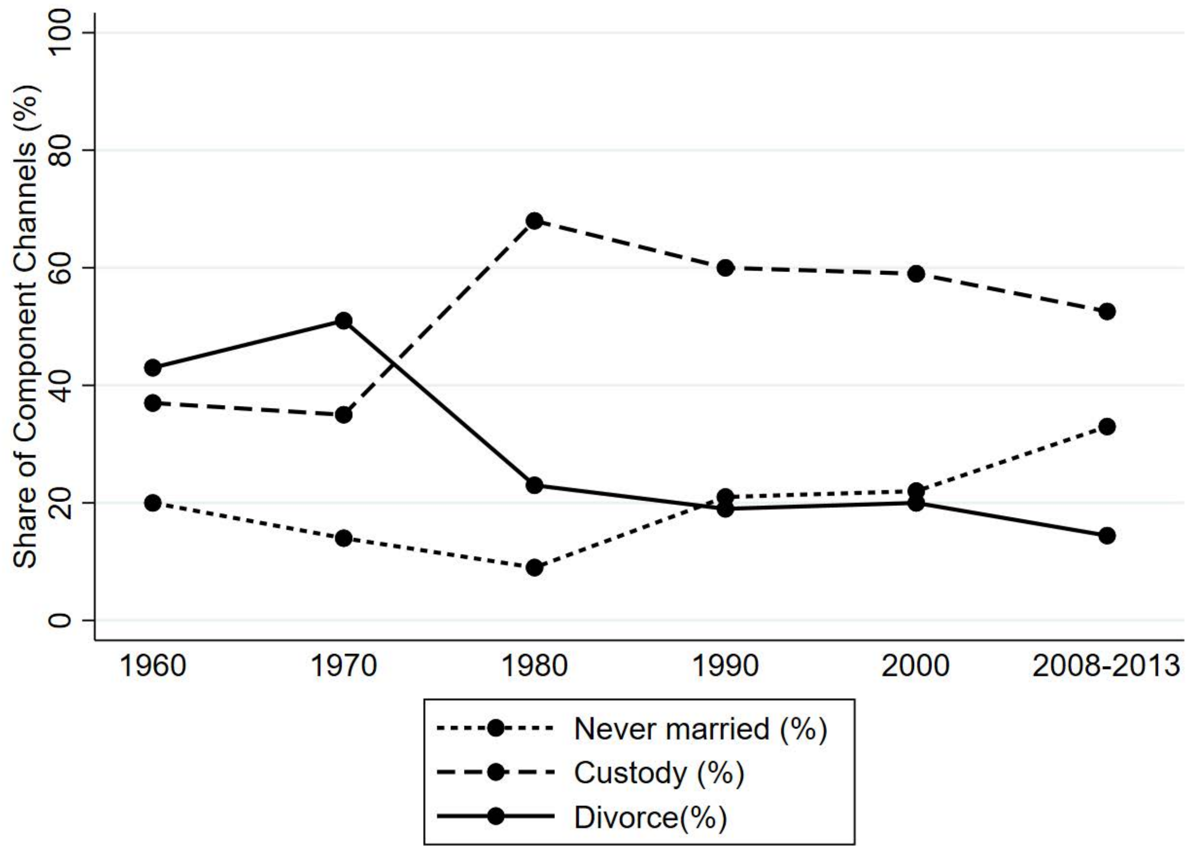
Figure 1: Effect of First Child Girl on the Probability of Living Without a Father by Year (Percentage Points)



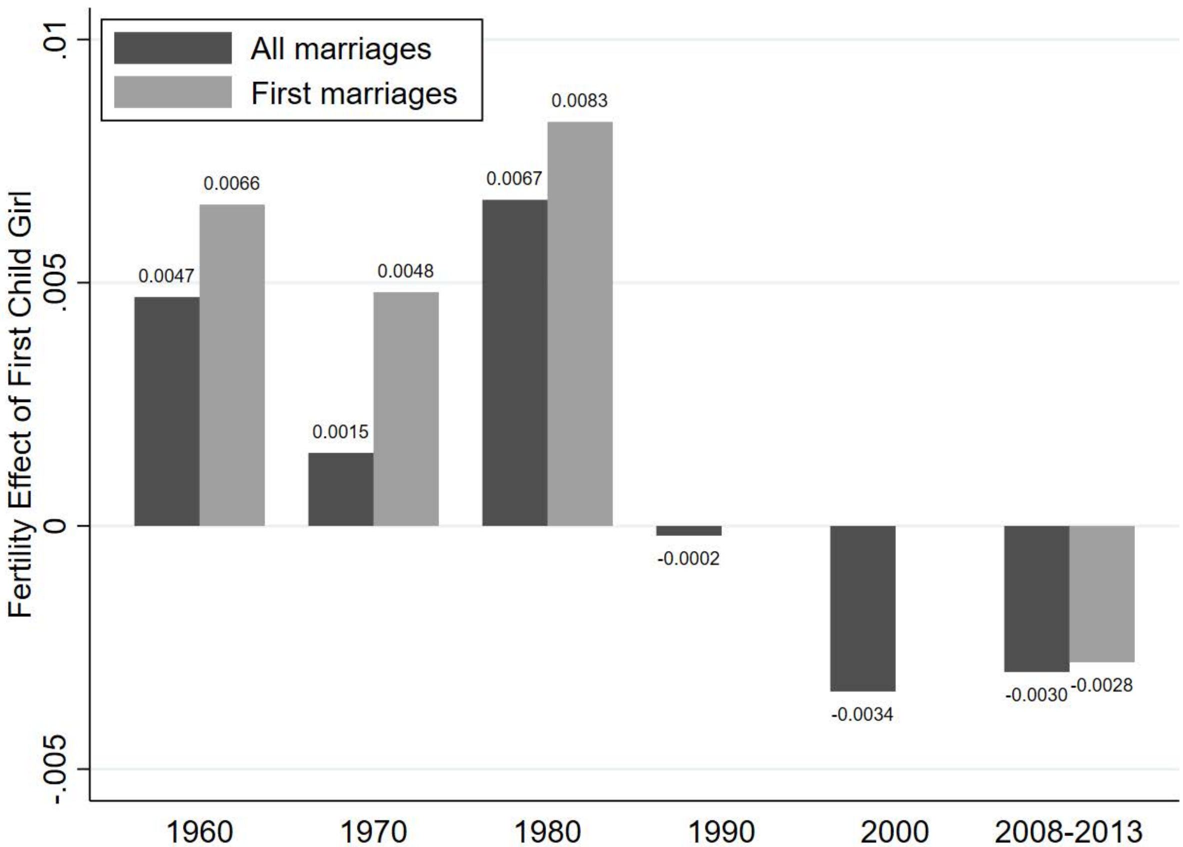
Notes: Estimates for 1960-2000 are based on DM (2008) Figure 1 and Table A1, with additional calculations using their data, available at https://econweb.ucsd.edu/ ~gdahl/sons-code.html. Estimates for 2008-2013 are calculated using the American Community Survey with the Dahl and Moretti (2008) sample restrictions and specification.

Figure 2: Share of Each Component in Accounting for the Effect of First Child Girl on the Prob. of Living without a Father, by Year (%)



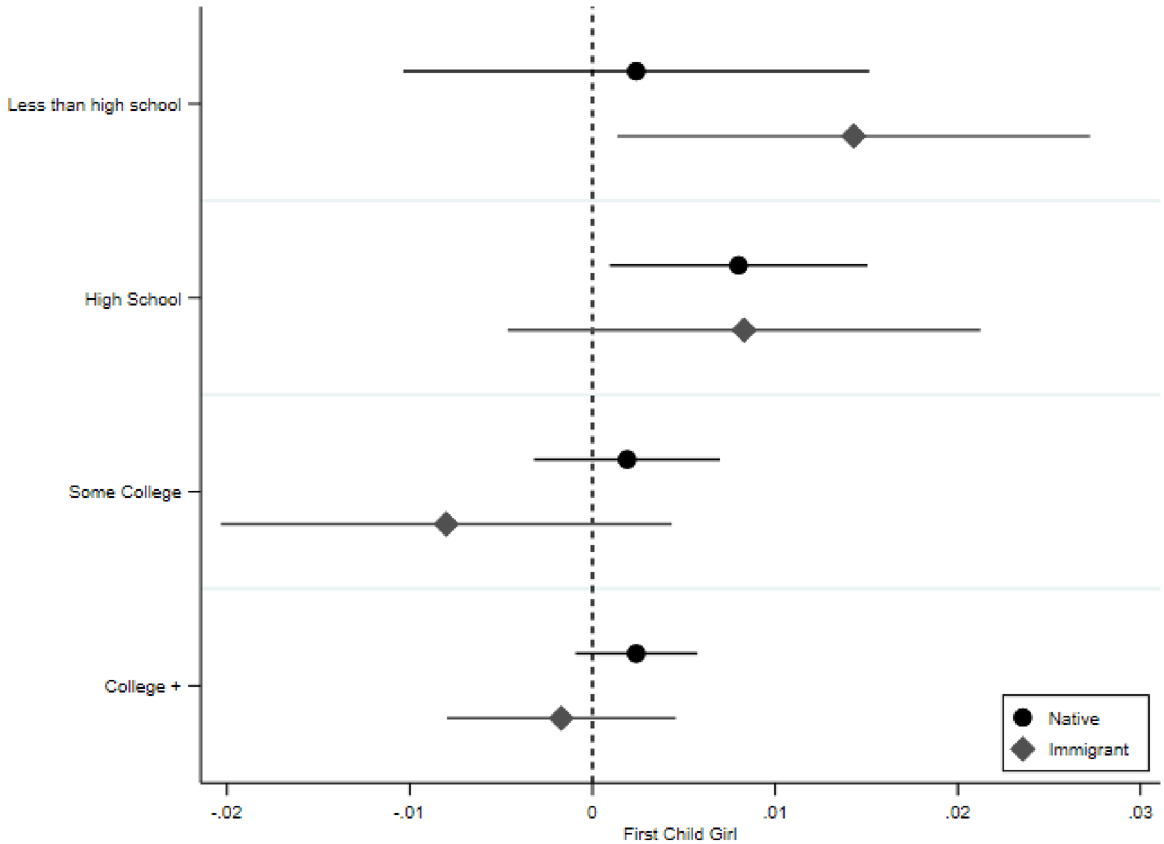
Notes: Decomposition of the overall effect of first child girl on the probability of living without a father into component channels: never married, divorce, and custody. For example, the figure indicates that, in 1960, 20% of the impact of first child girl on living without a father was due to its effect on the mother being never married. Estimates for 1960-2000 are based on DM (2008) Figure 1. Estimates for 2008-2013 are calculated using the American Community Survey with the Dahl and Moretti (2008) sample restrictions and specification.

Figure 3: Effect of First Child Girl on Fertility by Year



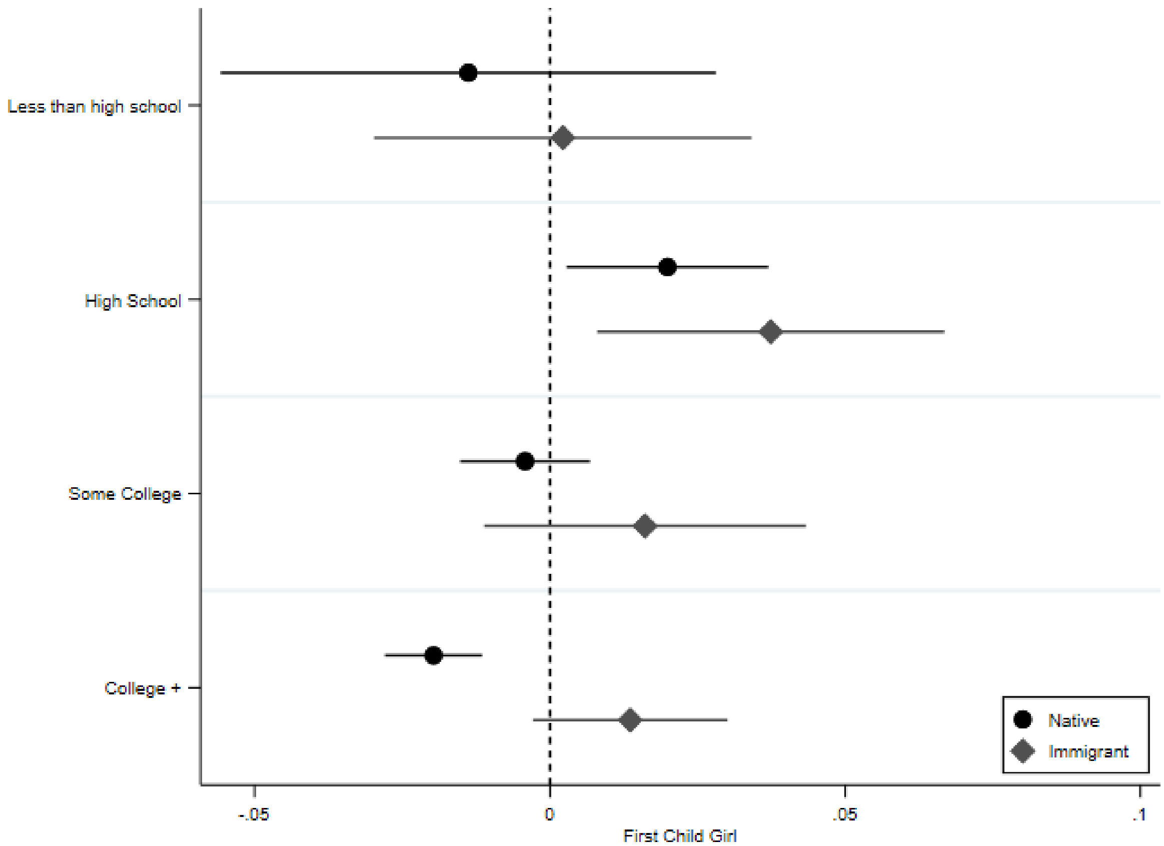
Notes: Estimates for 1960-2000 are calculated using Dahl and Moretti (2008) data, available at https://econweb.ucsd.edu/~gdahl/sons-code.html. Estimates for 2008-2013 are calculated using the American Community Survey with the Dahl and Moretti (2008) specification and sample restrictions. Information on first marriages is not available for 1990 and 2000.

Figure 4: Effects of a Female First Child on the Probability of Living without a Father by Education Level (Linear Probability Models)



Sample from the ACS 2008-2013, includes women, ages 18-40, who are the household head or spouse of the household head, with 1 or more children, where the oldest child is 12 or younger and all children are born in the US. Households with adopted, step or foster children are dropped. Same-sex couples, respondents living in group quarters, respondents born abroad to American parents, widows, as well as mothers with multiple children born in the same year and quarter are also excluded. The dependent variable is a binary equal to one if there is no father in a household. Controls include a cubic in mother's age as well as dummies for year, region (based on 9 Census categories), and race/ethnicity (based on White-nonHispanic, Black-nonHispanic, Asian-nonHispanic, Other-nonHispanic, and Hispanic). Regressions are weighted by normalized household weights that provide equal weighting for each sample year; the bars report the 95% confidence interval calculated with robust standard errors.

Figure 5: Effects of a Female First Child on Fertility by Education Group



Sample from the ACS 2008-2013, includes women, ages 18-40, who are married, spouse present, and who are the household head or spouse of the household head, with 1 or more children, where the oldest child is 12 or younger and all children are born in the US. Households with adopted, step or foster children are dropped. Same-sex couples, respondents living in group quarters, respondents born abroad to American parents, widows, as well as mothers with multiple children born in the same year and quarter, are also excluded. Controls include a cubic in both parents' ages as well as dummies for year, region (based on 9 Census categories), spouse's education (based on < HS, HS, Some College and College Degree), and both parents' race/ethnicity (based on White-nonHispanic, Black-nonHispanic, Asian-nonHispanic, Other-nonHispanic, and Hispanic). Regressions are weighted by normalized household weights that provide equal weighting for each sample year; the bars report the 95% confidence interval calculated with robust standard errors.

Table 1: Effects of a Female First Child on the Probability of Living Without a Father (Linear Probability Models)

	Prob. Living without a father					
	Core Sample			Extended Sample		
Sample	Native	Immigrant	Both	Native	Immigrant	Both
	(1)	(2)	(3)	(4)	(5)	(6)
First Child Girl	0.0031**	0.0027	0.0032**	0.0056***	0.0023	0.0050***
	(0.0014)	(0.0027)	(0.0013)	(0.0013)	(0.0027)	(0.0012)
N	551,325	111,854	663,179	686,996	130,256	817,252
Dep. Var. Mean	0.2901	0.1737	0.2681	0.3369	0.2059	0.3137
Pct Effects	1.0686	1.5544	1.1936	1.6622	1.117	1.5939

Notes: This table uses data from the ACS 2008-2013, includes women and men, ages 18-40, with 1 or more children, where the oldest child is 12 or younger and all children are born in the US. Same-sex couples, respondents living in group quarters, respondents born abroad to American parents, widows, as well as parents with multiple children born in the same year and quarter, are excluded. Men are included in the extended sample, but only if they are ever married and their children cannot be matched to a mother. The core sample additionally excludes men and women who are not the household head or spouse of the household head, as well as households with adopted, step or foster children. The extended sample excludes foster children, but not their families. The dependent variable is a binary equal to one if there is no father present. Controls include a cubic in parent's age as well as dummies for year, region (based on 9 Census categories), parent's education (based on < HS, HS, Some College and College Degree), and race/ethnicity (based on White-nonHispanic, Black-nonHispanic, Asian-nonHispanic, Other-nonHispanic, and Hispanic). Core sample regressions are weighted by normalized household weights that provide equal weighting for each sample year, extended sample regressions use normalized person weights; robust standard errors are in parentheses. ***significant at the 1% level; **significant at the 5% level; *significant at the 10% level.

Table 2: Effects of a Female First Child on Fertility

	Total # of Children			2 or more Children			3 or more Children			4 or more Children		
Sample	Native	Immigrant	Both	Native	Immigrant	Both	Native	Immigrant	Both	Native	Immigrant	Both
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Panel A: All Marriages, Core Sample												
First Child Girl	-0.0075**	0.0164**	-0.0023	-0.0054***	0.0020	-0.0038**	-0.0034**	0.0079***	-0.0009	0.0010	0.0043***	0.0017**
	(0.0032)	(0.0064)	(0.0029)	(0.0019)	(0.0037)	(0.0017)	(0.0015)	(0.0030)	(0.0014)	(0.0008)	(0.0015)	(0.0007)
N	409,567	92,982	502,549	409,567	92,982	502,549	409,567	92,982	502,549	409,567	92,982	502,549
Dep. Var. Mean	1.8657	1.8369	1.8596	0.6312	0.6138	0.6275	0.1848	0.1776	0.1833	0.0394	0.0371	0.0389
Pct Effects	-0.4020	0.8928	-0.1237	-0.8555	0.3258	-0.6056	-1.8398	4.4482	-0.4910	2.5381	11.5903	4.3702
Panel B: First Marriages, Core Sample												
First Child Girl	-0.0070**	0.0149**	-0.0023	-0.0056***	0.0013	-0.0041**	-0.0032*	0.0074**	-0.0010	0.0012	0.0042**	0.0018**
	(0.0035)	(0.0069)	(0.0032)	(0.0020)	(0.0040)	(0.0018)	(0.0017)	(0.0033)	(0.0015)	(0.0008)	(0.0017)	(0.0008)
N	339,290	78,278	417,568	339,290	78,278	417,568	339,290	78,278	417,568	339,290	78,278	417,568
Dep. Var. Mean	1.8815	1.8508	1.8748	0.6406	0.6237	0.6369	0.1894	0.1808	0.1875	0.0406	0.0379	0.0400
Pct Effects	-0.3720	0.8051	-0.1227	-0.8742	0.2084	-0.6437	-1.6895	4.0929	-0.5333	2.9557	11.0818	4.5000
Panel C: All Marriages, Extended Sample												
First Child Girl	-0.0075**	0.0182***	-0.0021	-0.0059***	0.0038	-0.0038**	-0.0030**	0.0081***	-0.0007	0.0007	0.0043***	0.0014**
	(0.0031)	(0.0062)	(0.0028)	(0.0018)	(0.0036)	(0.0016)	(0.0015)	(0.0029)	(0.0013)	(0.0008)	(0.0015)	(0.0007)
N	453,000	100,381	553,381	453,000	100,381	553,381	453,000	100,381	553,381	453,000	100,381	553,381
Dep. Var. Mean	1.8861	1.8356	1.8757	0.6360	0.6092	0.6305	0.1947	0.1793	0.1915	0.0437	0.0384	0.0426
Pct Effects	-0.3976	0.9915	-0.1120	-0.9277	0.6238	-0.6027	-1.5408	4.5176	-0.3655	1.6018	11.1979	3.2864
Panel D: First Marriages, Extended Sample												
First Child Girl	-0.0078**	0.0170**	-0.0026	-0.0065***	0.0034	-0.0044**	-0.0032**	0.0076**	-0.0009	0.0011	0.0040**	0.0017**
	(0.0034)	(0.0067)	(0.0031)	(0.0019)	(0.0039)	(0.0017)	(0.0016)	(0.0031)	(0.0014)	(0.0008)	(0.0016)	(0.0007)
N	364,943	83,972	448,915	364,943	83,972	448,915	364,943	83,972	448,915	364,943	83,972	448,915
Dep. Var. Mean	1.8893	1.8453	1.8799	0.6407	0.6174	0.6357	0.1940	0.1809	0.1912	0.0428	0.0383	0.0419
Pct Effects	-0.4129	0.9213	-0.1383	-1.0145	0.5507	-0.6922	-1.6495	4.2012	-0.4707	2.5701	10.4439	4.0573

Table 2: Effects of a Female First Child on Fertility (ctd)

Sample	Total # of Children			2 or more Children			3 or more Children			4 or more Children		
	Native	Immigrant	Both	Native	Immigrant	Both	Native	Immigrant	Both	Native	Immigrant	Both
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Panel E: All Women, Core Sample												
First Child Girl	-0.0080***	0.0115*	-0.0044*	-0.0060***	0.0019	-0.0046***	-0.0031**	0.0054**	-0.0015	0.0009	0.0030**	0.0013**
	(0.0029)	(0.0059)	(0.0026)	(0.0016)	(0.0034)	(0.0015)	(0.0013)	(0.0027)	(0.0012)	(0.0007)	(0.0014)	(0.0006)
N	551,325	111,854	663,179	551,325	111,854	663,179	551,325	111,854	663,179	551,325	111,854	663,179
Dep. Var. Mean	1.8174	1.8234	1.8186	0.5896	0.5993	0.5914	0.1778	0.1768	0.1776	0.0397	0.0382	0.0394
Pct Effects	-0.4402	0.6307	-0.2419	-1.0176	0.3170	-0.7778	-1.7435	3.0543	-0.8446	2.2670	7.8534	3.2995
Panel F: All Women, Extended Sample												
First Child Girl	-0.0073***	0.0115**	-0.0040*	-0.0056***	0.0020	-0.0042***	-0.0027**	0.0054**	-0.0012	0.0006	0.0029**	0.0010*
	(0.0026)	(0.0056)	(0.0024)	(0.0015)	(0.0033)	(0.0014)	(0.0012)	(0.0025)	(0.0011)	(0.0006)	(0.0013)	(0.0006)
N	664,308	126,788	791,096	664,308	126,788	791,096	664,308	126,788	791,096	664,308	126,788	791,096
Dep. Var. Mean	1.7915	1.8012	1.7933	0.5652	0.5804	0.5679	0.1746	0.1732	0.1744	0.0407	0.0383	0.0403
Pct Effects	-0.4075	0.6385	-0.2231	-0.9908	0.3446	-0.7396	-1.5464	3.1178	-0.6881	1.4742	7.5718	2.4814

Notes-For Panels A,B, C, and D, this table uses data from the ACS 2008-2013, includes women, ages 18-40, who are married, spouse present, with 1 or more children, where the oldest child is 12 or younger and all children are born in the US. Same-sex couples, respondents living in group quarters, respondents born abroad to American parents, as well as mothers with multiple children born in the same year and quarter, are excluded. The core sample additionally excludes households with adopted, step or foster children as well as women who are not the household head or spouse of the household head. The extended sample excludes foster children, but not their families. Controls include a cubic in both parents' ages as well as dummies for year, region (based on 9 Census categories), both parents' education (based on < HS, HS, Some College and College Degree), and both parents' race/ethnicity (based on White-nonHispanic, Black-nonHispanic, Asian-nonHispanic, Other-nonHispanic, and Hispanic). Core sample regressions are weighted by normalized household weights that provide equal weighting for each sample year, and extended sample regressions use normalized person weights; robust standard errors are in parentheses. For Panels E and F, women who are not married spouse present are additionally included in both the Core and Extended Samples (although widows are excluded), and the controls for spouse explanatory variables are excluded. ***significant at the 1% level; **significant at the 5% level; *significant at the 10% level.

Table 3: Effects of Source Country Characteristics on the Probability of Living Without a Father and Fertility, Foreign Born Sample

	Prob. Living Without Father			Total # of Children		
	(1)	(2)	(3)	(4)	(5)	(6)
Main Effects						
Total Fertility	-0.0084	-0.0075	-0.0131	0.0968***	0.1001***	0.1014***
	(0.0069)	(0.0092)	(0.0088)	(0.0238)	(0.0249)	(0.0278)
Log of GDP	-0.0173**	-0.0039	-0.0087	0.1132***	0.0942***	0.0954***
	(0.0086)	(0.0128)	(0.0135)	(0.0256)	(0.0224)	(0.0245)
Labor Force Part.	---	0.0742**	0.0648**	---	-0.0521	-0.0469
		(0.0316)	(0.0300)		(0.0896)	(0.0869)
Sex Ratio at Birth	---	---	-0.5179**	---	---	-0.0697
			(0.1984)			(0.5585)
Equity Index	0.4934***	---	---	-0.4580*	---	---
	(0.0767)			(0.2323)		
First Child Girl	0.0480*	0.0049	0.0482	0.1922*	0.0589**	-0.3701**
	(0.0279)	(0.0084)	(0.0582)	(0.1007)	(0.0242)	(0.1444)
Interactions						
Girl*Labor Force Part.	---	-0.0030	-0.0026	---	-0.0741*	-0.0808**
		(0.0141)	(0.0140)		(0.0434)	(0.0397)
Girl*Sex Ratio at Birth	---	---	-0.0408	---	---	0.4046***
			(0.0546)			(0.1424)
Girl*Equity Index	-0.0678	---	---	-0.2692*	---	---
	(0.0426)			(0.1567)		
N	101,854	109,604	109,604	85,523	91,333	91,333
Dep. Var. Mean	0.1648	0.1725	0.1725	1.8346	1.8356	1.8356

Notes: Sample from the ACS 2008-2013, includes women, ages 18-40, who are the household head or spouse of the household head, with 1 or more children, where the oldest child is 12 or younger and all children are born in the US. Households with adopted, step or foster children are dropped. Same-sex couples, respondents living in group quarters, respondents born abroad to American parents, widows, as well as mothers with multiple children born in the same year and quarter are also excluded. Fertility regressions are restricted to women who are married spouse present. The dependent variable is a binary equal to one if there is no father in a household for the living without a father analysis and total number of children for the fertility regression. Controls include a cubic in mother's age as well as dummies for year, region (based on 9 Census categories), mother's education (based on < HS, HS, Some College and College Degree), race/ethnicity (based on White-nonHispanic, Black-nonHispanic, Asian-nonHispanic, Other-nonHispanic, and Hispanic), years since migration, and years since migration squared. The fertility regressions additionally include controls for spouse's age, education, race, years since migration, years since migration squared, and an indicator for whether the spouse is an immigrant. Country characteristics are 2000-2007 averages, with the exception of the gender equity index which is a 2006-2007 average. Regressions are weighted by normalized household weights that provide equal weighting for each sample year; standard errors clustered at the source country level. ***significant at the 1% level; **significant at the 5% level; *significant at the 10% level.

Table 4: Effects of Female First Child on the Probability of Living Without a Father and Fertility by Immigrant Generation

	Prob. Living Without Father	Total # of Children
	(1)	(2)
Panel A: 3rd+ Generation (Respondent and Both Parents Born in the US)		
First Child Girl	0.0039	-0.0189***
	(0.0024)	(0.0057)
N	154,863	117,253
Dep. Var. Mean	0.2536	1.8582
Pct. Effects	1.5504	-1.0189
Panel B: 1st Generation (Respondent Foreign Born)		
First Child Girl	0.0065	0.0413***
	(0.0045)	(0.0111)
N	30,981	25,618
Dep. Var. Mean	0.1561	1.8196
Pct. Effects	4.1846	2.2703
Panel C: 2nd Generation (At Least One Parent Foreign Born)		
First Child Girl	-0.0116	0.0281
	(0.0079)	(0.0179)
N	14,647	10,945
Dep. Var. Mean	0.2426	1.8626
Pct. Effects	-4.7721	1.5095

Notes: Sample from the March CPS 1995-2014, includes women, ages 18-40, who are the household head or spouse of the household head, with 1 or more children, where the oldest child is 12 or younger and all children are born in the US. Households with foster children are dropped. Same-sex couples, respondents living in group quarters, respondents born abroad to American parents, widows, as well as mothers with multiple children born in the same year are also excluded. Fertility regressions are restricted to women who are married spouse present. The dependent variable is a binary equal to one if there is no father in a household for the living without a father analysis and total number of children for the fertility regression. Controls include a cubic in mother's age as well as dummies for year, region (based on 9 Census categories), mother's education (based on < HS, HS, Some College and College Degree), and race/ethnicity (based on White-nonHispanic, Black-nonHispanic, Asian-nonHispanic, Other-nonHispanic, and Hispanic). The fertility regressions additionally include controls for spouse's age, education, and race. Regressions are weighted by normalized household weights that provide equal weighting for each sample year; robust standard errors are in parentheses. ***significant at the 1% level; **significant at the 5% level; *significant at the 10% level.

Table 5: Effects of Source Country Characteristics on the Probability of Living Without a Father and Fertility, Second Generation Sample (At Least One Parent Foreign Born)

	Prob. Living Without a Father			Total # of Children		
	(1)	(2)	(3)	(1)	(2)	(3)
Main Effects						
Total Fertility	-0.0104	0.0086	-0.0009	0.1010**	0.0937**	0.0796*
	(0.0137)	(0.0148)	(0.0159)	(0.0404)	(0.0409)	(0.0460)
Log of GDP	-0.0316**	0.0075	-0.0013	0.0516	0.0365	0.0230
	(0.0143)	(0.0196)	(0.0206)	(0.0360)	(0.0319)	(0.0374)
Labor Force Part.	---	0.1814***	0.1890***	---	0.2478	0.2542
		(0.0520)	(0.0511)		(0.1665)	(0.1688)
Sex Ratio at Birth	---	---	-1.3762***	---	---	-1.4783*
			(0.3407)			(0.8551)
Equity Index	0.2874**	---	---	0.5613	---	---
	(0.1377)			(0.3459)		
First Child Girl	-0.1950	-0.0455	-0.8349***	0.7394**	0.2201**	-0.0681
	(0.1461)	(0.0471)	(0.2672)	(0.2953)	(0.0893)	(0.7532)
Interactions						
Girl*Labor Force Part.		0.0603	0.0466	---	-0.3024*	-0.3124*
		(0.0740)	(0.0736)		(0.1579)	(0.1666)
Girl*Sex Ratio at Birth	---	---	0.7513***	---	---	0.2766
			(0.2618)			(0.7626)
Girl*Equity Index	0.2750	---	---	-1.0459**	---	---
	(0.2143)			(0.4375)		
N	11,687	13,481	13,481	9,040	10,047	10,047
Dep. Var. Mean	0.2426	0.2426	0.2426	1.8626	1.8626	1.8626

Notes: Sample from the March CPS 1995-2014, includes women, ages 18-40, who are the household head or spouse of the household head, with 1 or more children, where the oldest child is 12 or younger and all children are born in the US. Households with foster children are dropped. Same-sex couples, respondents living in group quarters, respondents born abroad to American parents, widows, as well as mothers with multiple children born in the same year are also excluded. Fertility regressions are restricted to women who are married spouse present. The dependent variable is a binary equal to one if there is no father in a household for the living without a father analysis and total number of children for the fertility regression. Controls include a cubic in mother's age as well as dummies for year, region (based on 9 Census categories), mother's education (based on < HS, HS, Some College and College Degree), and race/ethnicity (based on White-nonHispanic, Black-nonHispanic, Asian-nonHispanic, Other-nonHispanic, and Hispanic). The fertility regressions additionally include controls for spouse's age, education, race, immigrant status, and second-generation immigrant status. Country characteristics are based on the woman's mother's birthplace if the mother was an immigrant and the woman's father's birthplace otherwise. Country characteristics are 2000-2007 averages, with the exception of the gender equity index which is a 2006-2007 average. Regressions are weighted by normalized household weights that provide equal weighting for each sample year; standard errors are clustered at the mother's birth country (or father's if mother is not an immigrant). ***significant at the 1% level; **significant at the 5% level; *significant at the 10% level.

Table 6: Boy/Girl Ratio, Second and Third Children

	All Married Women		Women in First Marriage		All Women	
	Natives	Immigrants	Natives	Immigrants	Natives	Immigrants
	(1)	(2)	(3)	(4)	(5)	(6)
A. Second Child						
First Child Boy						
Sex Ratio	1.057	1.053	1.056	1.045	1.055	1.053
95% Conf. Int.	[1.046,1.068]	[1.029,1.078]	[1.044,1.068]	[1.019,1.071]	[1.045,1.065]	[1.031,1.075]
Sample Size	134,605	29,150	113,360	25,004	169,979	34,207
First Child Girl						
Sex Ratio	1.045	1.007	1.053	1.008	1.039	1.012
95% Conf. Int.	[1.033,1.057]	[0.983,1.030]	[1.040,1.065]	[0.982,1.033]	[1.029,1.049]	[0.991,1.034]
Sample Size	126,441	27,820	106,394	23,810	160,359	32,798
B. Third Child						
First Two Children Boys						
Sex Ratio	1.052	1.016	1.048	1.051	1.054	1.021
95% Conf. Int.	[1.024,1.080]	[0.955,1.076]	[1.018,1.078]	[0.983,1.119]	[1.029,1.078]	[0.966,1.077]
Sample Size	22,093	4,319	18,827	3,698	28,212	5,180
First Two Children Girls						
Sex Ratio	1.040	0.983	1.047	0.993	1.034	0.977
95% Conf. Int.	[1.011,1.069]	[0.925,1.041]	[1.015,1.079]	[0.929,1.056]	[1.009,1.060]	[0.924,1.031]
Sample Size	19,537	4,366	16,643	3,740	25,233	5,177
First Two Children Mix						
Sex Ratio	1.039	1.086	1.047	1.103	1.046	1.056
95% Conf. Int.	[1.017,1.061]	[1.036,1.137]	[1.023,1.071]	[1.048,1.158]	[1.027,1.066]	[1.012,1.101]
Sample Size	34,346	7,158	29,278	6,197	44,668	8,703

Notes: Sample from the ACS 2008-2013, includes women, ages 18-40, who are the household head or spouse of the household head, with 2 or more children (Panel A) or 3 or more children (Panel B), where the oldest child is twelve or younger, all children are born in the US and no children were adopted, step, or foster children of the household head. Same-sex couples, respondents living in group quarters, respondents born abroad to American parents, widows, as well as mothers with multiple children born in the same year and quarter, are excluded. Means are weighted by normalized household weights that provide equal weighting for each sample year; 95% confidence intervals are in parentheses. Confidence intervals are based on the standard errors of the percentage of second (Panel A) or third (Panel B) children who are boys.

Table A1: Children in Sample Compared to Reported Live Births

	All Women, Ages 18-40		Women, Ages 18-40, With Sample Restrictions			
			Married		All	
	Number	Percent	Number	Percent	Number	Percent
A. All Groups						
Sample = Live	20514	83%	11670	92%	15704	91%
Sample Less Than Live	3106	12%	726	6%	1157	7%
Sample Greater Than Live	1232	5%	290	2%	419	2%
B. Natives						
Sample = Live	16570	83%	9605	92%	13199	91%
Sample Less Than Live	2445	12%	591	6%	974	7%
Sample Greater Than Live	1047	5%	238	2%	347	2%
C. Immigrants						
Sample = Live	3945	82%	2066	92%	2505	91%
Sample Less Than Live	652	14%	134	6%	183	7%
Sample Greater Than Live	192	4%	52	2%	72	3%
D. Asian Immigrants						
Sample = Live	897	89%	552	94%	596	94%
Sample Less Than Live	75	7%	24	4%	26	4%
Sample Greater Than Live	39	4%	12	2%	13	2%
E. Hispanic Immigrants						
Sample = Live	2128	79%	1004	91%	1299	90%
Sample Less Than Live	448	17%	74	7%	108	7%
Sample Greater Than Live	102	4%	26	2%	39	3%
F. Second Generation Immigrants						
Sample = Live	1516	87%	889	94%	1237	94%
Sample Less Than Live	161	9%	44	5%	58	4%
Sample Greater Than Live	56	3%	16	2%	25	2%

Notes: Sample from the 2008, 2010, and 2012 June CPS. Number of live births is based on the CPS variable frever, which measures the number of live births the woman ever had. The unrestricted sample includes women ages 18-40 who are the household head or spouse of the household head and who have at least one child. The restricted sample is further limited to families where the oldest child is 12 or younger, where all children are born in the US, where no children are step or adopted, and where, for married women, mothers were listed as married spouse present. We exclude same-sex couples, widows, respondents living in group quarters, households with foster children of the head, and mothers with multiple children born in the same year.

Table A2: Boy/Girl Ratio, First Child

	Natives	Immigrants
A. All Married Women		
Ratio	1.0556	1.0472
95% Confidence Interval	[1.0492, 1.0621]	[1.0338, 1.0608]
Sample Size	409567	92982
B. Married Women in First Marriage		
Ratio	1.0567	1.0486
95% Confidence Interval	[1.0496, 1.0638]	[1.034, 1.0634]
Sample Size	339290	78278
C. All Women		
Ratio	1.047	1.0429
95% Confidence Interval	[1.0415, 1.0525]	[1.0307, 1.0552]
Sample Size	551325	111854

Notes: Sample from the ACS 2008-2013, includes women, ages 18-40, who are the household head or spouse of the household head, with 1 or more children, where the oldest child is 12 or younger and all children are born in the US. Households with adopted, step or foster children are dropped. Same-sex couples, respondents living in group quarters, respondents born abroad to American parents, widows, as well as mothers with multiple children born in the same year and quarter are also excluded. Means are weighted by normalized household weights that provide equal weighting for each sample year; 95% confidence intervals are in parentheses. Confidence intervals are based on the standard errors of the percentage of first children who are boys.

Table A3 : Probability that First Child is a Girl (Linear Prob Models)

Variables	Immigrants (1)	Immigrants (2)	Natives (3)
<HS	0.0104*	0.0092***	0.0032
	(0.0057)	(0.0033)	(0.0041)
Some College	-0.0033	-0.0053	-0.0003
	(0.0057)	(0.0055)	(0.0025)
College+	-0.0024	-0.0025	0.0006
	(0.0056)	(0.0054)	(0.0025)
Black	-0.0079	-0.0039	0.0103***
	(0.0087)	(0.0117)	(0.0028)
Hispanic	-0.0009	0.0021	0.0028
	(0.0059)	(0.0081)	(0.0031)
Asian	0.0017	0.0030	-0.0061
	(0.0058)	(0.0077)	(0.0070)
Other	-0.0025	0.0156	0.0053
	(0.0180)	(0.0119)	(0.0071)
Middle Atlantic	-0.0099	-0.0129	-0.0035
	(0.0094)	(0.0112)	(0.0047)
East North Central	-0.0252**	-0.0271**	-0.0032
	(0.0104)	(0.0119)	(0.0045)
West North Central	-0.0210	-0.0203	0.0012
	(0.0144)	(0.0193)	(0.0051)
South Atlantic	-0.0233**	-0.0266**	-0.0025
	(0.0093)	(0.0114)	(0.0045)
East South Atlantic	-0.0112	-0.0207	0.0042
	(0.0161)	(0.0182)	(0.0051)
West South Atlantic	-0.0151	-0.0166	-0.0017
	(0.0098)	(0.0139)	(0.0047)
Mountain	-0.0233**	-0.0235**	-0.0006
	(0.0110)	(0.0100)	(0.0050)
Pacific	-0.0214**	-0.0246**	0.0004
	(0.0090)	(0.0113)	(0.0047)
2009	-0.0002	0.0010	-0.0039
	(0.0060)	(0.0058)	(0.0028)
2010	-0.0051	-0.0040	0.0010
	(0.0060)	(0.0047)	(0.0028)
2011	0.0001	-0.0013	-0.0037
	(0.0064)	(0.0106)	(0.0030)
2012	-0.0069	-0.0057	-0.0020
	(0.0063)	(0.0051)	(0.0029)
2013	-0.0046	-0.0054	-0.0038
	(0.0063)	(0.0057)	(0.0029)

Table A3 : Probability that First Child is a Girl, Ctd (Linear Prob Models)

Variables	Immigrants (1)	Immigrants (2)	Natives (3)
Age	-0.0086	-0.0103	-0.0058
	(0.0334)	(0.0387)	(0.0143)
Age^2	0.0003	0.0004	0.0002
	(0.0011)	(0.0013)	(0.0005)
Age^3	-0.0000	-0.0000	-0.0000
	(0.0000)	(0.0000)	(0.0000)
Years Since Migration	-	0.0010	-
		(0.0007)	
Years Since $Migration^2$	-	-0.0000	-
		(0.0000)	
Total Fertility	-	-0.0046*	-
		(0.0025)	
Equity Index	-	-0.0754	-
		(0.0517)	
Log GDP	-	-0.0012	-
		(0.0041)	
F-stat	1.0240	3.4105	1.3471
Prob > F	0.4289	0.0000	0.1232
N	111,854	101,854	551,325

Notes: Sample from the ACS 2008-2013, includes women, ages 18-40, who are the household head or spouse of the household head, with 1 or more children, where the oldest child is 12 or younger and all children are born in the US. Households with adopted, step or foster children are dropped. Same-sex couples, respondents living in group quarters, respondents born abroad to American parents, widows, as well as mothers with multiple children born in the same year and quarter are also excluded. Country characteristics are 2000-2007 averages, with the exception of the gender equity index which is a 2006-2007 average. Regressions are weighted by normalized household weights that provide equal weighting for each sample year. Results in columns 1 and 3 implement robust standard errors, while in column 2 they are clustered at the source country level. ***significant at the 1% level; **significant at the 5% level; *significant at the 10% level.

Table A4: Effects of a Female First Child on the Probability of Being in the Labor Force: Married Women and All Women (Linear Probability Models)

	Married Women			All Women		
Sample	Native	Immigrant	Both	Native	Immigrant	Both
	(1)	(2)	(3)	(1)	(2)	(3)
First Child Girl	0.0029	-0.0058	0.0011	0.0026*	-0.0040	0.0014
	(0.0017)	(0.0038)	(0.0016)	(0.0015)	(0.0035)	(0.0014)
N	409,567	92,982	502,549	551,325	111,854	663,179
Dep. Var. Mean	0.7093	0.5552	0.6765	0.7425	0.5965	0.7150

Notes: Sample from the ACS 2008-2013, includes women, ages 18-40, who are the household head or spouse of the household head, with 1 or more children, where the oldest child is 12 or younger and all children are born in the US. Households with adopted, step or foster children are dropped. Same-sex couples, respondents living in group quarters, respondents born abroad to American parents, widows, as well as mothers with multiple children born in the same year and quarter are also excluded. The dependent variable is a binary equal to one if the women is in the labor force during the survey week. Controls include a cubic in mother's age as well as dummies for year, region (based on 9 Census categories), mother's education (based on < HS, HS, Some College and College Degree), race/ethnicity (based on White-nonHispanic, Black-nonHispanic, Asian-nonHispanic, Other-nonHispanic, and Hispanic), years since migration, and years since migration squared. The regressions for married women are restricted to women who are married spouse present and additionally include controls for spouse's age, education, and race/ethnicity. Regressions are weighted by normalized household weights that provide equal weighting for each sample year; robust standard errors are in parentheses. ***significant at the 1% level; **significant at the 5% level; *significant at the 10% level.