Having a Son Promotes Clean Cooking Fuel Use in Urban India: Women's Status and Son Preference

AVINASH KISHORE
International Food Policy Research Institute

DEAN SPEARS
Centre for Development Economics, Delhi School of Economics

I. Introduction

Despite the profoundly negative health consequences of indoor air pollution, about half of the households in the world cook using solid biomass fuels (Smith 2002). In India, many households use firewood or dung cakes as the primary source of energy for cooking (NSSO 2007). Burning these unprocessed biomass fuels in traditional open fire burners (called *chulhas*) results in very high levels of indoor air pollution and an estimated 450,000–550,000 premature deaths and nearly 500 million cases of illness each year (Smith 2000). Such indoor air pollution is behind only open defecation and malnutrition as a major cause of disease and death in India. Women and children suffer disproportionately, as they spend more time indoors and women do essentially all of the cooking (ESMAP 2003).

A switch to cleaner cooking fuels such as kerosene, liquid petroleum gas (LPG), or biogas would save many lives and reduce suffering from indoor air pollution. However, despite the increased availability of cleaner fuel, households' transition from traditional cooking fuel has been slow. What explains the transition to clean fuel use?

Certainly there are many important factors relevant to households' use of biomass fuels, including price, accessibility, and a low opportunity cost of time

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spent collecting wood. This article focuses on the causal role of one factor: the sex of a household's first child. Economists have long recognized that households in developing countries (and elsewhere) are not unified economic actors (e.g., Udry 1996). There are at least two mechanisms by which child sex could affect fuel choice: by influencing the intrahousehold status of women, who bear more of the costs of traditional fuels, or by presenting an opportunity to invest in children's health, in the context of a preference for healthier boys.

Urban Indian households with a male first child are approximately 2 percentage points more likely to use clean cooking fuel than comparable households with a female first child. If child sex is not selected for by biased abortion or other processes, then the sex of a first child has an exogenous causal effect on household fuel choice. We show that the association between fuel choice and child sex is not driven by intentionally or unintentionally terminated pregnancies or by household wealth or family size. Among a range of outcomes we study, the effect of child sex is unique to fuel choice; our finding that there is no effect on other assets indicates that it is unlikely that the result is confounded by real or subjectively anticipated wealth. In addition to the National Family Health Survey NFHS-3, the main data source studied, we approximately replicate the result using two additional surveys.

A. Why Might Child Sex Matter?

Indian households have a well-documented preference for sons over daughters (e.g., Sen 2003). Jeffery, Jeffery, and Lyon (1989) recount a birth in rural Uttar Pradesh. The baby turns out to be a girl: "after a boy's birth, there would have be celebrations, presents, and jollity, so this birth is a disappointment" (5). The new mother is "plaintive"; the midwife blames God. Afterward, the mother-in-law of the woman who gave birth offers a smaller-than-usual payment to the midwife, explaining "it would be different for a boy" (6).

This section summarizes two mechanisms by which a woman having a boy rather than a girl could increase the likelihood of clean cooking fuel use. First, it could increase her intrahousehold status, and women suffer the most from traditional fuel. Second, households may elect to invest more in the health of boys than girls, and eliminating unsafe smoke from biomass fuel could be one such investment. Ultimately, our data will not allow us to definitively distinguish these possible mechanisms; indeed, both could be simultaneously active. However, we can present evidence for a "first-stage" effect of child sex on women's status: women whose first child is a son have a greater body mass index (BMI) than women whose first child is a daughter.

1. Women's Intrahousehold Status

In general, "women and children are among the most deprived in the usual way an Indian household is run" (Bardhan 2011, 41). It is women cooks (and their children) who suffer most from indoor air pollution and have the most to gain from a switch to cleaner fuel. Therefore, one might expect that increasing the status of women within a rural Indian household would increase the chances of the household using clean cooking fuel, the benefits of which would accrue disproportionately to the women cooks. According to this hypothesis, a household's use of clean cooking fuel is less of an investment in health than it is a consumption good for those who would otherwise have to breathe the smoke and bear other costs.

Figure 1 presents evidence consistent with this view from the NFHS-3. Households in which women have or share final say over more household decisions are more likely to use clean cooking fuel, a result that is not eliminated by controlling for heterogeneity of wealth.[1]

Our article is not the first to consider the role of women in households' fuel choice. However, to our knowledge this is the first article to focus on the role of women's status, applying a strategy of causal identification. Hoddinott and Haddad (1995) find a positive relationship between a wife's share of income and the budget share of fuel in household data from the Ivory Coast, although the parameter estimate was not statistically significant. In her study of fuel choice in urban Bolivia, Israel (2002) finds female earned income to be associated with a lower probability of firewood use. However, she cannot rule out that this is because women with higher earnings have a higher opportunity cost of time, rather than more bargaining power. Duflo, Greenstone, and Hanna (2008) find suggestive evidence for our conjecture in a survey of 2,200 households in rural Orissa, where "households in which women may be more empowered—by virtue of being members of a savings group—are 2 to 3 percent more likely to use a clean stove" (73). Finally, Miller and Mobarak (2011) show in an experiment in Bangladesh that women are more likely than men to adopt improved stoves when they are free, but less likely when there is a charge, suggesting that women have a stronger preference for the stoves but do not control the resources to purchase them.[2]

[1] For more correlational and regression evidence of an association between women's intrahousehold status and clean cooking fuel use, please see the original working paper version of this article, Kishore and Spears (2012).

[2] Further, in this experiment, women are much less likely to pay for a health-improving stove when they are much younger than their husbands (a proxy for low rank). The difference in men's and women's behavior does not interact with the presence of children or with children's health, suggesting that a desire to invest in children was not driving Miller and Mobarak's (2011) results.

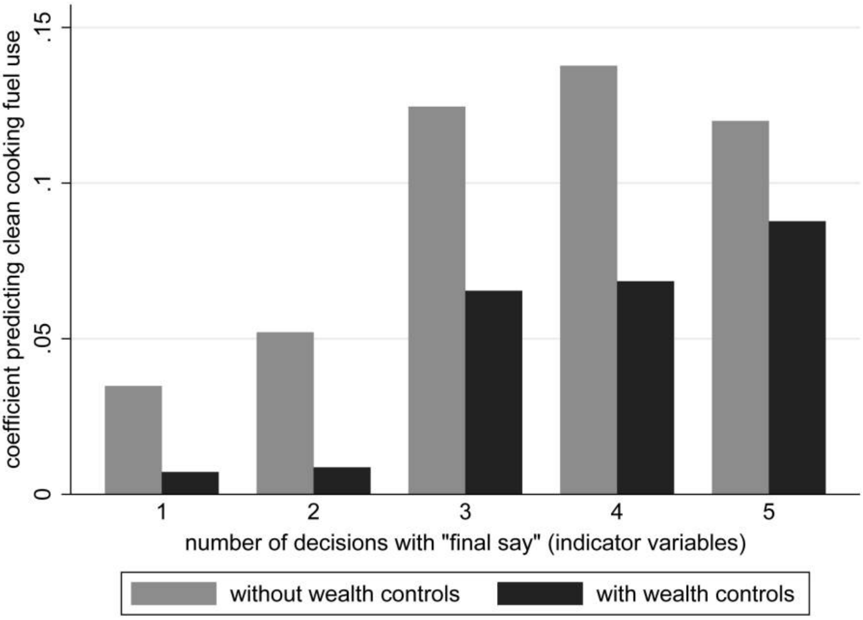

Figure 1. Clean fuel use by women's status, NFHS-3. Coefficients from two linear probability regressions, each an indicator of clean cooking fuel use on a set of five indicators for household decisions over which the woman reports having "final say." Urban sample with oldest child under 5; sampling weights used.

Having a firstborn daughter, relative to having a firstborn son, can entail loss of intrahousehold status to the mother; in other words, there is an "opportunity cost" of the status she would have gained by having a boy (e.g., Ahmed-Ghosh 2004; Silverman et al. 2011). Using as the dependent variable the same NFHS data about having say in decision making as is used in figure 1, Zimmermann (2012) documents that after having a boy rather than a girl, Indian women experience a temporary increase in intrahousehold decision-making status.[3] Similarly, Li and Wu (2011) find an effect of having a first son on mothers' bargaining power in China. If having a boy does cause an increase in decision-making status, and if women suffer disproportionately from traditional biomass cooking fuel use, then it may be the case that having a boy makes clean cooking fuel use more likely.

[3] Zimmermann emphasizes that this effect is small and that her results further suggest that this increase in decision-making status fades over time. If so, our results suggest that the benefits may have been "locked in" by making difficult-to-reverse changes during the period of increased status, although verifying this is beyond the scope of this article because our data give no information about when households switched fuel types.

Mechanism 1. Having a son, rather than a daughter, increases women's intrahousehold status. Because women bear more costs than men, on average, from traditional fuel use, this increases the probability of clean cooking fuel use.

A similar empirical strategy to ours was recently used by Pham-Kanter (2010), who finds that in the United States, women who have a firstborn son weigh more during the child's teenaged years than women who have a firstborn daughter. Remarkably, she, too, attributes this effect to son preference, apparently relevant and statistically detectable in the United States (Dahl and Moretti 2008). Pham-Kanter proposes that having a son rather than a daughter increases US women's intrahousehold bargaining power, allowing them to devote less attention to maintaining low weight. In Section III.D.2, we show that women whose first child is a son rather than a daughter have a larger BMI, which is consistent with higher social status within Indian households.

2. Investing in Sons' Health

In much of India, adult women live with their husband's family, so that households gain more of the benefits of investing in the early life health of sons. As Jeffery et al. (1989) quote from fieldwork in rural Uttar Pradesh, "raising a daughter is like watering a shade-giving tree in someone else's courtyard" (23).

There has been debate in the literature whether Indian girls receive less health investment than boys. Deaton (1997, 2003) finds little evidence that Indian families spend more on boys or are more likely to vaccinate boys. Yet, Barcellos, Carvalho, and Lleras-Muney (2012), studying time use data, find that households spend more time caring for very young first-child boys than for comparable girls.[4] If Indian households do wish to invest in the health of boys, then, given the danger of smoke from traditional fuel, they might invest in clean fuel.

Mechanism 2. Indian households prefer to invest more in the health of boys than in girls. Because clean cooking fuel is healthier, households with boys are more likely to adopt it.

However, even if boys do receive more investment in health than girls, there are reasons in the literature to doubt that intentional investment in boys' health is what explains the effect of child sex on fuel choice. Sometimes better

[4] They do not describe out of what alternative activities this time is substituted. Given the constraints faced by young Indian daughters-in-law, it is not inconceivable that child care time could be seen as a form of a consumption good for the new young mothers.

health for boys than girls is a by-product of pursuing other goals. For example, Jayachandran and Kuziemko (2011) find that Indian mothers breast-feed girls less than boys because, after girls, mothers are more quickly ready to have another child, in order to try to have a son. More broadly, Dupas (2011) reviews evidence that poor people worldwide spend a lot of money on health treatment but little in investment in health prevention; switching to clean cooking fuel for the purpose of a health investment would be an exception to this pattern.

B. Outline

Section II introduces the empirical strategy and the data. The main analysis uses the NFHS-3, but results are approximately replicated with the NFHS-2 and the DLHS-3. Section III presents the results, robustness checks, and a falsification test based on substituting "placebo" dependent variables. Section IV concludes and discusses implications of the results.

II. Empirical Strategy

Are households to whom boys have been born more likely to use clean cooking fuel than households to whom girls have been born? Like in Pham-Kanter's (2010) study of maternal weight gain, but unlike in Washington's (2008) study of US legislators, our empirical strategy compares firstborn sons to firstborn daughters, rather than studying the sex composition of all of a household's children. This is because son preference could create fertility stopping rules, whereby the decision to have more children depends on the gender composition of the children a household already has (Clark 2000). For example, a family might continue to have children until they have a boy. If these rules depend on, or their successful execution is correlated with, other relevant variables, then overall sex composition could be endogenous to wealth or fuel choice. Barring sex-selective abortion or other nonrandom selection into child sex, which we discuss below, the sex of the first child is more likely to be outside of households' control (Kishor and Gupta 2009).

A. Identification Strategy

Among households with at least one child, we estimate the linear probability regression

$$\text{clean fuel}_{is} = \alpha_s + \beta \text{boy}_{is} + X_{is}\theta + \varepsilon_{is}, \qquad (1)$$

where i indexes households, s indexes states, *boy* is an indicator that the firstborn child was male, and X_{is} is a vector of controls. Because the sex of the first child is like a random experiment, and because Freedman (2008) demonstrates that including covariates can bias experimental estimates, we present results with and without covariates and urban primary sampling unit (PSU) fixed effects.[5]

The vector of controls includes the Demographic and Health Surveys (DHS) wealth index as a quadratic polynomial; a set of three India-specific standard-of-living indicators from the NFHS; the height-for-age *z*-score of the oldest child; the number of living children of the child's mother as a quadratic polynomial; the number of household members; a set of indicators for the child's mother's 5-year age band; the mother's age at marriage; a set 12 of indicators for the child's mother's years of education; a set of five indicators for town or city size; and indicators for being a nuclear family, scheduled caste household, or scheduled tribe household. Among these, the number of children may be particularly important. In the sample of first children born in the last 10 years, households with a first boy have 0.11 fewer children ($t \approx 5$) on average. This is very likely to be largely a consequence of first child sex and is evidence of sex-biased fertility stopping.[6] Therefore, the total number of children is included as a quadratic polynomial.

1. Nonrandom Selection into Child Sex

Perhaps the most important threat to this empirical strategy would be if sex-selective abortion were more likely to be practiced by richer households, such that the apparent effect of a boy baby on clean fuel use were, in fact, a confounded effect of wealth. Arnold, Kishor, and Roy (2002, 762) report that "ultrasound typically costs between 500 and 1,000 rupees (about $10–20 at the current rate of exchange)."

Importantly, whether this is potentially a problem is a priori ambiguous. Sex selection of the first child would only cause endogeneity if it were correlated with omitted determinants of clean fuel use. Clark (2000, 95) finds that in India, "socially and economically disadvantaged couples . . . not only want but also attain a higher proportion of sons," so richer families may not be more likely to abort girls. Similarly, separating the effects of relative and absolute wealth using the NFHS-3, Gaudin (2011, 343) documents that "higher absolute wealth is strongly associated with lower son preference, and the effect is 20%–40% stronger when the household's community-specific wealth score [a control for local relative wealth] is included in the regression." Finally, in a

[5] Identifiers for district fixed effects, unfortunately, are not available in the NFHS-3, the central data source of the article; however, district fixed effects are used in the analysis of the NFHS-2 and the DLHS-3.

[6] Consistent with this interpretation, the coefficient is smaller and less statistically significant in the sample of first children born in the last 5 years ($t = 2.3$).

review of international trends in sex selection, Hesketh and Xing (2006) report that sex-selective abortion is more common at later birth orders; this study focuses only on the sex of firstborn children.[7]

A similar threat may be posed by any other mechanism that selectively prevents male or female fetuses from being born alive. For example, using evidence from a policy change in the United States in the 1970s, Sanders and Stoecker (2011) argue that exposure to air pollution is more likely to cause male fetal deaths than female fetal deaths in utero. If so, this could be an example of the Trivers and Willard (1973) hypothesis from evolutionary biology that unhealthy women are more likely to have female children than male children. If air pollution indeed has this effect, then causality could run in the opposite direction to that this article proposes: using traditional fuel would cause girl babies.

We use a simple technique to ensure that neither of these issues—or any other nonrandom selection of which pregnancies become live birth—is driving the result. We replicate every result excluding the approximately 15% of the sample that reports ever having had a pregnancy that terminated (intentionally or spontaneously) rather than resulting in a birth.[8] Note that this is an indicator for having had a terminated pregnancy by the time of the survey, not by the birth of their first child. Because our estimates and conclusions are never importantly different for this subsample, we conclude that nonrandom selection is not driving our results.

2. Further Analysis

In order to verify the evidence of a causal effect of child sex on clean cooking fuel use, we conduct two further analyses. First is a set of falsification tests. We substitute, instead of clean cooking fuel use, 18 alternative dependent variables: ownership of 13 other households assets (that do not have the same gender implications as cooking fuel); social group membership indicators for low caste, "tribal" status, and being Muslim; possession of a nominally poverty-targeted ration card; and a wealth index. This is methodologically similar to a test for balance in a randomized experiment but here highlights the unique-

[7] Using data from the DLHS-2, a nationally representative survey of approximately 500,000 ever-married women age 13–44 years in 2002–4, Rosenblum (2013) shows that sex of the firstborn child is random and that parents are not likely to selectively abort their first pregnancy. Rosenblum cites Portner (2010), Bhalotra and Cochrane (2010), and Jha et al. (2011), who show that the first pregnancy in India has a biologically normal male:female sex ratio in the range of 1.04–1.07.

[8] Underreporting of fetal death is an important problem in infant vital statistics (Zupan and Åhman 2006, 8), and the DHS is likely no exception. Results below showing that there is no "effect" of a newborn's sex (i.e., before there would be time for an effect to unfold) provide further evidence that heterogeneity in underreporting of fetal death is not driving our results.

ness of clean cooking fuel among household assets. If the association between child sex and clean cooking fuel were spuriously driven by a household confound such as wealth or socioeconomic status—or by perceived wealth, perhaps because of an effect of household wealth on anticipated family size or dowry—we might expect that having a boy child would similarly show an "effect" on some of these alternative dependent variables. Second, we approximately replicate the main result with two alternative data sets.

B. *Data*

The main data set used in this analysis is the third round of the National Family Health Survey (NFHS-3), India's version of the DHS, conducted in 2005–6. The NFHS-3 is widely used and is the highest-quality recent large demographic survey representative of India.

Our dependent variable is coded from a question about the household's "main cooking fuel" that allows one response. We classify as clean cooking fuel use of LPG or natural gas (53% of urban households with the oldest child under 5), kerosene (8%), electricity (1%), and biogas (less than 1%).

Table 1 presents tabulated summary statistics. Slightly more than half of first children born alive are boys. A little more than 40% of all households and a little less than 75% of urban households use clean cooking fuel. In urban and all India, with and without women with terminated pregnancies, clean cooking fuel is more common among households with male first children. The table includes Fisher's exact *p*-values for this cross-tabulation. These can be seen as an initial, nonparametric statistical test of this article's hypothesis (incorrectly assuming independence of observations).

In addition to presenting results for children born to mothers with and without terminated pregnancy, we use various subsets of the NFHS-3 data. The primary results of the article focus on urban India, where clean cooking

TABLE 1
SUMMARY STATISTICS, NFHS-3 OLDEST CHILD UNDER 5

	Urban Only		Urban and Rural	
	(1)	(2)	(3)	(4)
Ever terminated*	Yes and no	No	Yes and no	No
Clean fuel, boy first	.758	.756	.421	.416
Clean fuel, girl first	.732	.729	.409	.403
Fisher's exact *p*-value	.009	.013	.079	.073
Fraction with boy first	.511	.514	.508	.510
Fraction with terminated pregnancy	.155	.000	.146	.000
n	6,318	5,337	13,808	11,793

* Indicates reporting ever having had an intentionally or unintentionally terminated pregnancy.

fuel is relatively readily available for sale, to the exclusion of rural India, where binding constraints other than women's intrahousehold status (e.g., availability, poverty, or social norms and tastes) may prevent take-up of clean cooking fuel. For example, alternative fuels such as dung are not very costly to collect in many rural settings. In the published statistics from the 1998–99 and 2005–6 DHS surveys, 46.9% and 58.7% of urban households used LPG, but only 5.1% and 8.2% of rural households did, respectively.

The birth history in the NFHS-3 includes sons born since 1968; this is clearly too early to have had an effect on the fuel used in 2006, and some cutoff must be used. For robustness, we use two. First, we limit the sample to households whose oldest child is under 5. This is a natural cut point because more data are collected in the NFHS-3 about children under 5, including height for age, which we use as a measure of investment in children's health. Second, we limit the sample to households whose oldest child was born before 1996. This cutoff was chosen for being 10 years before the survey and approximately the time when market liberalization in India was beginning to make LPG more available for purchase.[9]

After presenting results with the NFHS-3, we show two replications with other data sets. The next prior Indian DHS, the NFHS-2, was conducted in 1999.[10] It includes similarly high-quality data but has fewer observations and clearly will only have a limited sample of births since 1996, the period under consideration. The third round of the District Level Health and Facilities Survey, or DLHS-3, was conducted in 2008 by the Indian government and the International Institute of Population Sciences in Mumbai. It asks questions similar to those of the DHS to a large sample of Indian households and includes district fixed effects.

The NFHS-3 found that about 75% of urban households used clean cooking fuel; the earlier NFHS-2 found that a little less than 70% did. However,

[9] In the published DHS statistics, only 33.4% of urban households used LPG in 1992–93, compared with 46.9% in 1998–99. In government data on LPG connections, the rate of new connections increases by about 1.2 million per year in the 1980s and by 1.7 million per year from 1990 to 1995. However, this more than doubles to 3.6 million new connections in 1996, 4.4 million in 1997, and reaching over 10 million new connections per year in 2000 (Ministry of Petroleum and Natural Gas 2001). What matters to our analysis is less the number of connections that exist than the cost of switching to clean cooking fuel, which is influenced by transaction costs and the thickness of the market. Many Indian households had long-standing LPG connections, but in the mid-1990s these connections became much easier to newly acquire.

[10] As discussed, we do not use earlier data because government regulation made it much more difficult to switch to clean cooking fuel use. If, however, the main regression specification is applied to the 1992 NFHS-1, looking at mothers of firstborn children under 5 years old, those with boys are no more likely to be using clean cooking fuel than those with girls (difference = 0.002; $t = 0.12$).

the DLHS-3 found that only about 63% of urban households use clean cooking fuel. The DLHS-3 reports a similar number of households using LPG (56%) but less than half as many urban households using kerosene (3.7%). The DLHS-3 results are robust to excluding households that report kerosene use from the analysis.

III. Results

Are households in urban India with a boy first child more likely to use clean cooking fuel than those with a girl first child? If so, is this result spuriously driven by nonrandom selection?

Figure 2 offers a more detailed answer than the cross-tabulations did. The figure plots local polynomial regressions of clean cooking fuel use on the wealth index included in the DHS, separately for households with male and female first children. Unsurprisingly, very few poor households use clean cooking fuel, and almost all rich households do. Importantly for our analysis, the line for households with first boys is above the line for households with first girls, suggesting that our finding is not driven by a correlation between child sex and wealth.[11] The upward slope of the plots also makes clear that child sex is only one determinant of clean cooking fuel use; affordability and access also matter. Finally, the similarity of figure 2*A*, which includes women who have ever had a terminated pregnancy, and figure 2*B*, which does not, indicates that our results are not driven by selection.

Before presenting regression results, we verify that households with first boys and first girls are similar on observable characteristics. Table 2 demonstrates that they are: in a range of demographic, social, and economic variables, there are no differences in either sample.[12] Section III.B will extend this verification with placebo tests for "effects" of first children on a range of economic assets.

A. Regression Results

Table 3 verifies the statistical significance and robustness of figure 2 by presenting estimates of equation (1) but without covariate controls. Panel A presents

[11] The gray bands are 95% confidence intervals for the conditional mean of clean cooking fuel use, given wealth. In the most densely populated middle of the wealth distribution, these confidence intervals do not overlap. If these confidence intervals never overlapped, then one could conclude that clean cooking fuel use were statistically significantly more common among households with first boys at all wealth levels, a stronger claim than this article's and an unlikely one, given the first-order effect of wealth.

[12] One difference is that households with first boys have fewer children on average; this likely effect of stopping rules is discussed in more detail below and controlled for in regressions.

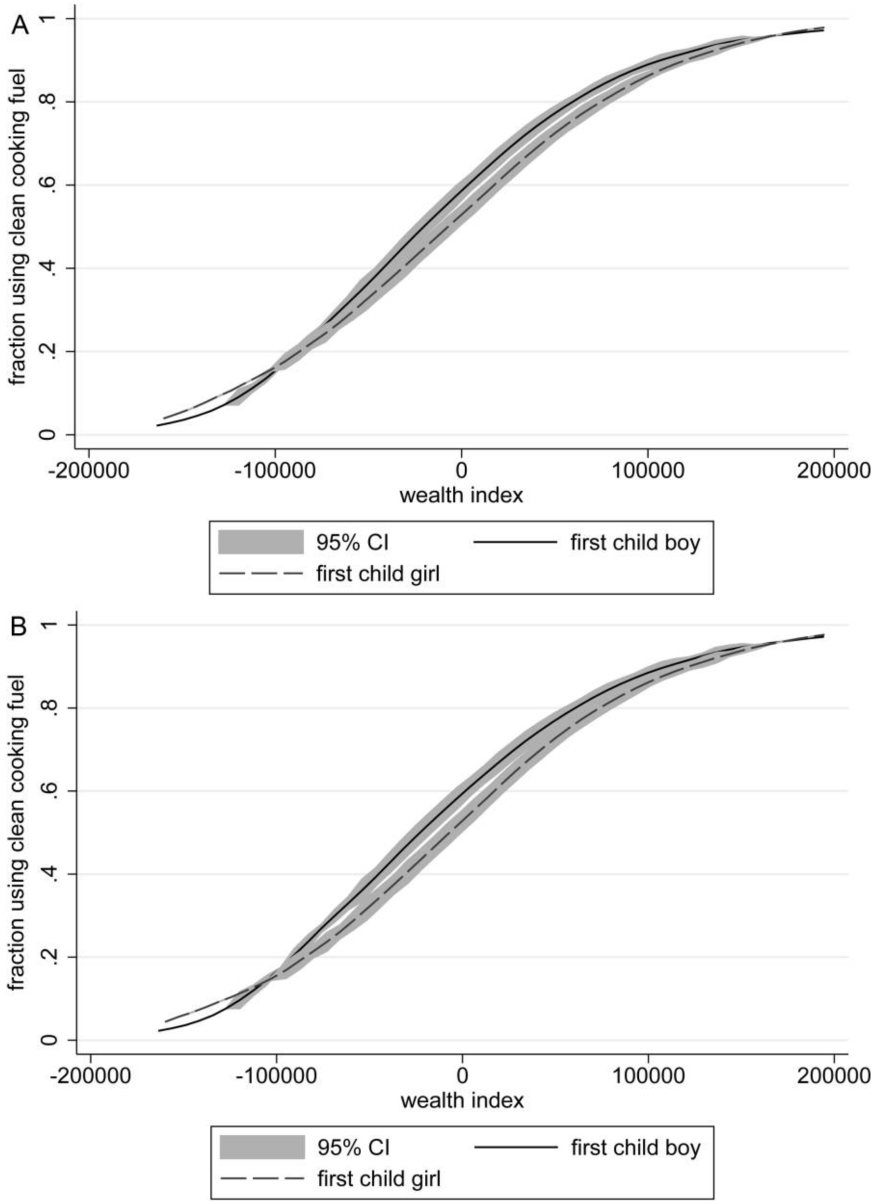

Figure 2. Clean fuel use by sex of first child: local polynomial regression, NFHS-3. *A*, Full sample; *B*, restricted sample: no terminated pregnancy. Urban sample with oldest child under 5; sampling weights used.

TABLE 2

BALANCE OF COVARIATES IS CONSISTENT WITH "RANDOMIZATION," URBAN NFHS-3

	Oldest Born since 1996			Oldest Child under 5		
	Girl First	Boy First	*t*	Girl First	Boy First	*t*
Poverty ration card	.14	.14	.53	.12	.13	.40
Open defecation	.24	.23	−.99	.26	.24	−1.22
Muslim	.16	.17	.17	.16	.16	−.46
Scheduled caste	.16	.16	−.47	.17	.17	−.04
Scheduled tribe	.03	.03	−.29	.03	.03	.73
High caste	.41	.41	−.67	.42	.40	−1.08
Mother's age at marriage	19.12	19.20	.91	19.53	19.65	.97
Mother's years of school	3.29	3.37	1.40	3.44	3.52	1.04
Mother illiterate	.22	.21	−1.05	.17	.18	.43
Age of household head	42.93	43.49	1.66	43.57	44.39	1.52
Household size	6.17	6.11	−.72	6.19	6.16	−.22
Household women	1.54	1.56	1.31	1.66	1.68	.79
Household children	1.37	1.26	−5.19	1.47	1.40	−2.36
Nuclear family	.43	.41	−1.44	.35	.33	−1.27

TABLE 3

OVERVIEW: SEX OF FIRST CHILD AND CLEAN COOKING FUEL USE, URBAN AND RURAL NFHS-3

	(1)	(2)	(3)	(4)	(5)
Ever terminated[a]	Yes and no	Yes and no	Yes and no	No	No
Place	All India	Urban	Rural	All India	Urban
A. Oldest child born since 1996:					
First boy	.0111†	.0254*	−.000938	.0168*	.0297*
	(.00628)	(.0113)	(.00484)	(.00695)	(.0126)
n	31,688	14,090	17,598	26,205	11,448
B. Oldest child under 5 at time of survey:					
First boy	.0153	.0355*	−.00163	.0164	.0380*
	(.00997)	(.0166)	(.00769)	(.0106)	(.0182)
n	13,808	6,318	7,490	11,793	5,337

Note. Standard errors (in parentheses) clustered by primary sampling unit; survey weights used; two-sided *p*-values.

[a] Indicates reporting ever having had an intentionally or unintentionally terminated pregnancy.

† $p < .10$.

* $p < .05$.

results using the larger sample of children born since 1996, 10 years before the survey, and panel B presents results using the sample of children born in the 5 years before the survey. Standard errors are clustered by PSU.

The main result is visible in table 3 columns 2 and 5: urban households with a boy first child are more likely to use clean cooking fuel by a little over 2 percentage points, an effect size that will be approximately consistent across specifications and data sets. In the larger subsample of panel A, this result is also statistically significant for all of India (cols. 1 and 4); in panel B, the result is quantitatively similar but imprecisely estimated. For rural India (col. 3), the coefficient is very close to zero; rural India will be dropped from the remainder of this article, which will focus on the main result for urban India.

Table 3 columns 4 and 5 check for a confounding effect of selection by excluding from the sample women who have ever had a pregnancy intentionally or unintentionally terminated, that is, other than in a birth. If anything, the coefficients are larger, suggesting that selection was not responsible for the result. If, instead, the whole sample is used and the indicator for having a boy first child is interacted with a demeaned indicator for never having had a terminated pregnancy, the coefficient on having a boy is unchanged ($\hat{\beta} = 0.035$, $t = 2.14$), and the interaction is not distinguishable from zero ($\hat{\beta} = 0.013$, $t = 0.31$).

1. Urban Results with Controls

Table 4 presents estimates of equation (1) including the vector of controls. The sample size decreases slightly from 14,090 to 14,075 due to 15 observations lacking some covariate data. As before, in all specifications household with a male first child are about 2 percentage points more likely to use clean cooking

TABLE 4

SEX OF FIRST CHILD AND CLEAN COOKING FUEL USE, WITH CONTROLS, URBAN NFHS-3

	Oldest Born since 1996		Oldest under 5 Years Old					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Ever terminated[a]	Yes and no	No	Yes and no	Yes and no	Yes and no	No	No	No
Boy	.0125†	.0147†	.0260*	.0256*	.0315*	.0303*	.0299*	.0252†
	(.00753)	(.00855)	(.0117)	(.0130)	(.0161)	(.0126)	(.0138)	(.0134)
Height for age[b]					−.00633	−.00865	−.00618	−.00241
					(.00465)	(.00699)	(.00511)	(.00494)
Height for age[b] × boy					.00446			
					(.00898)			
Controls	✓	✓	✓	✓	✓	✓	✓	✓
Urban PSU fixed effects								✓
n	14,075	11,433	6,307	5,220	5,220	5,326	4,377	4,377

Note. Standard errors (in parentheses) clustered by PSU (primary sampling unit); survey weights used; two-sided *p*-values. Controls are the DHS wealth index as a quadratic polynomial; a set of three standard-of-living indicators from the NFHS; number of living children of the child's mother as a quadratic polynomial; number of household members; a set of indicators for the child's mother's 5-year age band; a set of 12 indicators for the child's mother's years of education; mother's age at marriage; a set of five indicators for town or city size; and indicators for being a nuclear family, scheduled caste household, or scheduled tribe household.

[a] Indicates reporting ever having had an intentionally or unintentionally terminated pregnancy.

[b] Height-for-age *z*-score of the oldest child.

† $p < .10$.

* $p < .05$.

fuel. Excluding women who have ever had a terminated pregnancy does not eliminate this result; if anything, it becomes stronger. Results with urban PSU fixed effects in column 7 demonstrate that local geographic heterogeneity is not responsible for the correlation.

A particularly important control variable is the height-for-age *z*-score of the first child, computed using the 2006 World Health Organization reference population.[13] Child height is often used as a summary measure of long-term and early life health and nutrition. One hypothesis mentioned in the introduction was that the effect of having a boy on clean cooking fuel use reflected a desire to invest in the health of boys. If controlling for the height for age of the firstborn—as a marker of investment in the child's health—had importantly diminished the association between sex and clean fuel, then this might have been evidence for that hypothesis. However, there is essentially no change in the coefficient on child sex. This is also true in table 4 column 5, where the effect of height for age is allowed to interact with the child's sex; not only does this account for the different height trajectories of boys and girls, but the lack of an interaction suggests that the result is not found only in households that have otherwise invested in their boys' health.

2. Robustness Checks

Fraction of Children Who Are Male

Because of concerns about the endogeneity of sex composition among a household's children, we have focused on the sex of the first child.[14] However, the result is similar if the fraction of children who are male is used as the independent variable instead. Controlling nonparametrically with indicators for the total number of children, with or without controls for household wealth, across various specifications, changing the fraction of children who are sons from 0 to 1 is linearly associated with an approximately 3–5 percentage point increase in the likelihood of using clean cooking fuel (t-statistics from 2.09 to 2.90).

Of the approximately 14,000 households in the sample, about 6,000 of these have had exactly one child. For this smaller sample, the sex of the first child is identical to the fraction of the children who are male. Using only this

[13] Height is available only for children under 5 in the NFHS-3 (hence our emphasis of this subsample) but not for all of those, as the fall in sample size from cols. 3 to 4 and cols. 6 to 7 in table 4 shows.

[14] Ever having had a son would not be a suitable indicator: 82% of households in our sample eventually have had at least one son (and not all of the women in the data have finished having children). Note, unsurprisingly, that the first child has an effect on the total number of daughters: households with a first daughter rather than a first boy have about one more daughter on average, with or without controls for total number of children ($t \approx 60$–70).

subsample, we find essentially identical parameter estimates: 0.016 and 0.026 for the samples of children born in the last 10 and 5 years, respectively. Of course, with a sample of half the size, standard errors are much larger: 0.018 and 0.020. However, the similarity of the point estimates emphasizes the robustness of our result.

Interaction with Female Literacy

We might expect that literate mothers (45% of our sample) would have more exposure and connections and, perhaps, would be both more likely to have a sex-selective abortion and more likely to have access to clean cooking fuel. To test for this possibility, we interact the indicator for having a male first child with an indicator for the mother's literacy. If the association between child sex and clean cooking fuel use were seen only among literate women, then this sort of endogeneity could be a concern. However, the interaction is not statistically significant ($t = 0.89$) and has the opposite sign (the association is, if anything, greater for illiterate women). The coefficient on child sex is unchanged ($\hat{\beta} = 0.033$, $t = 2.01$).

Interaction with Urban PSU-Level Sex Ratio

As a further test to rule out a driving role of sex selection, we take advantage of geographic heterogeneity in desire for or access to sex-selective abortion. For each PSU, we compute the fraction of children under 5 born alive who are boys, a variable with a mean of 0.525. Then, we interact this PSU-level variable with the household-level indicator of sex of the first child. If the result were driven by sex-selective pregnancy termination, and if sex-selective pregnancy termination varied geographically at the PSU level, then we would expect the effect of sex on clean fuel use to be concentrated in places where more boys were born. However, the coefficient on boys is unchanged, if less precisely estimated due to multicollinearity ($\hat{\beta} = 0.022$, $t = 1.68$), and the interaction with the PSU-level fraction of boys is not statistically significant ($t = 0.13$).

Restriction to Mothers of Living Children

The sample of children born in the last 5 years is, by our design, also the sample of children for whom height is available, allowing the control for child's human capital in table 4. Therefore, these children survived until the survey. In contrast, the sample based on children born in the last 10 years contains all children, including those who have subsequently died. If, as a robustness check, children who died before the survey are excluded, the results are stronger, with the coefficient increasing from 0.025 to 0.030 ($t = 2.6$) without the controls and from 0.0125 to 0.0150 with them ($t = 2.0$). Thus, mortality selection is not responsible for our positive coefficient.

B. Falsification Tests: Placebo Assets

If there is a causal effect of having a boy first child on clean cooking fuel use, is it unique, or are there similar effects on ownership of other household assets? If our finding were driven by omitted heterogeneity in wealth or status for which we have not been able to account, then we would expect it to predict ownership of other goods. However, if the result indeed reflects the nonunitary household mechanism we describe, then we would expect clean cooking fuel to be special and other outcomes not to be "caused" by having a first son.

Figure 3 presents *t*-statistics on the indicator for a male first child from 19 "placebo" estimations of equation (1) (without controls). We substitute, instead of clean cooking fuel use, 18 alternative dependent variables, including the NFHS's wealth index computed from other assets.[15] The vertical position of the dependent variable names is not significant; they are separated merely for clarity. The *t*-statistics show an expected unimodal and symmetric distribution around zero.[16] Only clean cooking fuel use has a *t*-statistic greater than 2 in absolute value; the next largest is owning a computer at -1.37, that is, statistically insignificantly favoring households with a girl first child.

Figure 3*B* replicates these results excluding first children of mothers who have ever had a terminated pregnancy. The results are qualitatively similar. Some of the *t*-statistics change, as would be expected if they were simply random noise: across figures 3*A* and 3*B*, the *t*-statistics (excluding clean cooking fuel) have a Spearman's rank correlation of 0.75.

These null placebo results for other economic assets further rule out an alternative mechanism for a causal effect of a girl baby: making a household poorer, or making it feel poorer, either because it has more children in an attempt to have a boy or because it anticipates paying a future dowry when the girl gets married. Like a sex-selective abortion confound, this alternative mechanism would have caused the coefficient on boy_{is} to be positive in these placebos.

[15] The dependent variables are the wealth index, has ration card, scheduled caste, scheduled tribe, Muslim, mattress, pressure cooker, chair, cot or bed, table, electric fan, black-and-white TV, color TV, sewing machine, computer, water pump, thresher, tractor, and clean cooking fuel. This includes the full set of India-specific assets (items s47b through s47w) in the NFHS.

[16] Shapiro-Wilk ($p = .98$), Shapiro-Francia ($p = .77$), and skewness/kurtosis ($p = .68$) tests all fail to reject that the *t*-statistics have a normal distribution.

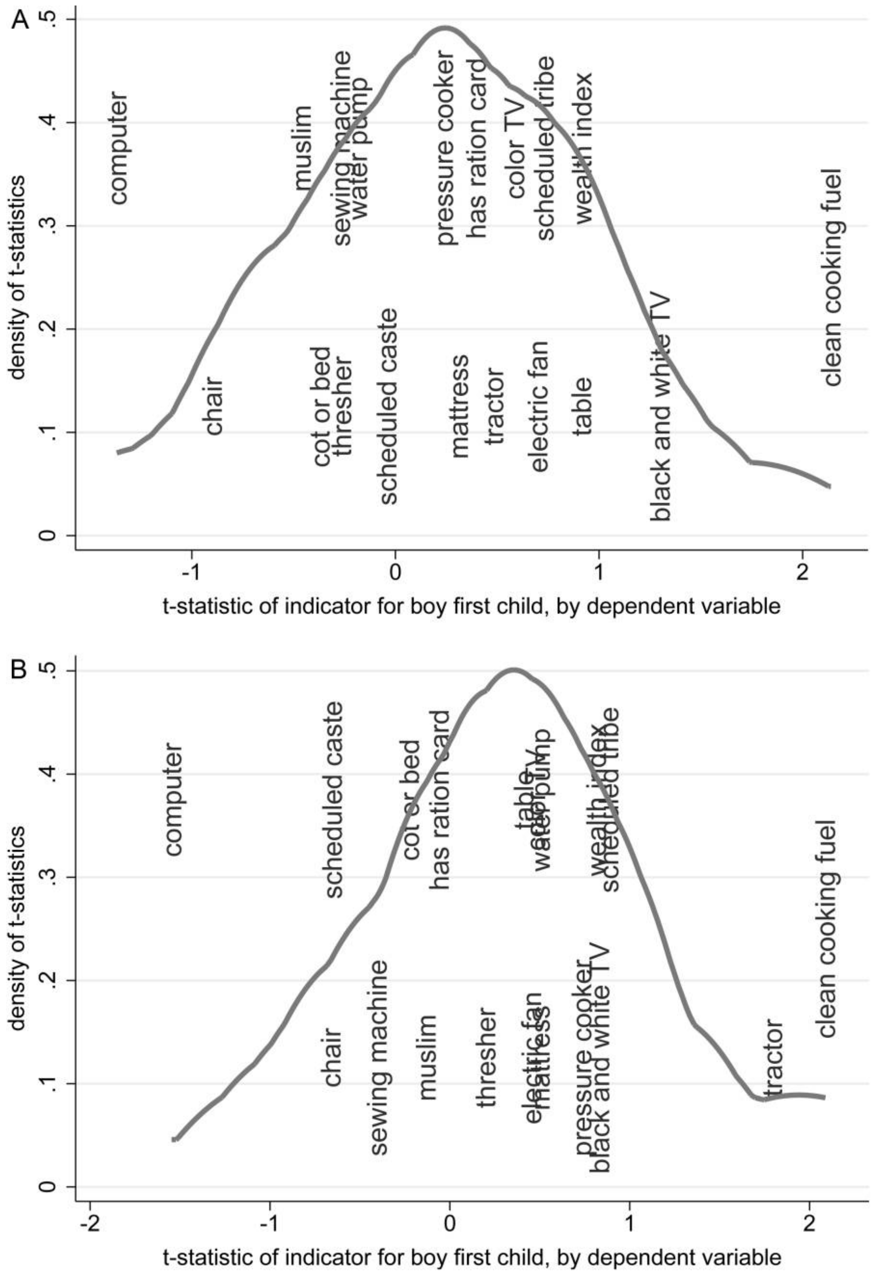

Figure 3. Falsification test: 18 alternative dependent variables, NFHS-3. *A*, Full sample; *B*, restricted sample: no terminated pregnancy. Each point is the *t*-statistic on a separate regression of the named dependent variable on an indicator for a boy first child. Urban sample with oldest child under 5; sampling weights used. Vertical position of dependent variable names has no significance; they are separated for clarity.

TABLE 5
REPLICATION: SEX OF FIRST CHILD AND CLEAN COOKING FUEL USE, NFHS-2

	Urban				All India	
	(1)	(2)	(3)	(4)	(5)	(6)
Ever terminated[a]	Yes and no	Yes and No	No	No	Yes and no	Yes and no
Boy	.0119	.0301†	.0103	.0239	.00736	.00822
	(.0174)	(.0155)	(.0200)	(.0175)	(.00977)	(.00951)
District fixed effects		✓		✓		✓
n	3,302	3,302	2,837	2,837	10,475	10,475

Note. Standard errors (in parentheses) clustered by primary sampling unit; survey weights used; two-sided p-values.
[a] Indicates reporting ever having had an intentionally or unintentionally terminated pregnancy.
† $p < .10$.

C. Replications with Other Data

This section presents two replications of the main result of the article. In all three data sets, among urban Indian households, having a male first child is associated with an approximately 2 percentage point increase in the likelihood of clean cooking fuel use, a result that is not driven by nonrandom selection of pregnancies into births.

1. NFHS-2 from 1999

The NFHS-2 is a similar data set to the main data of the article, collected 7 years earlier. Access to clean cooking fuel expanded substantially in the 1990s in India as a result of market liberalization. Because in this article we are limiting our analysis to children born since 1996 (10 years before the NFHS-3), this analysis has the smallest sample used in this article and will not be able to achieve similar levels of statistical precision. However, unlike the NFHS-3, the NFHS-2 permits district fixed effects.

Table 5 presents results of estimating equation (1) using the NFHS-2. In all urban cases, a similar coefficient—an effect of around 2 percentage points—is found, but only in column 2, with the largest urban sample and district fixed effects, is it estimated sufficiently precisely to be statistically significant.[17]

2. DLHS-3 from 2008

Results from the DLHS-3 are similar but reflect a differently constructed birth history. Unlike in the NFHS birth histories, the DHLS-3 birth history only goes back a few years, to 2004, restricting the usable sample. Table 6 col-

[17] As an informal statistical significance test, if these six coefficients were all zeros, and if they were independent (which they are clearly not), there is only a 0.015 probability of them all being positive.

TABLE 6
REPLICATION: SEX OF FIRST CHILD AND CLEAN COOKING FUEL USE, URBAN DLHS-3

	Born in 2004						Born in 2007
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Ever terminated[a]	Yes and no	Yes and no	Yes and no	No	Yes and no	No	Yes and no
Boy	.0202	.0215†	.0210†	.0189	.0248†	.0186	.00228
	(.0154)	(.0115)	(.0114)	(.0126)	(.0146)	(.0123)	(.0100)
Wealth		.231*	.222*	.223*		.203*	.227*
		(.00612)	(.00644)	(.00690)		(.00928)	(.00566)
Wealth^2		.0635*	.0565*	.0552*		.0593*	.0600*
		(.00290)	(.00301)	(.00333)		(.00464)	(.00272)
Controls			✓	✓		✓	✓
District fixed effects					✓	✓	
n	4,053	4,053	4,053	3,402	4,053	3,402	5,411

Note. Standard errors (in parentheses) clustered by primary sampling unit; survey weights used; two-sided p-values.
[a] Indicates reporting ever having had an intentionally or unintentionally terminated pregnancy.
† $p < .10$.
* $p < .05$.

umns 1–6 use the furthest-back possible sample: children born in 2004. A similar, if not always precisely estimated, effect is found: a first boy increases the likelihood of clean cooking fuel use by about 2 percentage points.[18] Controls suggest that this effect is not due to fixed heterogeneity across districts or household wealth or demographic differences. Coefficient estimates are very similar if women with terminated pregnancies are excluded (cols. 4 and 6) but lose statistical significance due to the smaller sample.

Table 6 column 7 changes the sample from children born in 2004 to households with first children born in 2007, immediately before the survey. No effect is seen. One interpretation is that this is random noise. Another, however, is that this is evidence of Granger causality of child sex on clean cooking fuel use. According to Granger causality, causes should precede effects. If child sex were truly causing clean cooking fuel use, the effect would take some amount of time to unfold. In contrast, if the results were due to selection—either via abortion or lethal smoke—then the correlation would be visible immediately upon birth.[19]

[18] As discussed in Sec. II.B, the DLHS-3 finds a different number of kerosene users than the DHS. The arguably main result in table 6 col. 2 is similar if kerosene-using households are excluded ($\hat{\beta} = 0.021$, SE $= 0.011$).

[19] If, instead of splitting the sample up by cohorts, the full set of 4 available years is used (all children born since 2004), the coefficient is positive (0.002) but not statistically significant ($t = 0.33$), which could reflect sampling noise or could reflect the effect taking time to unfold in this young sample.

D. Mechanism: Women's Status or Investing in Children?

Recall our observation that, theoretically, child sex could influence fuel choice through desire to invest in the child or through women's status. If the sex of the first child influences cooking fuel type because of a first-stage effect on women's social status, then this first-stage effect should be observable. However, as discussed above, women's status is well known to be difficult to measure using survey questions: different people could describe similar situations differently. This section considers both survey-reported and physically measured indicators of women's status, as well as other indicators of investment in children.

1. Survey-Reported Decision-Making "Say"

Recall that figure 1 reported that women who had decision-making power over more situations were more likely to use clean cooking fuel. Is survey-reported decision-making power influenced by the sex of a women's first child? If so, it is not detectable in our data: among five decision situations asked about in the NFHS-3, the women in our sample with a firstborn boy do not report having say on more of the decisions than do women with a firstborn girl (born in last 10 years: $t = 0.95$; born in last 5 years: $t = 1.03$). Nor are women with a firstborn son more likely to report having say on at least one of the five issues.

However, women's intrahousehold status is notoriously difficult to measure, especially with self-reporting in large surveys (Dyson and Moore 1983). Kahneman (2000) has described in detail how survey measures common in applied economics could be influenced by "hedonic adaptation": people only seem to notice changes in well-being for a short time, before adapting back to a baseline. Relatedly, Das, Hammer, and Sánchez-Paramo (2012) have shown that richer and poorer mothers in India answer survey questions about children's health differently, presumably in part due to different standards of "normal." Zimmermann (2012) may find a transient effect of having a boy on reported social status in part because hedonic adaptation makes the improvement less noticeable over time.

2. Women's BMI as a Measure of Intrahousehold Status

One solution is to find an objectively measurable indicator of women's status, even one imperfectly correlated with status. We follow Coffey, Khera, and Spears (2012) in studying adult women's body mass index (BMI) as an indicator of social status. They find that lower-ranking daughters-in-law in joint Indian households have lower BMI than higher-ranking daughters-in-law in the same households, an explanation for their main finding that women's status influences children's height and stunting.[20] BMI is a plausible indicator of social status because women of childbearing age in Indian households often eat last, limited to the food remaining after other household members have eaten, and expend much energy in household work (Jeffery et al. [1989] document these patterns in detail for rural India).

Here, we use the same urban sample as throughout this article, relying on the randomness of child sex for identification. Table 7 reports results of regressions of the form

$$\text{BMI}_{ic} = \alpha_c + \beta \text{boy}_{ic} + X_{ic}\theta + \varepsilon_{ic}, \tag{2}$$

where i are adult women in households in our sample; c are clusters, here urban PSUs; *boy* is an indicator that the firstborn child is a male; and X is a vector of controls. The controls include the women's height as a quadratic polynomial and a set of controls intended to rule out a direct effect of subsequent fertility: total children ever born as a quadratic polynomial, an indicator for only having one child, and an indicator for being currently pregnant.[21]

On average in this sample, women with a first boy have a BMI about 0.2 greater in the sample of women whose first child was born in the last 10 years.[22] This result does not appear to be a mechanical result of subsequent fertility.[23] Although it would be outside of the scope of this article to fully document and explore this association, the result is consistent with an effect of the sex of the first child on women's intrahousehold status.

3. Other Investments in Boys and Girls

Do households invest more in boys than in girls? This is almost certainly true in India to some extent (Jeffery et al. 1989), but the magnitude in survey data has been debated in the literature (e.g., Deaton 1997, 2003). Also, households might favor boys' education (e.g., Coffey [2012] finds that migrant households in a sample from rural north India are more likely to send boys to school) without necessarily investing in their health, especially if they do not fully

[20] They also find that these women spend less time outside of the home in a typical day.

[21] Height, in the denominator of BMI, primarily reflects early life health and net nutrition, not conditions after a woman's marriage.

[22] Note that the coefficient is 0.11 and not statistically significant ($t = 0.78$) in the sample of women with a child born in the last 5 years, consistent with a causal effect that unfolds over time.

[23] It is not inconceivable that using clean cooking fuel could raise a woman's BMI by reducing work requirements or the disease burden, although we find no evidence of this.

TABLE 7
WOMEN'S BODY MASS INDEX, URBAN NFHS-3 OLDEST BORN SINCE 1996

	(1)	(2)	(3)	(4)	(5)
Ever terminated[a]	Yes and no	No	No	No	No
First child boy	.207*	.230*	.229*	.190†	.230†
	(.104)	(.117)	(.117)	(.117)	(.127)
Height			.683†	.709†	.643†
			(.402)	(.401)	(.390)
Height^2			−.00216	−.00225†	−.00216†
			(.00133)	(.00132)	(.00128)
Only one child				−1.213***	−.492
				(.320)	(.378)
Children ever				−1.529***	−.187
				(.452)	(.567)
Children ever^2				.143*	−.0120
				(.0657)	(.0861)
Currently pregnant				−0.0879	.377†
				(.177)	(.193)
Urban PSU fixed effects					✓
n	14,040	11,341	11,341	11,341	11,341

Note. Standard errors (in parentheses) clustered by PSU (primary sampling unit); survey weights used; two-sided *p*-values.
[a] Indicates reporting ever having had an intentionally or unintentionally terminated pregnancy.
† $p < .10$.
* $p < .05$.
*** $p < .001$.

understand health inputs. In our urban sample of firstborn children under 5, we do not see consistent evidence of disproportionate investment in boys's health: boys are not more likely to have a health card (0.23) or polio vaccine ($t = -0.37$), but 92% of them have had a BCG vaccine, compared with 90% of girls, a statistically significant difference ($t = 2.76$). Of the 594 children in the sample whose mothers report them having had diarrhea in the past 2 weeks, boys are 3.7 percentage points (or 5.6%) more likely to have been given some kind of treatment, but this difference cannot be distinguished from zero ($t = 0.66$). Of course, clean cooking fuel is only likely to be chosen as an investment in health if parents believe that traditional smoke is harmful, a minimum requirement that is only true of 80% of women surveyed in the 2005 India Human Development Survey.

IV. Discussion and Conclusion

This article has documented that urban Indian households with a male first child are approximately 2 percentage points more likely to use clean cooking fuel than comparable households with a female first child. This correlation appears to represent a causal effect: it is not driven by nonrandom pregnancy termination or by wealth and is unique to clean cooking fuel use among assets.

This article is unable to definitively distinguish between greater desire to invest in the health of boys and increased status of women as mechanisms of the effect; both may be at work. However, there is evidence for a role of women's status: having a first son also increases a women's BMI; moreover, the effect of fuel choice is not concentrated in households that have also invested disproportionately in sons, as measured by the child's height. If women's status is part of the explanation, however, then this finding illustrates two important constraints on children's health in India.

First, clean cooking fuel would be another example of a preventive investment in health that is subjectively valued by households primarily as a consumption good. To the list of hand-washing soap (Burke 1996) and latrines (Black and Fawcett 2008), purchased as conveniences and symbols of modernity, would be added clean cooking fuel, appreciated for its convenience and comfort. Household construal of preventative health goods as consumption rather than investments could be one explanation for the puzzle of low preventative health investment described by Dupas (2011).

Second, women in India typically have both low intrahousehold status and most of the responsibility for child care. This pattern could help explain very poor child health outcomes in India (Ramalingaswami, Jonsson, and Rohde 1996). For example, Jeffery et al. (1989) describe young women in the household having low entitlement to food, with little difference even when they are pregnant (78–79). In this article, a by-product of Indian women's inability to cause households to switch to clean fuel is that children—while in utero or spending time with their mothers—are often exposed to dangerous levels of indoor air pollution.

However, whether our findings are explained by women's status or by intentional investment in sons, they are further evidence of inequality along gender lines within nonunitary Indian households. Moreover, indoor air pollution represents an additional input through which boys' health could be made better than girls'.

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