Intergenerational effects of women's status: Evidence from joint Indian households

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Abstract

The hypothesis that a woman's social status has intergenerational effects on the human capital of her children has featured prominently in development policy and social science. Our paper is the first to econometrically identify such an effect. We exploit an institutional feature of joint households in rural India: women married to the younger brother are assigned lower social rank beginning at marriage than women married to the older brother in the same household. Children of lower-ranking mothers accumulate less health and human capital, reflected in shorter heights and higher mortality, compared to children of higher-ranking mothers in the same household.

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1 Introduction

Economists and other social scientists have long debated the effects of social hierarchy on health. In particular, researchers have asked whether and how women's social status and empowerment shape the health and early-life human capital formation of their children (Thomas, 1990). This question is especially relevant for developing economies, where gender inequalities are often large, where women are often responsible for essentially all aspects of child care, and where low human capital is often a constraint on human development (Duflo, 2012). Many development programs and policies are built on the assumption that socially empowering women will improve their children's outcomes (World Bank, 2001). This view has substantially influenced the non-profit sector as well (Gates, 2014), such that in some cases promoting women's status is seen as a better approach to improving child welfare even than attempting to feed children directly.[1] Although a convincing estimate of an effect of a woman's social status on her children's outcomes therefore would be of clear importance, no prior paper in economics has isolated such an effect.

Several challenges combine to make causal effects of a woman's status on her children's health and human capital difficult to study. First, women of different social status often differ in other relevant dimensions of human capital, personal resources, or household wealth (Strauss and Thomas, 1995). Second, women's status and empowerment are difficult to measure. Survey-reported indicators of women's status used by economists, such as "say" in household decisions, can suffer from considerable biases.[2] Finally, programs or events that

[1]Melinda Gates (2014) writes in *Science* that "there are strong associations between women's empowerment and specific health and development outcomes" including "maternal, newborn, and child health [and] nutrition." Ruel et al. (2013) write in *The Lancet* that "women's empowerment" could be among "nutrition-sensitive interventions and programmes" that succeed where directly providing children with food has not. Supplementary Appendix A5 details further examples that women's empowerment is frequently targeted by development programs and policies, often by focusing on social status directly, rather than women's other resources.

[2]For example, in the 2005 Indian DHS, whether a woman agrees with the statement "wife beating could be justified" correlates positively with wealth. In this case, higher status women may be more likely to interpret an instance of maltreatment as "beating" than lower status women.

influence women's status may also have other direct effects on child outcomes. For example, Jensen and Oster (2009) estimate an effect of cable TV in India on women's status and on child education, but they note that, although isolating empowerment effects would be an important finding, they cannot attribute the effects on child schooling to the improvement in women's empowerment: TV could have influenced schooling through mechanisms other than women's status, such as perceived returns to schooling. Other studies of women's status and child outcomes in the economics literature face similar limitations.[3] Thus, although cleanly isolating an effect of a woman's social status on her children's outcomes is a clear goal of the literature, it remains unfulfilled.

This paper overcomes these three identification challenges by exploiting a unique institutional feature of joint patrilocal households in rural India. Joint patrilocal households are ones in which adult sons live together with their parents, their wives, and their children: in rural India, women move to their husbands' villages upon marriage, and typically into their husbands' parents' homes. In this context, at the time of marriage, women married to the older son are assigned higher social status than women married to the younger son. As a result, the lower-ranking woman may eat last and eat less, may engage in more strenuous work (even in pregnancy), may engage in work harmful to health (such as using an indoor traditional cookstove), and may face worse treatment by her in-laws who head the household in many small but palpable ways, such as being required to sit on the ground while other adults use chairs. Because this difference in women's status occurs within households,

[3]Similarly, Thomas (1990), in an important early study on this topic, shows an effect of a mother's economic resources on her children's health, but does not separate a special effect of a mother's ability to purchase goods from any social or bargaining consequences it may entail. Miller (2008) shows an aggregate effect of women's suffrage on child health, but does not study effects of a mother's status on her own children. Kishore and Spears (2014) find an effect of women in India having a son rather than a daughter on clean cooking fuel use, but cannot separate the mechanism of women's status from households' preference to invest in a son's health; they do not show consequences for children. Udry (1996) documents an effect of women's intrahousehold status on agricultural outputs, but does not study intergenerational consequences for children. Other research studies women's resources that are important but distinct from social status, such as money (Schaner, 2015; Duflo, 2003) or education (Breierova and Duflo, 2004).

we can identify effects on children of an objectively measurable source of variation in their mothers' status, while holding much about the environment to which children are exposed constant. Importantly, in this context, marriage markets do not sort and match women of different pre-determined "quality" into different status ranks; rather, status is assigned when a woman marries into a joint household. We empirically verify this with data on arranged marriage.

Using this strategy, we find that children of the lower-ranking mother are about a third of a height-for-age standard deviation shorter than their cousins born to the higher-ranking mother within the same joint household. This is an importantly large effect; to put it in context, it is nearly as large as the average difference between high and low caste children, and is large enough to expect long-term effects on human capital.[4] This is true irrespective of the heights of the child's parents and irrespective of the child's birth order in the joint household — that is, birth order among both its siblings and cousins. Thus our finding demonstrates, for the first time, an important effect on children's human capital of their own mother's social status, as distinct from treatments such as money transfers that may provide women with resources while simultaneously or as a consequence improving her social status.

In the second part of the paper, we document a plausible and independently important mechanism for our effect: that maternal net nutrition is a consequence of women's social status.[5] A mother's low intrahoushold rank – which in this context is associated with restricted

[4]Our effect of 0.35 height-for-age reference standard deviations is about three-fourths as large as the average difference in child height between high and low caste Indian children in the same data. It is about half as large as the effect of a widely-studied field experiment which provided early-life nutrition to some children in Guatemala (Maluccio et al., 2009), which has been followed up upon to reveal effects on adult human capital.

[5]This is an important but subtle difference between our study and those in the literature that consider programs that provide women with resources, such as cash transfers. A cash transfer program, for example, provides a resource to women that could have its own direct consequences for children and could influence her social status; in contrast, we study a difference in social status which, we show, has as a consequence a difference in maternal net nutrition.

access to food and intense physical work requirements, including during pregnancy – lowers her body mass after marriage relative to higher-ranking mothers, decreases her babies' birth weight, and consequently increases neonatal mortality. In data that we newly collected from a government hospital in a rural district of north India, we find that children of lower-ranked daughters-in-law weigh less at birth than children of higher-ranked daughters-in-law. Birth weight is highly dependent on maternal nutrition (Yaktine and Rasmussen, 2009), and is known to be a determinant of child height (Adair, 2007).

We provide a collage of evidence supporting the identifying assumption that the women who gain low status were not different prior to the status assignment, which occurs when the women co-locate with their sisters-in-law. First, we show that lower-ranking mothers and their husbands are both indistinguishable from their higher-ranking counterparts within the same household on height, education, and other dimensions of human capital predetermined before marriage. This balance is expected because studies of the institution of arranged marriage in India confirm that it does not sort bride quality by husband birth order. Second, we conduct an audit study of marriage advertisements following Banerjee et al. (2009), and we show that the complex and constrained optimization in the arranged marriages that we exploit considers many dimensions, but does not respond to a potential groom's age rank among brothers. Third, we provide evidence that arranged marriages would not respond to the groom's birth order even if it were revealed in marriage ads by conducting a survey experiment among Indian participants using hypothetical ads.

Fourth, we exploit the fact that many joint households split into nuclear households before children are born. We verify that the effect only occurs if the children and low-ranking mother actually live in a joint household at the time of birth. We perform a falsification test exploiting a separate data source that has longitudinally tracked households, and in which we observe households splitting. Consistent with our exclusion restriction, we observe no effect among paternal cousins who are not co-resident — that is, there is no effect among children born to the mothers who would otherwise be low-ranking by our definition, but who were born after the institution enforcing rank (meaning, the joint household) was dissolved.

This paper contributes to several literatures in economics. First, we address a significant open empirical question by providing the first well-identified estimates of an effect of a women's social status on her own children's health. Second, by linking the difference we identify in height outcomes with maternal nutrition and birth weight, we contribute to the active literature on very early life origins of economic disadvantage (Black et al., 2007; Maluccio et al., 2009; Aizer and Currie, 2014), and particularly to studies in economics on physical height as a measure of human capital (Case and Paxson, 2008; Steckel, 2009; Spears, 2012). Finally, we add to a growing literature that documents large effects of social and household institutions, especially in developing economies, where the non-unitary structure of households has important economic implications (Vogl, 2013; Bertrand et al., 2003). This paper is also relevant to the many development programs and policies, including those implemented and funded by leading organizations, that intend to improve children's outcomes by improving their mothers' social status.[6]

This paper is organized as follows. First, section 2 provides further background on the institutional features of joint households in rural India that we exploit. Then, section 3 presents our empirical strategy and data; section 3.2 verifies that children of higher- and lower-ranking mothers are balanced on their mothers' and fathers' pre-marriage human capital and other properties. Section 4 presents evidence of an effect of intrahousehold rank on other measures of women's status, in the spirit of a first stage. Section 5 presents our main result: within the same household, children of the lower-ranking mother are shorter than their cousins. This difference is seen only when adult brothers' children live together in a joint household, not when nuclear families live separately, as we show in section 5.2.

[6] Examples of such programs and policy-makers' beliefs that motivate them are detailed in Supplementary Appendix section A5.

Section 6 presents evidence that maternal nutrition is one important mechanism through which women's status affects child height in this context.

2 Background

2.1 Women's status in joint Indian households

The joint households that we study are ones in which a married woman lives with her husband, his parents, his brother, and his brother's wives and children. Figure 1 diagrams such a joint household. Approximately 7% of rural children under five live in joint households with two daughters-in-law, and about 8% live in households with two or more daughters-in-law. To put this in context, the absolute number of children living is such households is about seven million, which is one-third of the total number of children under 5 in the United States.

Joint households are characterized by patriarchy and by age-hierarchy: women are subordinate to men and younger members are subordinate to older members. In her husband's home, a young woman typically behaves in ways that both reflect and reinforce her low social position; Mandelbaum (1988) describes how a newly married woman is expected to "be most diffident, shy, and self-effacing...[keeping] her gaze lowered, her voice still, her features covered, and her whole presence unobtrusive" (5).

The status of a woman who marries into the household is derived in part by her husband's birth order (Singh, 2005), and this is reflected in expectations for her behavior. The demands of propriety are therefore even greater for a daughter-in-law who is married to a younger brother than for one who is married to the other brother. Jeffery et al. (1988) document that when a new daughter-in-law enters the joint household, the daughters-in-law who are already established in the household often "wield authority" over the new wife, "policing" her actions (30-31). Although relationships between the oldest brother's wife with her husband's younger brothers are often casual and friendly, the younger brother's wife is expected to send signals of respect and deference to all adult members of the household (Mandelbaum, 1988). These differences between the lives of higher- and lower ranked women in joint households have led Dyson and Moore (1983) to remark that "senior wives tend to dominate young in-marrying wives" (44).[7]

Not only does a woman's rank within the joint household influence the amount of stress she experiences, it also influences her food intake. In joint households, it is common for people to eat in the order of their social rank, with the household heads eating before their sons, who eat before the children, who eat before their mothers. Palriwala (1993) describes how, in the joint households she studied, "[t]he person who cooked and the youngest daughter in law, usually the same person, ate last. This acted against her...often there could be no vegetables or lentils left and she made do with a pepper paste and/or raabri. In a situation of deficit she went hungry when other household members did not have to " (60).

2.2 Marriage does not match bride quality to groom's birth order

An important feature of the institution that our identification exploits is the fact that, although arranged marriages are important decisions which incorporate many factors, arranged marriage markets do not match the bride quality to the groom's birth order. The 2005 India Human Development Survey found that 95% of marriages in rural India are arranged, meaning that the parents or extended family members of the bride and groom decide whether or not a couple will marry. Despite the fact that research by anthropologists and demographers on joint family life suggests that status differences between higher- and lower-ranking daughters-in-law are acknowledged in everyday life, it is nevertheless the case that the birth

[7] Dyson and Moore (1983) document that that gender hierarchies and intrahousehold rankings are more oppressive in north India than in south India. This implies that joint household hierarchies may not have the same effects on child health in south India. Supplementary Appendix table A2 tests whether the differences between the heights of children of higher- and lower-ranking mothers that we observe in the main sample of children living in joint households are more pronounced in the north; we find that this is indeed the case.

order of the husband-to-be is not an important factor for decision-makers in marriage markets. We verify this empirically with an audit study of marriage ads, and with a survey experiment that expands upon Banerjee et al. (2009). Extended-family decision-makers pay essentially no attention to a prospective groom's birth order in part because they see the low status of a daughter-in-law as temporary (the joint household in which the groom's age order matters eventually splits), relative to the other important factors governing arranged marriage. Indeed, it is true that the joint families that we study eventually dissolve; however, the differences in status that we exploit for identification have consequences for early-life health that a growing literature in economics suggests will have persistent implications for human capital. It is exactly during children's early-life critical period that households are likely to be joint. The fact that decision-makers are unaware of these lasting consequences of early-life health is analogous to other South Asian child rearing and marriage practices, studied in the economics literature by Jayachandran and Kuziemko (2011) and Vogl (2013)[8] respectively, in that all three institutions are widely recognized by South Asians to exist, in that the institutions have negative consequences for human capital, and in that these consequences are not understood by the people who practice them.

Marriage choices typically have more to do with economic and social incentives for the bride and groom's parents and extended families – who have final decision making authority – than with effects on the daily life of the bride-to-be. There is a large literature in economics that seeks to understand how marriages in India are arranged; in general, arranged marriages are highly constrained decisions, optimizing under many competing goals and constraints. For instance, Rosenzweig and Stark (1989) find that marriages to farther villages help families smooth consumption. Munshi and Rosenzweig (2006) discuss how marriage reinforces caste-based social networks that influence employment opportunities for men. In their research on

[8] Jayachandran and Kuziemko (2011) document that girls are breastfed less than boys, in order to promote the mother's fertility after a girl is born, in the context of son preference; Vogl (2013) finds that families of girls prefer to marry daughters in age order.

dowry in South Asia, Anderson (2003) and Rao (1993) both discuss the many characteristics that constrain the choice set of available grooms, including his caste, education, income, occupation, land holding, and city or village.

None of these papers – nor any other literature that we are aware of – claims or considers the possibility that decision-makers in the bride's family respond to the rank of the groom among his brothers. In particular, this factor is absent from a related literature in economics which documents the family's goals in marrying their daughter: to marry within sub-caste (Banerjee et al., 2009) and in the daughters' age order (Vogl, 2013). Indeed, in their study of bride and groom advertisements in Calcutta newspapers, Banerjee et al. (2009) note 38 characteristics that are considered important by families arranging marriages; the birth order of the groom is not among them.

2.2.1 Novel evidence on arranged marriages in India

To inform our analysis below, we begin by extending Banerjee et al.'s (2009) analysis of the content of marriage adds in India in two ways. First, we conducted our own catalog of classified ads for grooms in India, following that study. A table of characteristics included in the advertisements presented in Supplementary Appendix table A3 reveals the groom characteristics that matter in these marriage market. The distribution of characteristics mentioned in the ads we catalog is very similar to the advertisements reported by Banerjee et al. (2009). Importantly, the birth order rank of the husbands was *not mentioned in a single advertisement.*

Second, to further verify that age rank among brothers would not be important if it were included in marriage ads, we conducted a randomized survey experiment in which Indian participants chose which of two randomly generated ads describing grooms would be more attractive to brides' families. Among the many randomly assigned characteristics was whether the ad mentioned the groom's age rank among brothers, i.e. whether he was older or younger. Table 1 presents the results of the experiment. Although respondents' choices reacted to many other properties of the hypothetical marriage ads that are well-known to be important from the existing literature (education, skin tone, government employment, height, etc.), information on grooms' age rank among brothers was *precisely irrelevant* to respondents' choices. Supplementary Appendix section A3 and tables A4 and A5 present more details on the design of the experiment, on its sample, and on the results.

Our own fieldwork in rural Uttar Pradesh, a state where gender hierarchies are pronounced, suggests that the irrelevance of the groom's birth order to arranged marriage decisions may be because people see joint family life as temporary and as a small fraction of the duration of married life: joint families typically split into nuclear families after the household heads die (that is, the parents-in-law of the women we study). Children, however, tend to be born early in marriages before the groom's parents die, and marriage decision-makers are unaware of the lasting consequences for child human capital that we document. Considering this, it is not surprising that the balance table in section 3.2 finds that lower-ranking mothers are not disadvantaged in properties established before marriage – such as education or physical height. Marriage decisions do not sort on husband's birth order.

3 Empirical strategy and data

Because this paper draws upon several existing and newly collected data sources, table 2 summarizes the data used and its sources and purposes.

3.1 Data and regression specification

Our primary dataset is India's 2005 round of the Demographic and Health Survey (DHS); this is the most recent DHS from India, and is also known as the National Family Health Survey 3. In each surveyed household, all women between the ages of 15 and 49 were interviewed. Our analysis includes only children living in rural joint households with exactly two daughters-in-law, where both women have children under five years old.[9] The DHS defines households as groups of people who live together, pool resources, and specifically eat shared meals out of a common cooking pot. To ensure that we study households that indeed pool resources that influence children's health, we include in our sample only households that list the children's mothers' father-in-law or mother-in-law as the head of the household. Our main sample is limited to children under five years old because the DHS only measures the heights of children under five.

Our empirical strategy uses household fixed effects to compare cousins born to different mothers living in the same household. These household fixed effects are an important part of our identification strategy because they mean that we will be studying differences between children exposed to matching determinants of early life health and nutrition in many dimensions: identical local programs and policies; identical disease environment including local sanitation externalities; identical co-resident grandparents; and nuclear families eating out of the same cooking pot.[10]

Throughout the paper, for various outcomes y, we estimate regressions of the form:

$$y_{cmh} = \beta d_{mh} + X_{cmh}\theta + A_{cmh}\lambda + \alpha_h + \varepsilon_{cmh}, \tag{1}$$

where c indexes individual children; m indexes the two mothers in each household we study; h indexes households. α_h is a set of household fixed effects and X_{cmh} is a set of controls that

[9]In Supplementary Appendix table A8, we verify that our main result is robust to also including children in the small number of households with more than one daughter-in-law with children under 5. 88% of children living in a joint household where more than one daughter-in-law has children under 5 live in a household where precisely two do. We focus on these children for clarity, to concentrate on the most common case, and to permit binary comparisons between the pair of mothers and their children.

[10]Foster and Rosenzweig (2002), in an effort to emphasize the shared co-production and co-consumption in joint households in rural India — and drawing explicit constrast with the separate production in African households documented by Udry (1996) — describe adult brothers in these households as "co-workers who eat together;" these are the fathers of the children we study (p. 842).

vary depending on the application. We include a full set of 120 child age-in-months by sex indicators, A_{cmh}, ruling out the possibility of conflating our effects with those of age or its interaction with sex. Standard errors are clustered by survey primary sampling unit, which is the village in the rural sample.

Our independent variable of interest, d_{mh}, is an indicator with a value of 1 for the lower-ranking mother or her children and 0 for the higher-ranking mother or her children. We assign this indicator based on the ages of the two husbands; therefore, our measure of variation in women's status does not depend on self or subjective reports. In our main results, the outcome y_{cmh} will be a height-for-age z-score, scaled according to the 2006 WHO reference norms. In supporting regressions, we estimate the effect of *lower rank* on early-life mortality of children, and on properties of the children's mothers and fathers.

3.2 Balance and summary statistics

In this section we empirically verify the balance of the "treatment" of being the child of a low-ranking mother, in the standard sense of a verification that an instrument is not endogenously correlated with other observables. In particular, because we interpret our results as effects of intrahousehold social forces, it is important to verify that our results are not due to any differences that would have been established before marriage and formation of the joint household.

Table 3 shows that there is no difference in the pre-marriage observables of the mothers or fathers of the children we study that is likely to explain our main result.[11] These include age, education, literacy, human capital reflected in height, and other predetermined

[11] Following Cameron and Miller (2015), we cluster standard errors by survey primary sampling unit. However, because there are relatively few observations within cluster (we are only studying a subset of households) and because d would have very low intra-cluster correlation, we would not expect clustered standard errors to be very different from unclustered robust standard errors. Indeed, re-creating the balance table with non-clustered robust errors produces the same conclusion: the mean ratio of the standard errors across dependent variables is 1.04.

characteristics of mothers and fathers. Each coefficient $\hat{\beta}$ in Panel A is from a separate regression (equation 1) of the listed variable on an indicator for the lower-ranking mother and household fixed effects. To control for cohort differences and secular time trends, Panel B adds "time controls" for the century-month code month-of-birth (i.e., month number since January 1900) of the mother and the child.[12]

Our paper's main result below is that, within the same household, the children of lower-ranking mothers are, on average, shorter than their cousins born to higher-ranking mothers; therefore, any evidence that pre-marriage human capital is worse for lower-ranking mothers or their husbands would threaten our identification. However, we find no such differences: intrahousehold rank is balanced in these important determinants of a child's early-life health. In fact, although each of these differences is of small magnitude and not statistically significant, lower-ranking mothers are slightly *taller*, on average, than higher-ranking mothers; they are less likely to be stunted; they have more schooling.

We also find balanced characteristics among the fathers of the children we study. Table 3 shows that these two adult brothers, within a household, have indistinguishable height, education, and employment. Table 3 shows that the average child born to the lower-ranking mother has a father who is slightly and insignificantly taller

4 Household rank and women's status

Before proceeding to our main results, we show that intrahousehold rank is related to two measures of women's status. Although the measures of women's empowerment we study here are important, we do not claim that these are the only important dimensions of women's

[12]These cohort controls are included because lower-ranking mothers are younger, on average, than higher-ranking mothers. Because the construction of the rank variable is based on husband's age, their husbands are always younger than the husbands of higher-ranking mothers. Because of these cohort effects, the fathers of the children of lower-ranking mothers are statistically significantly more likely to have been to school (an advantage that would bias away from our findings) in Panel A. This association is not statistically significant with controls for time trends in Panel B.

status.

4.1 Reported decision-making

With the explicit purpose of eliciting a measure of women's empowerment, the DHS asks women whether they have final say in five decisions: the woman's own health care, making large purchases, making household purchases for daily needs, visits to family or relatives, and deciding what to do with money that her husband earns. We classify a woman as "having say" about that decision if she reports that she makes the decision herself or jointly with her husband. 51 percent of the children in our main sample have mothers who report having say on 1 or 0 of these decisions.

In our sample of children, using no household fixed effects or other controls, each additional decision on which a child's mother reports having say is linearly associated with the child being 0.102 height-for-age standard deviations taller [$s.e. = 0.032$; $t = 3.16$]. The identification challenge to which this paper responds is that this simple correlation includes confounding heterogeneity across mothers and households,[13] in addition to any causal effect of women's status.

To examine the causal effect, table 4 reports estimates of equation 1, in which indicators for say in each decision are dependent variables. It also shows results of a regression in which the count of decisions in which a woman has say is the dependent variable. Observations are daughters-in-law of the head of the household living in rural households in which there are exactly two such women. Lower-ranking women report having say in 0.28 fewer decisions than higher-ranking women in the same households, on average. The coefficient is negative in each case, and statistically significantly negative in three cases that may be especially important

[13] Bias of this unconditional correlation could go in either direction: women's status is typically more restricted in higher-caste households, which tend to have more economic and social resources. In our sample, the lowest-caste mothers report having say on 0.35 more issues than higher-caste mothers [clustered $p = 0.051$, comparing *dalit* with others, excluding *adivasis*]. Caste is absorbed by our household fixed effects.

for child nutrition: decisions about health care, about how to spend money, and about daily purchases such as food. Panel A includes no controls beyond household fixed effects; Panel B adds a control for the ages of the individual women, to separate intrahousehold rank from the fact that the lower-ranking daughter-in-law is younger, which may have a direct effect on decision-making.

It is important for our identification strategy that these differences in social status result from intrahousehold rank, rather than reflect fixed, individual properties of the women who become daughters-in-law of different rank. Figure 2 provides evidence for this; it plots the count of situations in which a woman reports having say at each duration of marriage, separately for higher- and lower-ranking mothers of the children we study in section 5. The figure is centered vertically at zero because residuals are plotted after a regression of say on household fixed effects; this permits us to see how the say of the two mothers evolves relative to one another.

The figure clearly shows that average reported say is initially similar between the two women before or very early in marriage, and separates with duration of marriage. This may reflect a time-path of actual decision power, or of perceived or reported decision power after accumulating exposure to household institutions. Either way, what matters for our identification is that this trend appears not to reflect pre-existing differences in personality or preferences that women bring into marriage at different ranks.

4.2 Time outside the home

In the literature on women's status in India, women's mobility outside the home is a measure of her status and empowerment (Kabeer, 1999; Rahman and Rao, 2004). In this section, we use data from the India Time Use Survey (ITUS) to ask whether lower-ranking women in joint households spend less time outside the home than higher-ranking women in joint households on a typical day.[14]

The ITUS, collected in 1998-9, interviewed the members of rural households in six states in India.[15] Each adult member was asked to report his or her activities on a typical day prior to the interview. In addition to reporting the activities in which he or she was engaged, respondents also reported how long he or she was doing the activity, and whether it took place inside or outside the home. We restrict our analysis to the 156 rural households with exactly two daughters-in-law of the household head.

Figure 3 presents results: lower-ranking daughters-in-law spend an average of 27 minutes less time outside the home on a typical day [$p = 0.08$, clustered by household]. Although the small sample does not permit much further investigation, Panel A confirms that a difference is apparent at all ages, and is not therefore due to differences in the average age of the women. Panel B shows that the distribution of time outside the home for higher-raked women stochastically dominates the distribution for lower-ranked women [Kolmogorov-Smirnov $p = 0.07$]. This difference in time spent outside the home is therefore consistent with an effect of intrahousehold rank on women's status.

5 Main results: Effect of mother's rank on child height

In this section, we present our main results: an effect of mothers' social status on their children's height. Children of lower-ranking mothers are shorter, on average, than their cousins from the same joint households who born to higher-ranking mothers. Figure 4 documents this simply and visually. This figure presents summary statistics using our main sample, without household fixed effects. It plots height-for-age z-scores by age-in-months for children of higher- and lower-ranking mothers and finds that, at all ages, there is a visible

[14] The ITUS classifies days as "normal," "weekly variant" (such as Sunday), and "abnormal" (such as holidays); we follow other papers that use the ITUS (*e.g.* Barcellos *et al.*, 2012) in restricting our analysis to normal days.

[15] These are Harayana, Madhya Pradesh, Gujarat, Tamil Nadu, Orissa, and Meghalaya.

difference between the heights of children of higher-ranking mothers and those of lower-ranking mothers. The presence of a difference at birth is consistent with *in utero* causes related to net maternal nutrition, which we discuss in section 6. The rest of this section quantifies and confirms the robustness of this demonstration of an effect of women's status.

5.1 Main result

Table 5 presents results of our fixed effect regression equation 1, estimated from India's Demographic and Health Survey using our main sample of children under five living in joint rural households with two daughters-in-law, in which both women have at least one child between zero and five. Column 1 shows that the average child of the lower-ranking mother is about one-third of a height-for-age standard deviation shorter than the average child of the higher-ranking mother in the same household. We include household fixed effects and 120 age-in-months by sex indicators, which is the level of disaggregation at which the WHO height reference charts are defined.[16] Columns 2 through 4 add controls to verify that this result is not driven by other dimensions of possible difference between these children or their nuclear families; all three sets of controls slightly increase the effect size, relative to the baseline specification in column 1.

Column 2 adds controls for demographic factors: an indicator for whether the child is first born to her mother, an indicator for whether the child is a single birth, and the child's birth order in the joint household. The joint household birth order is the child's birth rank among all children in the household, born to either mother. The average child of the lower-ranking mother is 0.432 positions later born [$s.e. = 0.106$; $t = 4.06$]. This control exploits

[16]If we had incorrectly included only household fixed effects and the indicator for a lower-ranking mother, without controls for the child's age, the coefficient would be positive. This erroneous result would be an artifact of the construction of height-for-age z-scores and the fact that, in a developing countries like India, where there is widespread stunting, height-for-age is well known to decline during first two years of life, as shown in figure 4 (Cummins, 2013). Because the average child of the lower-ranking mother is 8 months younger than the average child of the higher-ranking mother, age is an omitted variable that, if not controlled for, would positively bias the coefficient in column 1.

the fact that some children of the lower-ranking mother are born before some children of the higher-ranking mother, and verifies that our result is not merely due to a bias in favor of earlier-born children (to whichever mother).[17]

Figure 5 further documents average height differences between children of higher- and lower-ranking mothers, at each joint household birth order. The vertical axis plots residuals after height-for-age is regressed on household fixed effects and age-in-months by sex indicators. Although the results are somewhat noisy for higher-order births — joint birth orders 5, 6, and 7 together account for only 25 percent of the births — there is a difference between the heights of children of higher- and lower-ranked mothers at every birth order, which verifies that our result is not due to differences in intrahousehold birth order.

Column 3 of table 5 controls for properties of the child's mother: her height, indicator variables for her completed years of education, and her age at marriage. Column 4 controls for properties of the child's father: indicator variables for his education in categories and years, his category of occupation, and his age. Balance table 3 verified that children born to lower-ranking mothers are not disadvantaged in parents' pre-marriage characteristics, so it is no surprise that the coefficient on mother's rank does not decrease. Our result is not due to differences in pre-marriage human capital or parenting ability among parents, unless such differences are uncorrelated with the observable characteristics that we measure.

Figure 6 provides further evidence that differences in parents' heights — which would have been determined before marriage, in childhood — are not responsible for our results. Panel A uses our main sample of children to show that average height differences between children of higher- and lower-ranking mothers are present at all mothers' heights. Panel B shows a difference across fathers' heights; this verifies that differences in early-life treatment of the fathers, who are brothers of different birth orders, is not the mechanism driving the

[17] Controlling for within-mother birth order, rather than within-household birth order, produces very similar results.

results. The sample in panel B is smaller than the sample in panel A because the DHS only measured men's heights in a randomly selected sub-sample of households; panel B plots the heights of 408 children of 313 fathers. If, in the small sample of children under five in joint rural households whose fathers' heights were measured, we add mother's height and father's height as controls to our baseline specification, we find that, conditional on these variables and the demographic controls, children of lower-ranking mothers are 0.351 standard deviations shorter than children of higher-ranking mothers [$s.e. = 0.239$; $p = 0.14$], which is quantitatively very similar to our main result.

Figure 4, which plots average height-for-age against age-in-months separately for children of the higher and lower ranking mother, appears to suggest that our main result is larger for older children. This would be consistent with the well-known unfolding of stunting over early life as a result of environmental deprivation, rather than, for example, a fixed genetic average difference between the two mothers. In the supplementary appendix, table A7 tests for this interaction, and finds that it is indeed the case that the gap in child height by mother's rank is larger, on average, for older children within our sample. Supplementary Appendix table A8 confirms that our result is robust to including the small number of children in joint households with three or more daughters-in-law with children with measured height.

These effect sizes are importantly large, but plausibly sized, relative to the economics literature on child height. Our effect of one-third of a height-for-age reference standard deviation is about three-fourths as large as the average difference in child height between high and low caste Indian children in the DHS. It is about half as large as the effect of a widely-studied field experiment which provided early-life nutrition to some children in Guatemala, which has shown long-run effects on adult human capital (Maluccio et al., 2009).

5.2 Evidence from longitudinal data on household splits

We interpret our main result to be due to differences in status *within* joint households, not differences in pre-marriage characteristics between parents. In this section, we exploit an additional data source to show that there is *no* difference in the average heights of paternal cousins who *do not* live together in a joint household. This analysis supports our main finding in two ways. First, it demonstrates that the difference in child height that we observe is not due to endogenous household formation; that is, it is not the case the older brothers with less healthy children differentially move out of joint households, or that younger brothers with more healthy children differentially move out of joint households. Second, it supports our argument that the difference in child height we observe arises due to status differences at work *within* joint households.

The 2005 India Human Development Survey (IHDS) longitudinally followed a sample of rural households from the Human Development Profile of India (HDPI), a survey conducted in 1993-4, including following members of nuclear households who had lived in joint households in the 1994 survey.[18] The 2005 IHDS therefore allows nuclear households of adult brothers to be linked to the household in which they were living in 1994. This means that we can compare height differences among cousins living in joint households with height differences among cousins living in split households, where two brothers have formed new households with their wives and children. If height differences among children were due to endogenous household formation, we would expect child height in the split sample to show an opposite effect of mother's rank as the one we find in joint households, rather than no effect.[19] Alternatively, if height differences among children were only due to individual differences between their fathers – perhaps due to fathers' poor early-life health, human capital,

[18]The IHDS data are publicly available and are described by Desai et al. (2009).

[19]In other words, if out of the full set of children of a pair of brothers it is the case for any reason that pairs where the younger brother's children are relatively disadvantaged are more likely to remain as joint households, then the split households would necessarily be the remaining complement, those where the older brother's children are relatively advantaged. But this is not what we see.

or economic productivity – then we would expect to see the same effects of being the child of a younger brother among children living in split or joint extended families. However, we observe a height difference only among children living in joint households, and no effect among children living in split households, suggesting that height differences are due to exposure to the social institutions of joint households.

Table 6 reports results from the IHDS 2005 data. Observations are children living in households in which exactly two brothers with children under five in 2005 either live jointly, or lived jointly in 1994. For children living in joint households in 2005, we look only at rural children, to match our DHS sample. For children living in split households in 2005, we include children in the regression if at least one brother's children live in a rural place.[20] We use the log of child height in centimeters, instead of height-for-age z-scores, as the dependent variable in these regressions because reliable information on children's age in months is not available in the IHDS. We control for age in years, separately for girls and boys. In the IHDS data, among children under five, there is an effect on cousins' height of mother's social rank in joint households (column 2), but not in split extended families (column 3). This difference is statistically significant as an interaction (column 4). In columns 5 and 6, we replicate our main result from table 5 using the DHS sample and log of child height as the dependent variable, to show that the results from the IHDS data are quantitatively comparable to what is found in DHS data when the specifications are harmonized.

This result indicates that the effect of mother's rank on child height is due to intra-household processes – not due to marriage market sorting, endogenous household splits, or differences by age rank in the human capital of adult brothers. Additionally, the IHDS data provide evidence against the possibility that Indian joint households endogenously split or stay together not due to economic or social factors, but as a *consequence* of child health.

[20] If we instead focus on children whose fathers' households were classified as rural in 1994, regardless of urban/rural classification in 2005, we find nearly identical results.

Further evidence for this claim is that among the 675 studied young children living in split families, in all but 6 cases the household has not moved since before the child was born.

6 Mechanism: Maternal net nutrition

Children of higher-ranking mothers are taller than children of lower-ranking mothers, a result which is not due to differences among their parents from before marriage, nor to endogenous formation of households. The difference in child outcomes that we observe accompanies a difference in mother's social status, apparent in her reported say in household decisions and time spent outside the home. In this section, we consider mechanisms for the robust height difference between cousins in joint rural households. We find several indications that the maternal net nutrition to which children are exposed *in utero* and during breastfeeding is different for children of higher- and lower-ranking mothers. Effects of maternal net nutrition on child stunting are highly plausible in India; average maternal nutrition indicators are very poor, and maternal nutrition is known to impact birth weight, a leading indicator of child height (Ludwig and Currie, 2010; Adair, 2007). Coffey (2015b) has recently shown that maternal nutrition is correlated with other dimensions of women's social status.

6.1 Effect of rank on mother's body mass

We propose that one important mechanism linking a mother's intrahousehold status to her children's height is her net nutrition. By "maternal net nutrition," we mean that what matters for child health outcomes is both the nutrition that mothers consume and the energy that is expended in work; lower-ranking mothers work hard and have less claim on household food, as described in section 2. Maternal stress is a potential further channel through which intrahousehold rank may affect maternal net nutrition.

Although height is determined by a combination of a woman's genetic potential height and her own *early-life* health and net nutrition, her weight is subject to more recent influences.[21] Balance table 3 shows that lower-ranking mothers are not shorter than higher-ranking mothers; this is consistent with similarity of lower-ranking and higher-ranking mothers before marriage. In this section, we ask whether higher-ranking mothers weigh more than lower-ranking mothers.

Table 7 reports household fixed effects regressions of the form of equation 1, with the mother's body mass index (BMI) at the time of the survey as the dependent variable. Columns 1 (without household fixed effects) and 2 (with household fixed effects) show that lower-ranking mothers have about one-third of a BMI point less body mass than higher-ranking mothers, on average. Columns 3 and 4 take seriously the possibility that this may be due to the two women being at different points in their childbearing or breastfeeding careers. However, the average BMI difference between higher- and lower-ranking mothers increases when we control for current pregnancy, how many children the woman has had, the woman's year of birth, whether the she is breastfeeding and the age of her youngest child. The coefficient is essentially unchanged when we control more flexibly for the amount of additional energy a woman requires to breastfeed by interacting the age of her youngest child with whether the child is breastfeeding. The effect of intrahousehold rank on BMI that we identify is consistent with net maternal nutrition being a mechanism for the effect we document on child height.

In support of our interpretation of our estimate as an effect of *women's* status and about *maternal* nutrition, supplementary appendix section A4 and table A6 report results of a similar test for differences in the BMI of adult brothers in joint rural households. These are analogous to the husbands of the women we study, although they are not the precise identical men, because of the sampling of the men's recode. Across a range of specifications and samples, we find no evidence that the BMI of younger adult brothers is lower, on average,

[21] Thus, a diet can cause an adult to lose weight, but not to lose height.

than the BMI of their older brothers within the same household.

6.2 Effect of rank on very early life mortality

If the differences we observe in child height have their origins in differences in maternal net nutrition, then we might also expect to see effects of mother's social rank on other indicators of very early life health and mortality. Differences across places and times in population average height are correlated with differences in infant mortality rates, because the same disease and poor net nutrition that causes infant death also stunts the growth of survivors (Bozzoli et al., 2009; Hatton, 2013). Infant mortality (IMR), or death in the first year of life, is divided into neonatal mortality (NNM) in the first month of life, and post-neonatal mortality (PNM) in the second through twelfth months of life. Although studies of height and early-life mortality in Europe have emphasized correlations between height and post-neonatal mortality, which is thought to in part reflect the infectious disease environment, Coffey (2015a) finds that neonatal mortality is an important predictor of average height across state-cohorts in India, arguably due to extremely poor maternal nutrition.

Table 8 investigates whether mothers' intrahousehold rank is associated with early-life mortality. Estimates use the same 2005 round of the Indian DHS as our main result and include all births in the 10 years before the survey to women living in joint rural households with two daughters-in-law.[22] Compared to their cousins born to higher-ranking mothers, children of lower-ranking mothers are less likely to survive their first year of life. The effect is almost entirely because of survival differences during the neonatal period. Differences in neonatal mortality are consistent with differences in maternal nutrition, which leads to low birth weight.[23] Therefore, these differences in neonatal death suggest the possibility

[22]Infant mortality is observed for more children in the DHS than is height because it is survey-reported, not measured. Because of the long ten-year period under consideration, year fixed effects are included for the passage of time; younger women's children will be born in later years, on average, and therefore into healthier environments.

[23]Another determinant of neonatal mortality could be differences in medical care at birth. However,

of differences in birth weight, caused by worse net maternal nutrition and health *in utero*, which would also result in childhood height differences. We next present direct evidence of this channel.

6.3 Effect of rank on birth weight in a district hospital

The differences that we have documented in maternal BMI at the time of the survey and in neonatal mortality are highly suggestive of differences in maternal nutrition between higher- and lower-ranked mothers, leading to differences in the birth weights and subsequent heights of their children. The Indian DHS, our main data source, does not observe birth weight. Therefore, this section analyzes original birth weight data collected by the authors between February 2013 and September 2014 in a government district maternity hospital in a largely rural district of Uttar Pradesh, a poor state in north India.

Six days a week, for a randomly selected subset of vaginal births in the previous 36 hours, we measured the birth weights of newborns and the heights and weights of their mothers. Of 2,376 observed births, 631 were to mothers in joint households with two daughters-in-law; we were able to collect anthropometric data on 549 of these births.[24] Note that because we observe births in the hospital, we are unable to use household fixed effects with these data; very few households would have had two daughters-in-law giving birth in this one facility in the window of time when we were collecting data.

Figure 7 presents results for child birth weight, and figure 8 presents results for mothers' post-partum BMI.[25] Lower-ranking mothers and their newborns both weigh less than higher-ranking mothers and their newborns, respectively. These differences are not due to the mother's age nor to her number of prior births. On average, babies born to lower-ranking

results available in a working paper version provide evidence that lower-ranking mothers do not receive worse medical care at birth.

[24]Some mothers refused data collection; 3.5 percent of births are recorded as dead at birth; 4.4 percent were recorded to have been born alive but died before the first-day-of-life survey.

[25]Post-partum BMI is a standard indicator of whether a woman was well nourished during pregnancy.

mothers weigh 84 grams less than babies born to higher-ranking mothers [$p = 0.051$]. Lower-ranking mothers have 0.37 BMI points less body mass within 36 hours of delivery [$p = 0.086$] and weigh 1.04 kilograms less [$p = 0.085$], on average, than higher-ranking mothers. It is noteworthy that, as in our main DHS sample, there is no difference in the mother's heights [-0.36 cm; *s.e.* = 0.48; $p = 0.45$] in these data.[26] Controlling for whether the mother already has a living boy child does not change the result; the children born to low-ranking mothers are still at a disadvantage of 75 grams [$p = 0.081$]. The results from these data further confirm that disadvantages in net maternal nutrition, especially during pregnancy, could partially explain the effects we document of intrahousehold social rank on child height.

7 Conclusion

Economists and scholars throughout the social sciences have asked whether women's social status impacts their children's health. In particular, the claim that this is true has had considerable influence over the design of development policies and programs. As one IFPRI policy report on child outcomes summarizes, "many development programs that aim to alleviate poverty and improve investments in human capital consider women's empowerment a key pathway by which to achieve impact... Despite this, women's empowerment dimensions are often not rigorously measured and are at times merely assumed." Supplementary Appendix section A5 describes programs and quotes reports from UNDP, the Gates Foundation, IFPRI, CARE, and other organizations that emphasize the widespread importance for policy of the assumption that women's social status — above and beyond their material resources — matters for children's outcomes, including nutrition.

Although many development practitioners assume that such a relationship exists, and

[26] Nor are there differences between higher- and lower-ranking mothers in age at marriage [0.05 yrs; *s.e.* = 0.88; $p = 0.96$], or in the probability they have ever been to school [0.04 percentage points; *s.e.* = 0.04; $p = 0.36$], all of which would have been determined before marriage.

although such a relationship is intuitive and plausible, omitted variables and measurement problems make it particularly difficult to find situations where we can be confident that mother's status and its consequences have a causal effect on child health. This paper exploits the unique social institution of joint households in rural India, in which women married to the younger brother have lower social rank than women married to the older brother, but women are not sorted by quality into these social roles.

We find evidence that the focus of development practitioners on women's status can be well-motivated. With the important caveats that we are studying a particular determinant of social status in the Indian context, and that maternal nutrition may be less important in populations where overall net nutritional deprivation is less severe, our study supports the possibility that programs to enhance women's social status will have good effects on their children, and that recognizing and designing programs around women's status may be an important consideration in development policy.

The children of lower-ranking mothers are shorter, on average, than their cousins in the same household. We present evidence that this difference is an effect of intrahousehold social status: it only appears in joint households, not among the children of brothers who live separately. Effects of intrahousehold rank on women's reported empowerment emerge within the household environment. One important mechanism for the effect of mothers' social status on their children's well-being is maternal nutrition: although they are not shorter, lower-ranking mothers have less body mass than higher-ranking mothers, including immediately after delivery; their children have lower birth weights; and their children are more likely to die in the first month of life, a period in which good maternal nutrition is particularly important for survival.

Because the number of children under five living in the type of joint households that we study is one-third the total number of children under five in the United States, and because many children in developing countries are raised by mothers with low social status, the effect we document is of considerable general importance. These findings also have at least two specific implications for the design of specific programs and policies. First, despite the focus of development policy debates about how best to allocate food subsidies or nutritional transfers *to* households (Currie and Gahvari, 2008; Hidrobo et al., 2014; Muralidharan et al., 2016), allocation *within* households can additionally remain a binding constraint on net nutrition reaching children in early-life critical periods. Second, in India, home to one-fifth of births worldwide, women's status and maternal nutrition remain intersecting constraints on early life human-capital development (Coffey, 2015b).

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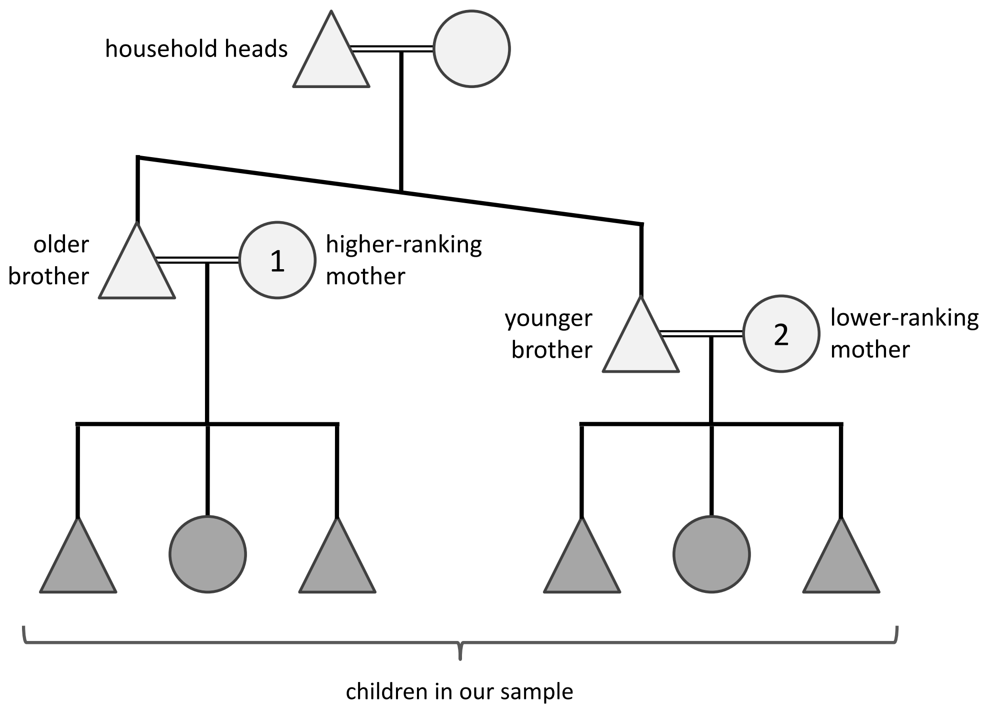
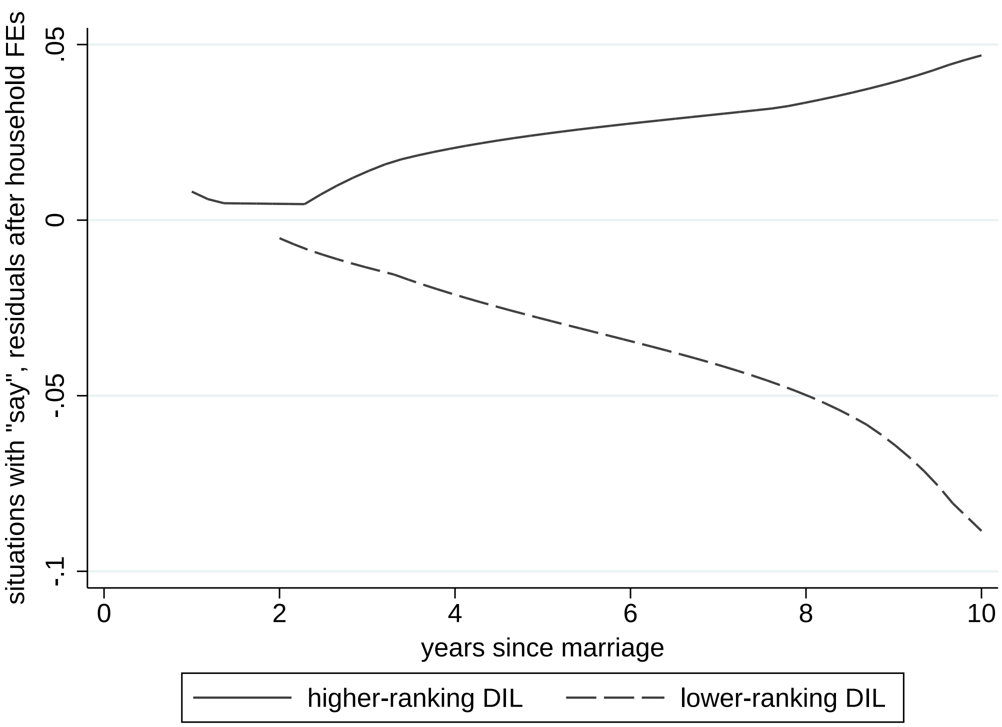
Figure 1: Structure of joint Indian households: Mother assigned social status at marriage



Males are represented as triangles and females as circles; marriage relationships are denoted as double lines. The slanted connection of the nuclear families reflects the difference in social rank that our identification strategy exploits. Although we only study households with two daughters-in-law, the particular number and sex composition of the children is an arbitrary illustration.

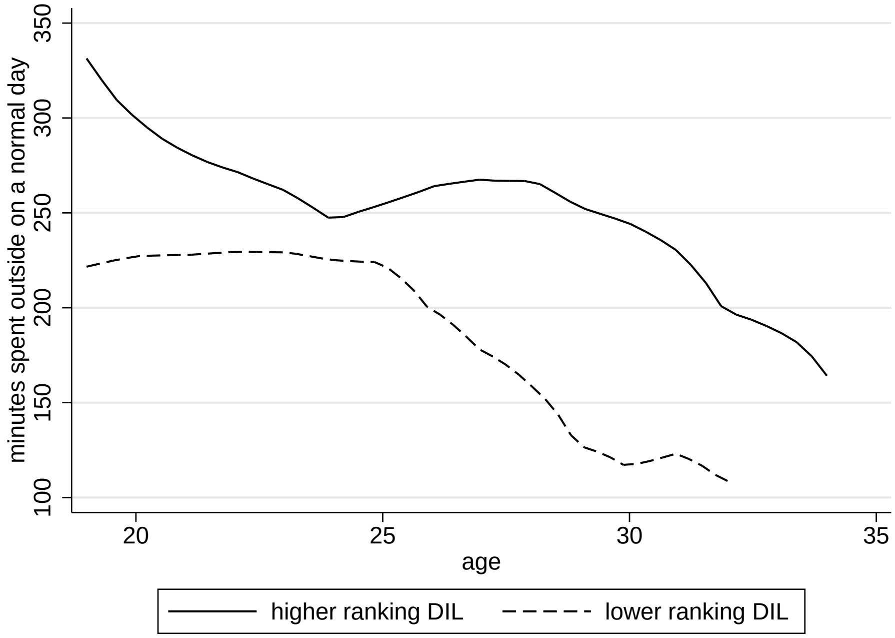
Figure 2: Background: Differences in status unfold after marriage



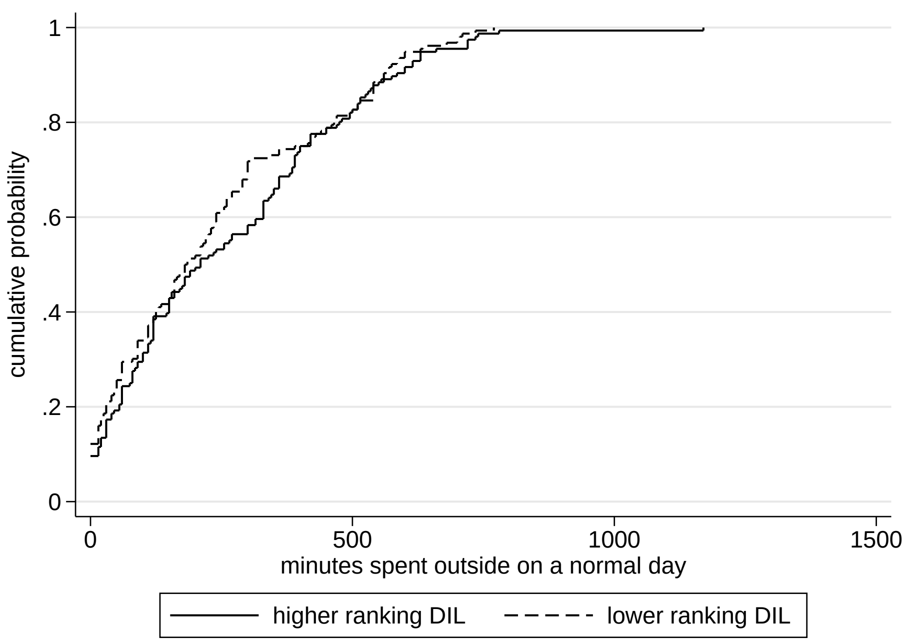
Data: India 2005 DHS. "Say" is the count of five situations in which the woman reports having say rather than not having say, either alone or jointly with her husband. Residuals are after a regression on household fixed effects. Local regression; bandwidth is three years. Observations are the mothers of the children in our main sample, in table 5.

Figure 3: Background: Time spent outside the house of a normal day, by woman's rank

(a) time spent outside on a normal day

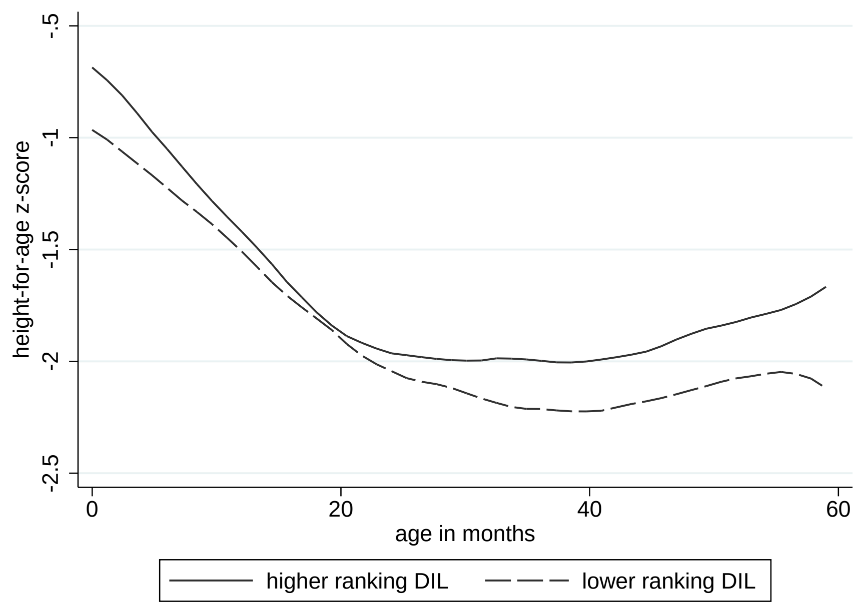


(b) CDF of time spent outside on a normal day



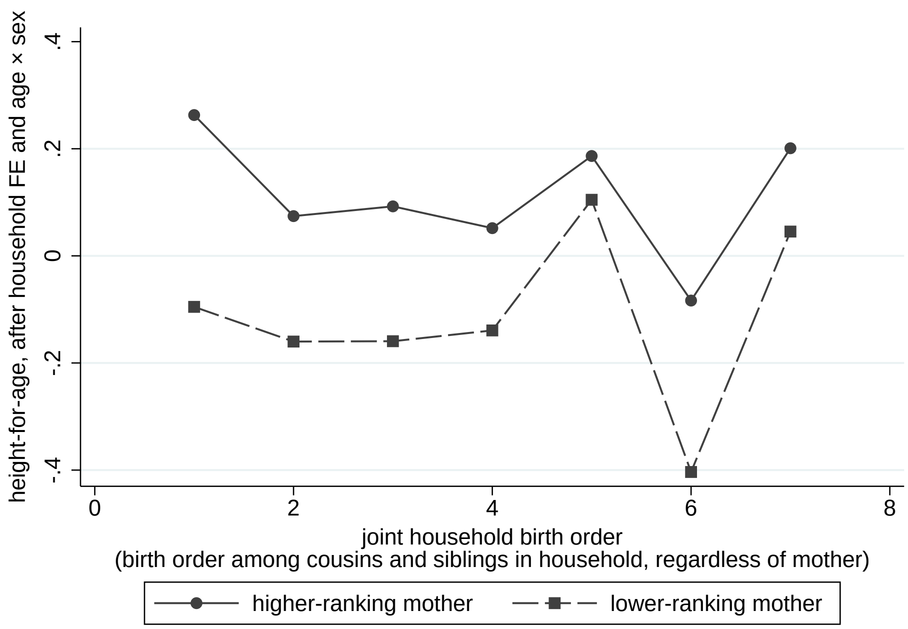
Data: India Time Use Survey (ITUS), 1998-9. Panel A plots minutes spent outside on a normal day by age using local polynomial regression with an epanechnikov kernel and a bandwidth of 1.75 ($p = 0.08$). Panel B is the cumulative distribution function of minutes spent outside on a normal day ($p = 0.07$). n=312. DIL is daughter-in-law. Observations are daughters-in-law of the head of the household living in rural households in which there are exactly two such women.

Figure 4: Main result: Height-for-age of children is lower for lower-ranking mothers



Data: India 2005 DHS. Local regression with an epanechnikov kernel and a bandwidth of 6 months; no controls or household fixed effects are included. $n = 1078$, identical to main sample in table 5. Height-for-age is computed according to WHO 2006 reference norms.

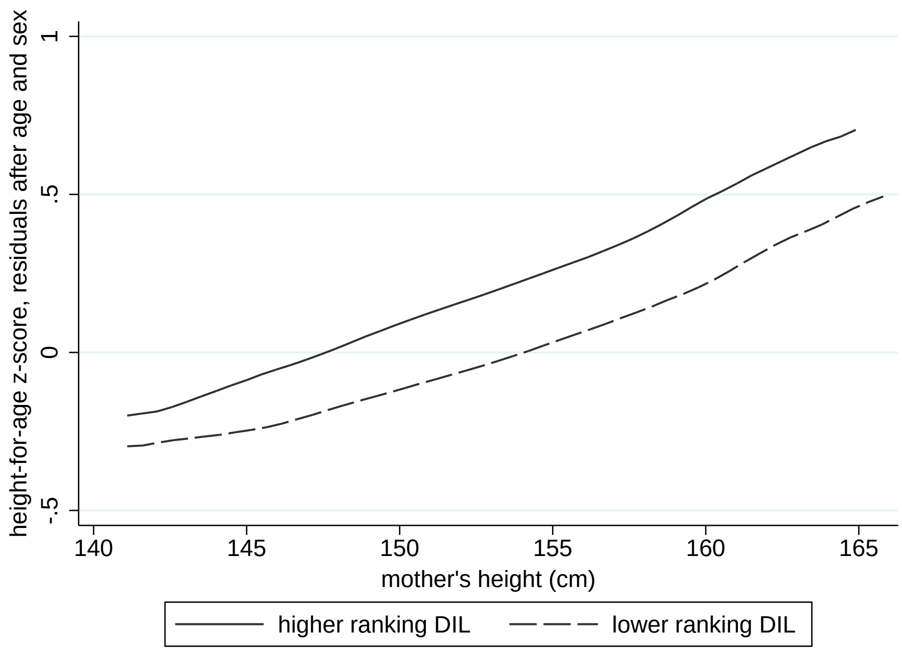
Figure 5: Robustness: Height difference is present conditioning on birth order among *cousins and siblings* within the joint household



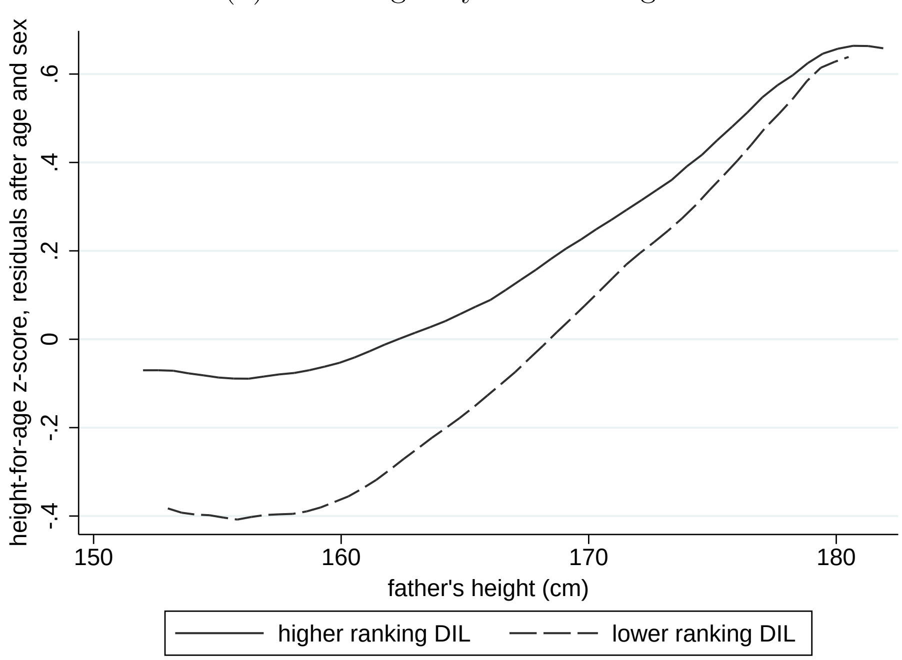
Data: India 2005 DHS. The dependent variable is the height-for-age residual after controlling for household fixed effects and age-in-months by sex indicators. Joint household birth order is age rank among children of both mothers – that is, birth order among all cousins and siblings in the household, regardless of mother. n = 1078, identical to main sample in table 5.

Figure 6: Robustness: Children of lower-ranking mothers are shorter at all heights of parents

(a) child height by mother's height



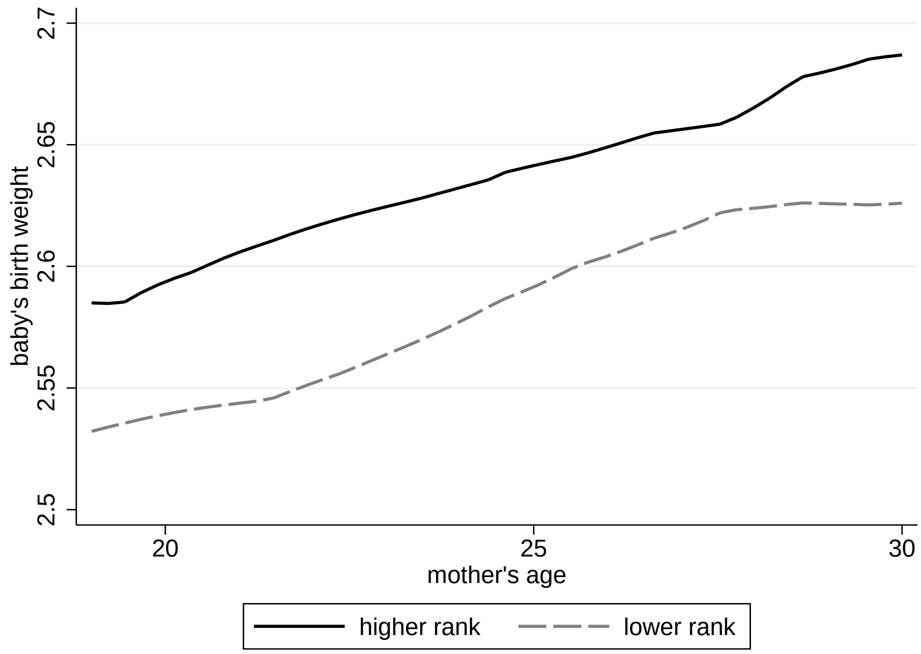
(b) child height by father's height



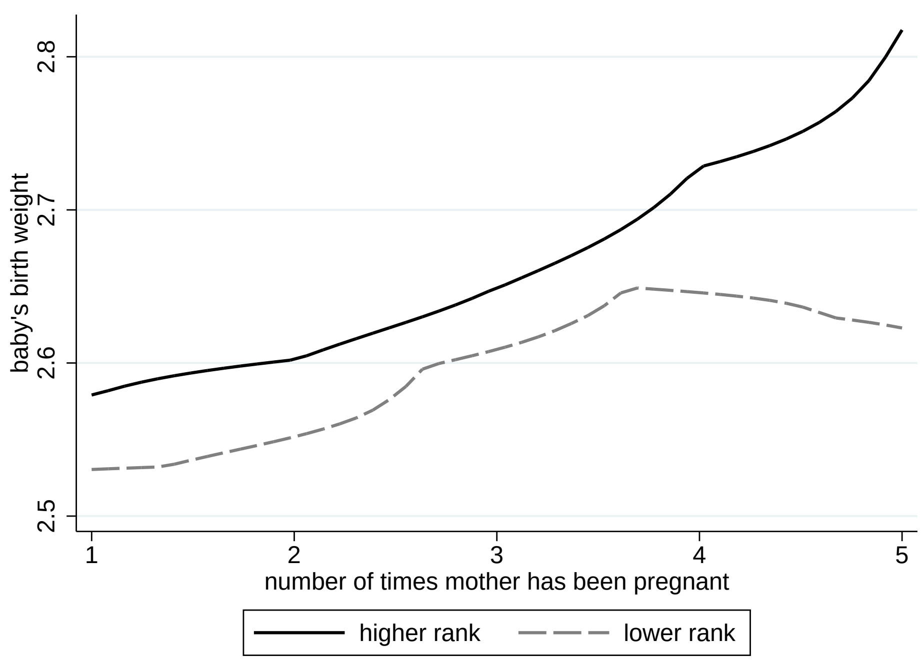
Data: India 2005 DHS. Local polynomial regression with an epanechnikov kernel and a bandwidth of 5cm. In panel A, n = 1078, identical to main sample in table 5. In panel B, n = 408, because the DHS only measured a sub-sample of men. DIL = daughter-in-law, the mothers of the children we study.

Figure 7: Mechanisms: *In utero* effect of mother's rank reflected in child's birth weight

(a) birth weight, conditional on age



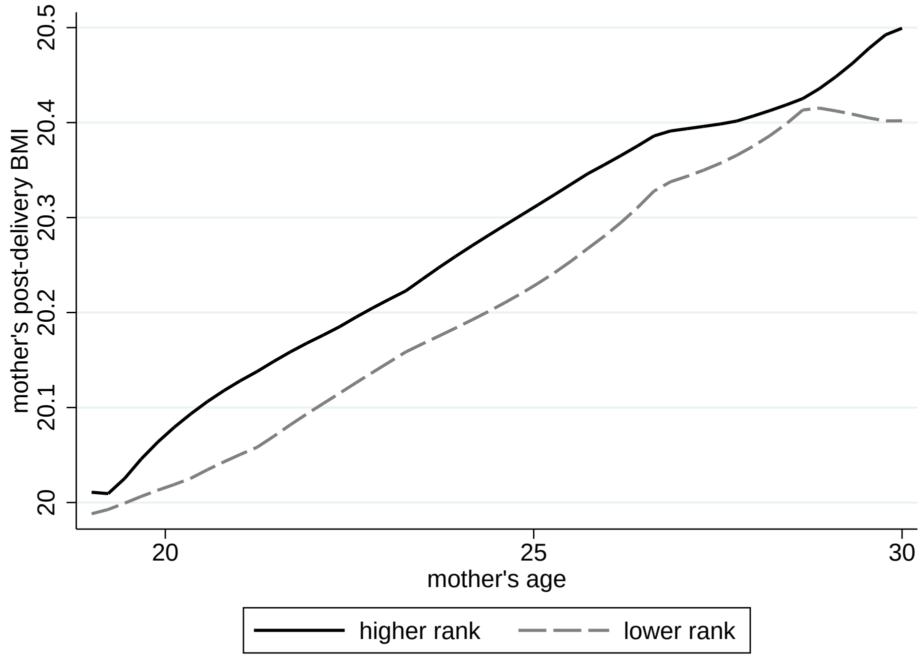
(b) birth weight, conditional on pregnancy count (parity)



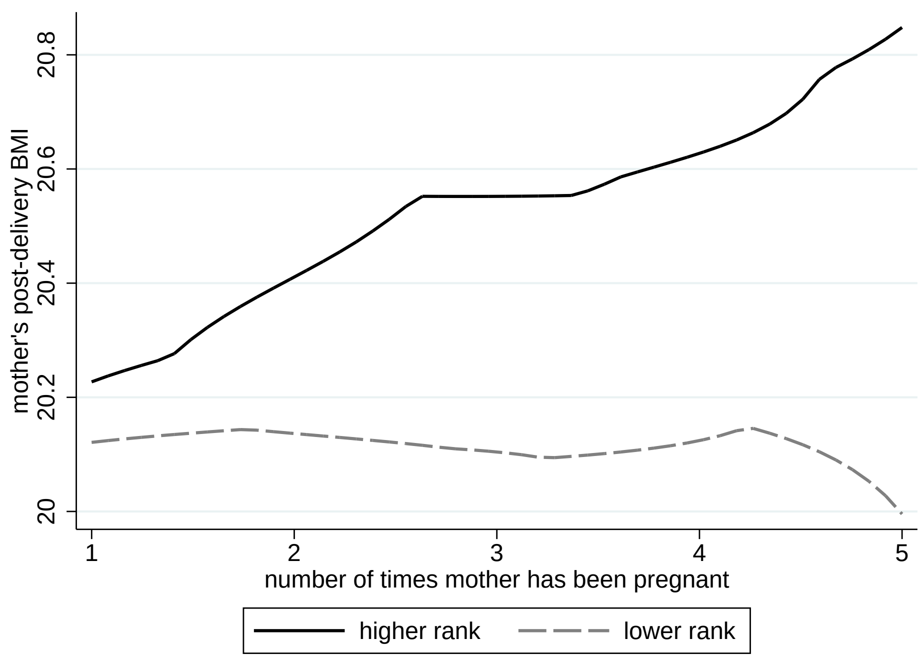
Data are from the authors' collection in a district maternity hospital in Uttar Pradesh. Birth weight was measured within 36 hours of birth. To match our main sample, data are restricted to women in households with exactly two daughters-in-law; note that because data are collected at the hospital, not in a household survey, household fixed effects cannot be used.

Figure 8: Mechanisms: Mother's rank and mother's postpartum BMI

(a) mother's BMI, conditional on age



(b) mother's BMI, conditional on pregnancy count (parity)



Data are from the authors' collection in a district maternity hospital in Uttar Pradesh. Post-partum BMI was measured within 36 hours of birth. To match our main sample, data are restricted to women in households with exactly two daughters-in-law; note that because data are collected at the hospital, not in a household survey, household fixed effects cannot be used.

Table 1: Background: In a survey experiment, groom's age rank among brothers is ignored by participants

	(1)	(2)	(3)	(4)
	linear probability chooses groom ad A			
is older brother	0.0135			
	(0.0160)			
is younger brother	0.00326			
	(0.0155)			
older minus younger		0.00493	0.00494	-0.00316
		(0.00872)	(0.00872)	(0.00756)
age				0.00919***
				(0.00188)
inches tall				0.00773***
				(0.00174)
high school				0.112***
				(0.0156)
college				0.239***
				(0.0145)
dark skin				-0.0620***
				(0.0126)
government job				0.216***
				(0.0145)
farmer				-0.121***
				(0.0151)
constant	0.512***	0.511***	0.515***	0.506***
	(0.0103)	(0.0103)	(0.0114)	(0.00884)
n (choices between two ads)	2,358	2,358	1,910	2,358
sample	full	full	brother info	full

Data: survey experiment conducted by authors; see Supplementary Appendix section A3. In support of our identifying assumption, we conducted a survey experiment in which Indian participants made 2,358 binary choices between fictional groom ads. Ads randomly varied properties of hypothetical grooms, including rank among brothers in a joint household. All coefficients are differences between the ad labeled A having that property and the ad labeled B having that property; so, for example, "farmer" is 1 if ad A claims the groom is a farmer and B does not, is -1 if ad B claims the groom is a farmer and A does not, and is 0 if the two ads match on this characteristic. Column 3 includes only choices in which at least one of the ads claims that the groom is either the older or younger of two brothers.

Table 2: Data sources used in paper

data source	purpose
Demographic and Health Survey (NFHS 2005)	**main result:** difference in child height-for-age **background:** difference in mothers' reported "say" **mechanism:** difference in IMR, mothers' BMI
India Time Use Survey (ITUS 1999)	**background:** intrahousehold rank predicts mobility of daughters-in-law
India Human Development Survey (IHDS 2005)	**main result & falsification:** replicate difference in height, but only within joint households (not cousins living separately)
birth anthropometry from a district hospital (collected by authors)	**mechanism:** differences in birth weight and post-partum BMI
categorization of real marriage ads (collected by authors)	**identifying assumption:** marriage markets do not consider husband age rank
survey experiment on marriage ads (collected by authors)	**identifying assumption:** marriage markets do not consider husband age rank

Table 3: Balance of nuclear family, child, and pre-marriage characteristics

		Panel A: no time controls			Panel B: with time controls			
	mean	$\hat{\beta}$ low rank	s.e.	t	$\hat{\beta}$ low rank	s.e.	t	n
mother's height (cm)	152.7	0.471	0.379	1.242	0.451	0.495	0.911	1,075
mother's height < 145 cm†	0.084	-0.022	0.024	-0.909	-0.012	0.034	-0.341	1,075
mother's age at marriage‡	18.2	0.157	0.231	0.680	-0.241	0.305	-0.792	795
mother married before 18‡	0.447	-0.036	0.044	-0.818	0.043	0.047	0.930	795
mother no education	0.340	-0.023	0.023	-0.982	0.033	0.032	1.044	1,078
mother secondary education	0.498	0.049	0.026	1.866	0.033	0.037	0.883	1,078
mother literate	0.623	0.017	0.028	0.614	-0.013	0.036	-0.366	1,078
mother native Hindi speaker	0.604	-0.003	0.003	-1.342	0.000	0.004	-0.063	1,078
mother's desired children	2.385	-0.054	0.044	-1.222	0.045	0.057	0.781	1,061
father's height* (cm)	165.8	0.778	0.561	1.387	1.181	0.961	1.230	408
father no education	0.127	-0.043	0.020	-2.119	0.019	0.027	0.687	1,075
father secondary education	0.766	0.033	0.026	1.291	-0.020	0.031	-0.661	1,075
father years of school	8.31	0.185	0.230	0.806	-0.542	0.317	-1.712	1,064
father absent for work	0.172	-0.006	0.025	-0.238	0.012	0.032	0.363	1,072
father works h.h. land, if owned	0.822	0.067	0.054	1.257	0.099	0.067	1.479	353
father does agricultural work	0.377	-0.023	0.026	-0.880	-0.007	0.038	-0.177	1,073
father does white collar work	0.067	-0.014	0.013	-1.011	-0.005	0.023	-0.218	1,073
child female	0.458	0.000	0.000	0.522	0.000	0.000	-0.128	1,078
child month of birth	6.84	0.133	0.212	0.627	0.165	0.288	0.574	1,078
child summer birth	0.229	0.023	0.025	0.921	-0.004	0.035	-0.119	1,078

Data: India 2005 DHS. Each coefficient $\hat{\beta}$ in Panel A is from a separate regression (equation 1) of the listed variable on an indicator for the lower-ranking mother and household fixed effects. To control for cohort differences and secular time trends, Panel B adds "time controls" for the century-month code month-of-birth of the mother and the child. Standard errors are clustered by survey PSUs. †The India DHS survey manual (2005) uses 145cm as a threshold for low height among adult Indian women. ‡Many women do not report an age at marriage. *By design, the DHS measured height only of a subset of adult men. Observations are children in the main sample in table 5.

Table 4: Background: Intrahousehold rank predicts reported decision-making "say" within households

dependent variable:	(1) sum (total of 5)	(2) own health	(3) large purchases	(4) daily purchases	(5) visit family	(6) money
sample mean:	1.64	0.41	0.22	0.25	0.33	0.43
Panel A: Without control for woman's age						
lower-ranking woman	-0.281*	-0.1000**	-0.0154	-0.0825*	-0.00906	-0.0737†
	(0.111)	(0.0363)	(0.0330)	(0.0344)	(0.0332)	(0.0382)
joint household fixed effects	✓	✓	✓	✓	✓	✓
n (daughters-in-law)	1395	1395	1395	1395	1395	1395
Panel B: With control for woman's age						
lower-ranking woman	-0.298†	-0.0418	-0.0324	-0.140**	-0.0418	-0.0423
	(0.162)	(0.0542)	(0.0448)	(0.0487)	(0.0493)	(0.0544)
woman's age at the time of the survey	-0.00473	0.0157	-0.00461	-0.0155†	-0.00885	0.00846
	(0.0282)	(0.0107)	(0.00869)	(0.00907)	(0.0105)	(0.0116)
joint household fixed effects	✓	✓	✓	✓	✓	✓
n (daughters-in-law)	1395	1395	1395	1395	1395	1395

Data: India 2005 DHS. Standard errors clustered at the PSU level are shown in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. For each dependent variable, a woman is said to have "final say" if she says she has final say on her own, or jointly with another family member, most commonly her husband. The dependent variables are: (1) the number of decisions (out of five) in which a woman has final say; (2) a dummy variable for whether she has final say in decisions about her own health; (3) a dummy variable for whether she has final say about large household purchases; (4) a dummy variable for whether she has final say about daily purchases; (5) a dummy variable for whether she has final say about visits to her family/friends; (5) a dummy variable for whether she has final say about how to spend money earned by her husband. Observations are daughters-in-law of the head of the household living in rural households in which there are exactly two such women.

Table 5: Main result: Effect of mother's rank on child's height-for-age

	(1)	(2)	(3)	(4)
dependent variable:	height-for-age z-score			
lower-ranking mother	-0.357**	-0.377**	-0.382**	-0.422**
	(0.127)	(0.133)	(0.136)	(0.159)
mother's age at birth	-0.0308	-0.0241	-0.0177	-0.0102
	(0.0220)	(0.0233)	(0.0289)	(0.0391)
mother's height			0.0190	0.0241†
			(0.0143)	(0.0142)
joint household fixed effects	✓	✓	✓	✓
age in months×sex controls	✓	✓	✓	✓
demographic controls		✓	✓	✓
mother specific controls			✓	✓
father specific controls				✓
n (children in joint households)	1,078	1,078	1,075	1,069

Data: India 2005 DHS. Standard errors clustered by PSU. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$. Controls added: (1) age-in-months by sex dummies; (2) a dummy for whether the child is first born to her mother, whether she is a single birth, the child's birth order in the joint household, whether the mother is currently pregnant; (3) mother's height, dummy variables for years of education completed by the mother, mother's age at marriage; (4) dummy variables for level of father's education, father's age at the time of the survey. z-score computed according to WHO 2006 method.

Table 6: Falsification test: No observed effect in dissolved joint households that split before birth of child

	(1)	(2)	(3)	(4)	(5)	(6)
dependent variable:	ln(height in centimeters)					
data source:	IHDS	IHDS	IHDS	IHDS	DHS	DHS
households included:	joint & split	joint	split	joint & split	joint	joint
purpose:		replication	falsification		main result	main result
lower-ranking mother	-0.0176*	-0.0368**	0.00228	-0.0345**	-0.0142**	-0.0220**
	(0.00863)	(0.0119)	(0.0127)	(0.0118)	(0.00475)	(0.00475)
lower-ranking × is split				0.0370*		
				(0.0170)		
joint household fixed effects	✓	✓	✓	✓	✓	✓
age-in-years × sex	✓	✓	✓	✓	✓	✓
DHS demography controls						✓
n (children)	1,659	984	675	1,659	1,078	1,078

Data: IHDS is the India Human Development survey, which tracked households longitudinally; height-for-age z-scores are not used because the IHDS did not measure age in months. DHS is the Demographic and Health Survey, our main data source.

Standard errors clustered at the PSU level are shown in parentheses. †$p < 0.1$, *$p < 0.05$, **$p < 0.01$, ***$p < 0.001$. Columns 5 and 6 use the same DHS sample as our main results in table 5. "Joint" households are as defined in our main sample; "split" households are cases where two adult brothers are married with children, but live separately and are recorded as separate households.

Table 7: Mechanisms: Lower-ranked mothers have lower BMI than their sisters-in-law

	(1)	(2)	(3)	(4)
dependent variable:	mother's body mass index (BMI)			
lower-ranking mother	-0.362*	-0.345†	-0.439†	-0.401†
	(0.142)	(0.177)	(0.233)	(0.237)
pregnant			1.503**	1.119*
			(0.551)	(0.509)
breast feeding			0.277	
			(0.372)	
age of youngest child			-0.00819	
			(0.0112)	
age of youngest child × breastfeeding				✓
mother's year of birth dummies			✓	✓
dummies for number of children ever born			✓	✓
household fixed effects		✓	✓	✓
n (daughters-in-law)	810	810	804	804

Data: India 2005 DHS. Standard errors clustered at the PSU level are shown in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Age of youngest child × breastfeeding is a vector of indicator variables for the age of the woman's youngest child in months interacted with whether or not she is breastfeeding that child. The sample is the set of mothers of children in our main result in table 5.

Table 8: Mechanisms: Effect of mother's rank on child's early-life mortality

	(1)	(2)	(3)
dependent variable:	IMR	PNM	NNM
lower-ranking mother	28.72†	4.791	26.43*
	(15.44)	(10.51)	(11.03)
girl	-12.08	-5.561	-10.04
	(12.72)	(8.981)	(9.804)
year of birth fixed effects	✓	✓	✓
household fixed effects	✓	✓	✓
constant	61.05	12.19	52.03
	(39.35)	(20.47)	(33.76)
n (live births in the last 10 years)	3,227	3,095	3,703

Data: India 2005 DHS. The dependent variable is an indicator for child death, scaled to 0 or 1000 for comparability to published mortality rates. IMR = infant mortality rate (first 12 months); NNM = neonatal mortality rate (first month); PNM = post-neonatal mortality rate (months 2-12). Sample sizes differ because the exposure period is different for these three types of mortality: to be exposed to NNM a baby must have been born at least 1 month before the survey; to be exposed to IMR a baby must have been born at least 1 year before the survey; to be exposed to PNM a baby must have been born at least one year before the survey and have survived the neonatal period. Standard errors clustered at the PSU level are shown in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

A1 Later-born adult brothers are no shorter than earlier-born brothers

Table 1 verified the balance of nuclear family characteristics, and of pre-marriage characteristics of parents, for children of higher- and lower-ranking mothers in joint households in rural India. It also found that among fathers of the children we study, younger brothers are slightly taller than their older brothers, on average. However, this difference was not statistically significant, and was observed in a small sample because the DHS only measured the heights of men in a sub-sample of Indian households.

Table A1 extends this analysis to three larger samples of adult brothers in the 2005 India DHS. The table presents simple regressions of height in centimeters on an indicator for being the older brother, along with a linear control for age or a set of year indicators. Regressions are shown with and without household fixed effects. Panel A uses a sample that includes all adult men who are sons of the household head, and who are living in joint rural households in which there are exactly two such men, whether or not they have children. In different specifications, males under 20 and under 22 are excluded to ensure that younger brothers do not appear shorter merely because they have not yet finished growing to their adult heights. Successive panels B and C restrict the sample from panel A to include only men with children, and then only men with children under 5. The observation that older brothers are, if anything, shorter than younger brothers, on average, appears to be robust to all of these different samples and respecifications, and would be consistent with literature showing that later-born siblings tend to have slightly greater birth weight. To facilitate comparison with height-for-age z-scores in the main results, note that the standard deviation of height for the sample in panel A is 6.9 centimeters.

A2 Evidence from a regional difference in women's social status

India is a diverse country with well-studied regional differences. The status of women is well-documented to be better in the south than in the rest of India. The better level of women's social status in the south of India can be verified in our main DHS dataset. Using the same measure of "say" in household decision making as in our first stage results, we can compare average say among rural women in the states we classify as southern with average say in the rest of rural India. Rural women in southern states report say in 0.11 more decisions [*s.e.* = 0.04], relative to an India-wide average of 2.2 decisions; this advantage increases to 0.19 more decisions [*s.e.* = 0.04] controlling for the standard DHS wealth quintiles and the woman's education. In our main analysis, such state-level differences are absorbed in household fixed effects.

Intrahousehold institutions that rank daughters-in-law and create the unequal allocations of status, work, and nutrition that we document are also well-known to be weaker in the south of India than in the north (Dyson and Moore, 1983). Therefore, if our result is indeed due to such social institutions, it may be absent or diminished in the south of India. Table A2 tests this prediction, returning to our main DHS sample from our main results table 3. We operationalize the south of India as the states Andhra Pradesh, Karnataka, Goa, Kerala, and Tamil Nadu. In the south, there is no evidence of an effect of intrahousehold rank; the coefficient for the rest of India in column 2 is essentially unchanged from table 3. Column 4 suggests that this interaction may be statistically detectable, if differences in mothers' heights (determined before marriage) are used to improve the precision of estimates. This result is consistent with intrahousehold rank being important for child height because of women's social status, rather than for some other, perhaps coincidental, reason.

A2.1 Distribution of sample into Indian states

Because table A2 shows that only a small fraction of our main sample of 1,078 children is from south India, a reader may wonder if it is useful to describe our paper as representative of rural India. Indeed, only a minority of children even in rural India live in joint households. However, as figure A1 shows, the distribution of children into Indian states in our sample closely matches the distribution of children into states in rural India, with an R^2 of 93%. The large northern state of Uttar Pradesh is slightly over-represented, but the neighboring state of Bihar is slightly under-represented. The principal reason that table A2 shows a small fraction of the sample in the south of India is that a small fraction of births in India happen in the south of India. These five states are all among the nine with the lowest fertility rates. In the full rural DHS, only 13% of children under 5 with measured height are from these five states.

A3 Survey experiment on marriage ads for grooms

Our identifying assumption is that differences in the outcomes of children of higher and lower ranking mothers reflects social forces within joint households – not matching processes that assign mothers of higher and lower quality to brothers of different age ranks in marriage markets. In the text, we discussed evidence that real ads for grooms in arranged marriages do not mention the groom's age rank among brothers. Here, we report a survey experiment that we conducted with Indian participants in which we experimentally manipulated ads to include information on hypothetical grooms' rank among brothers.

In a forced-choice survey experiment, we paid participants $0.10 to choose which of two groom-seeking-bride marriage ads would be more attractive to brides' families. Participants were recruited on Amazon Mechanical Turk (mTurk), an online labor market that is widely used in survey experiments in economics (Kuziemko et al., 2015). Ads took the form such as:

> Parents seek suitable BRIDE for GROOM. He is 26 years old, 5'1". fair complexion. Lives in village. 10th pass education. Working in government job. Has one older brother. Caste no bar. Horoscope not required.

Elements of the ad were randomly varied, including the adjective before "BRIDE;" the age, height, complexion, education, and job of the groom; the two short sentences at the end; and the family structure, either "Has one older brother," "Has one younger brother," or neither of these. 2,358 binary choices between groom ads were made between 1,158 distinct pairs of ads; in pairs that were asked about twice, respondents agreed 65% of the time.

Table 1 in the main text of the paper presents results of the experiment: there is absolutely no evidence that the age rank among brothers influenced respondents' decisions. As column 4 shows, this is not because respondents were simply not attending to the content of the ads: other properties proved quite important to respondents' decisions. For example, grooms with a dark complexion were 6 percentage points less likely to be selected, on average. Note that this experiment omits four of the most important dimensions of marriage market matching – religion, caste, location, and dowry price – suggesting that real decisions would be even more constrained than these, so grooms' age rank would likely be even more irrelevant.

As a verification that the randomization was successful in this experiment, table A4 confirms that the random assignment to experimental treatments does not predict survey-reported properties of respondents. Respondents were mTurk users and therefore use computers and understand English. However, other mTurk studies have found clear expression of traditional and discriminatory attitudes in this population, such as discrimination by caste name in charitable giving (Deshpande and Spears, 2016) or opposition to inter-caste and inter-religious marriage (Coffey and Thorat, 2016). Table A5 uses the small sub-sample of respondents claiming to live in rural India to verify that results are the same.

A4 No difference in the BMI of adult male brothers

Table 5 in the main text shows that lower-ranking women have less body mass than their higher-ranking sisters-in-law within the same household, on average. We interpret this as evidence for maternal nutrition as a mechanism linking women's social status to their children's outcomes. In this section, we verify that no similar difference is present for adult male brothers, who would be the husbands of the women we study. This is a verification that our results are not driven by general net nutritional differences between the two nuclear families within the joint household; this result is expected because households in the DHS are defined to share cooking and eat out of a common pot.

Table A6 presents results of regressing adult men's BMI on an indicator for being the younger brother. We construct two samples to account for the fact that only small subsamples are used in the men's recode: men in households with two married adult brothers, and men in households with two married adult brothers both of whom have children. There is no evidence from any specification that younger brothers tend to have lower BMI than their older brothers within the same household.

A5 Policy relevance: Many development programs target women's social status to promote child health

The evidence of this paper is important because the effects of women's social status have long been an important question in economics and in social sciences more generally, and because child height is an important marker of health and early-life human-capital accumulation. But it is further the case that the evidence of this paper is relevant to development programs and policies that explicitly target women's social status as a technique to promote children's outcomes. For example, the Ministry of Women and Child Development of the Government of India has a website listing sixteen "Women Empowerment Schemes." This section provides quotations and examples to demonstrate (1) that many organizations base policy on the belief that a woman's social status has a causal effect on her own children's outcomes, and (2) many programs are designed to attempt to change social status (such as norms around decision-making power) directly, rather than by distributing material resources such as money or food to women (which, of course, many programs also do).

As evidence of the policy relevance of this belief, it is worth presenting a full quotation from a high-profile and widely cited International Food Policy Research Institute (IFPRI) report, dedicated to the claim that women's status matters for child nutritional outcomes, such as height (Smith et al., 2003). These are the opening words of the foreword of the report, in the name of the Director General of IFPRI:

> Many researchers in the international development field have been startled to note that although child malnutrition is rampant in both Sub-Saharan Africa and South Asia, it is much more widespread in South Asia. According to other Millennium Development Goal indicators, children in South Asia should be in better shape. What lies behind the so-called Asian Enigma? One hypothesis holds that regional differences in women's status — their power relative to men — account for much of the regional differences in children's health and nutrition. In this report, Lisa Smith and her co-authors examine data from these regions, and from Latin America and the Caribbean, to show that a mother's ability to make decisions at home and in her community not only affects the care she receives and thus her own nutritional well-being but also enables her to provide better care and nutrition for her children. (page ix)

Although it is not the purpose of our paper to discuss the Asian Enigma, this document was remarkable as a high-profile statement from a leading food policy organization (which often emphasizes agriculture) that social status of mothers is an important cause. The main contribution of the report is a detailed set of observational regressions using DHS data, documenting partial correlations between the exact same child height-for-age and decision-making say variables that we use in our investigation.

A further IFPRI report (Smith et al., 2013), published a decade after the report quoted above, emphasizes both continuity of this belief and a lack of new evidence: "Many development programs that aim to alleviate poverty and improve investments in human capital consider women's empowerment a key pathway by which to achieve impact and often target women as their main beneficiaries. Despite this, women's empowerment dimensions are often not rigorously measured and are at times merely assumed." They emphasize the intra-household mechanisms at the core of our study: "In addition to being an end goal in itself, women's empowerment is also considered as a means by which to achieve other important development outcomes, such as improvements in child nutritional status. As women are often the primary caretakers in a household, intrahousehold dynamics that determine allocation of resources and their impact on individuals' well-being are increasingly a subject of analysis."

Further examples are readily available:

- The United Nations Development Programme writes on its webpage: "In addition to improving the lives of individual women and girls, gender equality improves the prospects of families, communities, and nations. When gender inequalities are reduces, more children go to school, families are healthier..." "By enhancing women's control over decision-making in the household, gender equality also translates into better prospects and greater well-being of children, reducing poverty of future generations."

- In an evaluation of a Mercy Corps program which "focuses specifically on women's empowerment as a vehicle for improving family nutritional statuses," Scantlan and Previdelli (2013) write "There is a growing body of research that promotes improving women's empowerment as an important point of intervention for improving nutrition. This hypothesis is potentially significant for Timor-Leste as the country dedicates more attention toward alleviating malnutrition. However, the evidence backing the effect of women's empowerment on malnutrition is inconsistent and more research is needed" (page E1).

- Ruel et al. (2013), describing the need for programs targeting the nutritional status of children to move beyond direct provision of nutrients, include "women's empowerment" and further write: "A third way to enhance the nutrition sensitivity of programmes is to engage women and include interventions to protect and promote their nutritional wellbeing, physical and mental health, social status, decision making, and their overall empowerment and ability to manage their time, resources, and assets."

The introduction to our paper drew a distinction between effects of women's *resources* such as cash transfers, and effects of their social status. Many development interventions focus on the latter: "UNDP focuses directly on gender equality and women's empowerment." Such programs often take the form of "behavior change communication," such as through women's meetings or media such as posters, signs, or songs. Many of these meetings and materials explicitly discuss the sort of intra-household structures (such as between mothers-in-law and daughters-in-law) that our empirical strategy exploits. For example, two of the authors (Coffey and Spears) have visited and attended meetings of a program in central Uttar

Pradesh by Rajiv Gandhi Mahila Vikas Pariyojana (RGMVP — "mahila vikas" is Hindi for "women's development"), an Indian NGO whose funders include the Bill and Melinda Gates Foundation. RGMVP's central activity is a program of village "Self-Help Group" (SHG) meetings of different ages and social ranks. Presenting RGMVP's "Theory of Change," their materials claim that "SHGs provide unifying platforms through which a diverse group of women can demonstrate strong bonds of mutual concern. Our Model enables women to disrupt deep hierarchies, which we believe are the source of poverty, inequality and exclusion. Beyond impacts on **health**, livelihoods and incomes, women experience increased levels of self-confidence and sustained agency" (emphasis added).

A similar program, operated in Bangladesh by the international NGO CARE, was studied by Smith et al. (2013): "the SHOUHARDO acronym stands for 'Strengthening Household Ability to Respond to Development Opportunities' and also means friendship or amity." The program included a focus on women's empowerment in order to "address deeper structural causes" of poor child outcomes such as stunting. The approach to women's social status resembled RGMVP's: "The central intervention designed to do so was Empowerment, Knowledge and Transformative Action (EKATA) groups, which established a recognised and accepted forum for women to meet and express themselves in a public role. The groups, comprised of twenty women and ten adolescent girls, provided a platform for empowering women and girls through education, solidarity, group planning, and rights advocacy." Smith, et al., in a USAID-funded mixed-methods (quantitative and qualitative) evaluation, conclude that "the project's women's empowerment interventions were found to have a strong independent impact on stunting." Melinda Gates (2014), describing the strategy of the Gates Foundation, explicitly cites this CARE program after writing "There are strong associations between women's empowerment and specific health and development outcomes. For example, women's control over resources is associated with better outcomes in family planning; maternal, newborn, and child health; nutrition; and agricultural development."

It is not a goal of this paper to evaluate any specific program or policy designed to improve women's social status. However, these examples make clear that many development policymakers operate under a belief that women's social status is a meaningful determinant of their children's outcomes, and that they attempt to influence it not only though providing cash transfers or marketable skills, but also through social interventions — sometimes focused on intra-household rank — that are intended to manipulate women's status directly. Several of the quoted authors emphasize that the evidence base for these programs and policies remains importantly incomplete, including in recently published reports. This paper contributes well-identified empirical evidence to this important question.

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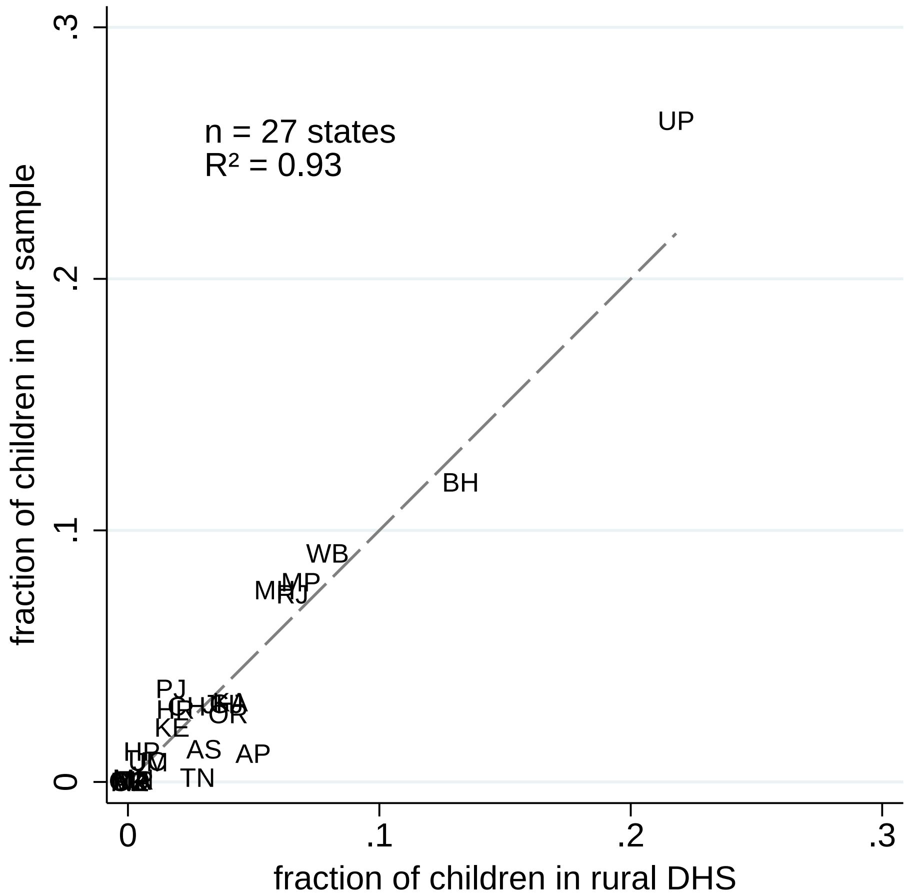
Figure A1: The distribution of children into states in our sample is similar to rural India



Each observation is a state included in India's DHS. The vertical axis plots the fraction of the 1,078 children in our main sample that are from each India state. The horizontal axis plots the fraction of all rural children with measured height (so, under 5 years old) that are from each Indian state. The 45° line is included for reference. In a simple linear regression, the 95% confidence intervals for the intercept and slope include 0 and 1, respectively.

Table A1: Heights of adult brothers, in centimeters

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
included ages:	all ages	over 22	over 22	over 20	over 20	over 20	over 20
year of birth FEs	✓	✓		✓		✓	
household FEs						✓	✓
Panel A: Adult brothers living in joint households							
older brother	-0.248*	-0.393*	-0.416*	-0.429**	-0.444**	-0.208	-0.147
	(0.121)	(0.164)	(0.162)	(0.141)	(0.139)	(0.236)	(0.211)
age (linear)			-0.00815		-0.00950		-0.0371
			(0.0179)		(0.0153)		(0.0345)
n (adult men)	9,939	5,292	5,292	6,770	6,770	6,770	6,770
Panel B: Adult brothers living in joint households, with a child							
older brother	-0.538	-0.477	-0.537	-0.499	-0.524	-0.563	-0.425
	(0.365)	(0.369)	(0.362)	(0.367)	(0.359)	(0.675)	(0.658)
age (linear)			0.00799		0.0116		0.0331
			(0.0293)		(0.0285)		(0.0900)
n (adult men)	1,388	1,347	1,347	1,381	1,381	1,381	1,381
Panel C: Adult brothers living in joint households, with a child under 5							
older brother	-0.466	-0.383	-0.362	-0.415	-0.350	-1.017	-1.675
	(0.427)	(0.434)	(0.420)	(0.429)	(0.414)	(0.861)	(1.047)
age (linear)			-0.00665		0.000352		0.258
			(0.0423)		(0.0403)		(0.185)
n (adult men)	1,029	988	988	1,022	1,022	1,022	1,022

The dependent variable is height in centimeters. Data are from the men's recode of the India DHS, 2005. Standard errors clustered by survey PSU. Two-sided p-values: †$p < 0.1$, *$p < 0.05$, **$p < 0.01$, ***$p < 0.001$.

Table A2: No effect of mother's intrahousehold rank in south India

	(1)	(2)	(3)	(4)
dependent variable:	height-for-age z-score, WHO 2006			
sample:	south only	non-south only	full sample	full sample
rank	0.113	-0.275*	-0.282**	-0.324**
	(0.280)	(0.108)	(0.107)	(0.109)
rank × south			0.457	0.593†
			(0.363)	(0.358)
interaction p-value			$p = 0.209$	$p = 0.098$
household FEs	✓	✓	✓	✓
age-in-months × sex	✓	✓	✓	✓
mother's height				✓
n	97	981	1,078	1,075

Standard errors clustered at the PSU level are shown in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. The sample is the same as our main sample in table 3. "South" is operationalized as Andhra Pradesh, Karnataka, Goa, Kerala, and Tamil Nadu. India, 2005 DHS data.

Table A3: Rank among brothers is never mentioned in ads about grooms seeking brides

	Fraction of ads mentioning characteristic	
characteristic:	Uttar Pradesh	Banerjee, et al. (2009) Table C4
rank among brothers	0	not reported
number of brothers	0	not reported
number of siblings	0	not reported
age	1.00	0.98
religion	1.00	1.00
caste	1.00	0.97
education	0.93	0.78
height	0.91	0.90
type of job	0.88	0.61
income	0.44	0.22
appearance	0.28	
language	0.02	
n	148	8,038

See discussion in section 2.2 of the main text. We extended the methodology of Banerjee, et al. (2009) by cataloging 148 advertisements, randomly selected from 5 newspapers in Uttar Pradesh, describing grooms seeking brides for arranged marriage. Our sample of ads appears to match Banerjee, et al.'s on the dimensions that they observed, suggesting that the ads we cataloged were not unusual. Not a single ad mentioned the groom's rank among his brothers – or any other fact about his siblings – which is evidence for the fact that marriage markets do not importantly consider this factor.

Table A4: Balance of observed properties of respondents in survey experiment

	(1)	(2)	(3)	(4)	(5)
property of respondents:	age	urban	male	no high school	college graduate
F-test of balance:	0.723	1.632	1.612	0.180	0.456
p-value:	0.577	0.164	0.169	0.949	0.768
ad A is older brother	-0.276	0.00452	0.0132	0.0123	-0.00109
	(0.264)	(0.0176)	(0.0236)	(0.0176)	(0.0264)
ad A is younger brother	-0.304	-0.0224	-0.0506	0.00972	-0.0272
	(0.272)	(0.0190)	(0.0243)	(0.0175)	(0.0263)
ad B is older brother	0.238	-0.00483	0.0253	0.00749	-0.00554
	(0.282)	(0.0188)	(0.0239)	(0.0175)	(0.0263)
ad B is younger brother	-0.00341	0.0230	0.00388	0.00648	0.00638
	(0.281)	(0.0170)	(0.0232)	(0.0171)	(0.0258)
constant	29.64	0.881	0.761	0.113	0.573
	(0.211)	(0.0139)	(0.0175)	(0.0128)	(0.0194)

Each column is a separate regression in which the listed property of the survey respondent is the dependent variable. The F test is the test that all four of the randomized properties (older and younger brother, ad A and B) jointly fail to predict the property of the respondent.

Table A5: Survey experiment results are robust to restricting to rural sub-sample of respondents

	(1)	(2)
sample:	full	reported rural only
is older brother	0.0135	0.000829
	(0.0160)	(0.0611)
is younger brother	0.00326	0.00819
	(0.0155)	(0.0576)
constant	0.512***	0.513***
	(0.0103)	(0.0377)
n	2358	183

Column 1 is the same sample as is used in the full results of the experiment, in table **??**. Column 2 restricts the sample to respondents who report living in rural India; there is no evidence that respondents incorporate information on the age rank among brothers of potential grooms in evaluating these hypothetical ads.

Table A6: No difference in the BMI of adult male brothers

	(1)	(2)	(3)	(4)
	husband's body mass index (BMI)			
Panel A: Married adult men in rural households with two such men				
younger brother	0.0923	-0.0147	0.557	0.561
	(0.174)	(0.244)	(0.421)	(0.421)
height (cm)				-0.0268
				(0.0340)
n (adult brothers)	954	954	954	954
Panel B: Married adult men with children in rural h.h. with two such men				
younger brother	0.160	0.0883	0.718	0.719
	(0.220)	(0.315)	(0.537)	(0.539)
height (cm)				-0.00979
				(0.0403)
n (adult brothers)	647	647	647	647
household FEs		✓	✓	✓
man's year of birth dummies			✓	✓

Data are rural observations from the 2005 DHS men's sample. Standard errors clustered by PSU. † $p < 0.1$.

Table A7: Main result effect is larger, on average, for older children

	(1)	(2)
	height-for-age	
lower-ranking mother	-0.288**	-0.459**
	(0.101)	(0.158)
lower rank × age	-0.0143*	-0.0147*
	(0.00587)	(0.00578)
household FEs	✓	✓
age-in-months × sex	✓	✓
extended controls		✓
n children	1,078	1,069

Identical DHS sample as main result. Age-in-months is demeaned to preserve comparability with main result. "Extended controls" is the identical full set of demographic, mother, and father controls as in the main result. Standard errors clustered by survey PSU. † $p < 0.10$; * $p < 0.05$; ** $p < 0.05$.

Table A8: Main result is robust to including joint households with 3 or more daughters-in-law

	(1)	(2)	(3)
dependent variable:	height-for-age	height-for-age	height-for-age
sample:	2 DIL	2+ DIL	2+ DIL
is lower ranked of two	-0.245*		
	(0.102)		
rank number (linearly)		-0.203**	
		(0.0777)	
is ranked #2			-0.280**
			(0.104)
is ranked #3+			-0.410*
			(0.200)
joint household fixed effects	✓	✓	✓
age-in-months × sex	✓	✓	✓
n (children)	1,078	1,219	1,219

Column 1 is the identical DHS sample as main result. $1,078 \div 1,219 = 88\%$ of children with measured height in households with two or more such daughters-in-law live in a household with precisely two. Column 2's independent variable is the rank number: 1, 2, 3,. . . . In column 3 highest-ranking is the omitted category. Standard errors clustered by survey PSU. † $p < 0.10$; * $p < 0.05$; ** $p < 0.05$.