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

Diane Coffey, Reetika Khera, and Dean Spears*

September 25, 2015

Abstract

We study the intergenerational effect of a woman's social status on the health of her children. We overcome limits to causal identification in existing literature by exploiting an institutional feature of joint households in rural India: women married to the younger brother are assigned lower social rank than women married to the older brother. Studying differences among cousins within households, we find that children of lower-ranking mothers are shorter than children of higher-ranking mothers. We further argue that this effect is in part due to differences in maternal nutrition: as evidence of this mechanism, we document effects of being a lower-ranking mother on post-partum BMI and birth weight in newly collected data, and on neonatal mortality in national data. Because the relatives who arrange the marriages that we study do not take the groom's age rank among brothers into consideration, more disadvantaged mothers are not more likely to be married to younger brothers. We show that women whose marriages assign them lower social rank are not disadvantaged in health, height, or human capital before marriage and that the difference in social status that we exploit emerges after marriage. We verify that our results are not due to pre-marriage differences between fathers or mothers; are not caused by endogenous household dissolution; and are only present among children of higher- and lower-ranking mothers living together in a joint household, but not among the children of brothers living in separate households.

*Coffey & Spears: Indian Statistical Institute - Delhi and Woodrow Wilson School, Princeton University. Khera: Indian Institute of Technology, Delhi. Corresponding author: Diane Coffey, Economics and Planning Unit, Indian Statistical Institute, New Delhi 110 016, India; coffey@princeton.edu. We appreciate comments and suggestions from James Berry, Anne Case, Angus Deaton, Jean Drèze, Andrew Foster, Michael Geruso, Jeffrey Hammer, Ravinder Kaur, Leigh Linden, Will Masters, Sara McLanahan, and audiences at the PAA Economic Demography Workshop, NEUDC, the Delhi School of Economics, UT Austin, Tufts, and Princeton. All errors are our own.

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). Although a convincing estimate of an effect of a woman's social status on her children's outcomes would be of clear importance, no prior paper in economics has identified 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 identify. 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.[1] Finally, programs or events that 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 child education, but they note that they cannot rule out direct effects of TV on children's education, through mechanisms other than women's status.[2]

[1]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.

[2]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

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, 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), 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, 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, marriages do not sort women of different quality into different 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 quarter of a height-for-age standard deviation shorter than their cousins born to the higher-ranking mother within the same joint household. In support of this main result, we verify "first-stage" institutional consequences of intrahousehold rank. Women married to the younger son indeed have observably lower social status than women married to the older son: we show that lower-ranking daughters in law report less decision making power than their higher-ranking counterparts and, in a time use survey, spend less time outside the home, a standard measure in the literature of autonomy among the women we study. We are also able to document one plausible and independently important mechanism for our effect: maternal nutrition. A

cooking fuel use, but cannot separate the mechanism of women's status from households' preference to invest in a son's health. 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).

mother's low intrahoushold rank – well-documented to be associated with restricted 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.

We directly address the central threat to the identification of our main result: the possibility that differences in child outcomes may be due to pre-marriage differences in the quality of mothers or fathers. In the spirit of verifying that randomization balances treatment assignment, we show that the rank of the mothers of the children we study is balanced with respect to characteristics of mothers and fathers that are fixed before marriage. 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. This balance is expected because the institution of arranged marriage in India does not sort bride quality by husband birth order; indeed, following Banerjee et al. (2009) we show in a supplementary appendix that advertisements offering grooms never mention age rank among brothers. We emphasize that there is no systematic difference within households in the height of the fathers of the children we study, who are brothers: we demonstrate this in the text, and provide further evidence in supplementary appendix A1. Finally, we observe that the difference in heights of the children of higher- and lower-ranking mothers is seen at all heights of their mothers and of their fathers.

Similarly, we verify that the effect on child height that we identify only occurs among children exposed to the institutions of joint households. This is because the variation in social hierarchy and rank that we study is particularly imposed and enforced in joint households, especially by the coresident grandparents. In a falsification test exploiting a separate data source that has longitudinally tracked households, we see that the effect of this measure of mother's status on child height is only found for children who are born into joint households; there is no difference in height between the children of adult brothers who live separately.

This falsification test verifies that our results are not due to the possibility that households that are joint at the time of the survey are endogenously formed, dissolved, or selected in ways that drive the results.

In the second part of our analysis, we show that a mechanism for the effect of mother's social status is maternal net nutrition. This evidence is important because it clarifies and emphasizes the causal chain linking status to child outcomes. We find that lower-ranking mothers have worse nutritional status than higher-ranking mothers. A woman's body mass, which depends on her food consumption and energy expenditure, influences her ability to nourish her child *in utero* and while breastfeeding. Although lower-ranking mothers are no shorter than higher-ranking mothers (height is determined before marriage, in childhood), they have less body mass (which reflects recent net nutrition). Further, we find that children of lower-ranking mothers are more likely to die in the neonatal period, their first month of life, when infant mortality is particularly dependent on birth weight and maternal nutrition, but are equally likely to die in their second to twelfth months of life, when infant mortality in India is more importantly determined by the infectious disease environment. Finally, in newly collected data 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).

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 (Bertrand et al., 2003; Vogl, 2013).

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. Section 6 presents evidence that maternal nutrition is one important mechanism through which women's status affects child height in this context. Section 7 concludes.

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 brothers, and his brothers' wives and children. Figure 1 diagrams such a joint household. Approximately 8% of rural children under five live in joint households of this type. 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).[3]

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).

[3]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.

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 arranged marriage markets do not match the bride quality to the groom's birth order. Although 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 order of the husband-to-be is not an important factor for decision-makers in marriage markets. The 2005 India Human Development Survey found that 95% of marriages in rural India are arranged (Banerji et al., 2013), meaning that the parents or extended family members of the bride and groom decide whether or not a couple will marry. As we will show, these decision-makers pay essentially no attention to a prospective groom's birth order. This institution is analogous to other South Asian child rearing and marriage practices, studied in the economics literature by Jayachandran and Kuziemko (2011) and Vogl (2013) 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 externality 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 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.

We extended their analysis by conducting our own catalog of classified ads for grooms in India; a table of characteristics included in the advertisements is presented in Supplementary Appendix table A3. The distribution of characteristics mentioned in our ads is very similar to the advertisements reported by Banerjee et al. (2009). The birth order rank of the husbands was *not mentioned in a single advertisement*.

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

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. 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.[4]

Our empirical strategy uses household fixed effects to compare cousins born to different mothers living in the same household. Throughout the paper, for various outcomes y, we estimate regressions of the form:

$$y_{cmh} = \beta \textit{lower rank}_{mh} + X_{cmh}\theta + A_{cmh}\lambda + \alpha_h + \varepsilon_{cmh}, \tag{1}$$

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 vary depending on the application. We include a full set of 120 child age-in-months by sex

[4] In supplementary results (available online in the working paper version), we demonstrate that our results are robust to instead using a more inclusive sample of children living in households with two daughters-in-law where one or both has children under five years old. Although children without a cousin under 5 in such a sample would not contribute directly to estimating the main regression coefficient of interest, due to household fixed effects, they could change the coefficients on regression controls.

indicators, A_{cmh}. Standard errors are clustered by survey primary sampling unit, which is the village in the rural sample.

Our independent variable of interest, *lower rank*, 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 or random treatment assignment 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 1 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. The table is produced by estimating regression equation 1 with no further controls in Panel A, and with cohort controls for mother's and child's year of birth in Panel B.[5]

Our paper's main result is that, within the same household, the children of lower-ranking

[5]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.

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 1 shows that these two adult brothers, within a household, have indistinguishable height, education, and employment. Table 1 shows that the average child born to the lower-ranking mother has a father who is slightly and insignificantly taller;[6] however, men's height is from a small sample of 408 children and 313 fathers. Section A1 of the Supplementary Appendix uses larger samples of adult brothers living in joint households, who may or may not have children under five, to provide further evidence that there is no difference in brothers' heights that could be driving our results. Therefore, within joint households, the parents of the children we study are balanced in their height, in their human capital, and in the economic resources that fathers may be able to direct to their own children.

4 First stage: Household rank and women's status

Before proceeding to our main results, we show that intrahousehold rank is related to women's status. These results are comparable to "first stage" results verifying that an experimental treatment was implemented as intended, or that an instrumental variable is indeed related to the factor of interest. First, we use survey self-reported say in household decision making included in the DHS, which the DHS describes as measures of women's

[6]Note that, on average, there are no systematic heritable difference in genetic potential height between brothers of different birth orders.

empowerment. We show that this difference does not reflect pre-marriage differences in personality or preferences that women bring into marriage at different ranks: rather, the difference emerges with duration of marriage. Then, using a separate data source on time use, we show a difference between higher- and lower-ranking women in time spent outside the home; in this cultural context, women's mobility outside the home is a measure of her social status (Rahman and Rao, 2004; Kabeer, 1999).

4.1 Reported decision-making

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, in addition to any causal effect of women's status.

Table 2 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 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 say.

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.

As the diverging lines show, 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.[7]

The ITUS, collected in 1998-9, interviewed the members of rural households in six states in India.[8] 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. For more information on the ITUS, see MOSPI (1999).

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 a first-stage effect of intrahousehold rank on women's status.

5 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

[7]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.

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

children of higher- and lower-ranking mothers and finds that, at all ages, there is a visible 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 will 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 3 presents results of our fixed effect regression equation 1, estimated 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 0.25 height-for-age standard deviations 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.[9] 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, the mother's age at the time of the child's 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

[9]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.

[$s.e. = 0.106$; $t = 4.06$]. This control exploits 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).[10]

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 – 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 3 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 1 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

[10]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.

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.[11] 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

[11]The IHDS data are publicly available (Desai et al., 2008), and are described by Desai et al. (2009).

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. 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, 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 4 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.[12] 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 3 using the DHS sample and log of child height as the dependent

[12] 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.

variable, to show that the results from the IHDS data are quantitatively comparable to what is found in DHS data.

This result indicates that the effect of mother's rank on child height is due to intrahousehold processes. 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. 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.[13] Balance table 1 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 5 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

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

we identify is consistent with net maternal nutrition being a mechanism for the effect we document on child height.

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 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 6 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.[14] 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

[14]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.

low birth weight.[15] Therefore, these differences in neonatal death suggest the possibility of differences in birth weight, caused by worse net maternal nutrition and health *in utero*, which would also result in childhood height differences.

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.[16] Note that because we observe births in the hospital, we are unable to use household fixed effects with these data.

Figure 7 presents results for child birth weight, and figure 8 presents results for mothers' post-partum BMI.[17] 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 mothers weigh 84 grams less than babies born to higher-ranking mothers [$p = 0.051$]. Lower-

[15] Another determinant of neonatal mortality could be differences in medical care at birth. However, results available in a working paper version provide evidence that lower-ranking mothers do not receive worse medical care at birth.

[16] 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 survey.

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

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.[18] 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. This belief has had considerable influence over the design of development policies and programs. Although many development practitioners assume that such a relationship exists, and 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 has 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.

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. First-stage-type effects of intrahousehold rank on women's reported empowerment

[18] 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.

emerge over time 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 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 importance.

References

Adair, Linda S. 2007. "Size at birth and growth trajectories to young adulthood." *American Journal of Human Biology*, 19(3): 327–337.

Aizer, Anna and Janet Currie. 2014. "The intergenerational transmission of inequality: Maternal disadvantage and health at birth." *Science*, 344(6186): 856–861.

Anderson, Siwan. 2003. "Why dowry payments declined with modernization in Europe but are rising in India." *Journal of Political Economy*, 111(2): 269–310.

Banerjee, Abhijit, Esther Duflo, Maitreesh Ghatak, and Jeanne Lafortune. 2009. "Marry for what: Caste and mate selection in modern India." working paper 14958, National Bureau of Economic Research.

Banerji, Manjistha, Steven Martin, and Sonalde Desai. 2013. "Are the Young and the Educated More Likely to Have "Love" than Arranged Marriage? A Study of Autonomy in Partner Choice in India." *India Human Development Study Working Paper*(8).

Barcellos, Silvia, Leandro Carvalho, and Adriana Lleras-Muney. 2014. "Child gender and parental investments in India: Are boys and girls treated differently?" *American Economic Journal: Applied*, 6(1): 157–89.

Bertrand, Marianne, Sendhil Mullainathan, and Douglas Miller. 2003. "Public policy and extended families: Evidence from pensions in South Africa." *World Bank Economic Review*, 17(1): 27–50.

Black, Sandra, Paul Devereux, and Kjell G Salvanes. 2007. "From the Cradle to the Labor Market? The Effect of Birth Weight on Adult Outcomes." *The Quarterly Journal of Economics*, 122(1): 409–439.

Bozzoli, Carlos, Angus Deaton, and Climent Quintana-Domeque. 2009. "Adult Height and Childhood Disease." *Demography*, 46(4): 647–669.

Breierova, Lucia and Esther Duflo. 2004. "The impact of education on fertility and child mortality: Do fathers really matter less than mothers?"Technical report, National Bureau of Economic Research.

Case, Anne and Christina Paxson. 2008. "Stature and Status: Height, Ability, and Labor Market Outcomes." *Journal of Political Economy*, 116(3): 499–532.

Coffey, Diane. 2015a. "Early life mortality and height in Indian states." *Economics & Human Biology*, 17: 177–189.

——— 2015b. "Pre pregnancy body mass and weight gain during pregnancy in India & sub-Saharan Africa." *Proceedings of the National Academy of Sciences*, 112(11): 3302–3307.

Cummins, Joseph. 2013. "On the Use and Misuse of Child Height-for-Age Z-score in the Demographic and Health Surveys." working paper, UC Riverside.

Desai, Sonalde, Amaresh Dubey, B Joshi, Mitali Sen, Abusaleh Shariff, and Reeve Vanneman. 2008. "India Human Development Survey (IHDS) [Computer file]. ICPSR22626ev1. University of Maryland and National Council of Applied Economic Research, New Delhi [producers], 2007." *Ann Arbor, MI: Inter-university Consortium for Political and Social Research.[distributor].*

Desai, Sonalde, Amaresh Dubey, Brij Lal Joshi, Mitali Sen, Abusaleh Shariff, and Reeve Vanneman. 2009. "India Human Development Survey: Design and Data quality." *IHDS technical paper*, 1.

Duflo, Esther. 2003. "Grandmothers and Granddaughters: Old-Age Pensions and Intra-household Allocation in South Africa." *The World Bank Economic Review*, 17(1): 1–25.

——— 2012. "Women Empowerment and Economic Development." *Journal of Economic Literature*, 50(4): 1051–79.

Dyson, Tim and Mick Moore. 1983. "On kinship structure, female autonomy, and demographic behavior in India." *Population & Development Review*, 9: 35–60.

Hatton, Timothy. 2013. "How have Europeans grown so tall?" *Oxford Economic Papers*, 66: 349–372.

Jayachandran, Seema and Ilyana Kuziemko. 2011. "Why do mothers breastfeed girls less than boys?: Evidence and implications for child health in India." *Quarterly Journal of Economics*, 126(3): 1485–1538.

Jeffery, Patricia, Roger Jeffery, and Andrew Lyon. 1988. *Labour Pains and Labour Power: Women and Childbearing in India*: Zed Book Ltd.

Jensen, Robert and Emily Oster. 2009. "The power of TV: Cable television and women's status in India." *The Quarterly Journal of Economics*, 124(3): 1057–1094.

Kabeer, Naila. 1999. "Resources, agency, achievements: Reflections on the measurement of women's empowerment." *Development and Change*, 30: 435–464.

Kishore, Avinash and Dean Spears. 2014. "Having a son promotes clean cooking fuel use in urban India: Women's status and son preference." *Economic Development and Cultural Change*, 62(4): 673–699.

Ludwig, David S and Janet Currie. 2010. "The association between pregnancy weight gain and birthweight: A within-family comparison." *The Lancet*, 376(9745): 984–990.

Maluccio, John A, John Hoddinott, Jere R Behrman, Reynaldo Martorell, Agnes R Quisumbing, and Aryeh D Stein. 2009. "The impact of improving nutrition during early childhood on education among guatemalan adults*." *The Economic Journal*, 119(537): 734–763.

Mandelbaum, David. 1988. *Women's seclusion and men's honor*: University of Arizona Press.

Miller, Grant. 2008. "Women's suffrage, political responsiveness, and child survival in American history." *The Quarterly Journal of Economics*, 123(3): 1287.

MOSPI. 1999. "Time Use Survey: Summary of findings." `http://mospi.nic.in/Mospi_New/site/inner.aspx?status=3&menu_id=91`. Ministry of statistics and programme implementation, Government of India.

Munshi, Kaivan and Mark Rosenzweig. 2006. "Traditional institutions meet the modern world: Caste, gender, and schooling choice in a globalizing economy." *American Economic Review*, 96(4): 1225–1252.

Palriwala, Rajni. 1993. "Economics and patriliny: Consumption and authority within the household." *Social Scientist*, 21(9/11): 47–73.

Rahman, Lupin and Vijayendra Rao. 2004. "The determinants of gender equity in India: Examining Dyson and Moore's thesis with new data." *Population & Development Review*, 30(2): 239–268.

Rao, Vijayendra. 1993. "The rising price of husbands: A hedonic analysis of dowry increases in rural India." *Journal of Political Economy*: 666–677.

Rosenzweig, Mark R and Oded Stark. 1989. "Consumption smoothing, migration, and marriage: Evidence from rural India." *The Journal of Political Economy*: 905–926.

Schaner, Simone. 2015. "Do opposites detract? Intrahousehold preference heterogeneity and inefficient strategic savings." *American Economic Journal: Applied Economics*, 7(2): 135–174.

Singh, J.P.. 2005. "The contemporary Indian family." in Bert Adams and Jan Trost eds. *Handbook of World Families*: Sage, Chap. 6: 129–166.

Spears, Dean. 2012. "Height and cognitive achievement among Indian children." *Economics & Human Biology*, 10: 210–219.

Steckel, Richard. 2009. "Heights and human welfare: Recent developments and new directions." *Explorations in Economic History*, 46: 1–23.

Strauss, John and Duncan Thomas. 1995. "Human resources: Empirical modeling of household and family decisions." *Handbook of Development Economics*, 3: 1883–2023.

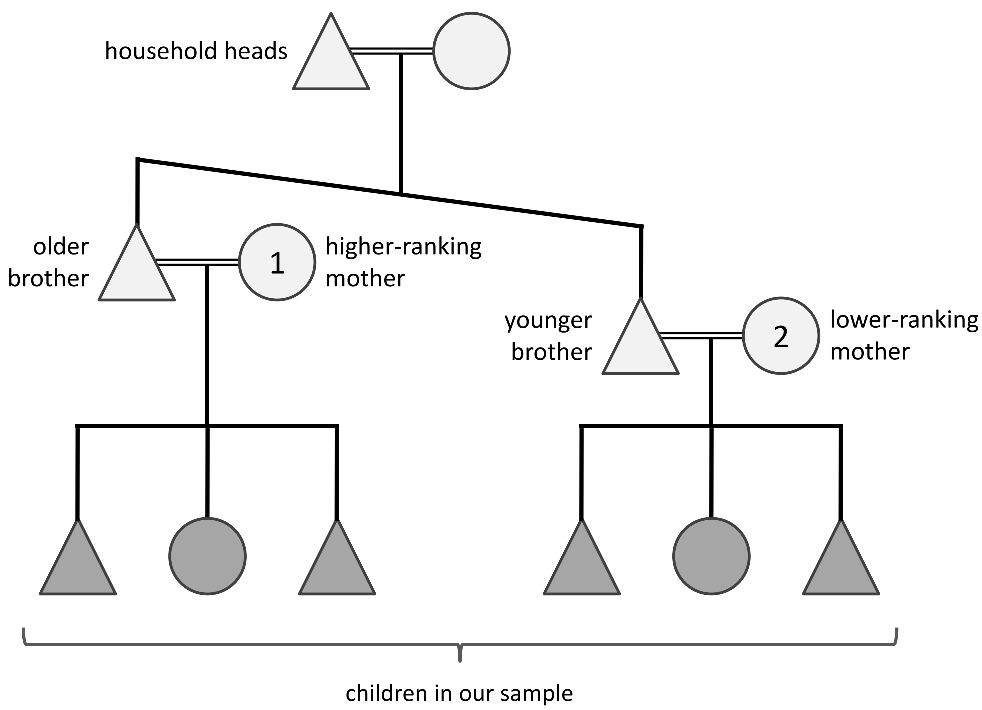
Thomas, Duncan. 1990. "Intra-household resource allocation: An inferential approach." *Journal of Human Resources*: 635–664.

Vogl, Tom. 2013. "Marriage institutions and sibling competition: Evidence from South Asia." *Quarterly Journal of Economics*, 128(3): 1017.

World Bank. 2001. *Engendering development: Through gender equality in rights, resources and voice*: World Development Report.

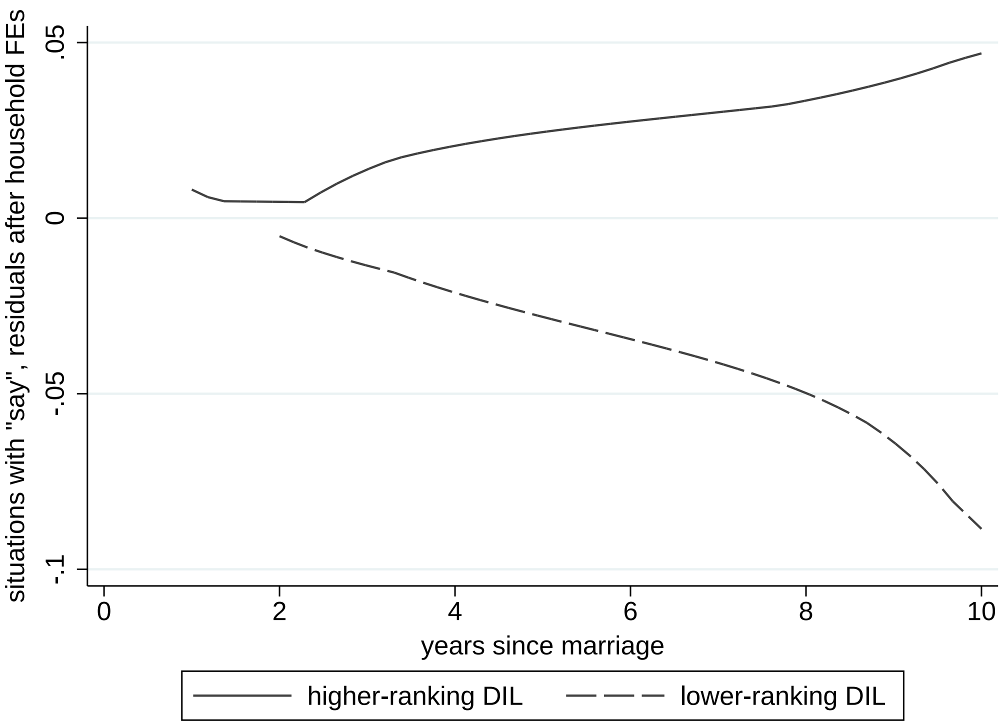
Yaktine, Ann L and Kathleen M Rasmussen. 2009. *Weight Gain During Pregnancy: Reexamining the Guidelines*: National Academies Press.

Figure 1: Structure of the joint Indian households we study



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.

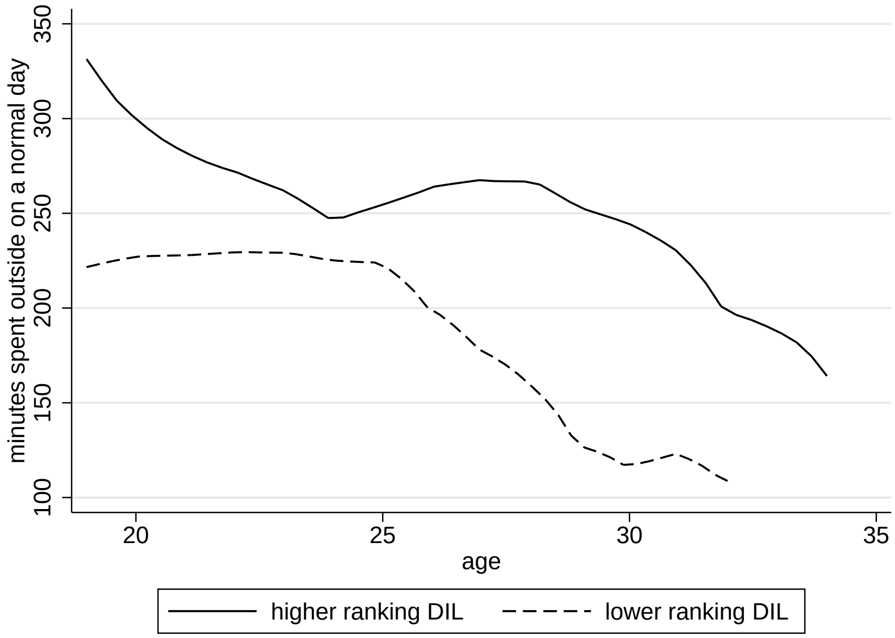
Figure 2: Differences in status unfold after marriage



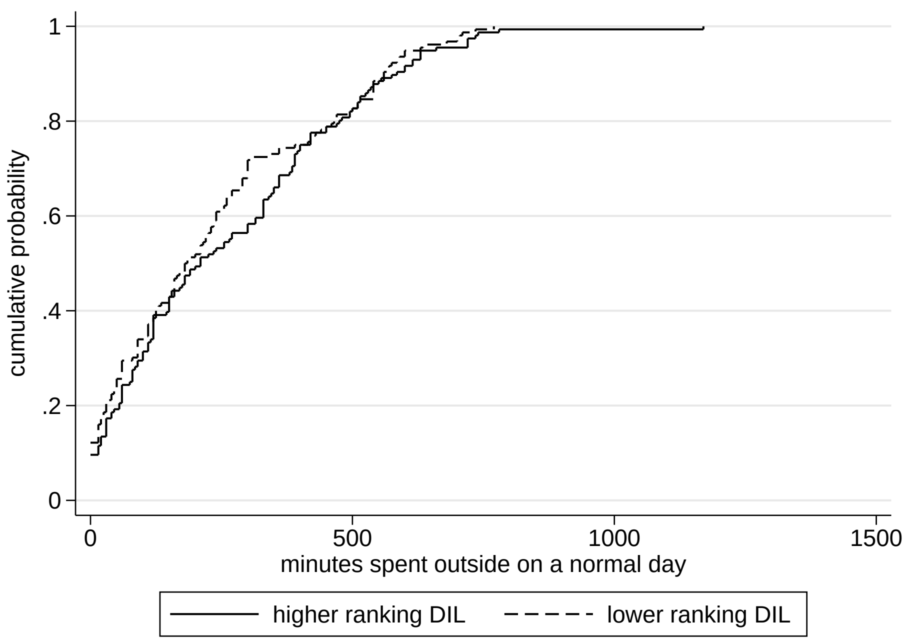
"Say" is the count of five situations in which the woman reports having say rather than not having say, either alone or jointly with her husband. Residuals are after a regression on household fixed effects. Local regression; bandwidth is three years. Data: India 2005 DHS. Observations are the mothers of the children in our main sample, in table 3.

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

(a) time spent outside on a normal day

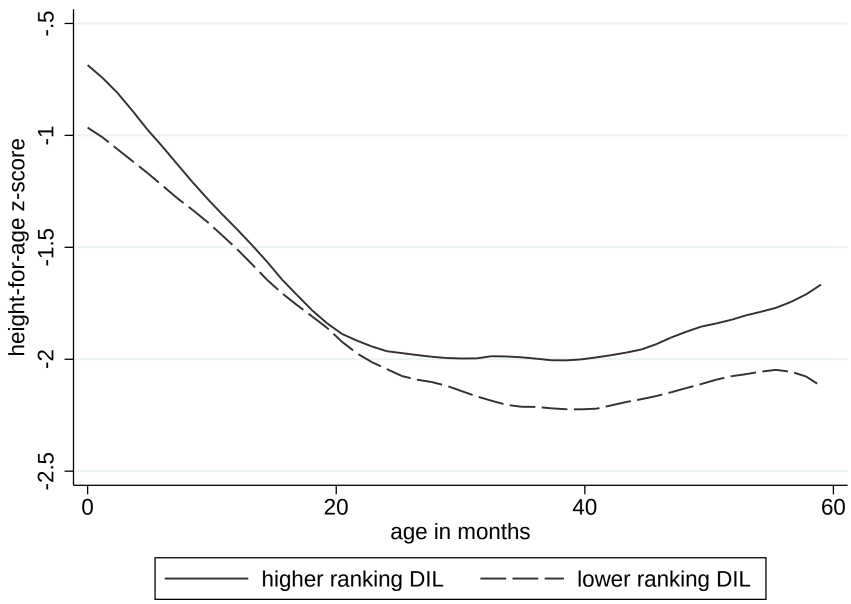


(b) CDF of time spent outside on a normal day



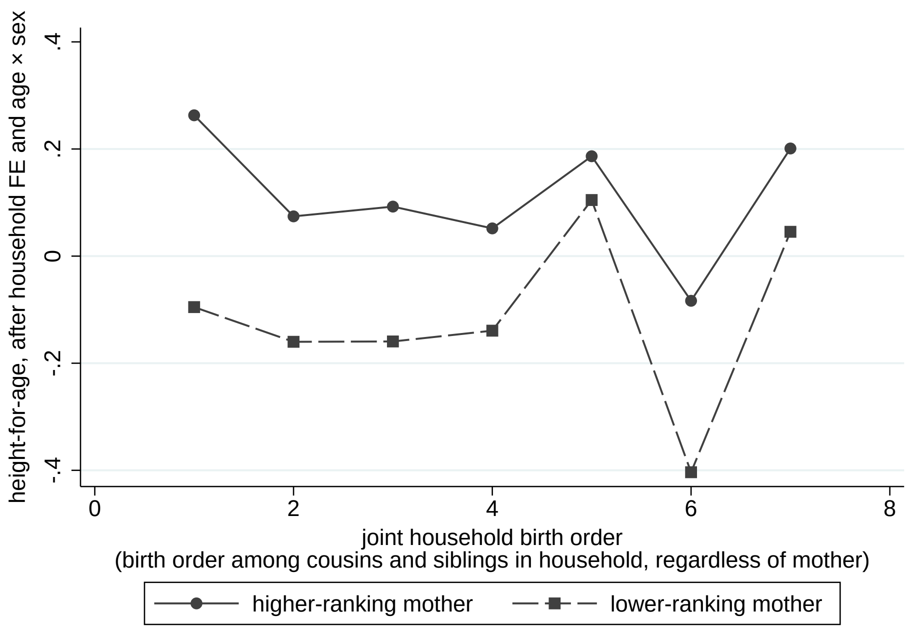
Panel A plots minutes spent outside on a normal day by age using local polynomial regression with an epanechnikov kernel and a bandwidth of 1.75. Panel B is the cumulative distribution function of minutes spent outside on a normal day. n=312. DIL is daughter-in-law. Data: India Time Use Survey (ITUS), 1993. 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: Height-for-age of children of higher- and lower-ranking mothers



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 3. Height-for-age is computed according to WHO 2006 reference norms. Data: India 2005 DHS.

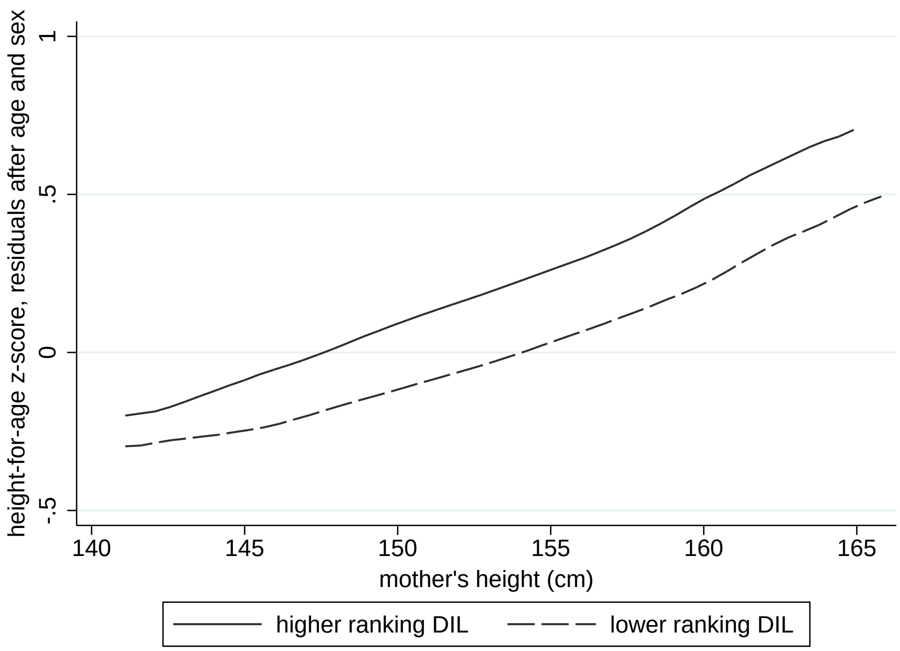
Figure 5: Height difference is present at all joint household birth orders



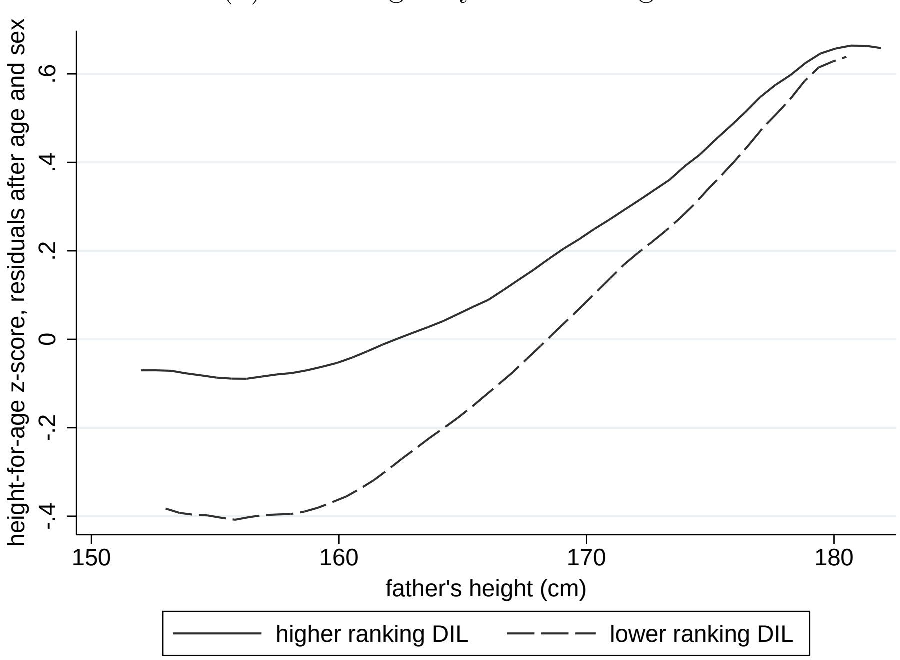
The dependent variable is the height-for-age residual after controlling for household fixed effects and age-in-months by sex indicators. Joint household birth order is age rank among children of both mothers – that is, birth order among all cousins and siblings in the household, regardless of mother. $n = 1078$, identical to main sample in table 3. Data: India 2005 DHS.

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

(a) child height by mother's height



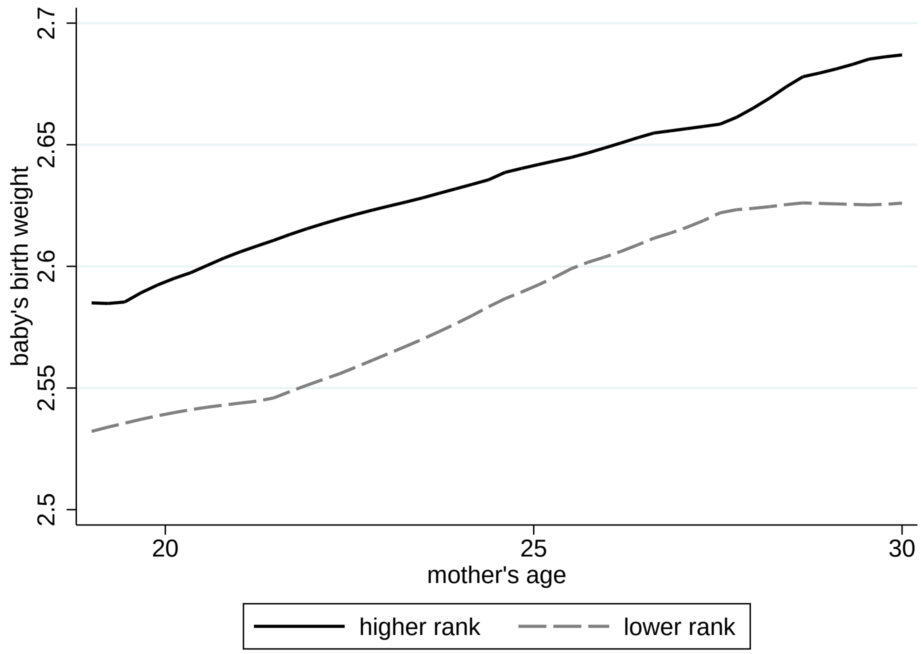
(b) child height by father's height



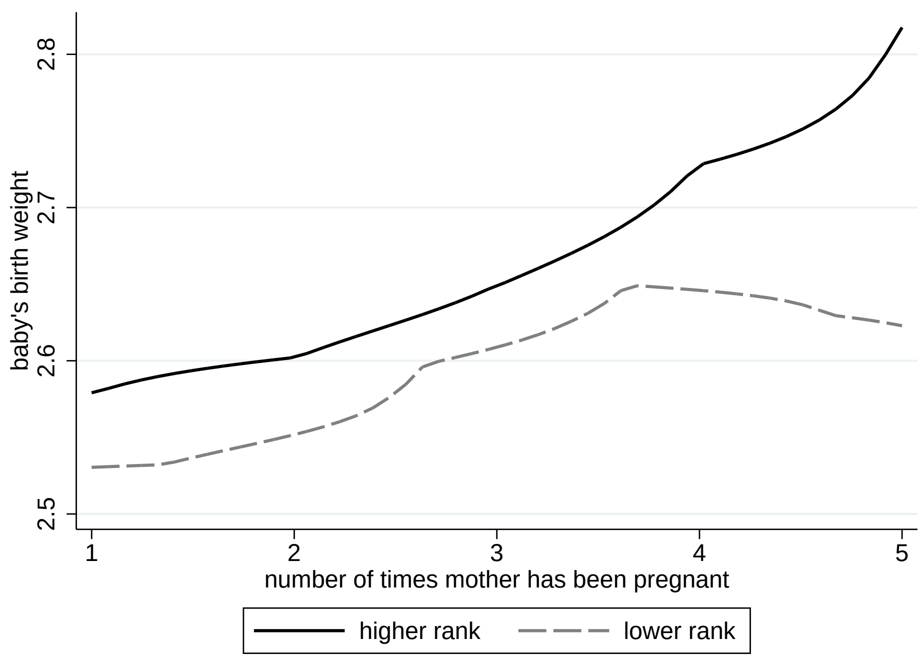
Local polynomial regression with an epanechnikov kernel and a bandwidth of 5cm. In panel A, $n = 1078$, identical to main sample in table 3. In panel B, $n = 408$, because the DHS only measured a sub-sample of men. DIL = daughter-in-law, the mothers of the children we study. Data: India 2005 DHS.

Figure 7: Intrahousehold rank and birth weight in a district hospital

(a) birth weight, conditional on age



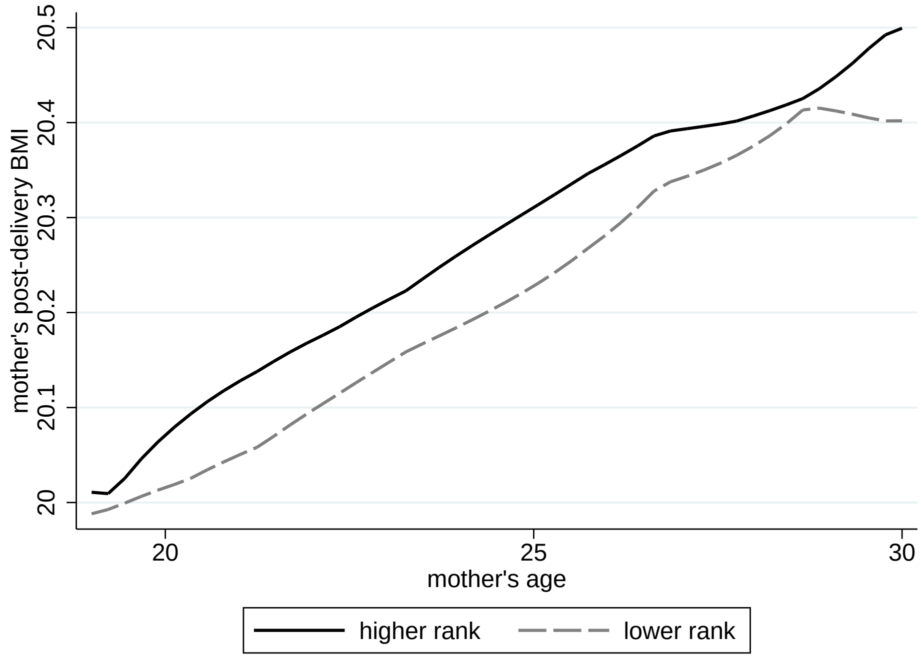
(b) birth weight, conditional on pregnancy count (parity)



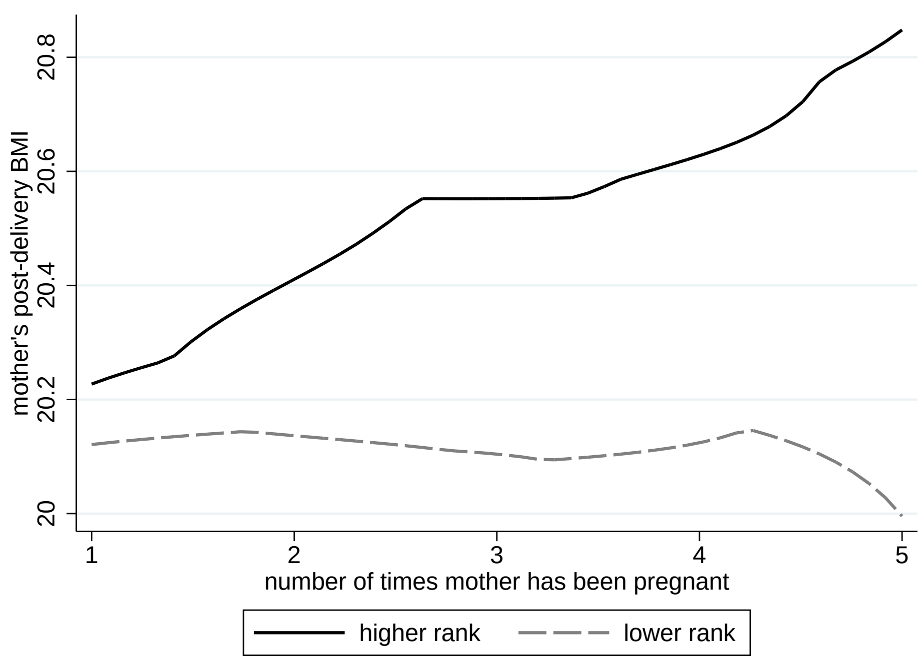
Data are from the authors' collection in a district maternity hospital in Uttar Pradesh. Birth weight was measured within 36 hours of birth. To match our main sample, data are restricted to women in households with exactly two daughters-in-law; note that because data are collected at the hospital, not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

Figure 8: Intrahousehold rank and post partum BMI in a district hospital

(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: 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

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. Data: India 2005 DHS. Observations are children in the main sample in table 3.

Table 2: First stage: Intrahousehold rank predicts reported decision-making "say" within households

	(1)	(2)	(3)	(4)	(5)	(6)
dependent variable:	sum (total of 5)	own health	large purchases	daily purchases	visit family	money
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

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. Data: India 2005 DHS. Observations are daughters-in-law of the head of the household living in rural households in which there are exactly two such women.

Table 3: 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	-0.245*	-0.377**	-0.382**	-0.422**
	(0.102)	(0.133)	(0.136)	(0.159)
mother's age at birth		-0.0241	-0.0177	-0.0102
		(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

Standard errors clustered at the PSU level are shown in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Controls added by column: (1) age-in-months and sex dummies and their interactions; (2) a dummy for whether the child is first born to her mother, whether she is a single birth, her mother's age at the time of her birth, the child's birth order in the joint household; (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. Data: India 2005 DHS.

Table 4: Effect of rank on child height only seen in joint, not split, households

	(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
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

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 3. “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. 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.

Table 5: Mechanism: 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

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. Data: India 2005 DHS. The sample is the set of mothers of children in our main result in table 3.

Table 6: Mechanism: 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

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). Standard errors clustered at the PSU level are shown in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Data: India 2005 DHS.

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. 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 section **??**, 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.

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.

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.