A child feeding a child: early mother's marriage and offspring stunting in India

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Abstract

It is well known that Indian children are among the most stunted in the world. In this paper we show that early marriage of mothers, which is also prevalent in India, has played a key role. Using the totality of India's national health surveys (1993-2021) and age at menarche as an instrument, we document that one year earlier marriage has decreased height-for-age for offspring by 0.20 standard deviations. Exploring mechanisms, we find that early-married mothers delay the introduction of complementary solids feeding and that they breastfeed for less time. We provide a theoretical model and associated evidence that adverse feeding practices are mediated by a mix of knowledge gaps and preference issues. Amid slowly shifting marriage norms, supporting the nutrition efforts of early-married mothers can have profound impacts on intergenerational wellbeing.

Keywords: Early marriage, child stunting, health behavior, rural development

JEL Codes: I12, I15, I18, J12

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1 Introduction

India, home to 440 million children, continues to have one of the highest child stunting rates in the world. Stunting, defined as height-for-age which is at least two standard deviations below the World Health Organization's standard median, stands at 35.5% for children under five in the country. It is particularly prevalent in rural areas - where the majority of the population lives - with some regional variation as demonstrated in **Figure 1**. Being short is not a *genetic* attribute of the Indian population: Indian children of immigrant parents in advanced economies grow to the height of their peers even while their parents are shorter (Alacevich and Tarozzi, 2017). Rather, stunting is a major signal of malnutrition early on in childhood, and for this reason stunted children are not only too short for their age but also potentially face irreversible delayed cognitive development later on in life (Alam et al, 2020).

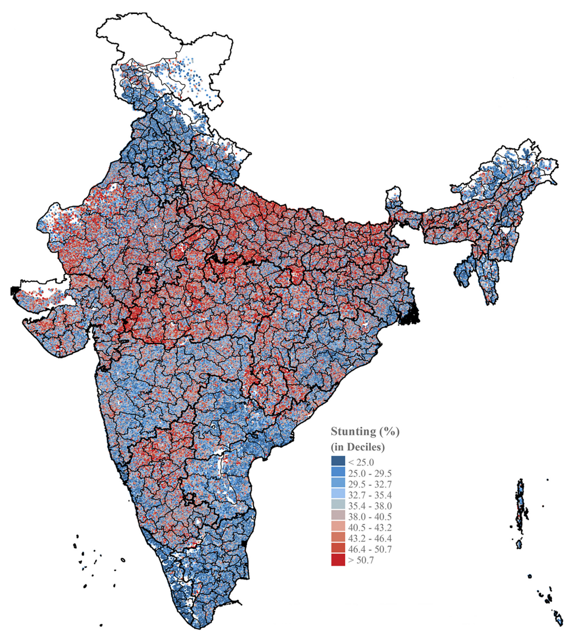


Figure 1: Stunting rates for children under five in rural India

Source: Reproduced from Kim et al (2021).

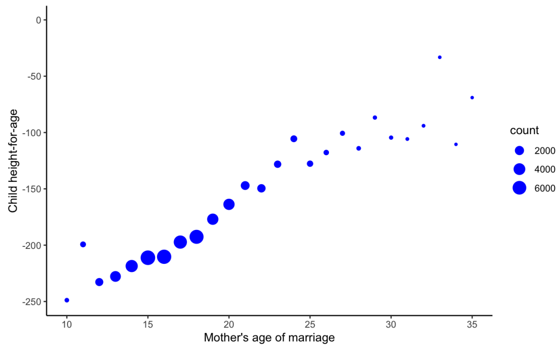
Improvements in child stunting in India have advanced only slowly over the past decade despite large-scale government feeding programs, and a plethora of research has attempted to pinpoint the deeper socioeconomic causes of malnutrition and stunting in the country. Previous empirical work explored the causes of Indian children's height disadvantage relative to sub-Saharan African children, identifying the contribution to the height gap of firstborn son preference in India as well as nutrient loss from exposure to open defecation (Jayachandran and Pande, 2017; Spears, 2020). Meanwhile, evidence of a more correlational nature abounds, documenting associations between child stunting and several facets of socioeconomic disadvantage including extreme poverty, rural residence, and parental illiteracy (Khan and Mohanty, 2018; Chaudhuri et al 2023).

In this paper, we explore the contribution to child stunting of a socioeconomic phenomenon hitherto not studied in-depth in the stunting literature: widespread *early marital age* of the mother in India. India is home to 223 million out of 650 million women worldwide married as minors, and the country's median age of marriage for women, currently nineteen, is among the lowest globally (UNICEF, 2021). Mothers play a key role in early child nutrition and therefore in child growth, and, as noted by an Indian journalist, India may have a problem of "a child mothering a child" (Bhattacharya, 2019). While the health literature has documented negative correlations between early marriage of mothers and child health in various settings, age of marriage is correlated with many family of origin characteristics (Paul, 2019; Ahonsi et al, 2019), making it is difficult to draw conclusions from these studies or to understand underlying mechanisms. The closest to our paper, Nguyen et al (2019), documents an association between early pregnancy and childhood undernutrition in India using 2016 data, and finds a range of other correlated socioeconomic variables.

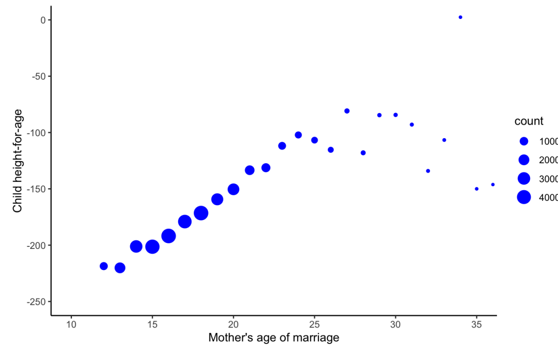
To motivate our investigation, we begin by documenting two empirical regularities which arise in all of India's National Family Health Surveys (NFHS), containing half a million mother-child observations from 1993 to 2021. First, the majority of mothers in India had married young: nine out of ten ever-married women were married by the age of 24, with a median age of 18 over the period. Therefore, half of them married as minors.[1] By 2021, median age at marriage had only etched to 19 nationally. Second, for mothers in India married by 24, i.e. the vast majority, there is a persistent correlation over time between age at marriage and subsequent child height-for-age in the surveys, with children being more stunted on average when born to earlier-married mothers. Conversely, for mothers married after their mid-twenties, this correlation

[1]For comparison, note for instance that this was the case in China in 1950, whereas today average women's marital age in China stands at about 26.

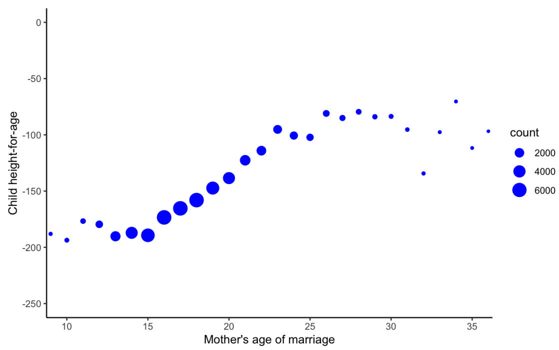
with their children's height disappears. This is shown in **Figure 2**.



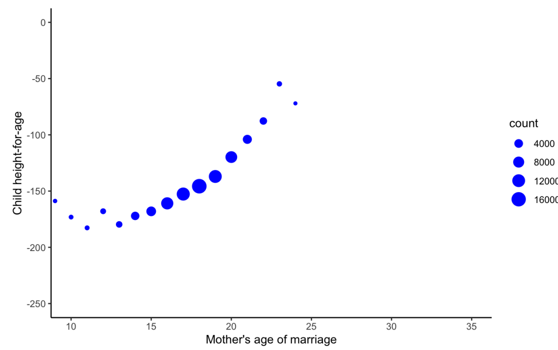
(a) NFHS-1, 1993



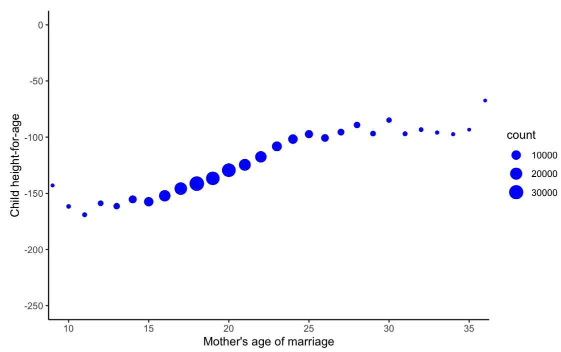
(b) NFHS-2, 1999



(c) NFHS-3, 2006



(d) NFHS-4, 2016



(e) NFHS-5, 2021

Figure 2: Correlations across surveys

For each of the five surveys, Figure 2 plots each mother age at marriage (x-axis) and, for that age, average offspring height-for-age in '00 standard deviations (y-axis). Circle size represents the count density of mothers married at this age in the survey. Note that the NFHS-4 did not collect data on mothers older than 24, hence the sharp stop in plot (d).

To investigate whether early marital age in India has driven offspring stunting or whether the correlations reflect socioeconomic confounders of early marriage, we use the woman's age at her first menstrual period - also known as *menarche* - as an instrument for her age at marriage. The intuition for the instrument, which was first proposed by Field and Ambrus (2008), is that in societies with early marriage norms, the timing of marriage can be influenced by the onset of puberty, while the latter is largely biologically determined and independent of potential confounders.

Before employing the instrument, we show that earlier menarche drives earlier marriage and that the instrument is strong, with the first stage F-stat surpassing the threshold for IV inference set in Lee et al (2022). We also argue that it survives concerns about exogeneity. In particular, we highlight the evidence that the strongest contribution to age at menarche variation is genetic, and that while too-late menarche may be driven by severe undernutrition in childhood, variation in menarche is otherwise largely stochastic. Accordingly, we exclude from the analysis outlier menarche ages as they may be driven by earlier nutritional problems and family-of-origin dynamics, show that for the remaining age range there is no correlation between age at menarche and mother's height as proxy for her earlier nutrition, and check the robustness of the results to including the latter as a control.

For the analysis we use data from three of the five waves of the National Family Health Survey (NFHS-1 in 1993, NFHS-4 in 2016, and NFHS-5 in 2021), which are all the waves containing data on age at menarche, and which also contain information on anthropometric measures and nutrition of children under the age of five.[2] Exploiting the entirety of the available data allows us to maximize power and to check for trends over time, the latter being important in the context of potentially evolving norms around marriage and nutrition. In the analysis, we focus on mothers who were at most 24 years old by each survey. As we explain, this limits the potentially serious problem whereby older women may use age of marriage to recall age at menarche, creating false correlation between the instrument and endogenous variable.[3]

With over 162,000 mother-child observations in the subsequent analytical sample, we first document *large* adverse effects of (instrumented) early marriage on the child's height-for-age, with one year earlier marriage reducing offspring height-for-age by 0.20 standard deviations. Our specification controls for, among other things, mother birth

[2]The second and third rounds (1999 and 2006) did not ask about age at menarche.

[3]It also ensures that the analysis is focused on observations which drive the correlations in the data (mothers married before mid-twenties)

year and the child's birth order, so that we compare same birth-order children of same age mothers who married at younger versus later ages. We find no impact of early marriage on the birth weight or probability of premature birth, signalling that the stunting is driven by postnatal - and not prenatal - mechanisms.

To put the magnitude of the results into perspective, a crude linear estimate would suggest that, had the median age of marriage for women been 25 instead of 18 over the period under study, the median child in India would have had near average height-for-age by World Health Organization standards, almost entirely eliminating the stunting skew in the height distribution of Indian children.[4]

We then turn at length to underlying mechanisms. First, we explore whether the results are mediated by the potential effect of early marriage on age at conception or on educational attainment for the mother. We find that (instrumented) earlier marriage precipitates time at conception *and* reduces educational attainment, but only the former appears to act as a link to greater offspring stunting.[5] In other words, the key to the intergenerational impact on child growth appears to be that the early married mother also becomes a young mother, and less so that she is prone to being pulled out of school as a result of early marriage.

This begets the question of what it is that early married - and as a result younger - mothers *do* which precipitates higher child stunting rates, particularly during the crucial growth window in the first years of life. On growth, there is widespread medical agreement that, for the first four to six months of life, breastmilk provides the optimal mix of nutrients for physical and cognitive development (Lessen and Kavanangh, 2015), and that complementary feeding of mushy solids - such as mashed vegetables and carbohydrates - is critical for infants starting at six months old, as milk alone is no longer sufficient for growth after this point (Fewtrell et al, 2017). There is also evidence that protein intake generally is important for growth, although this evidence is more robust for preterm infants (Tonkin et al, 2014) than in general (Millward, 2017).[6] In addition to direct intake, loss of nutrients through exposure to

[4]This is found by adding 7*0.20 standard deviations to the median height-for-age for children in India over the period, which was -1.50 standard deviations.

[5]The main results become insignificant when conception time - but not when education - is controlled for.

[6]The latter study notes that, among animal-based foods, only milk has been consistently demonstrated to contribute to linear growth.

disease, primarily diarrhea, can expose children to stunting (Nasrin et al, 2023).

Accordingly and to study nutritional mechanisms, we investigate whether age at marriage - instrumented by menarche - impacts breastfeeding for at least four to six months and/or the introduction of complementary solids feeding thereafter. We also study how early marriage affects child intake in the past day of protein-rich foods including dairy, legumes, and eggs, as well as recent occurrence of child diarrhea. Because the availability of this information changes between the earliest and latest surveys,[7] we examine these channels separately by survey, and check whether they are consistent with the survey-specific estimates on stunting.

We find evidence that early marriage of the mother delays complementary feeding of infants and contributes to suboptimal breastfeeding practices. One year earlier marriage reduces the probability that a child was introduced to complementary feeding during the critical age of 6-8 months by 14.4%, and it delays the time at which complementary feeding begins in general by 0.36 months. It also reduces the probability of breastfeeding for at least six months by up to 7.1%.[8] This evidence is also consistent with the results on stunting by survey: the impact of instrumented early marriage on both infant nutrition and child stunting is driven by the NFHS-1 and NFHS-4. In the NFHS-5, by contrast, the coefficients on both are smaller and insignificant, also matching the weaker descriptive correlation between early marriage and child stunting in the survey (**Figure 2e**).

Given the evidence on suboptimal feeding practices, we ask if early-married mothers simply do not *know* how to best feed their children, or if something else is at play. We provide further evidence - as well as a theoretical model of mother behavior - that mothers' knowledge *and* preferences mediate the results. Differentiating the main results by child birth order and gender, we find that the impact of marital age on stunting is higher for girl children than for boy children, and that it is higher for firstborns than for laterborns when it comes to boy children. By contrast, early marriage continues to reduce girl children's height strongly regardless of birth order.

An accompanying model shows these findings are consistent with a situation in

[7]NFHS-1 has information on introduction of complementary feeding but not on consumption of a range of protein nutrients, while NFHS-4 and NFHS-5 have information on the latter, not the former.

[8]In contrast, we find little impact on the child's prior-night consumption of other protein rich foods or on the prevalence on diarrhea.

which (i) mother's early marital age adversely affects child nutrition and growth, (ii) mother effort can mitigate this impact, (iii) effort is higher when the child is a boy, reflecting son preference, and (iv) later children can benefit from the mother's childbearing experience but also suffer from resource constraints, and the latter is more likely to dominate when effort is low. In sum, it appears that while early age at marriage for the mother contributes to poor infant nutrition and subsequent stunting, circumstances relating to the experience and/or burden from raising other children, combined with priorities or preferences in the household vis-a-vis children, can act to buttress or amplify the impacts on the child.

We conclude by offering robustness checks and by briefly addressing the recent improvement of the early marriage-stunting relationship that we observe in NFHS-5. We investigate if the latter can be explained by a nationwide scheme in India which began in 2018 and aimed to raise awareness among mothers about optimal child feeding practices. We find that the districts targeted by this scheme by survey time - comprising half of all observations and chosen in part on the basis of high initial rates of child stunting - did *not* experience a weakening of the early marriage-stunting relationship. Recent improvement therefore came from other districts, primarily in the South, which had *not* yet been targeted by the scheme and which had less prevalent stunting to begin with. What we find thus adds to the mounting evidence on divergent economic and human development in India (Bardhan, 2010), and we leave it to future research to more closely study and track these pockets of regional development.

Our paper speaks directly to the question of, and mechanisms behind, high rates of child stunting in the world's most populous country. Whereas previous work in the Indian context focused on Indian children's height *gap* relative to sub-Saharan African children, and therefore on factors that are prevalent in India beyond what would be expected with "only" underdevelopment (e.g. firstborn son preference and widespread open defecation), we identify for the first time the contribution of a practice - early marriage - which is prevalent in, but not in any way unique to, India. At the same time, our findings support some of the earlier findings particularly around the role of son preference in mediating stunting outcomes (Jayachandran and Pande, 2017).

The policy implications of our findings are nontrivial. Although the Indian government recently raised the legal age at marriage for women (The Hindu, 2023), it is doubtful that *practiced* age at marriage, which has flouted legal thresholds for decades, will rise quickly. Our results suggest that, in the presence of sticky marital age norms and weak enforcement, improving the nutritional practices of early-married mothers toward offspring is key, as is understanding why some policy schemes - such as the one we look into briefly - may not have been successful at doing this. Beyond India, and because early marriage is widespread in other parts of the developing world - including other South Asian nations and sub-Saharan Africa - our paper points to the possibility of an understudied connection between early marriage, feeding practices, and high rates of stunting elsewhere in the global South. We are not aware of other papers exploring causality and mechanisms behind these potential links.

On early marriage, we contribute to the growing literature using an instrumental variable approach to study the consequences of this practice, most of which focuses on outcomes for the mother (e.g. Roychowdhury and Dhajima, 2021; Carpena and Jensenius, 2021). The closest to our paper in this literature is Chari et al (2017), which quantifies the impact of early marriage in India on a different set of child health indicators and using the smaller India Human Development Survey of 2005.[9] Of methodological note is Asadullah and Wahhaj (2019), which has data on Bangladeshi sister-pair women that can be used for a menarche IV with household-of-origin fixed effects; there are minor or no differences when these fixed effects are used, suggesting that family of origin concerns are not plaguing the analysis.[10]

2 Data and Methods

2.1 Data and sample construction

The National Family Health Survey of India is a multi-round survey conducted by the Ministry of Health and Family Welfare, coming out in five waves beginning in 1992-1993 (NFHS-1) and continuing to 1998-1999 (NFHS-2), 2005-2006 (NFHS-3), 2015-2016 (NFHS-4), and 2019-2021 (NFHS-5).[11] A repeated cross-sectional panel, the survey provides the largest-scale representative source of information on popula-

[9]The authors look at the effect on antenatal care, home birth, vaccination completion, and breastfeeding duration. The IHDS sampled married women between 15 and 49 in 40,000 households.

[10]In particular, including sister fixed effects changes the impact of age of marriage on the woman's education slightly, but it does not alter the effect of age of marriage on the woman's age at first birth. The latter variable is particularly important in any study of intergenerational effects of age of marriage. In our paper, the NFHS data does not track sisters to allow for such fixed effects.

[11]For brevity we use the end date of the surveys when referring to them throughout the paper.

tion and family health in India. The majority of waves survey ever-married between the ages of 15 and 49 and ask about motherhood attributes as well as health information on children up to five years old. Between all five waves there are nearly half a million mother-child observations in total, making the NFHS the largest worldwide within the umbrella Demographic and Health Survey (DHS) program.

We construct our analytical sample by first focusing on the waves which elicited information on mother's age at menarche. This leads us to restrict our attention to the first, fourth, and fifth survey waves, as the second and third did not ask mothers about the age at which they reached puberty. Within these, we restrict on mother's age at interview time as well as age at menarche. Specifically, we select mothers who were aged under 25 by survey time, and who did not have menarche too late, calculated as an upper limit of third quartile menarche age plus one year.[12] The former aims to stem issues arising from possible recall bias, whereby older women may forget age at menarche and erroneously recollect it using remembered age at marriage,[13] while the latter aims to exclude outlier menarche ages possibly influenced by nutritional issues in childhood (see below). We also remove the minority of twin children in the data in order to be able to differentiate meaningfully by birth order.

The resulting analytical sample contains over 162,000 unique mother-child observations with information on mother age at marriage and age at menarche, child health and nutrition, and family characteristics relating to caste, religion, assets, and rural/urban status, among other variables.

Notably, despite covering the same range of topics, the different waves are not homogeneous in all the variables they elicit. For our purpose, a key difference is in the questions asked about the nutrition of the child under five. The NFHS-1 inquired about duration of breastfeeding as well as the time at which complementary feeding was introduced, for every child under five. It contains little information, however, on feeding of the child in the prior day and night. In contrast, the NFHS-4 and NFHS-5 did not collect information on time at which complementary feeding was introduced, for any children. They did collect information on breastfeeding duration but only for the last child born in the last five years, as well as on feeding of an array of nutrients

[12] This is subsequently age 16 for NFHS-1, and age 15 for NFHS-4 and NFHS-5.

[13] In email correspondence with NFHS officials, the potential severity of recall problems was cited as the reason for which information on mothers was collected *only* on mothers who were up to 24 by interview time in the NFHS-4.

for the last child in the prior day and night.

2.2 Empirical strategy

We are interested in the impact of mother's marital age on offspring stunting, captured in the child's height-for-age. A simple OLS regression would be as follows:

$$H_{ij} = \theta_0 + \theta_1 M_i + \mathbf{X}'\mathbf{\Theta} + \epsilon_{ij} \quad (1)$$

where H_{ij} is the height for age of child j of mother i, M_i is the mother's age at marriage, and $\mathbf{X}'\mathbf{\Theta}$ is the product of control by coefficient vectors. Mother-specific controls would include her current age while child-specific controls would include the child's birth order and gender, among other variables.Therefore, the specification would effectively be comparing the outcomes of children of a similar birth order and gender, born to similarly aged mothers, but one of whom married earlier than the other. The use of an age-adjusted outcome allows us to avoid controlling for child age, so that mother's marital timing variation can drive conception timing variation.

In the absence of selection problems, θ_1 in Equation (1) would identify the impact of marital age on the child's height-for-age, with $\theta > 0$ indicating that early marriage increases stunting, but it is difficult to argue that there is no such selection.

Early marriage of women is usually facilitated by the family, and the descriptive evidence is robust that earlier-married mothers tend to come from less educated as well as more rural families (UNICEF, 2023). Ethnographic research also suggests that a mix of social norms and incentives particularly among rural Indian families drive early marriage of daughters (Abbhi et al, 2013). These include the belief that the daughter cannot attract a good husband if she is older, an ability to lower the dowry paid (an illegal but still widespread practice) if she is younger, the desire to have one less mouth to feed, the desire to protect communal reputation around daughter chastity, and social judgment from neighbors the longer a daughter remains single.

To the extent that differences along family of origin and also along mother characteristics cannot be controlled for with observables, selection bias would exist. Family of origin unobservables may pressure θ_1 upwards, as better-off families drive up both daughter's age at marriage and her offspring's health, while mother unobservables may pressure θ_1 downwards, as potential health and other aspects of "desirability" may drive the daughter's marital age down and her offspring's health up.

To attenuate selection bias, we resort to a two-stage approach wherein we use variation in age at marriage driven by variation in age at menarche:

$$H_{ij} = \beta_0 + \beta_1 \widehat{M_i} + \mathbf{X}'\mathbf{\Psi} + \varepsilon_{ij} \tag{2}$$

and

$$\widehat{M_i} = \alpha_0 + \alpha_1 Z_i + \mathbf{X}'\mathbf{\Phi} + \nu_{ij} \tag{3}$$

where Z_i is the mother's age at menarche, $\widehat{M_i}$ are the fitted values for age at marriage derived from the first stage regression, and $\mathbf{X}$ is once again the vector of controls.

To the extent that the instrument is relevant and valid, then β_1 is a consistent estimator of the impact of marital age on child height-for-age. The next subsection explores instrument relevance and validity more closely.

2.3 Age at menarche as an instrument

While social acceptability, weak enforcement, and legal loopholes in India have allowed marital practices to remain somewhat detached from *de jure* minimum-age requirements, an alternative *de facto* lower bound on marital age has historically been age at puberty, which is usually between 11 and 15 for most girls. As the onset of fertility, puberty is viewed in various religious and cultural contexts as the age above which consummation of marriage may be sanctified (Haar and Duncan, 2023).

Figure 3 uses the NFHS data to plot age at first period, on the x-axis, and the average age at marriage for the women who had their first period at that age, on the y-axis. It illustrates that, while the average parent in India did not marry their daughter *as soon as* she reached puberty, puberty still influences the timing of marriage, with daughters who reached it earlier on average also marrying earlier. As another illustration, **Figure 4** plots the *distribution* of age at marriage for each age at menarche, and shows shows that this distribution shifts to the right - i.e. marriage is somewhat later - for each age at menarche.[14]

[14] The density plot has peaks at whole years due to the NFHS-1, where age at marriage is not reported in month increments but only in whole years.

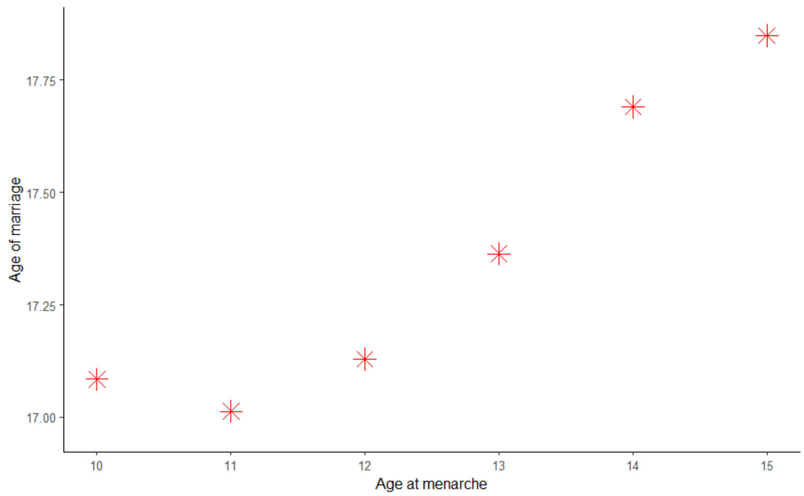


Figure 3: Menarche and mean age at marriage

Figure 3 plots age at menarche on the x-axis and the average at marriage for women who had their first bleeding at that age on the y-axis.

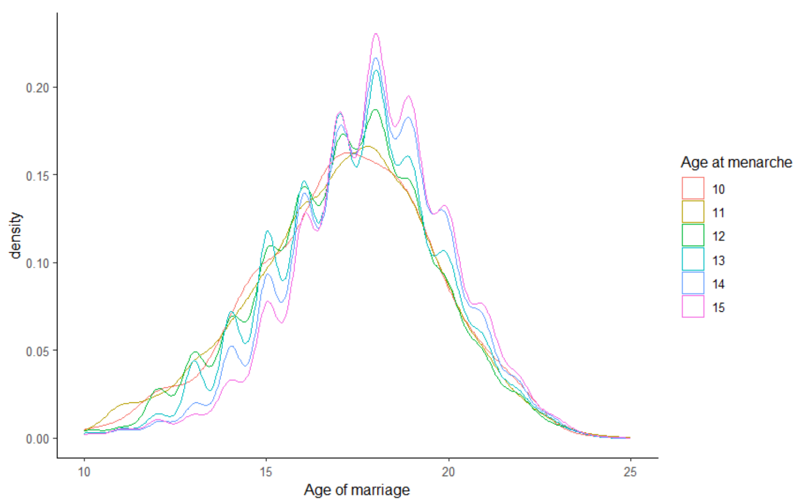


Figure 4: Menarche and distribution of age at marriage

Figure 4 plots the distribution of marital age for each age at menarche in the sample.

These correlations are a precursor to the strong first stage results in the analysis in Section 3 (below). Importantly, we show that while the correlation between age at menarche and at marriage is highest in the NFHS-1, it holds in the NFHS-4 and NFHS-5 as well: each of these surveys separately demonstrates a significant positive coefficient on the instrument and with an F-stat ≥ 104.7.

Regarding validity of the instrument, the age at which a girl's first bleeding occurs is understood to be a complex trait with a strong genetic component. Twin and familial studies indicate that up to 80% of variation in menarche timing is driven by hereditary genes involved in steroid - especially estrogen - metabolism (Dvornyk and Waqar-ul-Haq, 2012), and genome mapping projects show that timing of puberty is a "highly polygenic childhood trait" (Day et al, 2017) influenced by dozens of genes. Importantly, there is evidence that this estrogen-metabolizing genetic component contributes only very marginally to the woman's adult height,[15] making it unlikely to influence height-for-age of young children via inherited height-relevant genes.

Given a strong genetic component to age at menarche which is orthogonal to offspring outcomes, it remains to address the residual *environmental* factors that influence mother's age at menarche and that can be intergenerationally significant. Most importantly, menarche is linked to body fat increases which signal hormonally that the girl's body is ready for reproduction (Deardoff et al, 2014), making adequate nutrition in childhood relevant to age at menarche. For this reason, girls who have severe malnutrition during childhood may have their menarche delayed above and beyond what can be accounted for by genetic components (Soliman et al, 2014).

Validity would therefore be violated *if* the woman's age at menarche is driven to a significant extent by her childhood nutrition, and that this would bias our results toward a null coefficient. This is because, in India, higher fat intake during childhood is associated with higher socioeconomic groups (Mathew et al 2023), and therefore earlier menarche, and *earlier* marriage, would be associated with improved offspring height-for-age outcomes. Therefore, if the instrument suffers from such validity issues, β_1 in Equation (3) would be pressured downward even if the true population parameter is $\beta_1 > 0$, and we would interpret our results as *underestimating*, not overestimating, the impact of marital age on offspring height-for-age.

To examine this possibility more closely, we remove outlier menarche ages from our analysis in order to avoid including observations possibly influenced by severe

[15] There are few studies on the impact of this genetic component of age at menarche on the woman's adult height. The one comprehensive study (Onland-Moret et al, 2005) finds that although one year later age at menarche may cause the woman to grow taller by allowing for more leg bone growth, it would do so by only 0.35 cm per later year of menarche, and therefore with menarche age variation accounting for at most 1% of the height variation between adult women. Other later studies touching on this find no clear patterns of increased or decreased adult woman height with these estrogen metabolizing genes (Geczik et al, 2022).

nutritional problems, and we examine the relationship between age at menarche and the woman's adult height as a proxy for her childhood nutrition. **Figure 5** illustrates that the distribution of women's height by distinct ages at menarche is largely similar, and we certainly do not see the leftward shift (women being shorter for later menarche ages) that we would expect if later menarche was associated with malnutrition in childhood.[16] Despite the absence of a clear correlation between our instrument and the woman's height, we also include the latter as a control in the robustness checks.[17]

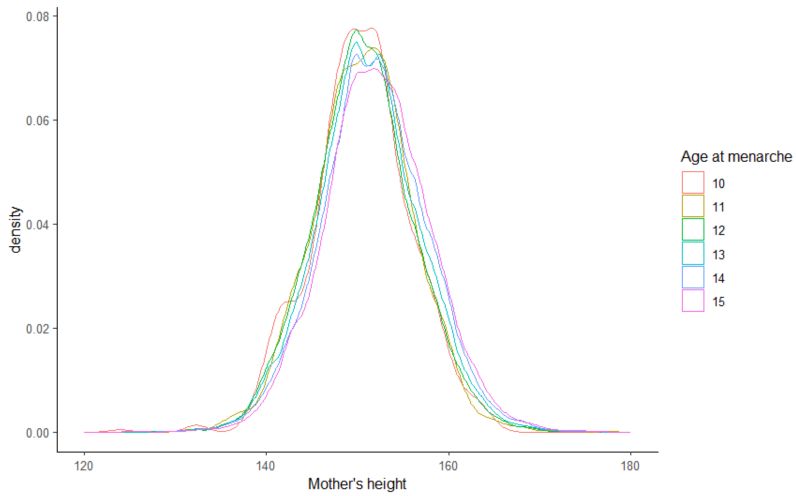


Figure 5: Distribution of mother height for each age at menarche

Figure 5 plots the distribution of the mother's height, for each age at menarche in the sample.

Before presenting the main results, we note a distinct selection issue, related not to the validity of the instrument but to the interpretation of the results, of possible selection in the *offspring*. This would occur if early marriage induces significantly higher child mortality, in which case the remaining pool of children would be the healthier survivors and possibly also with higher height-for-age. Once more, this would cause us to underestimate, not overestimate, the impact of marital age on offspring height-for-age. We show in the robustness checks that instrumented marital age does not

[16]The very small shifts to the right may be a consequence of the marginal genetic contribution of later menarche to woman bone leg growth (see prior footnote) in which case it would not be correlated with the socioeconomic confounders of concern.

[17]We do not use the woman's height as a control in the main results (pooled sample) because the NFHS-1 does not have data on the woman's height. In the robustness results, we check that the results for NFHS-4 and NFHS-5 datasets are robust to adding woman's height as a control.

appear to drive infant mortality (nor related factors like delivery complications).

3 Main results

Table 1 displays the results of the two stage least-squares analysis. In this and all subsequent tables, robust standard errors, clustered by mother year-of-birth and state, are shown below the estimated coefficients.

Table 1: Impact of marriage age on offspring stunting

	All (Pooled)	NFHS-1	NFHS-4	NFHS-5
	(1)	(2)	(3)	(4)
	PANEL A. Second stage. Outcome: child height-for-age			
Age at marriage	**0.198*****	0.200**	0.300***	0.169
	(0.046)	(0.078)	(0.041)	(0.122)
	PANEL B. First stage. Outcome: Age at marriage			
Age at menarche	0.122***	0.280***	0.110***	0.084***
	(0.013)	(0.041)	(0.019)	(0.011)
First stage F-stat	450.1	288.5	151.1	104.7
Observations	132,316	9,473	65,391	57,452
Mother, child, and HH controls	Yes	Yes	Yes	Yes

Note: *p<0.1; **p<0.05; ***p<0.01

Table 1 reports the results of 2SLS regression in Eq. (2)-(3), first for the whole pooled sample then by survey wave. Panel A demonstrates the results of the second stage whereas Panel B demonstrates the results of the first stage.

The 2SLS regressions include controls for mother's age and her year of (own) birth, child birth order and gender, and household religion, wealth, state of residence, and urban or rural status. All of these are entered as factors in our regressions, including dummies for mother age which allows for maximum flexibility in the way in which mother age can impact the outcome, so not necessarily linearly.[18] Because child age is left out due to an already age-adjusted outcome, children compared can be of different ages, i.e. conceived earlier or later by their same-age and same birth-year mothers.

In **Table 1**, Panel A estimates β_1 in Equation (2), i.e. the impact of instrumented age at marriage on offspring height-for-age. It estimates that one year later marriage increased offspring height-for-age by about 0.20 standard deviations over the period from 1993 to 2021, albeit with a tapering of the effect toward the end of the period (which we explore in Section 6). At the 95% confidence interval, we cannot exclude effects as large as 0.29 standard deviations of early marriage on height-for-age over the period.

Panel B estimates the first stage relationship, i.e. α_1 in Equation (3), showing that one year later menarche postponed marriage by about 0.12 years over the period. Examining the trends over time, the diminishing magnitudes from 1993 to 2021 show that age at menarche has played a smaller role in marital decisions in the past decade, but the instrument remains relevant nonetheless, with a significant positive coefficient and a high F-stat even in the fifth survey. Coincidentally, the lowest survey-specific F-stat, 104.7, is exactly the threshold that Lee et al (2022) find is needed (as opposed to $F > 10$) to enable valid t-ratio inference at the 95% confidence interval. In other words, our instrument is strong even by the more recent and stringent criteria of instrumental relevance.

4 Mechanisms

To explore the mechanisms behind these results, we first look into how early marriage may affect characteristics of the mother relevant to child health - particularly age at motherhood and education attained - and ask whether the impact on offspring is mediated by marriage-induced changes to these variables. The remaining discussion then turns to direct potential mechanisms, by examining the impact of early marriage on growth-relevant feeding and nutrient intake, which mothers largely oversee.

[18] As we show in Section 6, entering mother's age as a continuous variable yields nearly identical results, as does controlling for mother age *or* mother birth year alone.

4.1 Mother age and education channels

Early marriage can precipitate the time at which a woman becomes a mother for the first time and it can also result in limiting her educational intake, and there is evidence of both impacts on Indian women. As demonstrated in **Table A1** in Appendix A, one year earlier marriage - instrumented by early menarche - precipitates the age at which a woman conceives her first child by 0.7 years and, when comparing among same birth-order children, it precipitates conception time by a similar magnitude.[19] With regard to education, early marriage reduces the woman's total schooling attainment by a whole year, and it also results in marrying less educated husbands, likely through spousal matching on education characteristics.

To check whether motherhood age and/or formal schooling are mediating the results, we run the two-stage least square specification in Equations (2)-(3) with additional controls in vector $\mathbf{X}$ - separately - for conception and for education.

In columns (2) and (3) in **Table 2**, we control for age at first conception and for age at conception of the specific child, respectively. In column (4) we control for the woman's education years, while in column (5) we restrict the analysis to women who never completed primary schooling. Because the majority of the latter women were already out of school *before* being married, there is unlikely to be an educational channel for this subsample, so that the impact on child stunting would be much weaker *if* educational channels were largely responsible.

The results in **Table 2** suggest a main mediating role for age at motherhood, but *not* for mother education. The impact of instrumented early marriage on child stunting largely dissipates and becomes insignificant when we control for age at conception, but it remains strong and robust when we control for mother's education or when we restrict the analysis to mothers without a primary school education.

This begets the question of what it is that early married (and as a result younger) mothers *do* which precipitates higher child stunting rates, which we turn to next.

[19]This suggests there is little impact of early marriage on the *spacing* of conception between children.

Table 2: Conception age and education as mediating channels

	Outcome: child height-for-age				
	No controls	Conception controls		Education controls	
	(1)	(2)	(3)	(4)	(5)
Age at marriage	**0.198*****	0.151	0.042	0.164***	0.191***
	(0.046)	(0.170)	(0.101)	(0.048)	(0.063)
Age at first conception		0.104			
		(0.154)			
Age at conception			0.241		
			(0.101)**		
Mother schooling years				0.030***	
				(0.004)	
Observations	132,316	111,640	132,280	132,283	36,353
Only no-primary-education obs.	No	No	No	No	Yes
Mother, child and HH controls	Yes	Yes	Yes	Yes	Yes

Note: *p<0.1; **p<0.05; ***p<0.01

Table 2 reports the second-stage results for the 2SLS in Eq.(2)-(3) with additional controls in Columns (2)-(4). Column (2) controls for age at first conception of the mother, Column (3) for age of conception of the specific child, and Column (4) for mother schooling years. In column (5) the 2SLS is run with no additional controls, but only using a subsample of the mothers who did not complete primary schooling, i.e. five years of education. Only the second stage results are reported.

4.2 Nutritional mechanisms

As outlined in the Introduction, the nutritional mechanisms we investigate are complementary feeding after at most six months, breastfeeding duration, protein intake, and nutrient loss from disease. Since the survey waves do not all elicit all the relevant variables, we discuss mechanisms separately by survey, and compare survey-specific estimates on mechanisms with survey-specific main results (columns 2-4 in **Table 1**). **Table A2** in the Appendix first provides a descriptive overview of these nutritional variables, including the survey and number of observations on which we have data, as well as the mean in the data. A number of points stand out.

First, complementary feeding, on which we have information from the first wave, takes place too late for young infants in India. Only one in three infants between the ages of 6 to 8 months was introduced to mushy solids, and the average age of initiating complementary feeding was about 9.4 months, well above the World Health Organization's recommendation of 6 months. Second, about nine out of ten infants is breastfed for at least 4 to 6 months, and this is largely consistent over the period. Third, with the exception of non-breast-milk, most children below the age of five did not consume a source of protein within the last day and night including from dairy, legumes, eggs, or flesh. Fourth, the prevalence of recent diarrhea is also consistent, standing at about one in every ten surveyed children over the period.

To examine whether the mother's age at marriage exerts an influence on growth-relevant early child nutrition, we regress the latter on age at marriage instrumented by age at menarche. We do this by replacing child height-for-age in the second stage, i.e. Equation (2), with the nutritional variable for the child. While the test for complementary feeding start and breastfeeding duration follow the same specification on the RHS, we note that for prior day nutrition and on diarrhea, it is necessary to compare children of the same age. For example, it is not realistic to expect the intake of legumes or other food to be the same for a one year old as for a two year old. It is not possible to control for both child and mother age without omitting variation in age at conception. Therefore, in these regressions we omit mother age and birth year as controls and use child age as a control, comparing the results of same-age and same birth-order children born to mothers who married earlier versus later, and who may therefore be on average younger versus older mothers.

The identification assumption is similar to before: that age at menarche of the mother influences the child's nutrient intake through - and *only through* - influencing the mother's age at marriage. The results are presented in **Table 3** and illustrated in **Figure 6**.[20]

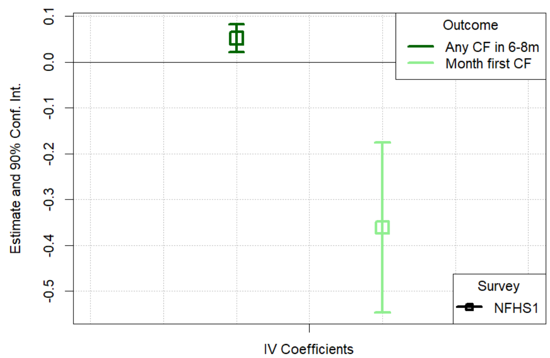
[20] Although only the second stage results are displayed in Table 3, all regressions demonstrate a strong first stage, with F-stats > 104.7.

Table 3: Impact of marriage age on nutrient intake and loss

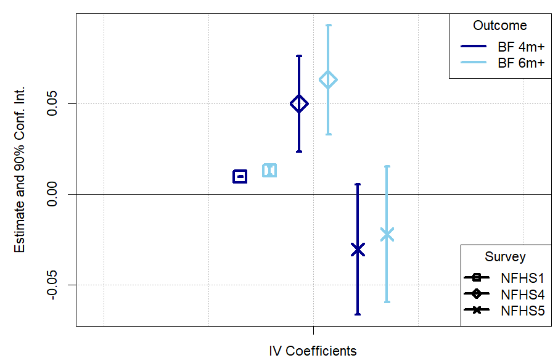
	NFHS-1	NFHS-4	NFHS-5
	(1)	(2)	(3)
Outcome	*PANEL A. Complementary feeding initiation*		
Any CF for 6-8 month olds (y/n)	0.052**		
	(0.019)		
Age CF introduced (months)	−0.36***		
	(0.112)		
	PANEL B. Breastfeeding duration		
At least 4 months, for ≥4m olds (y/n)	0.010***	0.050**	−0.030
	(0.001)	(0.016)	(0.022)
At least 6 months, for ≥6m olds (y/n)	0.013***	0.063***	−0.022
	(0.001)	(0.018)	(0.023)
	PANEL C. Prior day nutrition		
Any non-BF milk (y/n)	0.014	0.010	0.011
	(0.011)	(0.015)	(0.029)
Any legumes (y/n)		0.006	0.000
		(0.010)	(0.019)
Any dairy (y/n)		0.004	0.023
		(0.010)	(0.019)
Any eggs (y/n)		−0.017**	−0.019
		(0.008)	(0.018)
Any flesh (y/n)		−0.02*	−0.011
		(0.011)	(0.019)
	PANEL D. Diarrhea prevalence		
Diarrhea in last two weeks (y/n)	−0.003	0.000	0.000
	(0.007)	(0.013)	(0.013)

Note: *p<0.1; **p<0.05; ***p<0.01

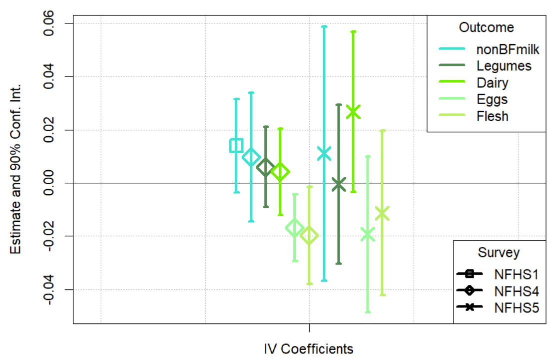
Table 3 reports the second-stage results of the 2SLS in Eq.(2)-(3) where the outcome is the nutritional variable for the child. The results are demonstrated by nutritional category and by survey.



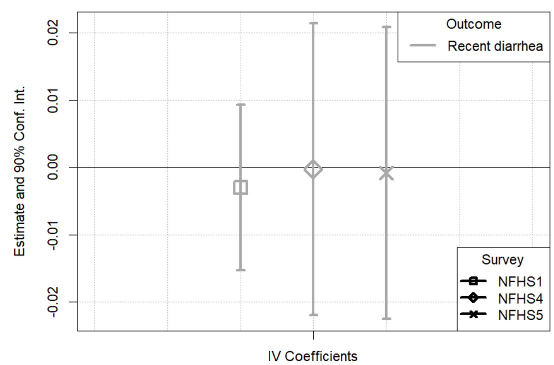
(a) Complementary feeding



(b) Breastfeeding



(c) Other proteins in last 24hr



(d) Recent diarrhea

Figure 6: Impact of marriage age on nutrition intake and loss

Figure 6 illustrates the results in Table 3, showing the point estimates for the impact of instrumented age at marriage on child nutrition intake and loss, and the 90% confidence intervals.

We find that early marriage postpones complementary feeding and reduces duration of breastfeeding of young infants, and that these results emerge *precisely* from the survey waves in which there is a detectable impact of early marriage on child stunting. One year earlier marriage reduced the probability that an infant aged between 6 and 8 months had been introduced to mushy foods by 5.2 percentage points in the NFHS-1, equivalent to a 14.5% decline, and it delayed the infant's introduction in general to foods by slightly over 10 days. In fact, descriptive evidence shows that only for mothers married above the age of 24 was the mean of complementary feeding introduction as low as the recommended six months old (omitted). As shown in **Table 3**, one year earlier marriage also reduced the probability that an infant was breastfed at least six months in both the NFHS-1 and NFHS-4, by up to 6.3 percentage points (equivalent to 7.1%) in the latter.

In contrast, we find no statistically significant impact of early marriage on any of the nutritional variables in the NFHS-5, in which the marriage-stunting relationship is also weaker. We also find little evidence in general for an impact of early marriage on prior day other-protein intake or on diarrhea incidence.[21] We do however interpret the latter results with some caution as we are unable to control for mother age while maintaining meaningful variation, whereas our preferred specifications control for mother age and use outcomes which are already adjusted for child age (such as height-for-age or infants of a narrow age range) or which can be independent of child age (such as minimum breastfeeding for all infants above a specific age threshold).

5 Underlying behavior

To the extent that early-married mothers feed their children less well, deeper questions still emerge about these mechanisms and underlying mother behavior. Do early-married mothers simply know less and are less well-prepared, or are their priorities and preferences also key to the outcomes on children? On knowledge, we note that while Section 4.1 suggests formal schooling is not key to the marital age-stunting relationship, this does not exclude the potential for beneficial *informal* knowledge acquired with age or maturity. Of course, informal knowledge and preparation may also be acquired with prior childbearing experience. At the same time, early-married mothers may not be behaving mechanistically in response to (lack of) age or experience. If there is a role for effort and attention in improving child nutrition and outcomes, then preferences and priorities would mediate the impact on children.

While the NFHS does not elicit questions directly on the extent to which mothers know how or when to feed their infants, nor on how they may view or prioritize child nutrition, we offer a simple theoretical framework delineating some of these possible channels and showing how they can be inferred about from differentiating results by

[21] Although the coefficients are unexpectedly negative on eggs and flesh in the NFHS-4, the base intake of these nutrients is low, at less than 7% of total children surveyed, as shown in the Appendix.

child characteristics. We then follow the model with relevant evidence from our data.

5.1 Theoretical model

Consider child j of mother i, where j refers to the child's birth order. Denote the child's height-for-age as H_{ij}, and let it be a positive function of a composite variable of healthful nutrition in the first years of life, denoted as $h_{ij} \in (0, 1]$.

$$H_{ij} = f(h_{ij}) \qquad f' > 0 \tag{4}$$

In turn, healthful nutrition h_{ij} is in part a positive function of the age at which the mother married, m_i. We explain this with reference to age at marriage placing a lower bound on conception age,[22] in line with the results in Section 4.1, but the model is in essence agnostic about this, so it could also be through educational or other channels as well.

Specifically, we assume that for mothers married after some threshold age $\bar{m}$, then healthful nutrition is at its maximum of 1. By contrast, if the mother married at an earlier age, then h is adversely affected by the distance between marital age and the threshold age. Combining these points with Eq. (4), we obtain:

$$H_{ij} = \begin{cases} f(1) & \text{if} \quad m_i \geq \bar{m} \\ f(1 - \beta(\bar{m} - m_i)); \quad \beta > 0 & \text{if} \quad m_i < \bar{m}; \end{cases} \tag{5}$$

with the subsequent restriction $0 < \beta(\bar{m} - m_i) < 1$. We note that if the model was based on conception age instead of age at marriage, the key results would hold.[23]

We augment this basic model in Eq. (5) as follows. First, in tending to her child j, the mother can exert some effort e_{ij} in order to mitigate adverse impacts on her child's

[22] In India, less than one percent of all children are born out of wedlock.

[23] To see this, suppose conception age is $c_{ij} = m_i + aj$, where a is the birth interval between children, and $\bar{c}$ is the threshold conception age. Then the model would become

$$H_{ij} = \begin{cases} f(1) & \text{if} \quad m_i + aj \geq \bar{c} \\ f(1 - \beta(\bar{c} - (m_i + aj))); \quad \beta > 0 & \text{if} \quad m_i + aj < \bar{c}; \end{cases}$$

The partial derivative of height for age on age at marriage would still be $\frac{\partial H_{ij}}{\partial m_i} = f'\beta$, the same as obtained from the setup in Eq. (5). The subsequent results, from the augmented model, would similarly be left unaffected.

nutrition, for example by seeking information about, or support for, good nutrition practices that she may not readily have, along with paying close attention to her child's cognitive and physical development. Second, once the mother has raised one child, she has gained experience E which she can also use, *if* she applies said effort, to improve nutrition practices for the current child; in a sense, this is a type of effortful learning by doing, where it is necessary to not only observe but to actively apply the learned experience. But having had a prior child can also exert a negative influence R on the current child's nutrition, by constraining the resources (such as attention) of the mother available; this too can be mitigated with the effort parameter.

The augmented model takes the following form:[24]

$$H_{ij} = \begin{cases} f(1) & \text{if} \quad m_i \geq \bar{m} \\ f\Big[1 - \Big(\beta(\bar{m} - m_i)\Big)\Big(1 + (1+E)^{j-1} e_{ij} - (j-1)\frac{R}{e_{ij}}\Big)^{-1}\Big]; & \text{if} \quad m_i < \bar{m}; \end{cases} \tag{6}$$

where the possibility of experience benefits and of resource constraints for later children imply $E > 0$ and $R > 0$, respectively. Note that if $E = 0$ and $R = 0$, the equation becomes the same for all children regardless of birth order j.

For illustrative purposes, and also because they comprise the majority of observations in our data (next section), we focus on implications for firstborn ($j = 1$) and second born ($j = 2$) children. The above becomes:

$$H_{ij} = \begin{cases} 1 & \text{if} \quad m_i \geq \bar{m} \\ f\Big[1 - \Big(\beta(\bar{m} - m_i)\Big)\Big(1 + e_{i1}\Big)^{-1}\Big]; & \text{if} \quad m_i < \bar{m} \quad \text{and} \quad j = 1; \\ f\Big[1 - \Big(\beta(\bar{m} - m_i)\Big)\Big(1 + (1+E)e_{i2} - \frac{R}{e_{i2}}\Big)^{-1}\Big]; & \text{if} \quad m_i < \bar{m} \quad \text{and} \quad j = 2; \end{cases} \tag{7}$$

In other words, for a firstborn, the mother's effort can improve healthful nutrition, mitigating the impact of early marriage, but there is no mother experience or current resource constraint from earlier siblings. A secondborn can (but does not necessarily, pending effort) additionally benefit from their mother's childbearing experience but can also be hurt by resources, including attention, spreading thinner.

The mother operates according to an implicit utility function in which effort is costly and its exertion is influenced by the perception (by her or possibly by the

[24] We use linear forms to derive simple closed-form solutions.

household more widely) of the importance of investing in the child's health and well-being. Drawing on a large amount of evidence on son preference in India (Pande and Malhotra, 2006), we assume this is at least tangentially related to the child's gender. For simplicity and letting effort take one of two values $e_{ij} \in \{\underline{e}, \bar{e}\}$, effort takes the former value if child j is a girl and the latter if it is a boy. With roughly equal gender distribution in the population, the average effort will be the midpoint between the two, which we denote as $\tilde{e}$, and we let $\Delta_e = \bar{e} - \underline{e}$.

From this model, we derive the following propositions, with proofs in **App. B**:

Proposition 1. *For early married mothers, both healthful nutrition for the child and height-for-age for the child are positively impacted by mother's (later) age at marriage.*

Proof. See Appendix. □

Proposition 2. *The impact of age at marriage on boy children is lower than its impact on girl children if $\Delta_e > 0$.*

Proof. See Appendix. □

Proposition 3. *The impact of age at marriage on firstborns is higher than its impact on secondborns if $\frac{R}{E} < \tilde{e}^2$. Differentiating by gender, this holds for firstborn versus second born boys if $\frac{R}{E} < \bar{e}^2$, and for firstborn versus second born girls if $\frac{R}{E} < \underline{e}^2$.*

Proof. See Appendix. □

The intuition for the first proposition is straightforward, hinging on later age at marriage of the mother driving more healthful nutrition for the child ($\beta > 0$), and on the latter improving child growth ($f' > 0$). The magnitudes are equivalent to the absolute size of the *negative* effects of early marriage on nutrition and growth.

The second proposition is also straightforward. Higher effort toward a child buttresses potentially negative impacts of early marriage on their nutrition, in turn translating into a smaller positive coefficient (of later marriage) on child outcomes.

For the third proposition, the intuition is as follows. A secondborn child can potentially both benefit and be hurt by being a second child, because the mother is now more experienced but also spread more thin in terms of resources or attention. To the extent that the potential scope for experience from previous childbearing is greater than the potential scope for resource constraints arising from it, then R/E is small. But this relative magnitude is not all that matters, as mother effort actually operationalizes the effects of both E and R on child nutrition and growth. Higher effort reduces the adverse impact of R on the second child while and augmenting the beneficial impact of E.

Therefore, a second child will be more protected from the impact of early marriage relative to a first child if effort is high enough to offset R/E, and this in turn would translate into a smaller effect of later marriage on the second child.

5.2 Evidence

Our findings so far, of a positive impact of later age marriage on child stunting (Section 3) as well as nutritional mechanisms (Section 4), can be interpreted as in support of Proposition 1. In a mechanistic sense, the impact we find of marriage age on nutritional variables supports $\frac{\partial hij}{\partial m_i} = \beta > 0$, while the impact of marriage age on child height-for-age supports $\frac{\partial H_{ij}}{\partial m_i} = f'\beta > 0$. Together, these also support $f' > 0$.

To provide evidence relevant to Propositions 2 and 3 - which touch on whether preferences, experiences and constraints mediate the marital-age stunting relationship - it is necessary to further distinguish the results by child characteristics. To test for the presence of a gender-based preference (Proposition 2) we differentiate the main results by gender, looking at the impact of marital age separately on boy versus girl children. To test for the possibility of a role for experience and resource constraints alongside preferences (Proposition 3), we differentiate the results by birth order generally, and then for boys separately, and for girls separately. We note that 92% of the children in our sample are first or secondborn children, so that focusing on these cohorts simplifies calculations without making the results less relevant.

The results are displayed in **Table 4** and illustrated in **Figure 7**. First, as shown in Columns (2) and (3) of Panel A in the Table and illustrated in **Figure 7a**, girl children bear a higher brunt of the burden of early marriage of the mother, with the coefficient being nearly twice as high and more significant than for boys. Second, as shown in Columns (4) and (5) and illustrated in **Figure 7b**, firstborns bear a higher brunt of the burden of early marriage than secondborns.

Third, differentiating by *a combination* of birth order and gender as shown in Panel B and illustrated in **Figure 7c**, childbearing experience *only* improves outcomes for secondborns who are boys, with the impact of early marriage disappearing entirely. We can show (omitted) that this is the case both when the secondborn is the first son, i.e. firstborn was a daughter, and when he is the second son, although the impact of early marriage is even slightly more diminished for the former group, evidence of some additional first son preference. In contrast, previous childbearing experience does not improve outcomes for girl children, with girls who are secondborns being even worse off in terms of the impact of early marriage than girls who are firstborn.

Interpreting these findings in light of the model, what we see suggests that

$$\Delta_e > 0 \tag{8}$$

$$E > 0 \tag{9}$$

$$R > 0 \tag{10}$$

$$\underline{e}^2 < \frac{R}{E} < \tilde{e}^2 < \bar{e}^2 \tag{11}$$

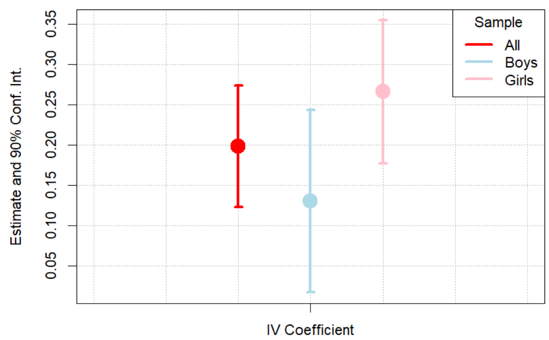
In sum, while early age at marriage for the mother contributes to poor nutrition of the infant and to subsequent stunting, it does not do so in a mechanistic manner. Rather, the mother's circumstances as relate to the experience and/or burden from raising other children, combined with her priorities or preferences, can act to buttress or amplify the impacts on her child.

Table 4: Differentiating by child birth order and gender

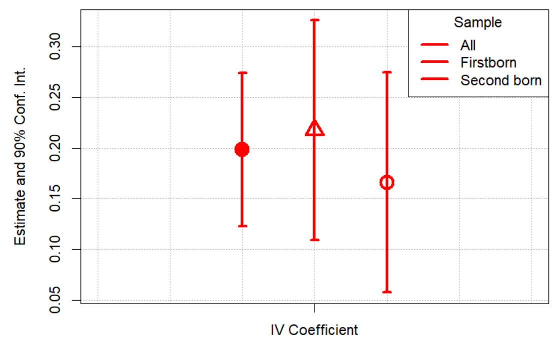
	(1)	(2)	(3)	(4)	(5)
	Panel A. By gender and birth order separately				
	All	Boys	Girls	Firstborn	Secondborn
Age at marriage	**0.198*****	0.130*	0.266***	0.218***	0.166**
	(0.046)	(0.069)	(0.054)	(0.066)	(0.061)
Observations	132,316	67,853	64,463	83,069	38,434
Mother, child, HH controls	Yes	Yes	Yes	Yes	Yes
	Panel B. By combination of the two				
	All	Boys		Girls	
And is:		Firstborn	Secondborn	Firstborn	Secondborn
Age at marriage	**0.198*****	0.173*	0.008	0.254**	0.348***
	(0.046)	(0.102)	(0.080)	(0.079)	(0.092)
Observations	132,316	42,485	19,715	40,611	18,683
Mother, child, HH controls	Yes	Yes	Yes	Yes	Yes

Note: *p<0.1; **p<0.05; ***p<0.01

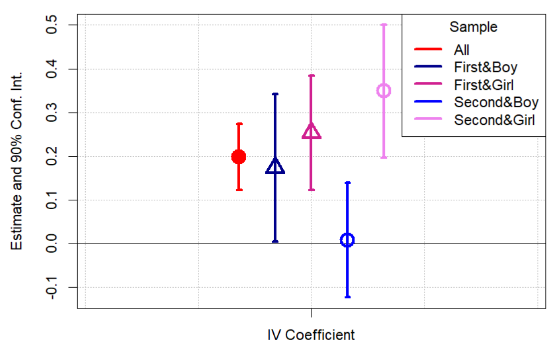
Table 4 reports the second-stage results of the 2SLS in Eq. (2)-(3), differentiated by gender and birth order of the child separately in Panel A, and differentiated by a combination of the two in Panel B. The outcome throughout is child height-for-age.



(a) By gender



(b) By birth order



(c) By combination of the two

Figure 7: Differentiating by child characteristics

Figure 7 illustrates the results in Table 4. It shows the point estimates for the impact of instrumented age at marriage on child height-for-age - differentiated by child characteristics - and the 90% confidence intervals.

6 Additional results

The following results are presented in **Appendix A**.

As mentioned above, mother's height may serve as a proxy for her own nutrition in childhood. Along with descriptive evidence in Section 2 that mother's height seems to be distributed independently and almost identically by age at menarche, in **Table A3** we additionally include mother's height as a control in the surveys with data on this variable, NFHS-4 and NFHS-5. The nature of the results remains similar - earlier age at marriage reduces height-for-age for the child in the fourth wave but not in the fifth wave - and the first stage estimates are nearly identical to before.

In **Table A4**, we examine more closely the choice of how to control for mother age. We re-run the main regression in Equations (2)-(3) with controls only for mother age in dummies (Column 1), only for mother birth year in dummies (Column 2), for a combination of the two with mother age as a continuous variable (Column 3), and for a combination of the two with mother age as a continuous variable *and* its square, to test for nonlinear effects (Column 4). The main results remain robust across all specifications.

In **Table A5**, we re-run the main regression but with age at conception age, instead of age at marriage, as the endogenous variable. This serves as an additional check on whether conception age is driving the results, in which case we would expect a significant first and second stage. In Column (1) we use age at first conception for the mother whereas in Column (2) we use age at conception of the specific child in the observation. In both cases, we find a strong first stage, with one year later menarche postponing conception by 0.07-0.08 years, as well as an impact of age at conception on child height-for-age of about 0.30 standard deviations.

To check the possibility that marital age impacts child stunting through *prenatal* mechanisms, in **Table A6** we test the impact of instrumented mother's age at marriage on preterm birth of the child, on birthweight, and on delivery complications in Columns (1)-(3) respectively.[25] To address the issue of offspring selection, i.e. the possibility that early married mothers have higher infant mortality and the remaining pool is healthier, we also test for impact of age at marriage on child mortality in Column (4). As shown, instrumented age at marriage does not predict any of these variables, suggesting no strong prenatal mechanisms nor offspring selection.

In addition to the robustness tests, we attempt to better understand the weakening of the age at marriage-stunting relationship observed in the fifth wave. The relationship is visibly attenuated both in the descriptive correlations (**Figure 2**) and in the analytical results (**Table 1**), and, for the analytical results, this occurs despite the instrument having a first-stage effect of similar magnitude to the fourth wave.

[25]A caveat is that information on the latter two is only available in some, not all, of the survey waves. Information on measured birthweight is available from NFHS-4 and NFHS-5 whereas information on delivery complications is available from NFHS-1.

Moreover, it can be shown that this is not due to a weakening of the link between marriage and conception age, as using conception age as the endogenous variable also yields survey-specific results in which we cannot reject a null of no effect (omitted).

Exploring in full the reasons underlying the divergence in the fifth wave lie beyond the scope of this paper, but we take a first step by examining whether a recent and potentially relevant national Indian policy scheme, titled *Poshan Abhiyaan*, has played a role. The scheme was rolled out beginning in 2018 with the aim of reducing malnutrition among children under five as well as pregnant and lactating mothers, and using tools such as "nutrition action plans", community events, and performance-based incentives for state functionaries (World Bank, 2018). Importantly, the scheme was *first* rolled out to districts that had high preexisting rates of child undernutrition in 2018 (Phase 1), accounting for nearly half of all districts in the country, followed later by a second and then third phase of districts, in descending order of malnutrition prevalence. Of these, only Phase 1 rollout happened fully in advance of the fifth wave.

If the scheme was largely responsible for the improvement we see, we would expect the weakening of the marital age-stunting relationship to be particularly pronounced in Phase 1 districts. In **Table A7**, we focus only on these districts, and compare results of the main 2SLS regression separately in the fourth and fifth waves, given little change in districting boundaries between the two. As shown, the impact of early marriage on child height-for-age is quite large and to very similar extents in Phase 1 districts in *both* survey waves, i.e. before and after rollout of the *Poshan Abhiyaan*. This is not directly a rejection of the potential effectiveness of the scheme later on, as it is possible that too little time had lapsed between rollout and demonstrable results. Nonetheless, it suggests that the recent weakening of the marriage age-child stunting relationship has come primarily from *other* parts of the country - concentrated in the South - which had lower rates of child stunting to begin with, and supports the notion that development in India is often regionally specific and divergent.

7 Conclusion

The majority of India's hundreds of millions of children are considered short for their age by global health standards, and one in every three children under five is short enough to be considered stunted. Stunting is strongly predicted by poverty and underdevelopment, and it is particularly predominant in the densely populated rural parts of the country, but more is needed to understand the ways in which socioeconomic conditions and practices translate into widespread child stunting.

In this paper we document the contribution to child stunting in India of a widespread socioeconomic practice which is driven by a mix of underdevelopment and social norms: early marriage of mothers. One in three of all women married as minors in the world reside in India, and the country's current median age at marriage, at nineteen, remains among the lowest worldwide. Given the key role of mothers in child nutrition, the paper essentially addresses the possibility that one of the pathways to child stunting lies in early marital age and in lack of appropriate nutritional practices for infants among early-married mothers.

Using age at menarche as an instrument for age at marriage and the totality of available data from India's National Family Health Surveys, we find that one year earlier marriage has decreased child height-for-age by 0.20 standard deviations over the past thirty years. Our results suggest the key is that marriage age places a lower bound on conception age, so that early married mothers become younger mothers, and that these younger mothers in turn introduce their children to growth-critical complementary feeding later and breastfeed for less time. We also find evidence that the impact of marital age on child stunting is mediated by gender preferences for boy versus girl children, as well as by the experience and burden from prior childraising.

Our findings shed light on a potentially devastating intergenerational effect of early marriage of women. Given slow evolution in marital age practices - itself an outcome of enduring underdevelopment and sticky social norms - successfully assisting young mothers in infant nutrition would help to stem lasting impacts on children's physical and possibly cognitive health. The policy challenges in doing so may be significant, but so may be the payoffs to human capital in today's most populous nation.

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APPENDIX

A Additional results

Table A1: Impact on parental variables

	Outcome: conception		*Outcome: education years*	
	Age at first conc.	Age at conc.	Mother	Father
	(1)	(2)	(3)	(4)
Age at marriage	0.696***	0.593***	1.004**	0.872***
	(0.057)	(0.056)	(0.131)	(0.132)
Observations	132,843	158,949	159,071	40,579
Mother and HH controls	Yes	Yes	Yes	Yes
Child birth-order control	No	Yes	No	No

Note: *p<0.1; **p<0.05; ***p<0.01

Table A2: Mean of nutritional variables

	NFHS-1	NFHS-4	NFHS-5
	(1)	(2)	(3)
Variable	*PANEL A. Complementary feeding initiation*		
Any CF for 6-8 month olds (y/n)	0.362		
	(1,656)		
Age CF introduced (months)	9.350		
	(11,862)		
	PANEL B. Breastfeeding duration		
At least 4 months, for ≥4m olds (y/n)	0.912	0.905	0.891
	(15,764)	(50,692)	(43,743)
At least 6 months, for ≥6m olds (y/n)	0.896	0.892	0.881
	(14,565)	(46,593)	(40,260)
	PANEL C. Prior day nutrition		
Any non-BF milk (y/n)	0.354	0.234	0.229
	(13,206)	(57,499)	(50,005)
Any legumes (y/n)		0.070	0.087
		(57,499)	(50,005)
Any dairy (y/n)		0.079	0.088
		(57,499)	(50,005)
Any eggs (y/n)		0.069	0.084
		(57,499)	(50,005)
Any flesh (y/n)		0.053	0.060
		(57,499)	(50,005)
	PANEL D. Diarrhea prevalence		
Diarrhea in last two weeks (y/n)	0.114	0.105	0.104
	(16,635)	(74,378)	(63,982)

Note: *p<0.1; **p<0.05; ***p<0.01

Table A3: Controlling for mother height

	NFHS-4	NFHS-5
	(1)	(2)
	PANEL A. Second stage. Outcome: child height-for-age	
Age at marriage	0.205***	0.064
	(0.035)	(0.118)
	PANEL B. First stage. Outcome: Age at marriage	
Age at menarche	0.110***	0.084***
	(0.019)	(0.014)
First stage F-stat	150.9	105.3
Observations	65,220	57,176
Mother, child, and HH controls	Yes	Yes

Note: *p<0.1; **p<0.05; ***p<0.01

Table A4: Variations of control for mother age

	Outcome: child height-for-age			
	(1)	(2)	(3)	(4)
Age at marriage	0.163*	0.189***	0.197**	0.199***
	(0.076)	(0.035)	(0.048)	(0.047)
Observations	132,316	132,337	132,337	132,337
Mother, child and HH controls	Yes	Yes	Yes	Yes

Note: *p<0.1; **p<0.05; ***p<0.01

Table A5: Age at conception as endogenous variable

	(1)	(2)
	PANEL A. Second stage. Outcome: child height-for-age	
Age at first conception	0.317***	
	(0.091)	
Age at conception		0.307***
		(0.077)
	PANEL B. First stage. Outcome: First/specific conception	
Age at menarche	0.068***	0.079***
	(0.006)	(0.006)
First stage F-stat	372.2	466.0
Observations	112,958	134,029
Mother, child, and HH controls	Yes	Yes

Note: *p<0.1; **p<0.05; ***p<0.01

Table A6: Impact on birth variables and child mortality

	Preterm (y/n)	Birthweight (kg)	Comp. (y/n)	Alive (y/n)
	(1)	(2)	(3)	(4)
Age at marriage	0.006	−0.025	0.018	−0.001
	(0.012)	(0.027)	(0.012)	(0.003)
Observations	158,949	119,841	17,871	159,129
Survey waves	All	4^{th}, 5^{th}	1^{st}	All
Mother and HH controls	Yes	Yes	Yes	Yes

Note: *p<0.1; **p<0.05; ***p<0.01

Table A7: Poshan Abhiyaan Phase I districts

	NFHS-4	NFHS-5
	(1)	(2)
	PANEL A. Second stage. Outcome: child height-for-age	
Age at marriage	0.487***	0.501***
	(0.138)	(0.183)
	PANEL B. First stage. Outcome: Age at marriage	
Age at menarche	0.103***	0.085***
	(0.021)	(0.020)
First stage F-stat	69.5	50.4
Observations	39,013	30,302
Mother, child, and HH controls	Yes	Yes

Note: *p<0.1; **p<0.05; ***p<0.01

B Proofs of propositions

B.1 Proposition 1

The proof of Proposition 1 is obtained by calculating the partial derivatives of H_{ij} and h_{ij} with respect to m_i.

Beginning with h_{ij}, then for $m_i \geq \bar{m}$, $\frac{\partial h_{ij}}{\partial m_i} = \frac{\partial 1}{\partial m_i} = 0$. In contrast if $m_i < \bar{m}$ then:

$$\frac{\partial h_{ij}}{\partial m_i} = \frac{\beta}{1 + (1+E)^{j-1} e_{ij} - (j-1)\frac{R}{e_{ij}}} > 0$$

given the restriction that the denominator is positive.

For H_{ij},

$$\frac{\partial H_{ij}}{\partial m_i} = \frac{\partial H_{ij}}{\partial h_{ij}} \frac{\partial h_{ij}}{\partial m_i}$$

Therefore, if $m_i \geq \bar{m}$, then $\frac{\partial H_{ij}}{\partial m_i} = f' * 0 = 0$. In contrast, if $m_i < \bar{m}$, then

$$\frac{\partial H_{ij}}{\partial m_i} = f' \frac{\beta}{1 + (1+E)^{j-1} e_{ij} - (j-1)\frac{R}{e_{ij}}} > 0 \tag{B.1}$$

since both $f' > 0$ and $\beta > 0$. This completes the proof for Proposition 1. $\square$

B.2 Proposition 2

The proof of Proposition 2 is obtained by calculating and comparing the partial derivative $\frac{\partial H_{ij}}{\partial m_i}$ for boy versus girl children. For mothers married $m_i \geq \bar{m}$, and following Proposition 1, $\frac{\partial H_{ij}}{\partial m_i} = 0$ for all children.

For mothers married $m_i < \bar{m}$, then for boy children

$$\frac{\partial H_{ij}}{\partial m_i}|_{boys} = f' \frac{\beta}{1 + (1+E)^{j-1} \bar{e}_{ij} - (j-1)\frac{R}{\bar{e}_{ij}}}$$

while for girl children

$$\frac{\partial H_{ij}}{\partial m_i}|_{girls} = f' \frac{\beta}{1 + (1+E)^{j-1} \underline{e}_{ij} - (j-1)\frac{R}{\underline{e}_{ij}}}$$

To see that the first part of the proposition holds, the former minus the latter yields:

$$\begin{aligned}
&\frac{\partial H_{ij}}{\partial m_i}|_{boys} - \frac{\partial H_{ij}}{\partial m_i}|_{girls} \\
&= f'\frac{\beta}{1+(1+E)^{j-1}\bar{e}_{ij}-(j-1)\frac{R}{\bar{e}_{ij}}} - f'\frac{\beta}{1+(1+E)^{j-1}\underline{e}_{ij}-(j-1)\frac{R}{\underline{e}_{ij}}} \\
&= \frac{f'\beta\big(1+(1+E)^{j-1}\underline{e}_{ij}-(j-1)\frac{R}{\underline{e}_{ij}}\big) - f'\beta\big(1+(1+E)^{j-1}\bar{e}_{ij}-(j-1)\frac{R}{\bar{e}_{ij}}\big)}{\big(1+(1+E)^{j-1}\bar{e}_{ij}-(j-1)\frac{R}{\bar{e}_{ij}}\big)\big(1+(1+E)^{j-1}\underline{e}_{ij}-(j-1)\frac{R}{\underline{e}_{ij}}\big)} \\
&= f'\beta\frac{(1+E)^{j-1}(\underline{e}-\bar{e})-(j-1)R(\frac{\bar{e}-\underline{e}}{\bar{e}\underline{e}})}{\big(1+(1+E)^{j-1}\bar{e}_{ij}-(j-1)\frac{R}{\bar{e}_{ij}}\big)\big(1+(1+E)^{j-1}\underline{e}_{ij}-(j-1)\frac{R}{\underline{e}_{ij}}\big)} \\
&= f'\beta\big(\underline{e}-\bar{e}\big)\frac{(1+E)^{j-1}+\frac{(j-1)R}{\bar{e}\underline{e}}}{\big(1+(1+E)^{j-1}\bar{e}_{ij}-(j-1)\frac{R}{\bar{e}_{ij}}\big)\big(1+(1+E)^{j-1}\underline{e}_{ij}-(j-1)\frac{R}{\underline{e}_{ij}}\big)} \\
&< 0 \quad \text{iff} \quad \Delta_e = \bar{e}-\underline{e} > 0
\end{aligned}$$

This completes the proof for Proposition 2. □

B.3 Proposition 3

The proof of Proposition 3 is obtained by calculating and comparing the derivative $\frac{\partial H_{ij}}{\partial m_i}$ for firstborns and secondborns. Once more, for mothers married $m_i \geq \bar{m}$ and following Proposition 1, $\frac{\partial H_{ij}}{\partial m_i} = 0$.

For mothers married $m_i < \bar{m}$, then for firstborns:

$$\frac{\partial H_{i1}}{\partial m_i} = f'\beta\frac{1}{1+e_{i1}}$$

For secondborn children,

$$\frac{\partial H_{ij}}{\partial m_i} = f'\beta\frac{1}{1+(1+E)e_{i2}-\frac{R}{e_{ij}}}$$

Since approximately half of all children regardless of birth order are boys and the other half is girls, average effort in both birth orders is $\tilde{e}$. Subtracting the derivative for firstborns minus secondborns yields:

$$\begin{aligned}
&\frac{\partial H_{i1}}{\partial m_i} - \frac{\partial H_{i2}}{\partial m_i} \\
&= f'\beta \frac{1}{\big(1+\tilde{e}\big)} - f'\beta \frac{1}{\big(1+(1+E)\tilde{e} - \frac{R}{\tilde{e}}\big)} \\
&= f'\beta \frac{\big(1+(1+E)\tilde{e} - \frac{R}{\tilde{e}}\big) - \big(1+\tilde{e}\big)}{\big(1+(1+E)\tilde{e} - \frac{R}{\tilde{e}}\big)\big(1+\tilde{e}\big)} \\
&= f'\beta \frac{\tilde{e}\big[E - \frac{R}{\tilde{e}^2}\big]}{\big(1+(1+E)\tilde{e} - \frac{R}{\tilde{e}}\big)\big(1+\tilde{e}\big)}
\end{aligned}$$

Given $f' > 0$ and $\beta > 0$, this difference will be positive if the expression in the square brackets is positive, i.e.

$$\frac{R}{E} < \tilde{e}^2$$

Otherwise, we would have $\frac{\partial H_{i1}}{\partial m_i} < \frac{\partial H_{i2}}{\partial m_i}$. Repeating the same arithmetic but for boys only, where $e = \bar{e}$, yields

$$\frac{\partial H_{i1}}{\partial m_i}|_{boys} - \frac{\partial H_{i2}}{\partial m_i}|_{boys} > 0 \quad \text{if} \quad \frac{R}{E} < \bar{e}^2$$

Similarly for girls only, where $e = \underline{e}$,

$$\frac{\partial H_{i1}}{\partial m_i}|_{girls} - \frac{\partial H_{i2}}{\partial m_i}|_{girls} > 0 \quad \text{if} \quad \frac{R}{E} < \underline{e}^2.$$

This completes the proof for Proposition 3. $\square$