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Measuring maternal autonomy and its effect on child nutrition in rural India

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Abstract

This paper examines the link between a mother's autonomy—the freedom and ability to think, express, make decisions and act independently—and the nutritional status of her children. We treat 'autonomy' as a latent variable, and design a novel statistical framework to measure this. This method allows us to separate the direct associations of maternal and family characteristics in our model for nutrition, from their indirect associations that work through maternal autonomy. Using data from India, we explore the sensitivity of our estimates to endogeneity caused by sample selection in the presence of son preference. We find: (i) a one standard deviation (SD) higher autonomy score is associated with a 0.16 SD higher height-for-age z-score (HAZ score); and (ii) a 10% lower prevalence of stunting (HAZ < -2 SD). The latter is equivalent to the prevention of approximately 300,000 children stunting, indicating the important role of maternal autonomy.

1 | INTRODUCTION

In addition to playing a pivotal role in increasing child mortality, poor nutrition in childhood causes irreversible damage to cognitive development and future health (Dreze 2004; Sumner *et al.* 2009; Saxena 2018; Victora *et al.* 2008). Child undernutrition is also strongly associated with shorter adult height, less schooling, reduced economic productivity, lower adult body mass index, and mental illness (see Victora *et al.* (2008) for a systematic review). Thus the importance of good nutrition during childhood cannot be emphasized enough. Recognizing this, the Government of India introduced the Integrated Child Development Services Scheme as far back as 1975 in selected blocks (administrative units) in the country. The scheme was eventually expanded to all administrative blocks in the country. In 2014–15, the Government of India had budgeted nearly Rs 181.95 billion to run the scheme through its 1.346 million Anganwadi (Mother and Childcare) Centres across all villages and towns of India.[1] The budget allocated for the scheme

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increased to around Rs 205.5 billion for the year 2023–24 (£1.94 billion).[2] These centres provide a vast range of health and nutrition services to children, adolescents, and pregnant and lactating women. Yet despite the massive spread of the policy and the huge investments in it, child undernutrition has been stubbornly high in India. India contributes to a third of the global burden of childhood stunting (a measure of chronic undernutrition).

Given the importance of proper child nutrition and the persistence of high levels of undernutrition in India, various studies in recent decades have attempted to understand the proximate and underlying factors relevant to child nutrition. These studies have helped to identify various factors that are crucial for child nutrition (exclusive breastfeeding in the first six months, maternal health, and so on) and have also contributed to policy-making. The focus of this paper, however, is on one factor whose relationship with nutrition is less understood in the literature—the role of maternal autonomy. For this, we need to overcome two important econometric challenges, which we discuss next.

The first challenge is how to define and measure autonomy. The term autonomy has often been confused with empowerment, though the latter is a process and the former the outcome (at least partly) of the process. We focus specifically on the outcome. Many different definitions and measurements exist in the literature. One strand of the literature assumes that 'autonomy' is a directly observed trait and measures it using an arithmetic average of binary answers to a set of questions that are elicited by surveys (Jensen and Oster 2009; Paul and Saha 2022),[3] or the answers to these questions directly in the equation (Dancer and Rammohan 2009; Imai *et al.* 2014).[4] For instance, Paul and Saha (2022) define 'autonomy' to be low/medium/high based on the values of the sum of dichotomous variables defined over nine different answers. Another strand of the literature uses definable and easily measurable variables such as education and health (e.g. Imai *et al.* 2014) as proxies for autonomy. In the same vein, Chilinda *et al.* (2021) use a composite score to capture decision-making power, tolerance of domestic violence, and financial independence as a proxy for maternal autonomy. Yet another strand of the literature, which is more related to what we do, assumes that the autonomy trait is essentially not observable, and the answers to a set of questions (to be listed later) give you some proxy mismeasured information about autonomy, and uses principal component analysis to construct/extract a measure of 'autonomy' (Chakraborty and De 2011).[5]

In this paper, we build on the last approach (from the ones mentioned above), which assumes that autonomy is an underlying latent trait that cannot be observed directly (like traits such as confidence or empathy) but expresses itself in how one thinks and acts. We create an index of 'autonomy' based on the following. (i) Autonomy *expresses* itself in several ways, such as having decision-making power, mobility, and command and control over resources. (ii) Socioeconomic factors play a role in enabling women to exercise autonomy. These factors can be captured through indicators such as the level of education and occupation of the woman and her husband, the woman's caste, religion, age, exposure to media, region, and so on. (iii) Autonomy is an unobserved trait that is fixed in the short term, and the answers given to the set of questions are *fallible measures* of autonomy. We use Bayesian shrinkage methods (Goldstein 2003) within a latent factor model to create an index of autonomy that controls for relevant socioeconomic factors. Our approach, which is new to this literature, allows us to separate the direct associations of maternal and family characteristics in our model for nutrition from their indirect associations that work through maternal autonomy.

The second major challenge is how to deal with possible biases due to the endogenous sex composition of children in the sample and the observed birth intervals (Yamaguchi 1989). The prevalence of 'son preference' in India can manifest itself in differences in nutritional status across children with similar backgrounds (Barcellos *et al.* 2014; Jayachandran and Kuziemko 2011). If a family sex-selects the second and subsequent children by using prenatal sex selection,[6] even the nutritional status of the firstborn will be affected by the presence of subsequent children in this family.[7] To mitigate this bias, our preferred model restricts our analysis to firstborns, and we check for sensitivity of our results to the different age compositions of these children used in the analysis.[8]

Our sample comes from the third round of the National Family Health Survey (NFHS-3, 2005–6; International Institute for Population Sciences (IIPS) and Macro International 2007a). The survey provides information on the three commonly used anthropometric indicators of the nutritional status of children: height for age z-score[9] (HAZ score), weight for height z-score (WHZ score) and weight for age z-score (WAZ score).[10] The survey indicates that 48% of children under 5 years of age in India are chronically malnourished (i.e. stunted), 43% are underweight, and 20% are acutely malnourished (i.e. wasted) (IIPS and Macro International 2007a; 2007b). These figures are extremely high since statistically, one would expect about 2–3% of the population of children aged under five to fall in the range below −2 standard deviations (SD).[11]

Our main results indicate that our autonomy index is a positive and significant predictor of better nutrition among rural children.[12] Among the sample of firstborn children aged under 18 months, the estimated association is positive and significant in the long-term child nutrition (i.e. HAZ score) equation, and negative on the probability of the child being stunted.[13] Based on these magnitudes, a back-of-an-envelope calculation indicates that a one SD higher autonomy score is associated with a 10% lower (0.3 million) number of stunted children among the firstborns (2.1 million) (30.5–27.3%).[14] These numbers indicate the important role of maternal autonomy in child nutrition. We do not find any differential effects of autonomy by sex of the child.

The paper is organized as follows. Section 2 discusses the relevant literature for India. Discussions on the data and sample selected are in Section 3. Section 4 sets out the methodology, while Section 5 provides some sample descriptive statistics. Section 6 presents the main results, and addresses the issue of sample selection in the presence of son preference. Finally, Section 7 presents a summary and conclusions.

2 | OVERVIEW OF THE LITERATURE

One of the earliest studies by Dyson and Moore (1983) on kinship structures and women's autonomy defines autonomy as the capacity to obtain information and make decisions about one's private concerns and those of one's intimates. In a similar vein, Safilios-Rothschild (1982), in the context of demographic change in the third world, defines autonomy as 'the ability to influence and control one's personal environment'. The essential elements of autonomy—namely, the ability and capacity to make decisions in a way that can influence one's environment—are reflected in other definitions, such as that by Jejeebhoy (2000), according to whom, autonomy is the 'extent to which women exert control over their own lives within the families in which they live at a given point of time'. As stated by Agarwala and Lynch (2006): 'These definitions assert a single construct that captures the multifaceted ability to gain control over the circumstances of one's life.'[15]

The nutritional status of a child is strongly related to the characteristics of the mother, as many studies have shown. Mother's education is associated with child survival (Murthi *et al.* 1995; Cleland 2010), and the nutritional status of a child (Borooah 2004; Frost *et al.* 2005). Mother's health is also reflected in health outcomes for children. At birth, one-third of Indian infants are underweight, and 20% are stunted because of poor intrauterine growth (Mamidi *et al.* 2011; Ramachandran and Gopalan 2011). Using the Indian NFHS-4, Paul and Saha (2022) find autonomy to be significantly associated with lower odds of a child being malnourished. Rahman *et al.* (2015) use the Bangladesh 2011 Demographic and Health Survey to find that children of mothers who are engaged in decision-making in the household are less likely to be undernourished. Carlson *et al.* (2015) review some of the current literature on the relationship between maternal autonomy, children's nutritional status, and child-feeding practices. The authors conclude that while enhancing maternal autonomy is important for improving children's nutritional status, gaps exist in the current knowledge that are confounded by complexities of how autonomy is measured.

In summary, autonomy has intrinsic relevance for a woman's own wellbeing and also contributes to enhancing the quality of life for the family. It determines largely her ability to make effective choices and to exercise control over her life. While most studies have looked at readily definable and easily measurable variables to proxy for autonomy, this paper treats 'autonomy' as a latent trait that is fixed in the short run, and only fallible measures of this trait are available to researchers.

3 | DATA AND THE SAMPLE

The data are from the third round of the National Family Health Survey (NFHS-3) for India, 2005, which is part of the Demographic and Health Survey (DHS). The DHS collects extensive information on population, health and nutrition, with an emphasis on women and young children. In addition, the data contain information concerning household decision-making, as well as answers to some questions relating to the 'autonomy' status of surveyed women.[16]

The NFHS-3 interviewed over 230,000 women (aged 15–49) and men (aged 15–54) from 29 Indian states, during the period December 2005 to August 2006 (IIPS and Macro International 2007a). All children aged under 60 months, and living in the household at the time of the survey, were weighed, and their heights were measured.

We select our sample based on the following criteria: (i) currently married women who are 'usual' residents and living in an area classified as rural; (ii) mothers who had at least one surviving child born in the past 60 months; (iii) no missing values for the main variables of interest. We keep our autonomy index fixed over our different sample cuts in the estimations of the nutrition equation. All our estimations account for the clustering of the error variance at the primary sample units.

4 | EMPIRICAL MODELS

We begin this section with a discussion of how we model the presence of 'autonomy' in the nutrition equation, and then discuss how we estimate the autonomy index, which is treated as a latent unobserved characteristic.

4.1 | Nutritional status

We focus on two outcome variables, height-for-age z-score (HAZ score) and 'stunted', which indicate long-term nutritional status. 'Stunted' is a binary indicator for HAZ less than −2 according to the World Health Organization (WHO) definition. All children in the family who were aged under 60 months at the time of the interview, and who had valid measurements for these variables, form the main sample.

All equations are specified as a linear regression model and estimated by ordinary least squares (OLS). All reported standard errors are bootstrapped.[17]

The equation for the measure of nutritional status y (HAZ score and stunted) for child k of mother i is specified as

$$y_{ik} = x'_{ik}\beta + z'_i\gamma + \delta a_i + \varepsilon_{ik}, \quad i = 1, \dots, n, k = 1, \dots, K, \tag{1}$$

where a_i is the mother's autonomy trait, which we assume to be *unobserved*, x_{ik} contains child-specific characteristics (age in months and indicators for sex and whether the child is a twin), and z_i contains the variables associated with mother i in her family, and hence also contains her partner's, and her family characteristics. More specifically, the characteristics included are as follows: (i) mother-level variables age, education, caste, religion, occupation, employment status in the last 12 months, some exposure to common media; (ii) father-level variables education and occupation; and (iii) family-level variables whether it is a nuclear family, household wealth index factor score,[18] and the state of residence.

The challenge here is to obtain a consistent estimator of equation (1) coefficients, where the parameter of interest is δ. Covariates that are correlated with the equation error term, for whatever reasons, will be 'endogenous' in the estimation equation. The question is: what is the reason for the correlation? One of the reasons for this correlation could be the existence of 'son preference', which will make the observed sex composition and birth intervals endogenous, hence the sample used in the estimation becomes crucial. We attempt to mitigate this by selecting the sample of firstborns who are aged under 18 months as our preferred specification. See Section 6 for further discussions on the issue of how 'son preference' may affect the observed sex composition of children and the observed birth intervals in the sample, and a discussion of sensitivity to different choices of the sample. Generally, the lower the chances of including a firstborn who has another sibling in the family, the lower the chances that the OLS estimator is biased and inconsistent.

We next discuss how we proceed with the measurement and the generation of the autonomy measure (a_i).

4.2 | Defining and measuring maternal autonomy

Based on the literature, we choose the answers given to the questions below in the NFHS as indicative of a woman's 'autonomy' to think, speak, decide and act independently.[19]

The following responses, categorized in terms of the three dimensions that we consider, are all coded as binary indicators.

- Related to physical autonomy: the woman is allowed to go *alone* to (i) the market (measure m1), (ii) the health clinic (m2), and (iii) places outside the community (m3).
- Related to decision-making autonomy: the woman decides *alone* on purchases for daily needs (m4); the woman decides *alone* or *jointly* with her husband on (i) her own healthcare (m5), (ii) large household purchases (m6), (iii) when they could visit family and friends (m7), and (iv) what to do with her husband's money (m8).
- Related to economic autonomy: the woman has money of her own that she can decide how to spend (m9).

We let the data tell us about the importance of the differential role of autonomy on different dimensions by using a latent factor model.[20] Since all measurements m_j ($j = 1, \dots, 9$) are binary, we use a logit model, specifying for woman i (conditional on her autonomy trait a_i):

$$\mathrm{Prob}\left(m_{ij} = 1 \,\middle|\, \eta_i\right) = \Lambda\left(\mu_j + \lambda_j a_i\right), \tag{2}$$

$$a_i \sim N\left(\theta' s_i, \sigma_a^2\right), \tag{3}$$

where Λ is the logistic distribution function, and μ_j and λ_j are the intercepts and factor loadings for measurement j, respectively, in equation (2).

Unlike in the literature, we control for maternal and family characteristics, and enabling factors, by allowing the trait to be correlated with a set of characteristics via the specification of s in (3)—that is, the mean of the distribution of a is allowed to be dependent on a set of characteristics relevant for woman i. These are the characteristics of the woman (age, caste, religion, education, occupation, employment status in the last 12 months, type of media exposure), her partner (education and occupation) and family (nuclear family, household wealth index factor scores, state of residence). Taking education as an example, the above specification (3) allows, on average, the autonomy trait to be different between women with different levels of education, *ceteris paribus*.

We impose the normalization that the first loading is 1 in equation (2) for identification. The model given by (2) and (3) is estimated jointly using maximum likelihood methods. We then use the estimated posterior conditional mean (empirical Bayes) $E(a_i|\text{data})$ of the latent variable a_i to construct our index of autonomy for every woman in the sample.[21]

In the language of item response theory (Carlson and von Davier 2013), the intercepts μ_j are called the item 'difficulty' and factor loadings (i.e. slope coefficients), and the λ_j are called item 'discrimination'. Comparing two intercepts, the smaller intercept implies a lower probability of saying yes to that question relative to the other one. In terms of the factor loadings, the probability of saying yes to the measurement that has a larger slope (factor loading) will be more sensitive to small changes in the autonomy trait compared to the one with the lower slope, hence measurements with higher slope coefficients are said to be more discriminatory. Hence (i) a larger intercept implies that women are more autonomous in this dimension, in general, and (ii) the larger the slope, the better the measurement in distinguishing autonomy traits between different women. In our model, even a small change in the autonomy index will be associated with a larger probability of the response to the autonomy question being positive. It is important to account for the differential role of autonomy in different dimensions.

The main advantage of our approach is that we are able to separate the direct associations of maternal and family characteristics in our model for nutrition conditional on 'autonomy', from their indirect associations that work through maternal autonomy. For identification of these separate effects, we do not require any additional variables as our index is a non-linear function of the variables s in the equation. Including additional variables provides extra variation in the index without relying on functional form for identification. This is similar to the issues in the two-step Heckman method for correcting for sample selection. Therefore we include an additional variable that measures the age difference between the partners. The assumption is that this variable affects the nutritional outcome of children only via a (*the index of autonomy*), and conditional on a, it is uncorrelated with the error term in equation (1). Findings from the literature generally support the idea that a smaller age gap is associated with greater autonomy (Hogan *et al.* 1999; Barbieri and Hertrich 2005; Sharma *et al.* 2021).[22]

5 | DESCRIPTIVE STATISTICS OF THE VARIABLES

The anthropometric information was collected on surviving children who were under 5 years old at the time of the interview. In the interests of space, we provide descriptive information for the main outcome variables and some child characteristics in Table 1. The full list of summary statistics of all the variables is in Table A1 of Online Appendix 1, which also contains the full set of estimation results. The sample consists of rural women who had at least one child aged under 60 months at the time of the interview, and had non-missing values for the nutritional status variables.

TABLE 1 Descriptive statistics.

	Mean (SD)
Panel A: Child covariates z (equation (1))	
Girl 0.48	
Age in months	30.2 (17.0)
Part of a multiple birth	0.01
Birth order	
1	0.27
2	0.26
3	0.18
4	0.11
5 or more	0.18
Preceding birth interval	
< 18 months	0.08
18–24 months	0.15
25–36 months	0.51
> 36 months	0.27
Panel B: Nutrition status variables	
HAZ score	−1.86 (1.66)
Stunted (children with HAZ < −2)	0.48
Panel C: Measurements used in construction of autonomy index	
Woman is allowed to go to	
market alone (m1)	0.48
health facility alone (m2)	0.45
places outside community alone (m3)	0.36
Woman has the final say alone on purchases for daily needs (m4)	0.29
Woman has the final say together on	
own healthcare (m5)	0.61
large household purchases (m6)	0.5
visiting family and friends (m7)	0.58
what to do with husband's money (m8)	0.62
Woman has money for her own use (m9)	0.36
Average total score (SD)	4.24 (2.48)
Mean of average scores	0.47
Median of average scores	0.44
Number of mothers	17,749
Number of children of 17,749 mothers	23,878
Proportion of mothers with one child in this sample	0.59
Mother has experienced at least one child death	0.06

Notes: The sample consists of women with children who were under 60 months old at the survey time and thus were eligible to contribute to the 'nutrition' analyses; see text for further details. The nutritional status variable definitions are based on WHO standards. All variables are binary except when a standard deviation (SD) is indicated in parentheses.

5.1 | Child characteristics

As per panel A of Table 1, 48% of children are female. The average age of children in the main sample is 30 months, and 53% are either firstborn or secondborn. It is interesting to note that around half of the births were 2–3 years after the previous birth. There are two possible explanations for this: either the births are properly spaced, or some families are engaged in prenatal sex selection. We will come back to this issue when we discuss the results in Section 6.

5.2 | Nutritional status variables

Summary statistics for the nutritional status variables are in panel B of Table 1. Here, 48% of children are stunted according to the WHO definition. This is a very large number compared to the predicted proportion of children who would be classified as stunted according to the WHO distribution. The HAZ score, the smoothed HAZ score by age in months, and the probability of being stunted by age, are provided in Figures 1, 2 and 3, respectively. Three points are noteworthy here. First, the distribution of HAZ score is shifted to the left relative to the WHO distributions. Second, HAZ scores deteriorate with age but stabilize after the child reaches approximately 24 months. Third, the proportion of children classified as stunted also increases rapidly with age until the age of about 36 months.

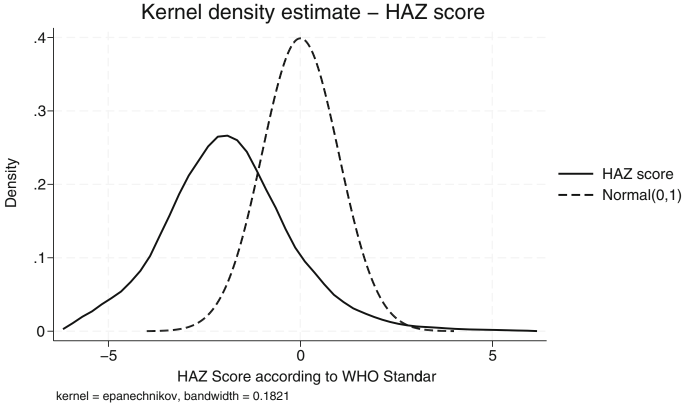

FIGURE 1 HAZ scores—rural children aged 0–59 months. *Notes*: Based on the authors' calculations from the sample used for the estimation of the model. The sample consists of rural women who had at least one child aged under 60 months at the time of the interview and had non-missing values for the nutritional status variables.

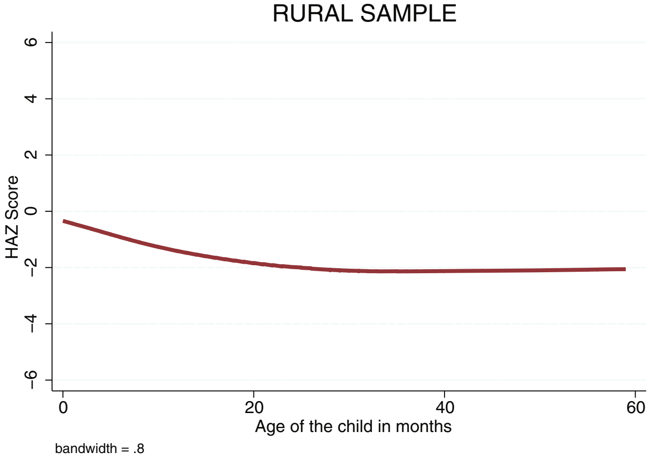

FIGURE 2 Smoothed plots of HAZ score, by age—rural sample. *Notes*: See Figure 1.

FIGURE 3 Proportion of children who are classified as 'stunted', by age.
Notes: See Figure 1.

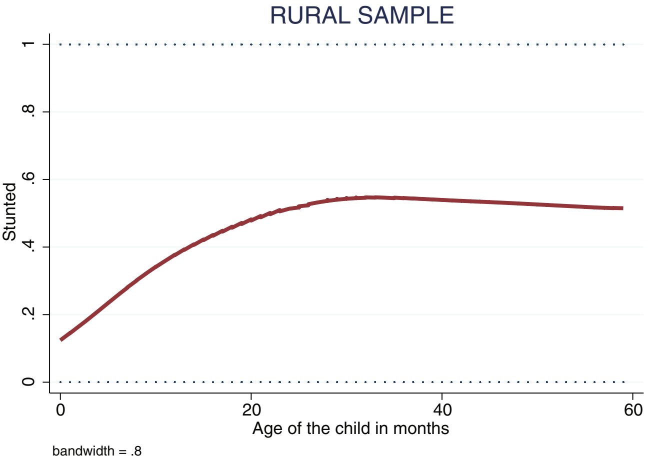

TABLE 2 Frequency distribution of the sum of the measurements (m1–m9) used in the construction of the autonomy index.

Sum	Number of women	%	Cumulative %
0	1454	8.2	8.2
1	1729	9.7	17.9
2	1589	9.0	26.9
3	1976	11.1	38.0
4	2800	15.8	53.8
5	2355	13.3	67.1
6	1784	10.1	77.1
7	2103	11.9	89.0
8	1564	8.8	97.8
9	395	2.2	100.0

Notes: See Table 1 for the definitions of the measurements. Number of women in the rural sample is 17,749. Sample average of the score is 4.2.

5.3 | Autonomy measurements

As shown in panel C of Table 1, the sample average score created by summing all the means of the autonomy-related measures is only 4.24. That is, on average, rural women have autonomy as indicated in only four of the measurements. The frequency distribution of the summed score is provided in Table 2. Around 8% of women do not have any autonomy at all according to our chosen measurements. Only a very small proportion of women (2.2%) are recorded as saying that they have full autonomy based on the answers to the selected questions.

6 | RESULTS

6.1 | Autonomy index

In the interests of space, we focus on the coefficient estimates of equation (2), which are provided in Table 3. The rest of the results are provided in Table A1 of Online Appendix 1.

TABLE 3 Estimates of equation (2) parameters (standard errors)—impact of women's autonomy on probability of positive response to the measurement question.

Measurements (binary indicators)	Factor loading (1)	Intercept (2)
Woman is allowed to go to market alone—intercept (m1)	1	−1.407*** (0.082)
Woman is allowed to go to health facility alone (m2)	0.886*** (0.026)	−0.005 (0.046)
Woman is allowed to go to places outside community alone (m3)	0.894*** (0.028)	−0.484*** (0.050)
Final say alone on purchases for daily needs (m4)	0.745*** (0.030)	−0.653*** (0.060)
Final say together on own healthcare (m5)	2.171*** (0.081)	−0.682*** (0.097)
Final say together on large household purchases (m6)	2.877*** (0.119)	−2.420*** (0.150)
Final say together on visiting family and friends (m7)	2.959*** (0.115)	−1.815*** (0.151)
Final say together on what to do with husband's money (m8)	1.463*** (0.055)	0.170*** (0.065)
Woman has money for her own use (m9)	0.023 (0.020)	0.758*** (0.085)
Estimated variance of woman-level heterogeneity	0.815*** (0.023)	
Estimated variance of PSU level heterogeneity	0.748*** (0.015)	
'Reliability' measure (%)		
Woman is allowed to go to market alone	16.8	
Woman is allowed to go to health facility alone	13.8	
Woman is allowed to go to places outside community alone	13.8	
Final say alone on purchases for daily needs	10.2	
Final say together on own healthcare	48.7	
Final say together on large household purchases	62.6	
Final say together on visiting family and friends	63.9	
Final say together on what to do with husband's money	30.1	
Woman has money for her own use	0.0	
Maximized log-likelihood value	−90,899	

Notes: The 'reliability' measure provides the percentage of variation attributed to the autonomy variable in the total variation observed in that *particular* measurement. A measurement with a larger 'reliability' measure is able to explain a larger proportion of the variability in the observed pattern of women's answers to that question relative to another measurement with a smaller reliability measure. PSU is the primary sampling unit. The rest of the results are provided in Table A1 of the Online Appendix.
***, **, * indicate $p < 0.01$, $p < 0.05$, $p < 0.10$, respectively.

The factor loading related to whether the woman is allowed to go to the market alone is normalized to 1, for identification.[23] As shown in Table 3, the decision-making measurements have high discriminatory power individually, relative to the reference category (i.e. they have a factor loading that is larger than 1). A higher factor loading indicates that a small difference in the autonomy trait is associated with a larger change in the probability of saying 'yes' to these questions relative to other measurements.

The 'reliability' measure calculated as the proportion of variance explained by the autonomy index in the total variation of the measures (m1–m9) *individually* is provided in Table 3.[24] The latent autonomy trait is estimated to explain more than 60% of the variations in the observed measure related to whether the woman has a role in the decisions concerning large household purchases, and visiting family and friends; the latent autonomy trait also explains nearly half of the variation in the woman's participation in decisions regarding her own healthcare. Unequal factor loadings estimated in this model reiterate the importance of allowing for different dimensions of autonomy to play different roles in the construction of the autonomy index; thus they illustrate why an index derived by simply averaging the measures might be problematic.

6.2 | Choice of the estimation sample

In order to facilitate the discussion of selecting the appropriate sample to address the possible issue of 'son preference', we first provide some relevant summary statistics in Table 4.

We first discuss how best to mitigate the effect of 'son preference' possibly biasing our estimator. As discussed in the Introduction, son preference is likely to lead to differential care and feeding practices, and hence to differential nutritional outcomes. That is, nutritional outcomes can depend significantly not only upon the sex of the child, but also upon the sex composition of existing children, and how this compares with parents' desired numbers of boys and girls.

There are two ways in which son preference may cause our sample to be selected endogenously. First, son preference may lead to sex-selective abortion and hence may lead to a lower proportion of girls at birth. Second, son preference is likely to have an impact on birth intervals and fertility choices. Parents may use a stopping rule for their fertility choice that depends on the numbers of girls and boys they already have (Barcellos *et al.* 2014). Additionally, the birth intervals between children also might depend on the sex of the previous child if the mother tries to conceive faster in the hope of having a boy after a girl (Jayachandran and Kuziemko 2011). Both these practices would imply that the number and sex of children in the sample are not determined randomly but depend upon various other observed and unobserved factors that may have been omitted from the specification—thereby causing the OLS estimator to be biased and inconsistent.[25]

While son preference could lead to sex-selective abortion of pregnancies at all birth orders, the likelihood of sex-selective abortion is negligible in first-order pregnancies but increases with higher-order pregnancies (Saikia *et al.* 2021). Jha *et al.* (2006) also use data from 1.1 million Indian households and find evidence of sex-selective abortion for second- and higher-order births, but no clear evidence for first-order births. Apart from being an emotional decision, childbirth is often a social and economic decision as well. Son preference also means that households may desire at least one or two male children. If there are resource constraints for raising children, then they may have to limit the number of children they can have to two or three. This is what often leads to sex-selective abortions. If the first child is a girl, then families often still have the hope that the second or the third child will be a boy. But if the second child also turns out to be a girl, then families may not wish to spend additional resources on raising a child they do not desire. Instead, they may choose to abort the foetus in the hope that the next child will be a boy. This leads to a higher probability of sex selection among higher-order births.

TABLE 4 Sample characteristics.

Panel A: Number of children aged <60 months

					Total (number)
Number of children	1	2	3	4	
% of mothers	68.5	28.6	2.82	0.08	17,749

Panel B: Distribution of birth order—column %

Birth order	1	2	3	4	5	6	≥ 7	Total
Girls	49.4	48.2	46.5	48.4	47.8	49.4	46.7	48.2
Boys	50.6	51.9	53.5	51.6	52.2	50.7	53.3	51.8
Total (number)	6434	6312	4219	2682	1758	1078	1395	23,878
Total—row %	27.0	26.4	17.7	11.2	7.4	4.5	5.8	100

Panel C: Firstborns with secondborn in the sample by year of birth of firstborn

							Total (number)
Year of birth	2001	2002	2003	2004	2005	2006	
Numbers	649	829	569	193	11	1	2252
%	28.8	36.8	25.3	8.6	0.5	0.04	100.0

Panel D: Age in months of firstborns

Age in months	0–15	16–17	18–23	24+	Total
Numbers	1696	241	657	3840	6434
%	26.4	3.8	10.2	60.0	100.0

Panel E: Age in months of firstborns with birth year >2004

Age in months	0–15	16–17	18+	Total
Numbers	1557	78	7	1642
%	94.8	4.8	0.4	100.0

Panel F: Firstborns without a younger sibling

Age in months	0–15	16–17	18–23	24+	Total
Numbers	1690	227	587	1698	4202
%	40.2	5.4	14.0	40.4	100.0

The data on nutrition were collected for children born within the last five years at the time of the interview and currently alive. In Table 4, we provide some descriptive statistics on the characteristics of the children observed in terms of age, birth order and year of birth, to help us with the selection of the sample for the estimation.

To shed some light on possible sex selection through abortion of female foetuses, we first look at the number of children born to the mothers in our sample. Here, 69% of mothers had only one child born during this time interval, and 97% of mothers contributed only one or two observations to the sample (panel A). Among all children, except for the firstborns, the sex imbalance is exacerbated (panel B).

The other issue (i.e. son preference affecting birth spacing, stopping rules, and care and feeding practices) is more complicated. If the firstborn is a girl, then the family may try to conceive sooner in the hope of having a boy. This would reduce the amount of time that the child can receive undivided care and attention (especially breast milk) (Jayachandran and Kuziemko 2011). Therefore the nutritional status of the firstborn may depend upon the parents' attitude (i.e. their son preference) as well as upon the birth interval, and the sex of the second child.

Panel C of Table 4 describes how many firstborn children were observed with a secondborn by the birth year of the firstborn. We find that 35% (2252 out of 6434) of firstborns have a second sibling in the sample. As one would expect, the older the firstborn at the interview time, the higher the chances of observing a second child in the sample. Since this pattern is dictated by the birth intervals, selecting a sample of firstborns without a sibling (panel F) will not deal with the problem of endogeneity caused by son preference as discussed earlier. This can be illustrated with an example. If the first child is a girl, then the mother may have the second child quickly in the hope of having a boy. On the other hand, if the first child is a boy, then the woman may delay the second pregnancy to allow the boy to receive full care and attention. Thus if we use this criterion, i.e. firstborns without a sibling, then *boys* may have a higher probability of inclusion into the sample (panel C). Therefore we need an additional criterion to restrict the sample of firstborns, in terms of either the birth year or the child's age.

In panels D and E of Table 4, we provide a breakdown of the ages of all firstborns and those born after 2004, respectively, as some may feel that a calendar-year-based criterion is likely to suffer less from endogenous selection bias. We focus on the year 2004 as the interviews took place in 2005 and 2006. As we compare these two panels, we find that the majority of the firstborns aged under 18 months were born in 2005 or 2006. The chance of finding a younger sibling becomes higher as we increase the calendar time interval (say choose 2003 instead of 2004).[26]

An additional issue that we will have to consider is whether there has been any child death in the family. If a child had died in a family because of severe malnutrition, then the sample of surviving children for whom we have valid nutritional information is selected endogenously. Some 6% of the mothers in the original sample had experienced a child death (Table 1). However, in the sample of firstborns aged under 18 months, only three mothers had experienced a child death. We do not expect this to be a problem.

In summary, based on the discussions above, our preferred specification is the one that restricts our sample to those firstborns who are aged under 18 months, that is, who are young enough that they are not very likely to be affected by the birth (and hence sex) of the second child. We have chosen to restrict our sample in this way rather than choosing the eldest children without a younger sibling because the choice of the 'only child' as a sample group will lead to endogenous selection if the mother conceives sooner after a girl (Barcellos *et al.* 2014). We also provide estimates based on the other sample selections as discussed in this section.

6.3 | Nutritional status

We next summarize the estimates of the parameter of interest—the coefficient of our autonomy variable—by different cuts of the sample used in the estimations, in Table 5. As discussed in the Introduction, the HAZ score measures the long-term nutritional status of the child. An additional interaction term between the autonomy variable and a girl child was included in the model to assess whether female children benefit more than male children when the mother is more 'autonomous', *ceteris paribus*. However, the interaction term was insignificant in all the regression models reported in Table 5.

The most important finding is that, in general, maternal autonomy is a significant positive predictor of a better HAZ score, and a negative one of stunting. We defer discussions on the magnitudes of these estimates until later in this section, and summarize the main results here (see Table 5).

(i) The autonomy index coefficient is not significant in the HAZ model estimated using all children aged 0–59 months, regardless of whether the sample contains all children or only

TABLE 5 HAZ score and 'stunted' regressions—coefficient estimate (standard error).

Variables	HAZ	'Stunting'	HAZ	'Stunting'
	(1)	(2)	(3)	(4)
Panel A	All birth order		Firstborns	
	Age 0–59 months		Age 0–59 months	
Autonomy	0.022	−0.011***	0.026	−0.008
	(0.015)	(0.004)	(0.025)	(0.008)
Constant	−1.880 ***	0.467***	−2.239***	0.623***
	(0.088)	(0.026)	(0.173)	(0.051)
R-squared	0.110	0.094	0.135	0.119
Number of children	23,788	23,788	6413	6413
Panel B	Firstborns		Firstborns	
	Age < 15 months		Age < 18 months	
Autonomy	0.146**	−0.029*	0.161***	−0.032**
	(0.058)	(0.015)	(0.051)	(0.014)
Constant	−1.218***	0.325***	−1.441***	0.382***
	(0.377)	(0.098)	(0.364)	(0.088)
R-squared	0.139	0.133	0.176	0.157
Number of children	1571	1571	1931	1931
Panel C	Firstborns		Firstborns with no	
	Birth year >2004		younger sibling	
Autonomy	0.158***	−0.033**	0.035	−0.010
Constant	(0.059)	(0.015)	(0.034)	(0.009)
	−1.129***	0.300***	−1.922***	0.503***
	(0.375)	(0.101)	(0.211)	(0.064)
R-squared	0.151	0.132	0.134	0.123
Number of children	1640	1640	4185	4185

Notes: The full set of variables included in the regression is in Table A2 of the Online Appendix. Age dummies (0–5 (base), 6–11, 12–17, 18–23, 24+), as well as the birth order dummies, were included where appropriate. Bootstrapped standard errors (allows for clustering at the district level with 500 replications) in parentheses. R-squared is valid when errors are homoscedastic. The number of observations used in the estimations, relative to those reported in Table 4, can differ due to the missing values in some of the included variables in the regressions.
***, **, * indicate $p < 0.01$, $p < 0.05$, $p < 0.10$, respectively.

the firstborn (panel A). As discussed, however, the estimator of the coefficients for this sample may be biased because the sex composition of children in this sample may suffer from endogeneity acting through 'son preference'.

(ii) In terms of the firstborn sample, a one SD higher autonomy score is estimated to be associated with about 0.16 higher HAZ score, depending on how we cut the sample (column (1) of panels B and C, and column (3) of panel B. When we focus only on the firstborns in terms of their age (our preferred sample), the longer the observation period (15 months versus 18 months), the higher the chances of another younger child in the family. However, these estimates are similar to whether we select the firstborns in terms of their age at the time of the interview (age under 15 or age under 18), or in terms of the birth year being greater than 2004. All three of these sample selections are used to reduce the probability of another younger sibling in the family (see panel C of Table 4,). In all three samples, the estimated effect of a one SD increase in women's autonomy index on the child's HAZ score is around 0.16 and is significantly different from zero.

(iii) We next turn to the results presented in column (3) of panel C, where we focus on the firstborns but without a sibling. The estimations carried out using this may suffer from 'endogenous' selection due to the reasons discussed earlier in this section. The estimated coefficient of the autonomy variable is not significant.
(iv) The association between women's autonomy and child stunting is generally significant, ranging from an estimate −0.01 to around −0.03. In particular, among the firstborn children aged under 18 months (or under 15 months, or born after 2004), the predicted probability of stunting when we shift our autonomy index by +1 SD is generally significant and lower by about 0.03 points (columns (2) and (4) of panel B, and column (2) of panel C). The model estimated using a logit specification using the sample of firstborns aged under 18 months produced an estimated average partial effect for this autonomy index of −0.034.

We therefore conclude that there is a positive association between the long-term nutritional status of the firstborn and maternal autonomy.

It is well known that the first two years of life are the most important 'window of opportunity' to make a long-term impact on children's nutritional status (UNICEF 2013), and their lifelong health and wellbeing. Thus the finding that more autonomous mothers can contribute to better health for their children specifically during this key window of time is crucial for policy purposes.

6.4 | Interpretation of the magnitude of the coefficient of autonomy

We next turn to the interpretation of the estimated coefficients of autonomy in the HAZ score, and the probability of stunting. As seen in Figures 1–3, relative to the HAZ scores of children younger than 6 months, the HAZ scores of older children become worse as they grow older; the probability of being stunted increases as well. These findings are reiterated in Table 6 (model results the same as columns (3) and (4) of panel B in Table 5), where we report a few coefficient estimates for the specification that uses a sample of firstborns aged under 18 months.

A 6–11-month-old child is estimated to have a HAZ score of about 0.3 SD lower than that of a child aged under 6 months, *ceteris paribus*. This deteriorates even more for a child who is between 12 and 17 months old. The observed average HAZ score and the proportion who are classified as stunted, for our sample of firstborns aged under 18 months, are −1.15 (SD 1.72) and 0.31, respectively. The estimated coefficient of the autonomy index for this sample is 0.161 (for HAZ score) and – 0.032 (for stunting) (see Table 6). Hence one SD higher (relative to the mean 0) autonomy index is associated with a higher HAZ score 0.09 (0.161 divided by 1.72), giving a new HAZ score – 1.06 and a new probability of stunting 0.28. In terms of the WHO distribution of HAZ scores, this is equivalent to a shift of a child from the 13th to the 15th percentile position. Interestingly, the effect of a change +1 SD of our autonomy index, both in our HAZ score and in stunted regressions, is about half the age effect for 6–11-month-old children, and about 15% for 12–17-month-old children, relative to those aged under 6 months. An estimated 22 million children aged under 18 months live in rural India (Census of India 2011). Given the sample proportion (30%), this translates to an estimated 6.6 million firstborn children in this age group; among them, approximately 2.1 million children (30.5%) would be classified as stunted. A one SD higher autonomy is associated with 300,000 fewer cases of stunting among firstborn *children aged under 18 months* (as evidenced by a decline from 30.5% of this population to 27.3%). As this group of children aged from under 5 months to the 6–11-month-old category, this level of increase in maternal autonomy would effectively halve the average deterioration in HAZ scores experienced.

TABLE 6 Estimates of equation (1)—firstborn rural children aged under 18 months.

	HAZ	'Stunted' (binary)
Maternal autonomy z-score	0.161***	−0.032**
	(0.051)	(0.014)
Child characteristics		
Age in months—binary (base <6 months)		
6–11	−0.318***	0.058**
	(0.100)	(0.025)
12–17	−1.032***	0.237***
	(0.099)	(0.027)
Girl	0.178**	−0.039**
	(0.071)	(0.019)
Part of multiple births	−2.392***	0.630***
	(0.442)	(0.162)
R-squared	0.176	0.157
Sample average of the dependent variable (SD)	−1.15	0.31
	(1.72)	
Number of children	17,749	11,187

Notes: This is the same as columns (2) and (3) of panel B in Table 5. A full set of estimates is provided in Table A1 of the Online Appendix. Bootstrapped standard errors (allows for clustering at the primary sampling unit level with 500 replications) in parentheses. ***, **, * indicate $p < 0.01$, $p < 0.05$, $p < 0.10$, respectively.

7 | SUMMARY AND CONCLUSIONS

This paper has attempted to address two important econometric challenges when exploring the link between maternal 'autonomy' and child nutrition: (i) the definition and measurement of 'maternal autonomy', and (ii) the estimation of this link using a sample of children born that might be selected 'endogenously' due to 'son preference' in India.

In order to address the first challenge, we start with the distinction between 'empowerment' and autonomy, and the premise that maternal autonomy is a latent trait, and hence can be inferred only through measures that are impacted by autonomy, such as decision-making, freedom of movement, etc. We then construct a measure of 'autonomy' allowing the data and the model to tell us about each measure or the relative extent to which autonomy is related to each measure. Unlike the models used in the literature, our methodology enables us to assess empirically the role of socioeconomic factors such as the age difference between the husband and wife, the level of education, employment, caste, religion, etc. of the partners, in informing us of their relationship with autonomy. However, some of these factors are also likely to have a direct impact on nutrition—through their impact on the availability of food, environmental conditions, health and nutrition practices, etc. Our methodology enables us to separate the direct associations of maternal and family characteristics in our model for nutrition, from their indirect associations that work through maternal autonomy.[27]

The second challenge was the possible endogenous selection of the sample of children used in the analysis, due to the existence of 'son preference' in India. We provide some sensitivity analysis to different selections of the estimation sample to address possible biases. Our preferred model estimates therefore restrict the sample to firstborns under 18 months old. The results are similar if we select the sample of firstborns less than 15 months old or firstborns born after 2004.

We conclude that there is a positive association between the long-term nutritional status of the firstborn and maternal autonomy. The finding that more autonomous mothers have an important role specifically during this key window of time is crucial for policy purposes.

While this paper establishes the links between maternal autonomy and the long-term nutritional status of children, it does not delve into the specific pathways that have been explored earlier.[28] Future research can utilize the proposed methodology to create an index of autonomy and to explore the impact and identify the effectiveness of various policies on women's autonomy. As this paper has argued, such policies would not only improve women's autonomy—a desirable outcome in itself—but also help to reduce persistently high levels of child undernutrition in India that acts via the effect on the mother's autonomy.

A limitation of the analyses presented here is the use of an earlier round of the survey—the NFHS-3 (2005–6) instead of the NFHS-4 (2015–16) (IIPS and ICF 2017; see note 16).[29] There were differences in the survey design, sample sizes and the nature of questions asked across the two surveys.[30] For example, only 20% of women who were selected for the state module were surveyed regarding the questions on 'empowerment', which we use. This does not mean that NFHS-4 cannot be used for the analysis. Rapid technological advancement over the years would, however, require a complete rethinking of how one defines 'women's autonomy'.[31] However, we believe that our methodological approach is unique and provides another way to estimate an index of autonomy where the concept is difficult to measure precisely.

One of the benefits of using the NHFS-3 is estimating a cleaner effect of the role of autonomy in shaping child nutrition, as not only new national and state-level policies have been introduced since the NFHS-3 was conducted, but some of the old policies also have been scaled up or reformed. For instance, the Indian Parliament passed the National Food Security Act in 2013, which made 75% of the country's population eligible for highly subsidized food grains, and also implemented policies for the nutrition of children and women, such as the Integrated Child Development Scheme and the Mid-Day-Meal Programme under the ambit of the law.[32] These programmes are likely to have varied and significant impacts on both women's autonomy and children's nutritional status. However, it is difficult to capture these impacts due to the unavailability of data and information on many of these measures, and to disentangle the effects of policies that affect autonomy from policies that were aimed at affecting nutrition directly. It is important to note that despite all the policies that have aimed to increase the levels of child nutrition, the levels of child nutrition are getting better only very slowly (see note 12). Our paper shows that policies to enhance maternal autonomy will have a positive impact on child nutrition.

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ENDNOTES

[1] Data obtained from the website of the Ministry of Women and Child Development, Government of India (9 March 2023).

[2] Data obtained from the Union Budget, Government of India, Demand for Grants, Demand No. 101 (https://www.indiabudget.gov.in/doc/eb/sbe101.pdf, accessed 15 February 2024).

[3] Some but not all, of the questions that are used by Jensen and Oster (2009) are similar to the ones that we use in this study, although the surveys are different. In contrast to our approach, the authors use autonomy as an *observable* variable, and measure it using the average of answers given to six questions/measurements with some overlap with our measures. Unlike our method, all answers were weighted equally in the construction of the index. Also, see the replication study by Iversen and Palmer-Jones (2014).

[4] This is similar to the literature that assumes that test scores measure unobserved ability (see, for example, Heckman *et al.* 2006).

[5] Principal component analysis is another data dimension reduction technique similar to what we use, and the first component is a linear combination of the observed data (measurements), which explains the largest variation in the observed measurements. Results from this and other methods of constructing an autonomy index are provided in Online Appendix 2.

[6] For example, Hu and Schlosser (2015) present some indirect evidence of possible prenatal sex selection in India.

[7] Information on nutritional status of children is usually collected for children born in the last 3 or 5 years of the survey.

[8] Barcellos *et al.* (2014) use the *first round* (1992) of the same data source (NFHS-1) as ours, to look at the effect of child sex on parental investments to avoid the issues related to sex-selective abortions. Since it is assumed that there was no prenatal sex selection in the early 1990s, their concern is regarding families possibly following a male-biased stopping rule. They address this by selecting a sample of last-born children aged under 15 months at the time of the interview, assuming that the family has not had time to react to the sex composition of the existing children.

[9] The z-scores are the numbers of standard deviations above or below a set of standard-deviation-derived growth reference curves by the Centers for Disease Control obtained from a reference population from the US National Center for Health Statistics, as recommended by the WHO (2006). The recommendations are based on evidence that differences in 'unconstrained growth' across children of different ethnic and racial backgrounds, socioeconomic status and feeding, are so minor for children under 5 years of age that it is appropriate to use a common reference.

[10] Children with a HAZ (WHZ, WAZ) score less than −2 SD are classified as stunted (wasted, underweight). Each index provides different information about the growth of a child. The HAZ score provides information about long-term nutritional status; it does not vary according to recent dietary intake. The WHZ score is an indicator of current nutritional status; a low WHZ score can indicate recent inadequate food intake or a recent episode of illness. The WAZ score, which reflects body mass relative to chronological age, is a composite indicator.

[11] Even in the latest NFHS-5 survey (2019–21), the percentage of children under 5 years of age who are stunted (wasted) is still high at 36% (19%) (IIPS and ICF 2021).

[12] Quality of education, dietary habits, access to nutrition, the concept of autonomy, etc., vary widely between rural and urban areas, hence pooling the sample for analysis does not make sense. Hence we have chosen to focus on rural women in this paper. In our sample, 61% of women reside in an area classified as rural. Results using the urban sample are provided in Tables A3 and A4 of Online Appendix 3 for comparison, but the focus here is on the rural sample estimates.

[13] The results are very similar when we select the sample of firstborns born after 2003 where the interviews were held during 2005 and 2006 (see columns (1) and (2) of panel C in Table 5). Essentially, this method also restricts the sample to younger children under 24 months.

[14] The autonomy index is normalized to have zero mean and unit variance.

[15] For other definitions, see, for example, Caldwell (1986), who defines opportunities for women to receive an education and work outside the home to proxy autonomy, while Mason (1986) uses control over household and societal resources to the same purpose.

[16] Currently, five rounds of the NFHS are available for India: NFHS-1 (1992–3), NFHS-2 (1998–9), NFHS-3 (2005–6), NFHS-4 (2015–16), and NFHS-5 (2019–21). The data collection for the latest round of the survey (NFHS-5) was interrupted during the Covid pandemic, and was released for public access only in July 2022. In this paper, we make use of data from NFHS-3 (2005–6). This is primarily because access to NFHS-4 became available only in 2018, after we had completed a significant part of our analysis. Although the main part of the questionnaires is similar across NFHS-3 and NFHS-4, there are slight differences—in particular, the sample selected for the empowerment questions, and the questions asked, are different. We comment on the benefits and costs of using NFHS-3 against NFHS-4 in the final section of the paper.

[17] Reported standard errors are bootstrapped to account for the fact that the autonomy measure is a 'generated regressor'. We chose to use a linear probability model instead of either a logit or a probit because of the generated regressor issue in the specification, as it does not make sense to calculate bootstrapped standard errors when the model specification is fully parametric. The average partial effect associated with maternal autonomy in the logit was very similar to that obtained from OLS.

[18] Provided with the dataset.

[19] We experimented with many more measurements, and found that the additional measures did not add significantly to the estimation of the autonomy index. The ranking of mothers in terms of their estimated autonomy status did not change with the addition of other measures.

[20] Results from the estimation of our preferred model using commonly used measures of autonomy are provided in Online Appendix 2.

[21] This is the Bayesian shrinkage estimator (see Goldstein 2003), which can be used for estimating unobserved individual specific heterogeneity (Train 2009, ch. 11) within a random effects model specification. Simply put, this estimator is $E(a|m_1, \dots, m_9) = \int a f(a|m_1, \dots, m_9)\, da = \int a f(m_1, \dots, m_9|a) f(a)\, da / f(m_1, \dots, m_9)$.

[22] The expectation is that the larger the age difference between the partners, the lower the autonomy of the wife *ceteris paribus* because the husband will be able to influence decisions more easily (Sharma *et al.* 2021). A smaller age gap has also been shown to be related to greater women's autonomy as it facilitates spousal communication and increases the woman's participation in reproductive decisions (Hogan *et al.* 1999). Similarly, Barbieri and Hertrich (2005) find that lower spousal age difference is associated with a higher probability of using contraception, which is indicative of higher woman's autonomy.

[23] We normalize on this factor loading since we expect autonomy to have a non-zero effect on this measurement.

[24] In our model, this is given by $\widehat{\lambda}_j^2 * \widehat{\sigma}_a^2 / \left[\widehat{\lambda}_j^2 * \widehat{\sigma}_a^2 + \widehat{\text{Var}}(\text{PSU} - \text{level unobservable}) + \pi^2/3\right]$ (where PSU means primary sampling unit; see equation (2)). A specific measurement with a larger 'reliability' measure is able to explain larger proportion of the variability in the observed pattern of answers to that specific question across the women relative to another measurement with a smaller reliability measure.

[25] The survey collected information on what were the ideal numbers of boys and girls the woman would like to have. We created a binary indicator for women who stated that they preferred a higher number of boys than girls. We do not report results with this variable included because of the possibility of it being highly correlated to the number of children already in the family and their sex composition.

[26] Barcellos *et al.* (2014) select a sample of last-born children under 15 months from the first round of the same survey, arguing that this will ensure that 'the parents have not had a chance to respond to the gender of the last child by having more children' (Barcellos *et al.* 2014, p. 187). Since our sample is drawn from a later round of the same survey, we are not able to assume that prenatal diagnostic tools were not widely available during the period covered by our data.

[27] As shown in Online Appendix 2, not only is the role of autonomy in child nutrition different when different autonomy indexes are used in the analysis, but it is also the case that the relative positions of women—in terms of the degrees of autonomy that the woman has—changes depending on which measure one uses.

[28] See, for example, Durrant and Sathar (2000) and Dyson and Moore (1983) for many pathways by which autonomy-related factors may influence child health outcomes.

[29] NFHS-5 was disrupted by the Covid pandemic, which caused havoc across the country as regular programmes stopped functioning, and state governments, local governments aand civil society organizations evolved their own unique strategies to ensure food and nutrition security while trying to control the spread of the pandemic and the consequent mortalities. Further, these policies, programmes and interventions took place at varying points of time in the pandemic (and in the survey).

[30] See Johri *et al.* (2021, Appendix 4, Section 1.1) for the differences in the sample design.

[31] Unsurprisingly, for example, the survey organizers felt that the question on decision-making regarding daily needs was no longer important, and it was dropped in NFHS-4. However, there were new questions that were added to capture the 'autonomy/empowerment' status of women.

[32] An example of another programme is the Pradhan Mantri Surakshit Matritva Abhiyan. The objective of this programme is to provide assured, comprehensive and quality antenatal care, free of cost, universally to all pregnant women on the 9th of every month. This was launched in mid-2016 during the survey data collection for the NFHS-4.

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SUPPORTING INFORMATION

Additional supporting information can be found online in the Supporting Information section at the end of this article.

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