Global Food Security 13 (2017) 12–20

Contents lists available at ScienceDirect

Global Food Security

journal homepage: www.elsevier.com/locate/gfs

Accounting for nutritional changes in six success stories: A regression-decomposition approach

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ARTICLE INFO

Keywords:
Nutrition
Bangladesh
Nepal
Ethiopia
Senegal
Zambia

ABSTRACT

Over the past two decades, many developing countries have made impressive progress in reducing undernutrition. We explore potential explanations of this success by applying consistent statistical methods to multiple rounds of Demographic Health Surveys for Bangladesh, Nepal, Ethiopia, Odisha, Senegal, and Zambia. We find that changes in household wealth, mother's education and access to antenatal care are the largest drivers of nutritional improvement, except for Zambia where large increases in bednet usage is the single largest factor. Other factors play a smaller role in explaining nutritional improvements with improvements in sanitation only appearing to be important in South Asia. Overall, the results point to the need for multidimensional nutritional strategies involving a broad range of nutrition-sensitive sectors.

1. Introduction

Between the mid-1990s and the 2010s, the five countries and the state of Odisha, India that are the focus of this special issue saw significant reductions in chronic undernutrition. These encouraging trends beg the main research question motivating this paper: what explains rapid and sustained progress in undernutrition across otherwise diverse settings? Addressing such a question is challenging. Experimental designs are generally not applicable to national level data, and are, in any case, not well suited to identifying the impacts of changes across multiple sectors. Qualitative approaches, as applied elsewhere in this issue, are essential for understanding the policy processes underlying national level change in nutrition, but stop short of identifying and quantifying which policies and programs have made a substantive difference to nutritional change. On the other hand, many quantitative observational analyses of nutrition have identified plausible determinants of nutritional differences across children, but stopped short of conducting more dynamic analyses of which factors may be driving nutritional change over time.

In this study we apply an alternative approach that is quantitative, dynamic and comparative. We build on the approach found in Headey et al. (2015), Headey and Hoddinott (2014) and Headey et al. (2016) (see Zanello et al. (2016) for the application of a quantile decomposition approach to Cambodia). The essence of this approach is: (1) to combine multiple rounds of comparative surveys, such as Demographic Health Surveys (DHS), to capture long term nutritional change; (2) to construct consistently measured explanatory variables over time; (3) to measure trends in these explanatory variables over time; (4) to estimate multivariate regression analyses to derive estimates of the marginal effects of these variables on nutrition outcomes; and (5) to apply decomposition techniques to estimate plausible changes in nutrition due to the changes in means observed in Step (3) and the marginal effects estimated in Step (4).

In this paper we go one step further by applying this approach to several countries spanning two continents and very diverse nutritional contexts: Bangladesh (1996/97-2014), Nepal (1996–2011), Odisha State, India (1992/93-2005/06, with an extension to 2011), Ethiopia (2000–2011), Senegal (1993–2011) and Zambia (2002–2014). We focus on linear growth, since this is widely regarding as the single most relevant indicator of overall nutrition with poor height-for-age z-scores (HAZ) causally linked to a whole host of adverse later life outcomes (Hoddinott et al., 2013). This exploratory approach is not without its limitations (we return to these caveats in our discussion section); mindful of these we emphasize that in the absence of experimental alternatives, it offers a transparent means of identifying a series of plausible explanations of nutritional change that more experimental studies can assess further.

The remainder of this paper is structured as follows. Section 2 outlines the data and methods used in the paper. Sections 3 and 4 present our results while Section 5 concludes.

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http://dx.doi.org/10.1016/j.gfs.2017.02.003
Received 30 September 2016; Received in revised form 1 February 2017; Accepted 1 February 2017

2. Data and estimation

We use DHS data from the five countries that are the focus of this special issue: Bangladesh, Nepal, Ethiopia, Senegal and Zambia as well as the state of Odisha in India. Details of these data sets and the surveys underlying them are found in ICF-International (2015). These multi-cluster surveys of ever-married women of reproductive age are generally well suited to our purposes, being high quality, nationally representative surveys that cover a broad range of the hypothesized drivers of nutritional change. We use the following DHS data sets:

- The 1996/1997, 1999/2000, 2004, 2007, 2011 and 2014 rounds of the Bangladesh DHS;
- the 1996, 2001, 2006 and 2011 rounds of the Nepal DHS;
- the 2000, 2005 and 2011 rounds of the Ethiopian DHS;
- the data on Odisha found in the 1992/1993 and 2005/2006 rounds of the India DHS;
- the 1992/1993, 2005 and 2010/2011 rounds of the Senegal DHS; and
- the 2001/2002, 2007 and 2013/2014 rounds of the Zambia DHS.

Our outcome variables are height for age (HAZ) z-scores for preschool children aged 0–59 months (Bangladesh, Ethiopia, Senegal, Zambia), 0–35 months (Nepal) and 0–47 months (Odisha) as measured against WHO growth standards that are described in de Onis et al. (2007). Table 1 presents HAZ means for the first and last round of data available to us (using sampling weights). Mean HAZ scores improved substantially from the 1990s to 2010s increasing by 0.90 in Bangladesh, 0.62 in Ethiopia, 0.57 in Nepal, 0.37 in Zambia, 0.33 in Odisha, and 0.29 in Senegal.

Our choice of drivers of change reflects the factors broadly outlined in the UNICEF (1990) and Lancet (Black et al., 2013) nutrition frameworks and in the existing economic and nutrition literatures on the underlying determinants of undernutrition. We discuss each in turn.

Table 1
Changes in the mean HAZ scores between earliest and latest DHS survey: Bangladesh, Nepal, Odisha (India), Ethiopia, Senegal, Zambia.
Source: Authors' calculations from Demographic and Health Surveys, using sampling weights.

Year	HAZ
	Bangladesh
1997	−2.28
2014	−1.38
Change	0.90
% Change	−39.6%
	Nepal
1996	−2.20
2011	−1.62
Change	0.57
% Change	−26.07%
	Odisha, India
1993	−1.98
2006	−1.65
Change	0.33
% Change	−16.59%
	Ethiopia
2000	−2.16
2011	−1.54
Change	0.62
% Change	−28.60%
	Senegal
1993	−1.44
2011	−1.14
Change	0.29
% Change	−20.35%
	Zambia
2002	−1.86
2014	−1.48
Change	0.37
% Change	−20.14%

Household socioeconomic status has long been identified as an important determinant of nutrition, whether measured as income or in terms of household assets (Haddad et al., 2003). Since the DHS does not collect income data, we used the asset/wealth index approach pioneered by Filmer and Pritchett (2001). This approach has been used in a large number of studies, though not without some criticism. Harttgen et al. (2013) note that the DHS data does not allow us to account for durables' age and depreciation when calculating asset indices. Other criticisms focus on the problems of using dichotomous variables in such an index, the lack of comparability between urban and rural areas; Filmer and Scott (2012) review these and other criticisms. Though cognizant of the limitations of this index, we include an asset index for each country – consistently measured across rounds in order to assess the role that changes in socio-economic status have played in these observed changes in nutritional status. The DHS asset module typically includes ownership of household durables (TV, radio, refrigerator, bicycle, motorcycle, and car) and housing characteristics (floor, wall and roof materials, number of bedrooms, and access to electricity) with the precise list of assets varying across countries. In our analysis, we use the set of assets that are recorded in all rounds for each country and state. We construct an asset index with the weights attached to different assets derived from a principal component analysis (PCA) for all the pooled rounds of data available for a given country(state). Hence, within each country(state) common weights are used across rounds to ensure consistent measurement of asset scores over time. The index is then scaled so that it varies between 0 and 10, with 10 being the maximum score observed across all rounds in a country and zero being the minimum. The values of these indices rise everywhere, but at rates ranging from 43% and 53% in Senegal and Zambia by more than 270% in Bangladesh and Nepal (Table 2).

Parental education has long been associated with child nutrition outcomes (Alderman and Headey, 2014; Behrman and Wolfe, 1984, 1987; Desai and Alva, 1998; Webb and Block, 2004). There are many hypothesized linkages, but parental education affects households' capacity to generate income and smooth shocks, knowledge of correct child care practices and, in the case of maternal schooling, bargaining power within the household. We include both mother's and father's education as measured by years of schooling. In all six case studies both maternal and paternal schooling rise over time with maternal schooling rising faster but from a lower initial value. There are large differences in schooling across case studies, with years of schooling higher in Zambia than elsewhere (where primary and lower secondary schooling are both free and compulsory). We note, however, that there is a difference between grades of schooling attained and years in school; the latter includes the impact of grade repetition and thus is an imperfect measure of educational attainment. Likewise years of schooling does not necessarily reflect schooling quality (Alderman and Headey, 2014).

Both work in nutritional sciences and economics stresses the importance of accounting for maternal height (see Behrman and Deolalikar (1988) for an early statement). Maternal height reflects genetics as well as the circumstances surrounding early life nutritional status of the mother. Since these are correlated with the circumstances surrounding investments in mother's schooling, failing to include maternal height will lead to upwardly biased estimates of the influence of maternal schooling on children's heights. However, maternal height changes little over the time period covered by these surveys (unsurprisingly).

The health environment surrounding the child also appears as an argument in reduced form demand functions for children's nutritional status. In this regard, work by Spears and co-authors has re-invigorated attention to sanitation (Hathi et al., 2014; Spears, 2013). In our regressions, we capture this as the proportion of households with no toilet at the village(cluster) level, with the exception of Zambia. This follows from the extant literature which shows that external bacteria

Table 2
Changes in mean values of drivers of change in HAZ between earliest and latest DHS survey: Bangladesh, Nepal, Odisha (India), Ethiopia, Senegal, Zambia.
Source: Author's calculations from Demographic and Health Surveys, using sampling weights.

Year	Asset Index (0–10)	Paternal educ. (yrs)	Maternal educ. (yrs)	Mother's height (cm)	Open defecation (%)	Private toilet (%)	Piped water (%)	4 or more antenatal visits (%)	Born in a medical facility (%)	Number of children	Households with bednets (%)
	Bangladesh (0–59 months)										
1997	1.52	3.6	2.4	150.3	25.7%		4.0%	6.8%	4.4%	3.4	
2014	5.97	5.7	6.2	150.8	2.9%		8.6%	31.4%	39.0%	2.2	
Change	4.5	2.1	3.8	0.5	−22.8%		4.6%	24.6%	34.6%	−1.23	
% Change	293.8%	59.6%	158.0%	0.4%	−88.7%		111.7%	361.0%	780.6%	−36.2%	
	Nepal (0–35 months)										
1996	1.0	3.6	0.8	150.5	88.3%		28.5%	7.2%	5.6%	4.1	
2011	3.9	5.0	2.7	151.0	52.4%		39.3%	40.0%	26.5%	3.4	
Change	2.8	1.4	1.9	0.5	−35.9%		10.8%	32.8%	20.9%	−0.7	
% Change	277.9%	40.4%	237.6%	0.3%	−40.7%		38.2%	454.6%	374.0%	−16.5%	
	Odisha, India (0–47 months)										
1993	1.0	4.7	2.4		90.5%		8.5%	19.0%	14.6%	3.1	
2006	1.9	5.7	4.3		81.0%		7.7%	37.6%	38.7%	2.6	
Change	0.9	1.0	1.9		−9.5%		−0.8%	18.6%	23.9%	−0.5	
% Change	81.5%	21.1%	79.5%		−10.5%		−9.9%	98.3%	165.2%	−17.0%	
	Ethiopia (0–59 months)										
2000	0.5	1.7	0.7	156.6	87.0%		12.1%	9.9%	3.8%	5.0	
2011	1.1	2.5	1.2	156.5	43.0%		26.5%	17.4%	8.3%	4.8	
Change	0.6	0.8	0.4	0.0	−44.0%		14.4%	7.5%	4.5%	−0.2	
% Change	112.5%	44.4%	60.6%	0.0%	−50.6%		117.9%	75.5%	120.5%	−4.0%	
	Senegal (0–59 months)										
1993	2.2	1.4	1.0	162.5	41.4%		39.6%	13.0%	44.7%	5.0	
2011	3.1	1.8	1.7	163.0	18.5%		67.8%	49.3%	72.0%	4.0	
Change	0.9	0.4	0.7	0.6	−22.9%		28.2%	36.3%	0.3	−1.0	
% Change	43.0%	29.3%	71.4%	0.3%	−55.3%		71.0%	278.6%	61.3%	−19.9%	
	Zambia (0–59 months)										
2002	1.5	7.3	5.2	157.8	30.1%	40.5%	29.4%	73.9%	43.4%	4.1	27.7%
2014	2.2	7.9	6.1	157.8	18.1%	49.8%	28.8%	56.8%	71.3%	4.1	75.5%
Change	0.8	0.6	0.9	0.1	−12.0%	9.3%	−0.6%	−17.1%	27.9%	0.0	47.8%
% Change	53.1%	7.6%	17.1%	0.0%	−39.9%	22.9%	−2.0%	−23.1%	64.2%	0.4%	173.0%

are more dangerous to an individual than the bacteria of fellow household members (Spears, 2013). However, in the case of Zambia our preliminary analysis found that the cluster level indicator had no explanatory power and alternative representations found that the percentage of households in the cluster with a private toilet was the only indicator associated with chronic undernutrition in Zambia. Previous studies provide some rationale for this, since sharing of toilets can lead to poor maintenance due to "tragedy of the commons" type problems (Gunther and Fink, 2010). While sanitation improves everywhere over time, the improvements in Ethiopia, Nepal and Bangladesh are especially notable; all three countries were pioneers of Community-Led Total Sanitation (CLTS), a subsidy-free approach based on behavioral change communications approaches combined with basic training on the construction of low cost latrines, usually pit latrines (Kar, 2003).

We also assess whether changes in access to clean water, specifically whether the household has a piped water supply, is associated with improvements in nutritional status. While a significant body of literature finds associations between piped water and reduced incidence of diarrheal diseases, associations between clean water and chronic undernutrition are more elusive (Gunther and Fink, 2010). Apart from Senegal, Table 2 shows that access to piped water is limited in the countries and states we study and only in Senegal was there a significant increase in access between the first and last DHS surveys. Another caveat is that the association between the source of water and the latent quality of the water is potentially ambiguous (Klasen et al., 2012).

Access to healthcare can be an important means of improving overall maternal and child health and also as a potential means for accessing nutrition-specific services, such as advice on child feeding and other care practices. For all countries(states) we focus on two variables: a dummy variable equaling one if the child's mother had four or more antenatal visits during pregnancy; and a dummy variable equaling one if the child was born in a medical facility. While these specific indicators are relevant to nutrition outcomes, we note that they may also be proxies for access to a range of other healthcare services, indicators of which were not available in all rounds (e.g. postnatal care, vaccinations, treatment of illness, access to nutrition knowledge). In all countries, the percentage of mothers who undertake four or more antenatal visits rises over time, albeit often from low bases. In five settings – Bangladesh, Ethiopia, Odisha, Nepal and Senegal – the earliest DHS surveys showed that fewer than 20% of mothers had given birth in a health care facility. In Bangladesh, Nepal, Odisha and Senegal this rises by 20% points or more. By contrast, in Zambia this falls from 73% to 56% between 2002 and 2014, which is likely related to a major decline in international support to Zambia's health sector (see Harris in this issue).

In addition to these health indicators, we include several additional indicators for Zambia, because (i) antenatal care access deteriorated over the timeframe in question despite nutritional improvements; and (ii) an existing literature documents large reductions in malaria in Zambia (Bhatt et al., 2015), a country in which malaria was previously endemic. We therefore included an additional variable in Zambia: the proportion of households in each village with bednets (since there are positive externalities that household bednet use can have on one's neighbors). The change in access to bednets is striking, which increased from 44% in 2002 to 69% in 2014. (Data on bednet usage is missing from one or more rounds from the other countries we consider.) And while there is surprisingly little research on malaria and child growth, there is indicative evidence that reductions in malaria could lead to improvements in birth outcomes (Eisele et al., 2012; Guyatt and Snow, 2004) and postnatal growth (Nyakeriga et al., 2004), although the relationship between malaria and nutrition is complex and potentially bi-directional.

A strand of work on household economics focuses on the trade-offs between the quantity and quality of child care (Becker and Lewis, 1973), since parents can devote more resources to each individual child when they have fewer children. There is also some evidence that lower order children are relatively neglected in some contexts (Jayachandran and Pande, 2013), and quasi-experimental evidence that having fewer children leads to reduced stunting rates (Jensen, 2012). Here we use the number of children ever born – achieved fertility – to capture these quality-quantity tradeoffs. The mean value for this variable falls markedly in Bangladesh, Nepal, Odisha and Senegal but much more slowly in Ethiopia and Zambia.

These variables and constructs – assets, maternal and paternal education, maternal height, sanitation, piped water, access to healthcare, and fertility – form the core of our analysis. In addition to these, we include four sets of control variables: child characteristics; maternal age; location effects; and time (survey round) effects. These additional controls are not "drivers of change" but their inclusion improves the precision of our estimates while reducing correlation between our drivers of change and the disturbance term. (Note that coefficients for these variables are not reported but results are available on request.) We include a dummy variable for child sex and monthly dummy variables for child age to capture the progressive growth faltering process that malnourished populations undergo until around two years of age (Shrimpton et al., 2001; Victora et al., 2010). We include a set of dummy variables denoting maternal age. We include regional dummy variables as well as a dummy variable equaling one if the child resides in a rural area. Lastly, we include survey round dummies to capture any factors generically improving nutrition outcomes at a national level (e.g. changes in food prices), and any systematic differences in survey characteristics.

We use linear regression models to assess the relationships between the height-for-age Z score (HAZ) for child i at time t ($H_{i,t}$) and the covariates (assets, maternal and paternal education, maternal height, sanitation, piped water, access to healthcare, fertility and the Zambia-specific variables) described above ($\boldsymbol{X}$), whilst controlling for child and maternal age dummies and location fixed effects (μ_i) as well as trend effects ($\mathbf{T}$). The vector of coefficients ($\boldsymbol{\beta}$) constitutes the set of parameters of principal interest. Adding in a white noise term ($\varepsilon_{i,t}$), we represent this relationship by Eq. (1):

$$H_{i,t} = \boldsymbol{\beta} \boldsymbol{X}_{i,t} + \mu_i + \boldsymbol{T} + \varepsilon_{i,t} \tag{1}$$

We use the estimated parameters from Eq. (1) to conduct a simple decomposition analysis described by Eq. (2), using only those variables which are statistically significant at the 10% level or higher. Standard errors account for the clustered nature of the sample.

$$\Delta \overline{N}_{i,t} = \boldsymbol{\beta} (\overline{\boldsymbol{X}}_{t=K} - \overline{\boldsymbol{X}}_{t=1}) \tag{2}$$

To calculate (2), we select the earliest DHS round ($t=1$) and the most recent round ($t=K$, where K is the number of DHS surveys used for each country). The decomposition then entails multiplying observed changes in the means of each explanatory variable over time by its regression coefficient. Doing so gives the predicted change in HAZ from each change in a nutrition-sensitive factor and thus shows the estimated contributions of each variable to changes in HAZ. For example, suppose that women's education rises by three grades between the first and last surveys for a given country, that is $(\overline{\boldsymbol{X}}_{t=K} - \overline{\boldsymbol{X}}_{t=1}) = 3$. Suppose also that the estimated coefficient on women's schooling, $\boldsymbol{\beta}$, equals 0.025. Multiplying these together yields 0.075. This indicates that the observed changes in women's schooling over time would predict a 0.075 standard deviation increase in HAZ. We can do analogous calculations for other potential drivers of nutritional change to gauge the extent to which each factor explains changes in HAZ over time, as well as how all the explanatory variables as a whole (i.e. the model) performs in explaining HAZ changes.

3. Results

Results of estimating Eq. (1) for all five countries as well as the state

Table 3
HAZ regressions by country.

Variables	(1) Bangladesh	(2) Nepal	(3) Odisha, India	(4) Ethiopia	(5) Senegal	(6) Zambia
Asset index (0–10)	0.0510***	0.0547***	0.0690**	0.0813***	0.0501***	0.0421***
	(0.00371)	(0.00627)	(0.0309)	(0.0130)	(0.0103)	(0.00944)
Father's education in single years	0.0238***	0.0133***	0.0287**	0.0197***	0.0152***	0.00708
	(0.00241)	(0.00391)	(0.0119)	(0.00481)	(0.00563)	(0.00530)
Woman's education in single years	0.0179***	0.0236***	0.0361***	0.0192***	0.0283***	−0.00319
	(0.00312)	(0.00480)	(0.0137)	(0.00630)	(0.00780)	(0.00547)
Mother's height	0.0554***	0.0553***		0.0416***	0.0418***	0.0462***
	(0.00164)	(0.00256)		(0.00229)	(0.00303)	(0.00253)
Households with no toilet, village proportion	−0.0992*	−0.176***	−0.329	−0.113*	−0.0979	
	(0.0538)	(0.0576)	(0.219)	(0.0616)	(0.0721)	
Households with private toilet						0.0759**
						(0.0306)
Households with piped drinking water	−0.0285	−0.0610**	−0.0383	−0.00605	0.0628	0.177***
	(0.0369)	(0.0304)	(0.136)	(0.0474)	(0.0473)	(0.0511)
4+ antenatal visits during pregnancy	0.0695***	0.0840**	0.177**	0.156***	0.147***	0.0737**
	(0.0218)	(0.0356)	(0.0877)	(0.0418)	(0.0430)	(0.0314)
Children born in medical facility	0.142***	0.233***	0.158	0.179***	0.163***	0.0720**
	(0.0230)	(0.0409)	(0.103)	(0.0518)	(0.0435)	(0.0348)
Total number of children ever born	−0.0430***	−0.0301***	−0.0133	−0.00899	−0.00686	−0.0202**
	(0.00689)	(0.00941)	(0.0274)	(0.00862)	(0.0109)	(0.00979)
Households with bednets, village proportion						0.272***
						(0.0817)
Constant	−10.13***	−9.521***	−1.259**	−7.395***	−7.863***	−8.375***
	(0.263)	(0.403)	(0.622)	(0.411)	(0.534)	(0.445)
Observations	28,043	10,608	2,540	14,390	7,727	13,174
R-squared	0.234	0.308	0.156	0.279	0.214	0.148

Note: Village-level clustered standard errors in parentheses. *, **, and *** indicate significance at 10%, 5%, and 1% levels, respectively. The regressions above include a number of time-variant controls, including regional fixed effects, year fixed effects, month-specific child age dummy variables, and dummy variables for various categories of maternal age.

of Odisha are shown in Table 3. Six regressions, each with approximately 10 regressors, produces 60 parameter estimates and their associated standard errors. To make sense of all these numbers, we begin by noting some general results before turning to country specific findings.

First, in all countries increased wealth is positively associated with HAZ and this association is statistically significant at the one percent level. There are however some differences in the magnitudes of these associations, with the parameter estimates ranging from 0.04 in Zambia to 0.08 in Ethiopia.

Second, in all countries except Zambia both paternal and maternal education are associated with higher HAZ and these associations are statistically significant. The parameter estimates for maternal schooling are sometimes slightly higher than those for paternal schooling but it is not obvious that these differences are functionally important. While the parameter estimate for maternal schooling in Odisha is slightly higher than that of other countries, we do not have data on maternal height in all survey rounds for Odisha. When we drop maternal height from the other country regressions, parameter estimates on maternal schooling tend to rise and so the slightly higher Odisha parameter estimate may reflect some omitted variable bias.

Third, holding other factors constant, taller mothers have taller children. We observe this in all countries for which we have data on maternal height. Moreover, coefficient magnitudes are reasonably similar, though the coefficients on maternal height in Bangladesh and Nepal are larger than those estimated for the sub-Saharan African countries.

Fourth, access to health care during pregnancy is associated with improved child height outcomes everywhere, and these associations are quite precisely estimated. To get a sense of the magnitude of the coefficients, we can add the coefficients on both variables (since they are dummies variables) to contrast the HAZ outcomes for a child whose mother who had four or more antenatal visits during pregnancy and who was born in a medical facility with a child with no exposure to these two services. These sums are 0.21 standard deviations (SD) in Bangladesh, 0.31 SD in Nepal, 0.33 SD in Odisha, 0.33 SD in Ethiopia, 0.31 SD in Senegal and 0.14 SD in Zambia. These are large changes in HAZ, suggesting that exposure to healthcare in the first 1000 days of life is a strong predictor of childhood nutrition outcomes.

Fifth, the sanitation variables are significant in four of the six case studies, with the Indian state of Odisha and Senegal being the main exceptions. However, the point estimate on open defecation is large in magnitude in Odisha (indeed, larger than the corresponding coefficients in the other samples), while the coefficient in Senegal is similar in magnitude to the coefficients estimated for Bangladesh and Ethiopia, but more imprecisely estimated. There is perhaps some suggestion in these results that the growth impacts of sanitation are conditional upon population density (Hathi et al., 2014), which is much higher in South Asia than in sub-Saharan Africa. The largest estimated coefficients are in Odisha and Nepal. However, other research from Odisha suggests that many households who nominally own toilets still defecate in the open (Barnard et al., 2013; Boisson et al., 2014), which may explain the imprecision of these estimates. Still another explanation of this imprecision is that Odisha has a much smaller sample size.

Sixth, the magnitudes of the estimated parameters of total number of children ever born are significant, negative and moderately large in Bangladesh and Nepal, smaller in magnitude in Zambia (but still significant), but insignificant in the three other samples. The largest association is found in Bangladesh and even there, a reduction of two ever born children would lead to an increase in HAZ of only 0.08 SD. Hence unless fertility rates were to drop precipitously, it is unlikely that this demographic change will be a major driver of nutritional improvement.

Finally, we turn to the additional explanatory variable used in the Zambian surveys, the proportion of households in each village with bednets. Strikingly, the coefficient on village bednet use is large, positive and highly significant. Moving from zero bednet use to full bednet use in a Zambian village is expected to increase a child's HAZ by 0.27 SD. We note that this coefficient likely masks parameter heterogeneity, since the effects of bednets should be larger (smaller) in locations where malaria is more (less) prevalent or endemic.

We now turn to the results of our decomposition analysis. This is shown in Fig. 1 as a stacked bar chart. The vertical axis is the percentage of the change in HAZ explained by the change in the covariates described in Section 2. Each segment of each bar for each country is calculated by multiplying the parameter estimates from Table 3 by the change in the value of that covariate found in Table 2. This product – a change in HAZ – is divided by the actual change in HAZ to estimate the percentage of actual HAZ statistically explained by changes in a given explanatory variable. The total height of the stacked bars for a given country tells us how much of the change in HAZ is explained by the combination of the associations reported in Table 3 and the changes in values reported in Table 2. So, for example, the parameter for the asset index in Bangladesh is 0.051 SD. The change in this variable between the first and last surveys is 4.45 (on the 0–10 scale) and so their product is 0.23 SD. The actual change in HAZ over this period in Bangladesh is 0.90 SD and so increased asset holdings is associated with 25% of the change in HAZ. Note that the healthcare segment in Fig. 1 represents two covariates — mothers with four or more antenatal visits and child being born in a medical facility — and the parental schooling segment represents paternal and maternal schooling variables. Other segments represent one covariate.

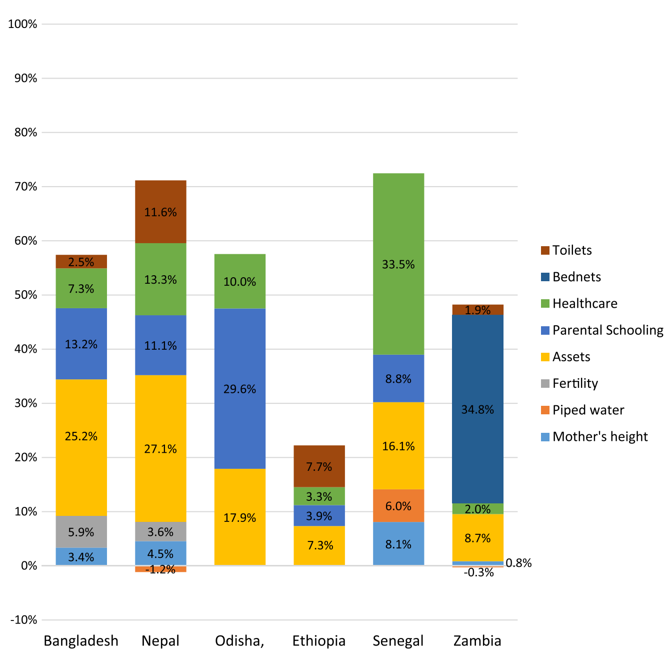

Fig. 1. Estimated contributions of selected factors to changes in HAZ: Bangladesh, Nepal, Odisha (India), Ethiopia, Senegal, Zambia. *Notes:* These are estimates changes in HAZ scores due to changes in the explanatory variables listed in the legend over the country-specific time frames. These estimates (or retrospective predictions) are based on a linear decomposition at means, in which changes in the mean of each explanatory variables are multiplied by the corresponding coefficient from Table 3.
Source: Authors' estimates from DHS data described in Section 2.

Also of interest is the explanatory power of the decomposition model as a whole, which is given by the total height of the stacked bars (which is analytically distinct from explanatory power of the regression model). For example, the corresponding value for Bangladesh is 57%, meaning that the decomposition explains just over half of the actual change in HAZ in Bangladesh between 1997 and 2014. The model for Odisha explains 58% of the change in HAZ, while the models for Senegal and Nepal explain 66% and 70%, respectively. The decomposition models for Ethiopia and Zambia have less explanatory power, accounting for 48% of the HAZ change in Zambia and just 22% in Ethiopia. This suggests that other factors not captured by our (common) model may be playing a role, or that measurement errors may be more important in some countries than in others.

In terms of the contributions of different factors, a striking result in Fig. 1 is the varying contribution of asset accumulation in many countries. Asset accumulation accounts for 16% or more of total HAZ change in Bangladesh, Nepal, Odisha, and Senegal, but less than 9% in Ethiopia and Zambia. Parental education is also an important factor in several countries, especially Odisha, as is healthcare, particularly in Nepal, Odisha and Senegal. In Zambia the health story is much more complex. Antenatal care declined and medical births improved very little. Overwhelmingly, however, the model suggests that increased bednet coverage were the large identifiable contributor to HAZ change in Zambia. Given the paucity of evidence linking anti-malaria interventions to child nutrition, this result suggests that the nutritional impacts of these interventions should be examined more closely.

These regression and decomposition results are based on pooling DHS data across years on the assumption that coefficients are stable over time. To test this, we estimated these models separately by year and formally tested to see whether coefficients were stable over time. In the vast majority of cases, we could not reject the null that there were no changes in parameter values. As a robustness check, we re-calculated our decompositions using coefficients from the first and last DHS surveys in each country; doing so does not change the pattern of results found in Fig. 1 (results available on request).

We also conducted additional robustness checks. We re-estimated Eq. (1) using stunting and severe stunting as the dependent variables. Doing so provides similar results to those reported in Table 2. Specifically, improvements in asset holdings, maternal schooling and reductions in open defecation all lead to statistically significant improvements in children's nutritional status and the magnitudes of these associations are similar across countries. We experimented with the functional form specification of our control variables, such as maternal age and schooling and child age, but doing so does not affect our findings. Adding other control variables, including access to vaccinations and additional fertility-related controls, does not materially affect our findings.

We also estimated separate regressions for younger (0–24 months) and older children (greater than 24 months). There are several good reasons to do so. First, younger children constitute an age group of particular interest because most growth faltering takes places in the first 1000 days of life. Second, some underlying factors, such as wealth and education, may affect growth right up to 24 months if not longer, implying that the inclusion of younger children that have not yet fully benefited from household wealth or parental education might lead to attenuation bias. Third, coefficients based on the inclusion of older children might lead to a different form of attenuation bias if there are recall errors. One example is antenatal care (parents might find it difficult to remember if a child 59 months old received antenatal care or not). Another is toilet use, since the survey asks about current ownership of toilets, and not about toilet use when the child was an infant. When we estimate separate regressions for children aged 0–24 months and children older than 24 months we find some evidence of these concerns. Generally speaking the coefficients on assets and parental education are larger for children older than 24 months, though we find no consistent pattern of results related to healthcare variables. The coefficients on sanitation vary but not in any consistent fashion, most likely because this indicator is measured at the cluster level. Overall, our core results might be slightly underestimating the contributions of assets and parental education, but the material inferences from those results are unchanged.

Lastly, we estimated models that excluded indicators of healthcare and demographic variables, on grounds that these might be potentially endogenous; for example, they might primarily be a function of assets and education. We find that excluding health and fertility variables leads to increases in the coefficients on the asset index and maternal education variable but that the magnitudes remain comparable across countries. If we exclude the health and demographic variables, the predicted changes in asset accumulation and parental education account for a little less than 40% of the actual change in HAZ across all countries.

4. Extension: Odisha

Because our DHS data for Odisha are only available for 1993/1994 and 2006, we cannot capture the drivers of the rapid fall in stunting that occurred after 2006. However, the 2011 India Census and District Level Health Survey contain data on state level stunting rates and state level means for many of the factors comparable to those used in the previous section. Small adjustments to the DHS indicators allow us to create a subset of indicators that are comparable across these different sources of data, which effectively allows us to make an out-of-sample prediction about the sources of nutritional change in more recent years.

Because only information on stunting, not HAZ, is available, our first step is to re-estimate Eq. (1) as a linear probability regression model with stunting is the dependent variable. We modify our regressors to match the variables in the 2011 census report. This entails replacing the asset index with access to electricity and ownership of a motorbike and TV, years of parental schooling with dummies for reaching secondary education or more, four or more antenatal visits with three or more antenatal visits, and village-level open defecation. All other variables remain the same. After estimating this regression we use the parameter estimates to conduct an out-of-sample decomposition analysis based on Eq. (2), using the earliest year as 1993 DHS round and replacing the most recent year (2006) with means from the 2011 census instead. Reflecting the recent changes in means allows us to make inferences about what has driven more recent rapid changes in stunting in Odisha.

Descriptive statistics are reported in Table 4 and the results of estimating an adjusted specification of Eq. (1) are found in Table 5. These are comparable to the previous model, with motorbike — a proxy for assets — and parental education having a negative and statistically significant association with stunting. The parameter estimates on healthcare and open defecation, which were previously not statistically significant or significant at 10% level, are more precisely estimated.

Fig. 2 summarizes the results of the decomposition analysis. We find that healthcare is the strongest driver of change, explaining more than 45% of the actual change in stunting rates. This is followed by parental schooling which was the strongest driver when using 1993–2006 data. Sanitation and water are modest contributors, as is expected by the fact that open defecation is still at 77% in Odisha. Overall, the extension analysis reveals that improvements in healthcare access has been the main contributor to the recent rapid reductions in stunting.

Table 5
Pooled regression using stunting as the dependent variable: Odisha (India).

Variables	Odisha, India Stunting (=1 if stunted)
Households with motorbike	−0.0942***
	(0.0350)
Households with TV	−0.0246
	(0.0303)
Households with electricity	−0.000317
	(0.0274)
Mother with secondary education or more	−0.103***
	(0.0306)
Father with secondary education or more	−0.0596**
	(0.0235)
Households with no toilet	0.0686*
	(0.0351)
Households with piped drinking water	−0.00911
	(0.0335)
3+ antenatal visits during pregnancy	−0.0665***
	(0.0212)
Children born in medical facility	−0.0620**
	(0.0283)
Total number of children ever born	0.00201
	(0.00709)
Constant	0.368***
	(0.110)
Observations	2,685
R-squared	0.175

Note: Village-level clustered standard errors in parentheses. *, **, and *** indicate significance at 10%, 5%, and 1% levels, respectively. The model includes a number of time-variant controls, including regional fixed effects, year fixed effects, month-specific child age dummy variables, and dummy variables for various categories of maternal age.

Table 4
Changes in mean values of stunting prevalence and drivers of change in child stunting between earliest DHS survey and 2011 India Census: Odisha (India).
Source: Author's calculations from 1993 Demographic and Health Surveys and 2011 India Census

Year	Stunting prevalence (%)	Households with electricity (%)	Households with TV (%)	Households with motorbike (%)	Mother with secondary educ. or higher (%)	Father with secondary educ. or higher (%)	3 or more antenatal visits (%)	Born in medical facility (%)	Open defecation (%)	Piped water (%)	Number of children
	Odisha, India										
1993	51.9%	34.9%	13.5%	5.5%	18.9%	38.6%	38.4%	14.6%	90.5%	8.5%	3.1
2011	38.0%	43.0%	26.7%	14.5%	40.3%	47.8%	78.5%	77.7%	76.6%	13.8%	2.4
Change	−13.9%	8.1%	13.2%	9.0%	21.4%	9.2%	40.1%	63.1%	−13.9%	5.3%	−0.64
%Change	−26.8%	23.1%	97.5%	162.4%	112.8%	23.7%	104.5%	321.6%	−11.2%	4.8%	−20.9%

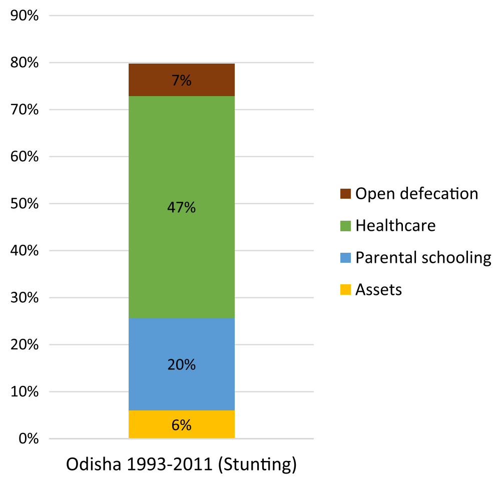

Fig. 2. Estimated contributions of selected factors to changes in stunting: Odisha (India). *Notes:* These are estimates changes in HAZ scores due to changes in the explanatory variables listed in the legend over 1993–2011. These estimates (or retrospective predictions) are based on a linear decomposition at means, in which changes in the mean of each explanatory variables are multiplied by the corresponding coefficient from Table 5.
Source: Authors' estimates from 1993 DHS data and 2011 India Census data.

5. Discussion

In this paper, we systematically quantify and compare the factors that might explain long-term reductions in child undernutrition in Bangladesh, Nepal, Odisha, Ethiopia, Senegal and Zambia. We use regression analyses to estimate marginal effects of different underlying determinants on child HAZ scores and use these and their historical changes in means of these determinants to account for changes in HAZ over time in each country. We find commonalities and differences across countries.

Consistent with previous studies (Headey et al., 2015, 2016; Headey and Hoddinott, 2015), we find that asset accumulation and parental education are important predictors of nutritional improvement in most countries, with Zambia the main exception. While many studies focus more on macroeconomic analyses of the impacts of economic growth on nutrition – often finding only moderate impacts – changes in GDP per capita are widely acknowledged to be an imperfect indicator of economic development both for conceptual and measurement reasons. The household asset index, in contrast, seems a robust predictor of HAZ outcomes, despite some of the criticisms levelled at this indicator (Filmer and Scott, 2012; Harttgen et al., 2013). There is generally more acknowledgement that parental education is important for child nutrition outcomes, and indeed the effects uncovered in this paper are broadly consistent with those estimated by Alderman and Headey (2014).

There is much greater variation in the roles of other determinants across the six locations. We generally find that sanitation has significant associations with HAZ, but the effect sizes are relatively modest in most instances, and large but imprecisely estimated in Odisha. In contrast, piped water seems to be of no direct importance to HAZ. Improvements in maternal height and reductions in fertility rates explain modest improvements in HAZ, but changes in access to health care are more complex. Antenatal and neonatal care are strong predictors of nutritional improvements in Nepal, Odisha and Senegal. However, in Zambia, where malaria is endemic, the rapid expansion of bednets is a very strong predictor of HAZ change. Strikingly, although there are hypothesized nutritional impacts of malaria during pregnancy (Eisele et al., 2012; Guyatt and Snow, 2004) and early childhood (Nyakeriga et al., 2004), there is very little research investigating these impacts. It might also be hypothesized that expansions in the treatment of HIV/AIDS in places like Zambia, where the disease is widespread, are also driving some improvement in nutrition. While variables directly representing the treatment of HIV/AIDS are not available in the DHS data, as a robustness check we added an indicator of maternal underweight prevalence as a proxy, which might also reflect improved food security or women's empowerment. We find a negative and significant association between maternal underweight prevalence and HAZ in Zambia. We therefore flag linkages between malaria, HIV/AIDS and nutrition as an important issue for future research in this part of the world.

Our study has limitations. First, our data are observational and we only use them to assess how accurately changes in potential determinants of nutrition *account* for nutritional change over time; how close this kind of accounting process comes to a causal process is highly uncertain. Second, our models vary substantially in their ability to account for aggregate HAZ change over time. The models for South Asian countries and Senegal account for at least half of the observed changes in HAZ over time, but the models for Zambia and Ethiopia perform relatively poorly, suggesting mis-measurement, misspecification and/or omitted variables are problems in these two countries. Third, potential drivers of nutritional change, such as changes in women's empowerment, are not captured in our data. Fourth, recent work by Semba et al. (2016a, 2016b) has suggested that an absence of essential amino acids and choline (both found in animal source foods) may play a role in limiting linear growth. For this reason, it would be helpful to complement our work with an analysis of the determinants of children's diets. Unfortunately, these data are absent from many of the DHS surveys we use (particularly in early rounds); this represents an important topic for future work. Fifth, past performance is no guarantee of future returns. Specific variables that have driven past improvements may play a lesser role in the future, either because the effects of some factors are non-linear, or because of scope for saturation. For example, open defecation in Bangladesh in 2014 was just 2.9%, implying sanitation upgrading is more important for the future. Lastly, we have not tried to explain why our explanatory factors have changed over time, particularly the role of policies and programs (and their impact on nutrition related services, another datum not available in most DHS surveys) in driving changes in nutrition-sensitive factors. That is the subject of the more qualitative research studies included in this special issue, and hence our paper should be read as a complement to those more in-depth case studies.

Set against these weaknesses are several study strengths, most notably the application of the same statistical techniques to national level data from a common data source from which we have extracted a set of consistently measured explanatory variables. Furthermore, our results are robust to a variety of checks, all of which consistently point to some common and very plausible drivers of nutritional change in six country(state) studies. The significant associations between these various explanatory factors and child HAZ confirm the multidimensional nature of nutritional change and the important role of nutrition-sensitive sectors, including broader economic development, education, WASH, health and family planning. The statistical results imply that rapid nutritional change at a national level requires substantial progress in most if not all of these sectors, along with nutrition-specific interventions that are clearly important (Black et al., 2013), but not measured in our model.

Acknowledgements

We thank Stuart Gillespie for encouraging us to work on this topic and Kenda Cunningham and Jody Harris for helpful comments on an earlier draft. This work has partly been supported by the Department for International Development through its funding of the Transform Nutrition Consortium. We also thank three anonymous reviewers.

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