ISER Working Paper Series

Parental Education, Gender Preferences and Child Nutritional Status: Evidence from Four Developing Countries

Rafael Novella
Institute for Social and Economic Research
University of Essex

No. 2013-06
May 2013

INSTITUTE FOR SOCIAL
& ECONOMIC RESEARCH

E·S·R·C
ECONOMIC
& SOCIAL
RESEARCH
COUNCIL

www.iser.essex.ac.uk

Non-Technical Summary

This paper explores whether the allocation of household resources to boys' and girls' health vary with parental gender preferences. I use a model that allows me to disentangle the effect of parental characteristics on the technology of child health production from the effect of parental preferences. This relationship is explored using a unique dataset of young children in four developing countries: Ethiopia, India (Andhra Pradesh), Peru and Vietnam.

In this study, I assume that more education indicates more power in the intrahousehold resource allocation process, and thus, women who are better educated than their husbands should be able to impose their preferences and allocate more resources towards commodities they care more about.

This paper analyses the effect that maternal bargaining power in the household has on two indicators of child health that reflect short-run shocks (weight-for-length) and long-run shocks (length-for-age), after taking into account characteristics of the child, parents, household and genetic components. Moreover, the estimates take into account potential unobserved factors that might influence the presence of a young boy or girl in the household as well as the household formation.

The estimated effects of maternal power on child health vary across countries and the indicator chosen. In Peru and Vietnam (and partially in Ethiopia), maternal bargaining power has a general positive effect on children's health but it also has a differential effect, which suggests that mothers prefer to allocate more resources to their daughters in comparison to their sons. In contrast in the Indian sample, girls living with mothers with more power in the household receive fewer resources than girls in other households.

I find evidence that among households located in rural areas of India and Peru there seem to be differences in the allocation of resources between boys and girls. Similarly, some evidence of competition for household resources affecting girls' health is also found in the samples from Peru and Vietnam.

Parental Education, Gender Preferences and Child Nutritional Status: Evidence from Four Developing Countries*

Rafael Novella†

This version: March 2013

ABSTRACT

This paper examines whether the distribution of bargaining power between parents affects permanent and transitory nutritional indicators in the early stages of boys' and girls' life. I use the Young Lives sample, which is a survey of young children living in poor households in Ethiopia, India (Andhra Pradesh state), Peru and Vietnam. By adopting a methodology to disentangle gender differences produced by technology and preferences, I find evidence that the allocation of household resources varies with the gender of the child and the gender of the parents. After accounting for the potential endogeneity of the indicator of power distribution within the household, related to assortative mating in the marriage market, I find that maternal power has larger effects on girls' health than on boys' health in Peru and Vietnam. In contrast, in India, maternal bargaining power has a negative effect on girls' health, whereas in Ethiopia no differential effect is found. Further analysis confirms that differences in parental behaviour drive the estimated effects and that these are robust to the inclusion of genetic information.

KEYWORDS: intrahousehold allocations, nutrition, children, developing countries

JEL CODES: D13, I12, J13, O57

* I am indebted to John Ermisch, Patrick Nolen and Mark Bryan for their constant guidance and to Arjun Bedi and Adeline Delavande for their very helpful comments and suggestions. I would also like to thank participants at the COSME-FEDEA workshop in Gender Economics, 2010; the "Health, Happiness and Inequality" conference at the University of Darmsdadt, 2010; the International Association for Feminist Economics (IAFFE) conference, 2010; the "Women's Bargaining Power and Economic Development" workshop at the ISS-Erasmus University, 2010; and the Latin American and Caribbean Economic Association (LACEA) conference, 2011. The data used in this publication comes from Young Lives, a 15-year survey investigating the changing nature of childhood poverty in Ethiopia, India (Andhra Pradesh), Peru and Vietnam (www.younglives.org.uk). Young Lives is core-funded by UK aid from the Department for International Development (DFID) and co-funded from 2010 to 2014 by the Netherlands Ministry of Foreign Affairs. The views expressed here are those of the author. They are not necessarily those of Young Lives, the University of Oxford, DFID or other funders.

† ISER, University of Essex. E-mail: grnove@essex.ac.uk; rnovella@iadb.org

1. Introduction

The positive effects of adequate nutrition during early childhood on children's physical and cognitive development and its long-lasting consequences on wages, economic growth and welfare have been largely found in the empirical literature (Barker, 1990; Miguel and Kremer, 2004; Schultz, 2005; Strauss and Thomas, 1998). Particularly in developing countries, socioeconomic and cultural factors place children at higher risk of growth retardation and malnutrition due to higher exposure to infections (mainly, diarrhoeal diseases) and inadequate nutrition (Martorell and Habicht, 1986; WHO, 1986). However, boys and girls are not affected in the same manner by growth retardation and malnutrition, which might be partially explained by gender biases in the allocation of household resources that contribute to gender gaps in later outcomes.

The primary objective of this paper is to analyse whether the household allocation of resources in early stages of children's lives is affected by parental gender preferences and the distribution of bargaining power within the household. I focus on allocations directly related to the health of children aged between 6 and 18 months, measured by two well-known health indicators: height-for-age and weight-for-height.[1] Because environmental factors are particularly important determinants of child health in early childhood, it is reasonable to suspect that a child's anthropometric outcomes are determined by indicators of long-run resource availability within the household, such as parental education, and the distribution of power between parents. As a second objective, I explore whether differences in health indicators between boys and girls attributable to differences in parental gender

[1] For children aged 0–24 months, the World Health Organization (WHO) corresponding measures are 'length-for-age' and 'weight-for-length', where recumbent length instead of standing height measures are considered. Henceforth, I will refer to these two indicators.

preferences are constant across countries with different cultural and economic backgrounds.

In developing countries, the interaction of high prevalence of infectious diseases, limited availability of food and other resources, and the relatively large importance of traditional beliefs and customs provide a particularly interesting setting to study the impact of parental gender preferences on child health. The Young Lives (YL) household survey is a particularly interesting sample for the goals of this study. YL surveyed mainly poor households with children aged between 6 and 18 months, in four developing countries with substantially different socioeconomic and cultural backgrounds: Ethiopia, India (Andhra Pradesh state), Peru and Vietnam.[2]

Measuring how much of the gender difference in child health outcomes is attributable to parental gender preferences is not straightforward. First, observed differences of health outcomes between boys and girls might be the result of differences in the impact of parental characteristics on the technology of child rearing (e.g. it might be more efficient for mother-daughter and father-son to spend more time together) and in parental gender preferences. Second, a large number of studies examining boy/girl discrimination fail to control for the potential endogeneity of child gender. If households follow the son-biased stopping rule (i.e. the likelihood to stop having children after giving birth to a boy is larger than after giving birth to a girl), after controlling for household size, girls are likely to live in households with higher preferences for girls in comparison to households where boys live (Barcellos et al., 2012; Yamaguchi, 1989). Third, the use of cross-section data might confound the effect of bargaining power and the determinants of household formation when

[2] It is worth mentioning that the sampling design followed by the Young Lives team and used in this paper, does not aim to be representative at the national level. In general, the Young Lives sample over-represents households living in poor conditions. From now on, I will refer to the Andhra Pradesh sample as India.

estimating the intrahousehold allocation of resources to children (Lundberg, 1988). For instance, a larger effect of maternal bargaining power on girls' health might reflect that, an increase in maternal bargaining power allow women to allocate more resources to girls, or, that men with non-traditional (unobserved) gender roles tend to marry more powerful women.

Differences between the health outcomes of boys and girls might reflect differences in the technology of child rearing and/or differences in preferences between parents. The methodological strategy for disentangling this is by considering a household production function of health and a collective model for the household decision process. Any (distribution) factor affecting the ability of household members to reach her or his preferences, and which is not in the production function of health, will reflect preferences rather than technology (Browning et al., 1994; Thomas, 1994). This study uses parental relative education (whether mothers have more years of formal education than their husbands) as an indicator of power distribution in the household. If maternal power is higher, the probability that she is able to assert her own set of preferences is higher, which would finally lead to an allocation of resources towards the commodities she cares more about. Growing empirical evidence points out that factors modifying households' power distribution in favour of women, are associated with larger improvements in child health and increases in childcare expenditure (Duflo, 2000 and 2003; Hoddinott and Haddad, 1995; Lundberg et al., 1997; Reggio, 2011; Thomas, 1990 and 1994).[3]

After testing for the random assignment of boys and girls to households and the presence of the son-biased stopping rule in the YL sample, the empirical strategy of this study consists of estimating two indicators of child health that reflect short-run

[3] Blundell et al. (2005) show that increasing maternal power improve child welfare when mothers' willingness to pay for child goods is more responsive to changes in resources than fathers', and not because mothers have a larger willingness to pay for child goods from their resources.

status (weight-for-length z-scores) and long-run status (length-for-age z-scores), controlling for child, parent (including genetic components), household and community characteristics. Moreover, in contrast to most of the literature using relative education as an exogenous measure of maternal power, this study tries to deal with its potential endogeneity, which is associated with the assortative mating of parents in the marriage market. To do this, I include in the regressions, a large set of observable controls directly related to child health and also a set of variables likely to be correlated with the household formation. I also explore whether the differential effect of maternal bargaining power on girls and boys varies by urban/rural location and by the presence of other young children in the household. Finally, robustness checks are shown to test whether the estimates are driven by behavioural differences between parents and other observable characteristics (paternal height).

Results from the empirical analysis show that mothers prefer to allocate more resources to improve their daughters' health in Peru and Vietnam. In contrast, in India, maternal bargaining power has a negative effect on girls' health, particularly on the short-run indicator. No evidence of significant differential effect of maternal bargaining power on boys versus girls is found in the sample from Ethiopia. Having found that maternal preferences affects child health, breastfeeding emerges as a mechanism through which the effects might occur.

This study contributes to the empirical literature concerning the effect of differences in parental behaviour related to the gender of the child in developing countries. In particular, some contrasting evidence is found with respect to the previous findings of the literature in the YL countries. When evaluating two social programmes in Ethiopia, Quisumbing (2003) finds that the impact on child nutritional status depends on child gender and the type of aid, and that increases in maternal bargaining power would lead to larger investments in boys. Gertler and Glewwe (1992) found evidence in Peru, of parents placing more value on sending boys to secondary education than on sending girls. In the case of Vietnam, little evidence of differences on investment related to the gender of the child is found (Duc et al., 2008; Haughton and Haughton, 1997). On the other hand, this study finds similar results to the previous literature which finds that girls in India are disadvantaged with respect to boys, in terms of less investment in health inputs and outcomes (Barcellos et al., 2012; Jayachandran and Kuziemko, 2011; Subramanian and Deaton, 1991).

Additionally, this study contributes to better designs of public policies providing new insights on intrahousehold dynamics. In developing countries, the study of intrahousehold dynamics and determinants of the resource allocation process is particularly relevant since children are particularly vulnerable and households are, among other factors, more likely to face credit constraints. When capital markets are imperfect, even altruistic parents might have to sacrifice investments in children human capital. In particular, this might occur when differences in expected labour market returns to nutrient investments in boys and girls exist (Behrman, 1988), or the parental control over future returns is imperfect (Parson and Goldin, 1989). Even slight differences in the distribution of scarce resources might have dramatic consequences on children's health and nutritional status. If gender matters in intrahousehold allocation of resources, policies aimed at increasing children's wellbeing, such as conditional or unconditional cash and/or in-kind transfer programmes,[4] need to account, also, for the gender of the parent receiving the

[4] In 2002 when the YL survey was collected, none of the four countries had a Conditional Cash Transfer (CCT) programme implemented, but presently Peru and India (Andhra Pradesh state) counts with them. Recently, in 2008, India implemented in Andhra Pradesh and six other states, a CCT call "Dhanalakshmi", which aims to increase the nutritional and educational levels of girls giving monetary incentives to the households that fulfil conditions on these areas and where girls remain unmarried until the age of 18 years. This policy is particularly targeted to vulnerable populations from the SC/ST

benefits, his/her power to influence the household's decision and the gender of the child targeted.

The paper is structured as follows; Section 2 discusses the theoretical framework. Section 3 describes some relevant indicators for each YL country and describes the YL data. Section 4 discusses the presence of the son-biased stopping rule in the sample and presents the empirical strategy to estimate gender biases in the household resource allocation process. Section 5 shows the main results and Section 6 shows further robustness checks. Finally, Section 7 concludes.

2. Theoretical Framework

Differences in health outcomes between girls and boys may reflect differences in the technology underlying the health production function, as well as differences in parental preferences in favour of boys or girls. To disentangle these two channels, it is necessary to depart from the traditional unitary model and consider an intrahousehold model.

This paper is based on the collective model suggested by Chiappori (1988 and 1992). In contrast to the unitary model (Becker, 1991), the collective model neither assumes identical parental preferences, nor their representation into a single household utility function. Furthermore, the collective model's main assumptions are the existence of a stable decision process, *whatever its true nature*,[5] which leads to Pareto-efficient allocations (Browning et al., 2011). Each allocation in the Pareto frontier corresponds to a different decision process involving different sets of individuals' (Pareto) weights. These weights summarize the intrahousehold decision

castes. The "Indira Gandhi Matritva Sahyog Yojana (IGMSY)" is another intervention launched in 2011 and targeted at improving the health and nutritional status of pregnant, lactating women and infants. Since 2005 the CCT "Juntos" in Peru aims at tackling chronic malnutrition and extreme poverty of rural households with pregnant women and/or children.

[5] Chiappori (1997). For instance, the decision process may be characterized by a bargaining model (Nash, Kalai-Smorodinsky). However, this is not important for the general collective model as long as its solution leads to a particular efficient outcome on the Pareto frontier.

process. Browning et al. (2011) argue that the Pareto weights have a natural interpretation in terms of decision powers. For instance, an increase in a wife's weight results in a move along the Pareto set in the direction of higher utility for her (and lower utility for her husband). In a pure economic sense, then, a larger weight would correspond to more power and better outcomes for the wife. As Blundell et al. (2005) argue, a main goal of the collective model is to analyse how these weights (and therefore the decision process) respond to changes in prices, incomes and other exogenous factors *(distribution factors)* and how these responses further affect household allocations.

Although the collective model has been widely used for analysing how changes in intrahousehold decision power affect household allocations, relatively few applications include home production and publicly-consumed commodities (for instance, children in the household). Chiappori (1997) and Apps and Rees (1997) include home production of marketable and non-marketable goods in their earlier works on collective labour supply. However, they consider the case of privately-consumed home-produced goods, which is not applicable to child welfare. Child welfare can be easily considered public consumption for parents in the household; it is expected to increase, possibly at different rates, both parents' utilities. Recent studies (Blundell et al., 2005 and 2007; Chiappori and Ekeland, 2009; Ermisch, 2003) discuss collective models including home-produced publicly-consumed commodities, such as child welfare. Their main conclusions point out the importance of having exogenous factors (*distribution factors* as defined below) affecting Pareto weights, but not individuals' preferences; and, of using time-use data.

Given that the theoretical debate about collective models with home production of public goods is still open, most empirical studies have focused on responses of adults' labour supply and demand of private goods to changes in intrahousehold decision power. Few papers (Duflo, 2000 and 2003; Haddad and Hoddinott, 1994; Thomas 1990 and 1994) explore effects on child health, particularly on anthropometric outcomes. These papers find evidence of an unequal distribution of household resources to boys' and girls' health, measured through anthropometric outcomes, and that these differences vary with the gender of the adult. Among these studies, Thomas (1990 and 1994) suggests an empirical application of a collective model including home production of child health that might be applied to the data used in this study.[6]

The theoretical model used in this study follows Ermisch (2003) and Thomas' (1990 and 1994) models, in which a reduced form for the child health demand function (solution to the household's optimization) is obtained from combining a health production function with a collective model of the household decision-making process. First, let θ_i be the health of child *i*. The biological health production function for each child is modelled as a function of a set of inputs I (such as nutrient intakes),[7] which in turn depends on purchased goods x_c, mother's time t_m, and father's time t_f for childcare, conditional on observable health-relevant characteristics of the child μ_i (such as age and gender), the parents μ_j, where *j*=*m(*mother*), f(*father*),* (such as education and health/genetics), the household μ_h (such as demographic composition, access to water and sewerage), and the community μ_v (such as urban/rural location). η_i represents individual unobserved heterogeneity in health (or unobservable

[6] Duflo (2000 and 2003) exploits a natural experiment in the pensions system in South Africa to evaluate how changes in intrahousehold decision power affect child health, while Haddad and Hoddinott (1994) estimate a non-cooperative model of the household.

[7] A *biological health production function* relates an individual's health to his/her consumption of nutrients, which in turn depend on his/her consumption of foods (this may be considered a nutrient production function, as in Behrman and Deolalikar, 1990), other health-related inputs (such as, non-food health inputs, household resources, observable health-relevant personal characteristics), and exogenous health endowments of the individual and the environment where he/she lives (Pitt and Rosenzweig, 1985).

endowments), part of which might be common across individuals within a household and community.[8]

$$\theta_i = \theta_i\left[I(x_c), t_m, t_f; \mu_i, \mu_j, \mu_h, \mu_v, \eta_i\right] \quad (1)$$

As Thomas (1994) points out, the technology underlying the health production function is likely to be different between boys and girls.[9] For instance, it may be the case that it is more efficient that mother spend more time with girls and fathers with boys. Although unlikely because of the young age of children in the sample (6-18 months), differential impact of the maternal and paternal characteristics, t_j and μ_j in (1), on boys' and girls' health may reflect different technologies.

Second, I now turn to discuss differences due to preferences and the intrahousehold model of resource allocation. In addition to the basic assumptions of the collective model presented at the beginning of this section, I further assume that preferences of household members are "caring" (i.e. members care for each other's utility, but do not care how, in terms of individual private and public consumption, a given level of utility is obtained) and parents are the only members taking decisions within the household (i.e. children are not decision-makers in the household and their preferences enter into the household utility only through their parents' utility).[10] Thus, assume each parent *j*, in household *h* and community *v* (both subscripts omitted for simplicity), has a utility function given by:

[8] The function $\theta(\cdot)$ exhibits constant return to scale in x_c, t_m and t_f (Ermisch, 2003). That is, the level of production of the public good θ is determined by preferences and the decision process; and the time allocation of domestic work between parents depends on technology (Blundell et al., 2005).

[9] It may also vary with the age of the child or with other socio-demographic characteristics.

[10] Recent studies (Dauphin et al., 2011; Hao et al., 2008; Lundberg et al., 2009) incorporate older children and adolescents as decision-makers in the household decision process. Even though the literature is not conclusive about the starting age for becoming a decision maker, given the age range (6-18 months) of children in the YL sample, it is plausible to assume that children are not able to directly influence the allocation of resources decided by their parents.

$$U_j(x_j, \theta) \quad (2)$$

where x_j represents parent j's private consumption of goods and θ is a vector of home-produced child health, for all children in the household. θ is a public good for the parents, which increases their utility, possibly at different rates and according to each parent's preferences.

Following Browning and Chiappori (1998), I assume that the cooperative household's efficient allocations on the Pareto frontier correspond to the household maximization of a weighted sum of each parent's utility:

$$W = (1 - \lambda)U_f(x_f, \theta) + \lambda U_m(x_m, \theta) \quad (3)$$

where $\lambda \geq 0$ is the Pareto weight that summarizes the intrahousehold decision, and indicates the location on the Pareto frontier and the power of the mother in the household.[11] In general, this weight is a function of prices, each parent's wage s_j and non-labour income y_j, as well as other environmental characteristics or distribution factors z, that might affect parent j's ability to assert her/his preferences in the household allocation process. In particular, the distribution factors are variables affecting household behaviour only through their impact on the decision process (weights), but neither affects preferences nor budget constraints.[12] As Browning et al. (2011) mentions, while changes in prices or income potentially change weights, they

[11] In a similar representation, an efficient allocation must maximize the utility of one parent $U_f(x_f, \theta)$ subject to the other achieving at least a given utility $U_m(x_m, \theta) \geq U_m^*$, to a budget constraint and to a production function for the home-produced good. In this case, λ corresponds to the Lagrange multiplier for the first constraint, the "efficiency constraint".

[12] The distribution factors can be easily related to the threat points (or reservation utilities) in bargaining household models. A distribution factor can be any variable that, in the bargaining models setting, potentially affects individuals' threat points, but not the household's budget constraint, and may affect the intrahousehold power distribution and household behaviour. Bargaining models also provide a clear idea about the direction of these effects. A change in a variable that increases the wife's threat point should always positively affect her Pareto weight (Browning et al., 2011).

also change the set of efficient allocations. On the contrary, changes in the distribution factors are informative about the intrahousehold decision process because they only influence weights and consequently, outcomes on the same Pareto frontier.

The household welfare function (3) is optimized subject to the health production function (1); to the following budget constraint, where p_c is the price of goods used in the production of θ (the price of parents-goods set to one), and T represents each parent's total time available; to $0 \leq t_j \leq T$; and, to $x_c \geq 0$:

$$\sum_{j=m,f}(x_j) + p_c x_c = \sum_{j=m,f}(T - t_j)s_j + \sum_{j=m,f}(y_j) \quad (4)$$

From this optimization process, household demands for parental goods x_j; and, for inputs x_c, and parental time t_j used in the production of child health θ, are obtained.[13] This study is particularly interested in one dimension of vector θ; anthropometric measurements h. Substituting these derived demands on the production function (1), anthropometric measurements h are obtained as a function of prices, wage rates, non-labour incomes, the observed health endowments in (1) (of the child, parents, household and community), the distribution factors z included into the Pareto weights, as well as unobserved heterogeneity ξ_i at the child level which in part may be common across individuals within a household and community:

$$h_i = h_i(p_c, s_j, y_j, \mu_i, \mu_j, \mu_h, \mu_v, z, \xi_i) \quad (5)$$

It is worth noticing that since the distribution factors z are included in the health demand function and not in the health production function (1), a differential effect of it on the health of boys and girls must reflect parental preferences and not differences

[13] As Ermisch (2003) mentions, this setting assumes *separability* between parental choices of their own private and public goods consumption and their choices of inputs x_c, and t_j in the production of θ.

in the technology of child rearing. This study uses parental relative education (whether mothers have more years of formal education than their husbands) as a distribution factor *z*.[14] I expect this indicator to affect household behaviour only through the weights associated with each parent's utility and neither through preferences nor budget constraint.

In collective models with consumption of a home-produced private good (Chiappori, 1997) or public good (Blundell et al., 2005), the existence of at least one distribution factor is necessary for identifying preferences and Pareto weights. If the home-produced good is publicly consumed (such as child welfare), Blundell et al. (2005) mention the importance of having time use data. In this sense, the absence of time use data imposes restrictions to the identification of the effect of parental time to child health production in this study. The effect of parental time allocation to home and market production is captured by parental education. However, the main interest of this study is exploring whether changes in a distribution factor (parental relative education as discussed below) affects the intrahousehold decision process and therefore, the household resource allocation to observed indicators of child health.

3. The Young Lives Countries

3.1 Socio-Economic Context

Using external data sources, this section describes relevant indicators of household wealth and child well-being for the four YL countries in the year 2000, just before children in the sample were born.[15]

In 2000, the United Nations' Human Development Index reports that out of 174 countries, the YL countries were ranked from the poorest to the least poor, as follows:

---

[14] Similar indicators have been previously used as distribution factors in the literature (Thomas, 1994; Beegle et al., 2001; Schady and Rosero, 2007; Gitter and Barham, 2008; among others).

[15] In the Appendix, Table A.1 shows the indicators mentioned in this section and Section A.2 gives more detailed information about the YL corresponding to each country.

Ethiopia 171; India 128; Vietnam 108, and; Peru 80.A similar pattern is observed when considering extreme poverty, measured in monetary terms. That is, the proportion of population living under one dollar per day, expressed in purchasing power parity exchange rate, was: 35 in India, 26 in Ethiopia and 18 in Peru and Vietnam (World Bank, 2004). According to the World Health Organization (WHO), the average life expectancy at birth follows a similar pattern: 54 years in Ethiopia, 61 in India, 69 in Peru and 70 in Vietnam.[16]

Marked disparities between urban and rural areas are a constant characteristic across the four countries. Households in rural areas have more precarious access to the basic infrastructure directly related to a hygienic home environment, such as drinking water, and sanitation.

The YL countries also differ in relation to child health and nutritional status. For children younger than 2 years, the proportion of chronic malnourished children increases with age in the four countries. The proportion of chronic and acute malnourished children is higher in Ethiopia and India than in Peru and Vietnam. In particular, in terms of acute malnutrition, India presents the highest prevalence, whereas Peru the lowest among the four. In Peru, less than 3 percent of children suffer acute malnutrition.[17] Finally, infant mortality rates are higher in Ethiopia and India than in Peru and Vietnam. Regarding gender differences in infant mortality rate, Srinivasan and Bedi (2011) show evidence that, for several biological factors, the mortality rate for girls is expected to be about 80 percent of the one for boys. Using this parameter as a reference point, evidence from this study suggests that there are no

[16] WHO's Statistical Information System (WHOSIS).
[17] WHO's Global Database on Child Growth and Malnutrition.

gender differences in infant mortality rate in Ethiopia. However, in India, Peru and Vietnam, the mortality rate for girls exceeds that of boys.[18]

Due to the absence of comparable statistics, the analysis in this section shows indicators for India as a whole and not for the Andhra Pradesh state in particular. However, it is worth mentioning, that in 1999–2000, Andhra Pradesh had the fourth lowest rural poverty and the fifth lowest urban poverty rates in India (Dev and Ravi, 2003). Despite Andhra Pradesh being one of the Indian states achieving a significant reduction in income poverty in rural areas, the difference in the proportion of underweight children between rural and urban areas is still significant: 41 and 29 percents respectively.

3.2 The Young Lives Data

Young Lives (YL) is an innovative longitudinal research project aimed at improving the knowledge of the causes and consequences of childhood poverty. YL tracks two cohorts of around 12,000 children in Ethiopia, Andhra Pradesh, Peru, and Vietnam over 15 years.[19] In this study I use only the sample corresponding to the 'younger cohort' of about 2,000 children in each country, aged around 1 year in the first round collected in 2002. The YL questionnaire contains detailed information for only one child in the relevant age group per household and additionally it contains household characteristics. YL oversampled poor households distributed in twenty non-randomly selected 'sentinel sites' in each country, which were chosen according to their poverty levels.[20] Within each sentinel site, 100 households with a child aged between 6 and 18

[18] WHOSIS. The mortality rate gap between girls and boys increases over time. Mortality rate for girls under 5 years is 97 per 1000 live births against 86 for boys.
[19] Information extracted from the Young Lives website: http://www.younglives.org.uk.
[20] According to the YL project, a 'sentinel site' corresponds to a geographical area where it is possible to collect individual, household, community, regional and national characteristics. 'Sentinel sites' are in addition, important for YL complementary thematic studies.

months were randomly chosen.[21] While the YL sample is not meant to monitor average child outcomes at national level, it is still a valid instrument for analysing child welfare in each of the countries.[22]

The sample considered in this study includes children of the four YL countries aged 6-18 months in round 1 (oldest children in the sample are 17 months old). The original YL sample has been restricted to better fit the objectives of this study, according to the following criteria. First, I restrict the sample to households formed by two adults responsible for the child, in particular I consider households where the two biological parents are present.[23] Second, to minimize the presence of measurement error in variables related to child characteristics at birth and maternal anthropometrics, I further restrict the sample to only those households where the biological mother is the main caregiver. Fortunately, most sections of the YL questionnaire are directed at the primary child caregiver, who is in most of the cases the biological mother.[24]

Even though this study uses cross-sectional data from the first round of the YL sample, some time-invariant variables were imputed from the YL round 2 (2006) and 3 (2009). In the particular case of Vietnam, the first YL round did not collect the years of education achieved by adults in the household, which complicates

[21] In contrast to this sampling procedure, known as ‘sentinel site surveillance system’, the YL team in Peru used a multistage, cluster-stratified and random-sampling procedure to select the two cohorts of children, enabling the randomization of households and of sentinel sites, and selecting sentinel sites with a pro-poor bias (Escobal and Flores, 2008).

[22] For further details about comparisons between estimates using the YL sample and national representative surveys see the YL technical notes in: http://www.younglives.org.uk/our-publications/technical-notes/. Escobal and Flores (2008) argue that certain methodologies, as the one called ‘raking’, could help reduce the biases due to sample design between the YL sample and national representative surveys for Peru. According to the authors, this might be extendable to the other YL samples.

[23] Restricting the sample to households where the two biological parents are present is not strictly necessary, but it rules out additional channels through which the biological relationship with the decision-takers within the household may affect investments in child human capital. In addition, the proportion of households where at least one of the two adults deciding about the child is not a biological parent, represent less than 1 percent of the whole sample for the four countries.

[24] In Ethiopia, the primary caregiver in 96.6 percent of all the households is the biological mother; 99.4 percent in Andhra Pradesh; 98.8 percent in Peru; and, 98.7 percent in Vietnam.

comparability across countries and makes less precise the measure of parental bargaining power. For only those parents who in the first round declared not being students, I impute the years of education collected in the YL second round. Some other time-invariant parental characteristics were imputed from the second and third to the first round data, such as: mother tongue; ethnic or caste group; indicators of migration status;[25] age at the first round; and, maternal and paternal height.[26]

Further reductions in sample size due to missing values reduce the samples to: 1,247 children/households in Ethiopia, 1,598 in Andhra Pradesh, 1,495 in Peru and 1,472 in Vietnam. For each of the four YL countries, Table 1 shows summary statistics of some relevant characteristics of the child, parents and household.

Table 1 shows two definitions of parental education. The first corresponds to dummy indicators for literacy, i.e. whether the parent has completed at least one year of formal education. The second one is a continuous variable corresponding to the number of years of formal education achieved. Two main patterns are shown. First, mothers are less educated than fathers and second, the education achievement of both parents is particularly low in Ethiopia and India where on average, parents do not complete the minimum years of education required for achieving primary education. In particular for these countries, the literacy indicator is more informative.

A dummy variable indicator for whether mothers have more education than their partners is included in Table 1, aiming at capturing maternal bargaining power within the household. In Ethiopia, about 13 percent of the mothers are more educated than their husbands; 14 percent in India; 24 percent in Peru and about 25 percent in

[25] The indicators of migration status corresponds to dummy variables (one for fathers and one for mothers) indicating whether the parent has lived all his/her life in the community surveyed.

[26] Peru was the only country collecting maternal height in the YL first round, but this information is available for the other three countries in round 2. In contrast, paternal height was only collected in Peru in the YL second and third rounds.

Vietnam. When mothers are more educated than fathers, they have -on average- 3 more years of schooling in Ethiopia, Peru and Vietnam and 5 more years in India.

Table 1: Summary Statistics YL 2002

	Ethiopia		India		Peru		Vietnam	
Child nutritional indicators								
Wasted	0.21	(0.01)	0.28	(0.01)	0.02	(0.00)	0.09	(0.01)
Stunted	0.17	(0.01)	0.10	(0.01)	0.06	(0.01)	0.03	(0.00)
Child characteristics								
Child age (in months)	11.53	(0.10)	11.69	(0.09)	11.52	(0.09)	11.67	(0.08)
Child sex (Male = 1)	0.53	(0.01)	0.54	(0.01)	0.51	(0.01)	0.51	(0.01)
Weeks of premature birth	0.18	(0.03)	0.19	(0.02)	0.57	(0.03)	0.28	(0.03)
Child ethnic group 1[+]	0.24	(0.01)	0.19	(0.01)	0.06	(0.01)	0.84	(0.01)
Child ethnic group 2[+]	0.21	(0.01)	0.16	(0.01)	0.02	(0.00)	0.06	(0.01)
Child ethnic group 3[+]	0.21	(0.01)	0.47	(0.01)	-	-	-	-
Birth order	3.38	(0.06)	1.69	(0.02)	2.48	(0.04)	1.87	(0.03)
Parental characteristics								
Maternal literacy	0.41	(0.01)	0.35	(0.01)	0.92	(0.01)	0.88	(0.01)
Paternal literacy	0.54	(0.01)	0.49	(0.01)	0.98	(0.00)	0.92	(0.01)
Maternal years of education	2.37	(0.10)	2.73	(0.10)	7.87	(0.11)	6.62	(0.10)
Paternal years of education	3.69	(0.12)	4.20	(0.12)	9.02	(0.10)	7.35	(0.10)
Maternal ed. > Paternal ed.	0.13	(0.01)	0.14	(0.01)	0.24	(0.01)	0.25	(0.01)
Years of more ed. mother	3.25	(0.16)	5.06	(0.19)	2.84	(0.10)	3.14	(0.10)
Maternal mother tongue[++]	0.30	(0.01)	0.80	(0.01)	0.70	(0.01)	0.85	(0.01)
Paternal mother tongue[++]	0.27	(0.01)	0.82	(0.01)	0.72	(0.01)	0.84	(0.01)
Father is non-migrant	0.68	(0.01)	0.75	(0.01)	0.36	(0.01)	0.63	(0.01)
Mother is non-migrant	0.51	(0.01)	0.20	(0.01)	0.29	(0.01)	0.49	(0.01)
Maternal height (cm.)	158.7	(0.17)	151.3	(0.15)	150.1	(0.14)	152.2	(0.15)
Paternal height (cm.)	-	-	-	-	162.0	(0.20)	-	-
Maternal age - Paternal age	-8.86	(0.17)	-5.79	(0.07)	-3.55	(0.13)	-2.93	(0.08)
Ethnic match	0.87	(0.01)	1.00	(0.00)	0.91	(0.01)	0.96	(0.00)
Household characteristics								
Number of brothers	1.25	(0.04)	0.33	(0.02)	0.78	(0.03)	0.40	(0.02)
Number of sisters	1.15	(0.03)	0.39	(0.02)	0.77	(0.03)	0.49	(0.02)
Members aged 0 - 5	1.70	(0.02)	1.31	(0.01)	1.58	(0.02)	1.34	(0.01)
Members aged 6 -1 5	1.59	(0.04)	0.60	(0.02)	1.13	(0.03)	0.64	(0.02)
Wealth index	0.16	(0.00)	0.36	(0.01)	0.41	(0.01)	0.43	(0.01)
Piped water	0.10	(0.01)	0.15	(0.01)	0.78	(0.01)	0.10	(0.01)
Sewerage	0.22	(0.01)	0.20	(0.01)	0.78	(0.01)	0.44	(0.01)
Good-quality floor	0.08	(0.01)	0.25	(0.01)	0.38	(0.01)	0.56	(0.01)
Urban	0.30	(0.01)	0.20	(0.01)	0.67	(0.01)	0.21	(0.01)
Observations	1247		1598		1495		1472	

Notes: Standard errors in parentheses. [+]The child's ethnic/caste groups correspond to Amhara, Oromo and Tigrian in Ethiopia (other ethnic groups is the excluded category); Scheduled Caste-SC, Scheduled Tribe-ST and Backward Communities-BC in India (Open Category-OC caste is the excluded category); White and other(Asian, Black, native of the Amazon) in Peru (Mestizo is the excluded category), and; Kinh and H'Mong in Vietnam (other ethinc groups is the excluded category). [++]The indicators of parental mother tongue correspond to Amarigna in Ethiopia, Telugu in India, Spanish in Peru and Vietnamese in Vietnam. The number of observations for the indicator of wasted is 1165 in Ethiopia, 1593 in India, 1487 in Peru and 1472 in Vietnam. The number of observations for paternal height in Peru is 995.

In addition, Table 1 shows further parental characteristics likely to be related to either assortative mating in the marriage market or child health outcomes, such as: maternal height, mother's tongue, migration status, an indicator for whether parents belong to the same ethnic-caste group and the difference in age between mother and father. The bottom part of this table presents some indicators of households' wealth. The wealth index consists of a simple average of a housing quality index (including an indicator of households' members per room and indicators of wall and roof made of good-quality materials), a consumer durable index (consisting of the number of assets owned by the household) and a services index (including an indicator for whether the dwelling has electricity and an indicator for the use of a non-traditional fuel for cooking). Three additional indicators related to a hygienic home environment are also included: access to piped water into the dwelling or yard, access to own toilet facility and an indicator if the floor is made of cement or is tiled or laminated. As I mentioned above, these indicators show similar patterns between the YL sample and the national samples showed in Section 3.1. Both the wealth index and the indicators of healthy environments show that Ethiopia and India are in worse conditions than Peru and Vietnam.

3.3 Child Nutritional Status in the Young Lives Data

The first two rows of Table 1 report the proportion of wasted and stunted children in each sample. Also in terms of the indicators of child nutritional status, the YL samples are similar to the national estimates presented in the previous section. It is worth mentioning that all the anthropometric measures in the YL survey were collected directly by well-trained fieldworkers, following a set of rigorous instructions.[27]

[27] For further details about the procedure see the YL's website: http://www.younglives.org.uk/.

In Table 1, a child is defined as wasted if his/her weight-for-length falls below two standard deviations from the WHO standards, while a child is defined as stunted if his/her length is less than the 90 percent of the median length of a well-nourished child of the same sex and age, based on the WHO standards. India shows the highest proportion of wasted children, whilst Peru shows the lowest rates. The proportion of stunted children is higher in Ethiopia and India than in Peru and Vietnam.

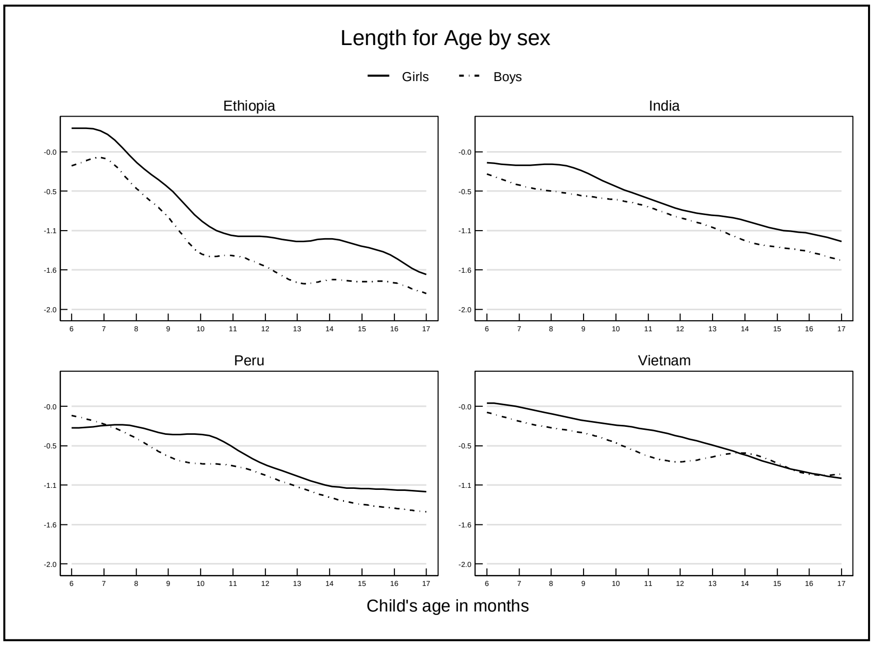
While stunting reflects cumulative episodes of stress that lead to reduced growth rates, wasting reflects the presence of episodes of stress at the time of measurement. The factors occasioning these episodes of stress might be chronic factors (e.g. seasonal scarcities of food and chronic diarrhoea) and/or acute factors (e.g. measles or weaning) (Martorell and Habicht, 1986). When facing a chronic factor, a child might become stunted but not necessarily wasted, but when a severe chronic factor is faced, it is likely that both stunting and wasting might be present. Martorell and Habicht (1986) argue that the interaction between acute and environmental factors might lead a child to become wasted but not necessarily stunted.

The distribution of malnourished children shows different patterns within each country according to which health indicator is chosen. In India and Vietnam, wasting or acute malnutrition (weight-for-length) is substantially higher than stunting or chronic malnutrition (length-for-age). The same occurs in Ethiopia, but the difference is smaller. In contrast, in Peru, chronic malnutrition has a relative higher incidence than short-run malnutrition: about 6 percent of the children are stunted while 2 percent are wasted.[28]

[28] The observed differences between wasting and stunting rates follow the same pattern in rural and urban locations. However, they intensify in rural areas.

Figures 1 and 2 show the estimates of nonparametric regressions for length-for-age and weight-for-length on child age in months by sex and country.[29] To allow comparability between boys and girls, and countries, standard deviation units (z-scores) are plotted. For each child in the sample, YL calculates z-scores (length-for-age; weight-for-age; and, weight-for-length) using the WHO Child Growth Standards for a population of young children facing a development under optimal environmental conditions.[30]

Figure 1: Length-for-Age Z-Scores



The length-for-age of children in the YL sample follows the traditional pattern in developing countries (Martorell and Habicht, 1986). Children's age in the YL sample coincides with the period where length-for-age declines steeply.[31] With respect to weight-for-length, the YL sample also follows the same patterns observed in different

[29] Figures A.1, in the Appendix A, shows the estimates of a nonparametric regression for weight-for-age on child age in months, by sex and country.

[30] For a complete discussion of the WHO child growth standards and a comparison with the NCHS standards, see WHO Multicentre Growth Reference Study Group (2006). According to this document "the standards depict normal early childhood growth under optimal environmental conditions and can be used to assess children everywhere, regardless of ethnicity, socioeconomic status and type of feeding." The Young Lives data provide these z-scores indicators.

[31] According to the authors, the length-for-age of children in developing countries stabilizes after the second year of life.

developing countries. Weight-for-length values fall very rapidly in infancy reaching their lowest values during the second year of life. Another interesting aspect is that children in Peru are in a similar situation to children in other Latin American countries. Their weight-for-length starts at relatively high levels (Martorell and Habicht, 1986). In the four YL countries girls are more similar than boys to well-nourished children. However, the pattern of growth retardation does not seem to significantly vary by sex. It is also clear from Figure 1 and 2 that in Ethiopia and India high levels of wasting coincides with high levels of stunting, while in Vietnam and Peru, fairly normal weight-for-length values coincide with moderate levels of stunting.

Figure 2: Weight-for-Length Z-Scores

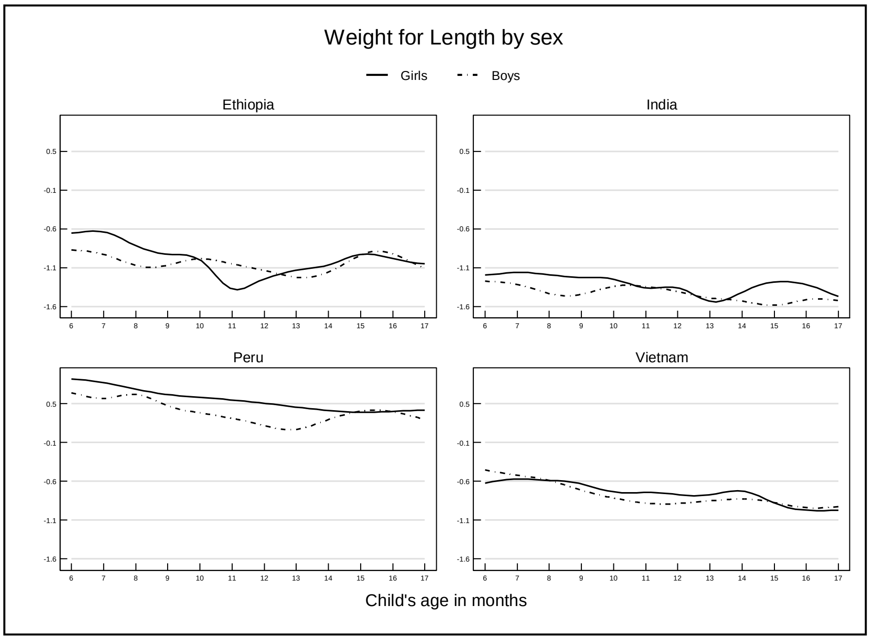


Figure A.2 in the Appendix shows the median values of the observed child length by sex and country. Children in the YL sample start near the fiftieth percentile of the WHO growth charts, but their length rapidly begins to fall as they grow up, finishing near the fifth percentile of the WHO growth charts. In developing countries, as Martorell and Habicht (1986) argue, ethnic differences in growth potential are minor prior to puberty, and so, differences in young children are mainly driven by socioeconomic factors (through nutrient intake and infections). According to the authors, the effects of socioeconomic factors on growth seem to be higher between 6 to 36 months. In developing countries, children under six-months are in a relatively good position compared to well-nourished children. After the sixth month their nutritional status deteriorates reaching its worst level during the second year of life and then it starts improving until the third or fourth year of life.

4. Empirical Strategy

4.1. Testing Random Assignment and Son-Biased Stopping Rule

Unfortunately, the YL data do not allow comparisons across girls and boys living in the same household, because the YL data only contain detailed information about one child living in the household.[32] To perform unbiased comparisons across households, I need to assure that intrinsic differences between households with a boy and households with a girl do not exist. This section discusses the bias in gender discrimination estimates related to the potential endogeneity of child sex and the son-biased stopping rule.

The son-biased stopping rule refers to situations where the probability to stop having children after giving birth to a boy is higher than after a girl, keeping everything else constant. Yamaguchi (1989) and Barcellos et al. (2012) show that if households follow the son-biased stopping rule child sex is no longer exogenous and is correlated with parental preferences for the gender composition of children. If this is the case, girls are more likely to live in households with higher preferences for girls compared to other households of the same size, which leads to biased comparisons.

[32] Only in its third round, YL has started collecting data of siblings in the household.

Assuming that the child' sex at birth is exogenous (i.e. there is no sex-selective abortion), Barcellos et al. (2012) argue that the characteristics of households that just had a boy and those that just had a girl are identical and, thus observed differences in child outcomes can be attributed to their sex. According to their study, this condition does not hold when samples with older children are included. If the son-biased stopping rule operates and older children are included, the probability that the youngest child in the household is a boy increases with the age of the child. In the presence of a son-biased stopping rule, households are more likely to stop having children after giving birth to a boy and thus they are no longer comparable with households alike (in observable characteristics) who stop having children after giving birth to a girl because the latter households have higher preferences for girls. Thus, child sex is no longer exogenous.

To overcome the possible endogeneity of child sex, Barcellos et al. (2012) choose a sample of households where the youngest child is 'young enough' to assure that the characteristics of those households with a boy and those with a girl are similar. This is an advantage of using the YL sample because, by sampling design, households with girls and boys at the same age group (6 – 18 months) are randomly selected.[33]

Despite this feature, I test whether the gender of the YL children is uncorrelated with household characteristics through the following linear equation:

$$b_i = X_i\beta + \varepsilon_i \tag{6}$$

Similar to Barcellos et al. (2012), in equation (6) b represents an indicator for whether the child i is a boy and X is a set of predetermined and observed characteristics. If $\beta = 0$, the child's sex is not jointly determined by X and thus the

[33] Following Barcellos et al. (2012), I restrict the sample to households where the YL child is the youngest and/or only child, which corresponds to 99 percent of households in the sample.

comparison of households with boys and girls is not biased. The main caveat of this test of random assignment is that it is based on observable characteristics and thus the role of unobservable characteristics related to $b$are not considered.

For each YL country, Table 2 reports a test of differences in mean between boys and girls for the characteristics considered. The p-values from the join test of significance, at the bottom of the table, confirm that in none of the four countries it is possible to reject the null hypothesis of $\beta = 0$. Therefore, comparisons between households with girls and boys using the YL sample will not lead to biased estimates.

The few significant differences in Table 2 are likely to be the product of chance and not of sex-selective abortion against girls or excess girl mortality. As in most previous studies, this study cannot rule out whether sex-selective abortion generates bias in the estimates. However, in 2000, when YL children were born, none of the YL countries show evidence that sex-selective abortion was a common practice. Even in Andhra Pradesh, the child sex-ratio at birth in 2001 does not show evidence of sex-selective abortion. In 2001, the northern states of Punjab, Haryana and Gujarat show the highest sex-ratio imbalance in India (Arnold et al., 2002; Guilmoto, 2007). Similarly in Vietnam, the 1999 census show no disparities in the sex-ratio at birth (Bélanger et al., 2003) even though more recent evidence shows an increase in the sex-ratio at birth (Guilmoto, 2009). If sex-selective abortion against girls and excess girl mortality represent a bias, it is more likely that these will reduce the estimated difference between boys and girls, because it is plausible to argue that surviving girls are more likely to live in households with higher preferences for girls or are relatively healthier.

Table 2: Testing for Random Assignment through Mean Differences by Gender

	Ethiopia	India	Peru	Vietnam
Child characteristics				
Child age (in months)	-0.01 (0.20)	-0.02 (0.17)	0.11 (0.18)	-0.02 (0.16)
Weeks of premature birth	0.02 (0.05)	0.08 (0.04)	0.13 (0.07)	0.07 (0.05)
Child ethnic group 1+	0.02 (0.02)	-0.02 (0.02)	0.01 (0.01)	-0.04 (0.02)*
Child ethnic group 2+	0.00 (0.02)	0.02 (0.02)	-0.01 (0.01)	0.02 (0.01)
Child ethnic group 3+	0.03 (0.02)	-0.03 (0.03)	- -	- -
Child is the youngest	-0.09 (0.11)	0.06 (0.05)	-0.04 (0.09)	0.00 (0.06)
Parental characteristics				
Maternal education	0.01 (0.03)	-0.02 (0.02)	0.15 (0.23)	-0.25 (0.21)
Paternal education	0.01 (0.03)	-0.03 (0.03)	0.25 (0.20)	-0.28 (0.21)
Maternal ed. > Paternal ed.	0.01 (0.02)	0.00 (0.02)	-0.02 (0.02)	-0.03 (0.02)
Maternal mother tongue++	0.01 (0.03)	-0.02 (0.02)	0.02 (0.02)	-0.03 (0.02)
Paternal mother tongue++	0.00 (0.03)	-0.03 (0.02)	0.02 (0.02)	-0.04 (0.02)*
Father is non-migrant	0.02 (0.03)	-0.04 (0.02)	0.03 (0.02)	0.00 (0.03)
Mother is non-migrant	0.03 (0.03)	-0.02 (0.02)	-0.04 (0.02)	-0.02 (0.03)
Maternal height (cm.)	0.08 (0.33)	0.24 (0.31)	-0.45 (0.27)	-0.35 (0.30)
Maternal age - Paternal age	0.17 (0.34)	0.15 (0.14)	0.47 (0.25)	-0.21 (0.17)
Ethnic match	0.02 (0.02)	0.00 (0.00)	-0.01 (0.01)	-0.00 (0.01)
Household characteristics				
Number of brothers	0.05 (0.07)	0.03 (0.03)	-0.04 (0.05)	-0.02 (0.03)
Number of sisters	-0.16 (0.07)*	0.03 (0.04)	-0.01 (0.06)	0.02 (0.04)
Members aged 0 - 5	-0.01 (0.04)	0.02 (0.03)	-0.03 (0.04)	0.01 (0.03)
Members aged 6 -1 5	-0.02 (0.08)	0.04 (0.05)	-0.00 (0.07)	0.01 (0.05)
Wealth index	0.00 (0.01)	-0.00 (0.01)	0.02 (0.01)	0.00 (0.01)
Piped water	0.01 (0.02)	0.00 (0.02)	0.02 (0.02)	0.02 (0.02)
Sewerage	-0.00 (0.02)	0.02 (0.02)	0.01 (0.02)	0.01 (0.03)
Good-quality floor	0.01 (0.01)	-0.00 (0.02)	-0.00 (0.03)	0.01 (0.03)
Urban	-0.00 (0.03)	0.01 (0.02)	0.05 (0.02)*	0.00 (0.02)
p-value joint significance	0.480	0.655	0.301	0.639
Observations	1247	1598	1495	1472

Notes: Standard errors in parentheses. Each cell corresponds to a separate linear regression of each characteristic on an indicador whether the child is a boy (reported) and a constant. The p-value for the joint significance of all the characteristics on a regression of child gender is reported at the bottom of the table. +The child's ethnic/caste groups correspond to Amhara, Oromo and Tigrian in Ethiopia (other ethnic groups is the excluded category); Scheduled Caste-SC, Scheduled Tribe-ST and Backward Communities-BC in India (Open Category-OC caste is the excluded category); White and other(Asian, Black, native of the Amazon) in Peru (Mestizo is the excluded category), and; Kinh and H'Mong in Vietnam (other ethinc groups is the excluded category). In Ethiopia and Vietnam, parental education corresponds to indicators whether parents are literate. In Peru and Vietnam corresponds to years of education. ++The indicators of parental mother tongue correspond to Amarigna in Ethiopia, Telugu in India, Spanish in Peru and Vietnamese in Vietnam. *Significant at 5%, **Significant at 1%.

4.2. Estimating Gender Differences

Considering the demand for child health (5), the following equation represents a general specification to assess whether girls and boys are treated differently based on parental preferences:

$$H_{ih} = \alpha_0 + \alpha_1 b_{ih} + \alpha_2 p_h + X_{ih}\beta + \alpha_3 (b_{ih} \times\ p_h) + \sum_{j=1}^{k} \gamma_j \left(x_{ih}^{j} \times b_{ih}\right) + \varepsilon_{ih} \quad (7)$$

where H is a vector of the child i's nutritional outcomes: length-for-age or weight-for-length z-scores. As discussed in Section 3.3, the first indicator reflects accumulated investments over the life of the child, including past and current deprivation or illnesses. In contrast, the second indicator reflects short-run responses to changes in the environment and nutrition (Martorell and Habicht, 1986).[34] b is a variable indicating whether the child is a boy, and p is a dummy for maternal bargaining power, indicating whether the mother's years of education are greater than the father's.

This paper focuses on one distribution factor z (in equation 5), which corresponds to whether the child's mother has more years of formal education than the child's father. If more education implies more bargaining power in the intrahousehold resources allocation, then better educated women than their husbands are more likely to impose their preferences and allocate more resources towards commodities they care more about.

A growing body of empirical evidence shows that factors modifying the distribution of bargaining power within the household, in favour of women, are associated with larger investments in child human capital and child expenditure, in comparison to situations where resources are directed to men (Duflo, 2000 and 2003; Hoddinott and Haddad, 1995; Lundberg et al., 1997; Reggio, 2011; Thomas, 1990 and 1994). It is reasonable to argue that having relatively more education, gives mothers more bargaining power in the household. As Beegle et al. (2001) and

[34] As it was mentioned also in Section 3.3, these indicators are expressed in z-scores using the WHO Child Growth Standards for a population of young children facing a development under optimal environmental conditions. The Young Lives data provide these z-scores indicators.

Thomas (1994) mention, having more education than their husbands reflects that mothers have more opportunities in the labour markets and at higher wages. It is also correlated with having acquired more 'modern skills' that might be applied when making decisions concerning the allocation of household resources. In addition, this indicator has been broadly used as a proxy for bargaining power (Basu and Ray, 2002; Beegle et al., 2001; Frankenberg and Thomas, 2003; Gitter and Barham, 2008; Thomas, 1990 and 1994), and represents the best alternative contained in the YL data.[35]

In equation (7), X is a matrix of child characteristics (age in months, weeks of premature birth, indicators of ethnic group/caste and birth order), parental characteristics (maternal and paternal years of education, maternal age, proficiency in main language and religion), household characteristics (number of child's brothers and sisters, wealth index, indicators for piped water, sewerage and good-quality floor), and finally to capture for potential phenotype and genotype effects on child health, maternal height is also included. ε contains unobserved characteristics related to the demand of child health, such as: parental labour income and time allocation to child rearing, traditional customs and beliefs, among others. The sixth term in the right hand side of (7) shows the interacted variables of X with the child's sex indicator to capture differences in child rearing technology for boys and girls.

In equation (7), the parameters α_2 and $(\alpha_2 + \alpha_3)$ represent the marginal effect of maternal bargaining power on girls' and boys' health, respectively, and therefore the differential impact that maternal bargaining power has on the nutritional outcomes of boys versus girls is captured in α_3.

---

[35] The YL data does not include other indicators of bargaining power used in previous literature, such as non-labour income (Thomas 1990 and 1994), resources brought to the marriage (Fafchamps and Quisumbing, 2002), among others.

If maternal relative education is considered to be exogenous in (7), as in most of the previous literature using similar indicators of bargaining power, the OLS estimate is unbiased. It is arguable, however, that households where mothers are better educated than fathers represent a random population. If the matching of parents in the marriage market is determined by unobservable characteristics, such as tastes for gender equality, that are also correlated with child health, the indicator of maternal power becomes endogenous and the OLS estimate biased.[36] For instance, a larger effect of maternal bargaining power on girls' health might reflect that; an increase in maternal bargaining power allow women to allocate more resources to girls; or, that men with non-traditional gender roles (unobserved) tend to marry more powerful women.

This is particularly relevant in studies using cross-sectional data, in which it is not possible to control for the permanent effect of unobservable (time-invariant) characteristics, such as those related to the assortative mating process. In the absence of experimental data, empirical strategies to overcome the possible endogeneity of maternal bargaining strength, consists of finding an instrumental variable (IV) or estimating an endogenous treatment model, considering the endogenous selection indicator (treatment) as a binary variable. Because I have neither longitudinal data nor credible instruments for family formation (i.e. the matching of more/less educated women with less/more educated men) to deal with this omitted-variable problem, the empirical strategy of this study attempts to recover the effect of maternal bargaining power on child health by including in the empirical model, the large set of observable

[36] While in Ethiopia and India (Andhra Pradesh) an important unobserved factor might be the extended practice of arranged marriages (Fafchamps and Quisumbing, 2005; NCTPE, 2003, for Ethiopia; Bloom and Reddy, 1986, for India), in Peru and Vietnam an important unobserved factor might be the submissive role of women in the society. However, Knodel et al. (2005) argue that the submissive role of women in traditional societies does not generally affect women control over the expenditure and distribution of food within the household.

controls mentioned above, but also a set of variables likely to be correlated with family formation. This strategy intends to washing out unobservable dimensions, such as tastes, opportunities and constraints affecting child health and being correlated with parental matching in the marriage market.

Subject to limitations of data on pre-marriage conditions in the YL survey, the set of covariates added to the vector X corresponds to: the age difference between mother and father; an indicator for whether mother and father belongs to the same ethnic group/caste; an indicator for whether parents share the same mother tongue; and, an indicator for whether parents have never migrated from the community.[37] Intuitively, I expect the age difference between mother and father to be positively related to mothers having more schooling than their partners. Moreover, for instance, single women who are getting closer to the end of their fertile life or in the vicinity of the 'traditional' age for marriage, might be more willing to marry men whose characteristics mismatch theirs (e.g. younger, less educated, smaller). Belonging to the same ethnic group/caste and sharing the same mother tongue are expected to be positively related to household formation. If parents have always lived in the same particular community, they are more likely to face the same marriage market, which might contribute to their union.

To explore whether the differential effect of maternal bargaining power on boys' and girls' health varies across (i) households in urban and rural areas and (ii) households with or without more young children, I introduce additional terms to equation (7). First, I include an indicator for urban location and its interactions with child gender and maternal bargaining power. Second, in a different specification, I

[37] Reggio (2011) includes the age difference between husbands and wives as an additional distribution factor (i.e. a factor affecting the intrahousehold distribution of power but do not affecting directly individual's preferences). The author finds that wives who are relatively much younger than their husbands have less decision power in the household.

include an indicator for the presence in the household of other children between 0-5 years and its interactions with child gender and maternal bargaining power.

Finally, I estimate a series of robustness checks. First, to check whether the results are driven by differences in parental behaviour (i.e. differential effects of parental education and bargaining power), I compare the estimates of a model of current child weight with those of a model of child growth in weight (both in kilograms). Assuming that parental characteristics do not have a differential effect on the weight of boys and girls at birth, if there are parental behaviour differences, these should be reflected both in current weight and growth in weight. Ideally, this robustness check should be made using an indicator of accumulative investments in health, but unfortunately weight at birth is the only variable included in the four samples. However, the sample of Peru includes child length at birth and thus, as a second robustness check, I estimate a similar model but for child height and growth in height (both in centimetres). Finally, even though all the previous specifications control for genetic endowments (maternal height), to further control for potential omitted variables, in a final specification for Peru, I also control for paternal height.

5. Results

Two groups of regressions are estimated in this section. First, using the complete sample for each country, I estimate OLS regressions, including the set of covariates in X, as listed in Section 4.2, and adding the set of variables related to household formation,[38] for child length-for-age z-scores (Section 5.1) and weight-for-length z-scores (Section 5.2). Second, for both nutritional indicators, I estimate OLS

[38] That is, parental age difference, an indicator of whether parents share the same mother tongue and belong to the same ethnic group/caste, and an indicator of whether parents have never migrated from the community.

regressions accounting for heterogeneous effects across different households (Section 5.3).

Since the main focus of this paper is assessing the differential effect of maternal bargaining power on child health outcomes, the tables below show only the most relevant coefficients.[39]

5.1 Length-for-Age Z-Scores

This section shows the estimates of (5) corresponding to the indicator of cumulative investments in child health. Table 3 reports for each YL country, the effect of maternal bargaining power on the z-scores of girls' and boys' length-for-age, under the specifications presented in the previous section.

In Table 3, models (I) show the most restricted set of covariates in X and models (II) include the additional covariates attempting to capture unobservable characteristics related to household formation. Adding these variables slightly improves models' fit,[40] but do not statistically change the point estimates from the first models.[41] Thus, the discussion of results focuses on models (II), which include the extended set of covariates.

The effects of maternal bargaining power on daughters' (coefficient on *MBP*) and sons' (sum of coefficients on *Boy* and *Boy*MBP*) length are, in almost all cases, positive, although not statistically significant. In Vietnam, where the effect of maternal bargaining power on boys' length is negative (and statistically insignificant), I also find that the effect of maternal bargaining power on girls' length is significantly larger than in boys' length. This gender difference (captured in the coefficient on

[39] Full regressions results are available on request.

[40] The chi-squared value of the likelihood ratio test is 19.06 in Ethiopia, significant at the 5% level; 10.61 in India, 5.36 in Peru, and, 10.86 in Vietnam, but no statistically significant.

[41] The null hypothesis of equality of interaction coefficients *Boy*MBP* across models is rejected in Ethiopia (p-value=0.67), India (p-value=0.79), Peru (p-value=0.94), and Vietnam (p-value=0.74).

*Boy*MBP*) reflects that mothers with more bargaining power in Vietnam prefer to allocate more resources to improving the health of their daughters than the one of their sons (-0.30 standard deviations, significant at the 10% level). Thomas (1994) finds similar results for children in Ghana. Maternal relative education has a significantly larger effect (1.18 standard deviations more) on girls' height than on boys' height.

Given that this indicator reflects cumulative investments on children, the lack of evidence on gender differences in the other three countries supports Duflo (2005) view, in which gender discrimination in allocation of resources mainly happens when households experience negative shocks.

Table 3: Effect of Maternal Bargaining Power (MBP) on Child Length-for-Age Z-Scores

	Ethiopia		India		Peru		Vietnam	
	(I)	(II)	(I)	(II)	(I)	(II)	(I)	(II)
Boy	-1.247	-1.595	0.399	2.430	-0.061	-0.004	1.724	1.699
	(3.045)	(3.062)	(2.489)	(2.870)	(1.964)	(2.004)	(1.622)	(1.624)
MBP	0.186	0.158	0.162	0.160	0.034	0.050	0.126	0.128
	(0.260)	(0.266)	(0.201)	(0.201)	(0.125)	(0.126)	(0.113)	(0.114)
Boy*MBP	-0.010	0.019	-0.057	-0.067	0.016	0.014	-0.294*	-0.300*
	(0.375)	(0.379)	(0.287)	(0.289)	(0.180)	(0.181)	(0.176)	(0.176)
Observations	1247	1247	1598	1598	1495	1495	1472	1472
R-squared	0.173	0.185	0.158	0.164	0.312	0.315	0.319	0.324

Notes: Models (I) correspond to OLS estimates of the variables shown on z-scores. The additional controls correspond to: child characteristics (age in months, weeks of premature birth, dummies for ethnic group/caste and birth order), parental characteristics (maternal and paternal years of education, maternal height, age, proficiency in main language and religion), household characteristics (number of child's brother and sisters, wealth index, indicators for piped water, sewerage and good-quality floor), and their interaction by child sex. Models (II) include, in addition to the covariates in models (I), a set of variables likely to be correlated with determinants of household formation (parental age difference, an indicator for whether parents belong to the same ethnic group/caste, indicators for whether parents share the same mother tongue, and an indicator for whether parents have never migrated from the community) and their interaction by child sex. Robust standard errors are in parentheses. *Significant at 10%, **Significant at 5%, ***Significant at 1%.

5.2 Weight-for-Length Z-Scores

I now turn to show the result of estimate equation (5) for the indicator of short-run nutritional status. Table 4 reports for each YL country, the effect of maternal bargaining power on the z-scores of girls' and boys' weight-for-length, under the specifications detailed in Section 4. Similarly to the estimations of length-for-age, using the set of additional covariates attempting to capture unobservable characteristics related with household formation slightly improves models' fit.[42] Differences between point estimates across models are not statistically significant either.[43] As in the previous section, the discussion of results also refers to models (II).

Table 4: Effect of Maternal Bargaining Power (MBP) on Child Weight-for-Length Z-Scores

	Ethiopia		India		Peru		Vietnam	
	(I)	(II)	(I)	(II)	(I)	(II)	(I)	(II)
Boy	-0.676	-0.765	-2.647*	-3.563**	1.051	1.098	0.083	0.127
	(2.334)	(2.361)	(1.459)	(1.681)	(1.853)	(1.878)	(1.534)	(1.570)
MBP	0.129	0.128	-0.265*	-0.269*	0.080	0.067	-0.048	-0.058
	(0.187)	(0.192)	(0.160)	(0.160)	(0.136)	(0.138)	(0.117)	(0.118)
Boy*MBP	-0.022	-0.033	0.294	0.317	-0.328*	-0.310*	-0.080	-0.067
	(0.293)	(0.298)	(0.219)	(0.220)	(0.182)	(0.183)	(0.178)	(0.178)
Observations	1174	1174	1600	1600	1488	1488	1472	1472
R-squared	0.174	0.177	0.069	0.076	0.102	0.104	0.064	0.066

Notes: Models (I) correspond to OLS estimates of the variables shown on z-scores. The additional controls correspond to: child characteristics (age in months, weeks of premature birth, dummies for ethnic group/caste and birth order), parental characteristics (maternal and paternal years of education, maternal height, age, proficiency in main language and religion), household characteristics (number of child's brother and sisters, wealth index, indicators for piped water, sewerage and good-quality floor), and their interaction by child sex. Models (II) include, in addition to the covariates in models (I), a set of variables likely to be correlated with determinants of household formation (parental age difference, an indicator for whether parents belong to the same ethnic group/caste, indicators for whether parents share the same mother tongue, and an indicator for whether parents have never migrated from the community) and their interaction by child sex. Robust standard errors are in parentheses. *Significant at 10%, **Significant at 5%, ***Significant at 1%.

The sample of Peru shows evidence of gender differences in the effect of maternal bargaining power. Maternal bargaining power has a larger effect (0.31 standard deviations, significant at the 10% level) on girls' than on boys' weight-for-length. That is, mothers with more bargaining power allocate more resources to their daughters than to their sons. Moreover, relative to other boys, boys living with mothers who have more bargaining power receive fewer resources (sum of

[42] In contrast to the length-for-age specification, using the extended set of variables significantly improves the fit of the model in India: the chi-squared value of the likelihood ratio test is 13.59, significant at the 10% level; and differences are not statistically significant in the other three countries.

[43] The null hypothesis of equality of interaction coefficients "Boy*Maternal bargaining power" across models is largely rejected in the Ethiopia (p-value=0.76), India (p-value=0.38), Peru (p-value=0.30), and Vietnam (p-value=0.38).

coefficients on *Boy* and *Boy*MBP*: 0.07-0.31=-0.24 standard deviations, significant at the 5% level).

In contrast, Indian girls living in households where mothers have more power are worse off than other girls (coefficient on *MBP* = -0.27 standard deviations, significant at the 10% level). This suggests that the preferences against girls of mothers with more power in the household are stronger than the ones of fathers with more power in other households with girls. In India, maternal bargaining power seems to also have a differential effect favouring boys relative to girls (coefficient on interaction *Boy*MBP* = 0.32 standard deviations, but statistically insignificant).[44] In the samples of Ethiopia and Vietnam, mothers with more bargaining power seem to allocate more resources to girls than to boys, but these coefficients are small and statistically insignificant.

A channel through which maternal preferences might be negatively affecting children's health in the short-run is related to breastfeeding practices. Particularly in India, children of mothers who are relatively more educated than their husbands are weaned sooner than children of relatively less educated mothers.[45] As Martorell and Habicht (1986) shows, in developing countries the most important problem of infection affecting child health is diarrhoeal diseases, which are more likely to occur during the period of weaning. The so called "weaning dilemma" in developing countries refers to the choice between incipient marasmus from continuing breastfeeding (with breast milk poor in quantity and nutrients but which offers antimicrobial protection against infections), or diarrhoea from starting to feed children with unhygienic and poor in proteins and in calories weaning food. The

[44] Significant at the 15% level.
[45] Children of relatively more educated women are weaned 0.28 months before other children in India (significant at the 10% level). In Peru, they are also weaned 0.10 months before other children but differences are not statistically significant.

evidence cited in Martorell and Habicht (1986) shows that in particular, in poorer areas in developing countries, breastfed children are more protected against infections and show better health indicators. Therefore, maternal preferences for weaning in the YL samples of India and Peru might put children at greater risks of getting infections, which are being captured by the short-run indicator ‘weight for height’.

Further evidence to help explain the larger effect of maternal bargaining power on boy’s health in India and girls’ health in Peru might be related to gender bias in breastfeeding practices. In India, boys are weaned later than girls, while in Peru girls are weaned later than boys. Jayachandran and Kuziemko (2011) and Chakravarty (2012) find that gender bias in breastfeeding practices, favouring boys in India and Africa, respectively, might be due to son-preference and fertility preferences. The gender bias in breastfeeding reduces when women achieve their ideal number of children and girls have an older male sibling. In Sub-Saharan Africa, when mothers have exceeded their ideal number of children, girls receive less breastfeeding than boys, which suggests that households prefer to allocate more to boys when resources are scarce (Chakravarty, 2012).

5.3 Heterogeneous Effects

This section explores whether the differential effect of maternal bargaining power on children health is homogeneous across (i) rural and urban households (Table 5) and (ii) households containing other young children (Table 6).[46] The former analysis is motivated by the disparities in socioeconomic indicators between urban and rural areas discussed in Section 3.1. The latter is motivated by the fact that children might compete for household resources (care, food) when other younger children live in the household.

[46] The specifications shown in this section control for the full set of variables in X and those correlated with household formation.

In general, Table 5 shows that the differential effect of maternal bargaining power on boys and girl health does not vary between urban and rural areas (coefficient on *Boy*MBP*Urban*). The significant differential effect of maternal bargaining power on girls' and boys' length-for-age found in Table 3 for Vietnam does not seem to be concentrated in urban or rural areas (left panel of Table 5). Similarly, the significant differential effect of maternal bargaining power on girls' and boys' weight-for-length found in Table 4 for Peru seems to be homogeneous across urban and rural areas (right panel of Table 5).

Table 5: Effect of Maternal Bargaining Power (MBP) by Urban/Rural Location

	Length-for-Age Z-Scores				Weight-for-Length Z-Scores			
	Ethiopia	India	Peru	Vietnam	Ethiopia	India	Peru	Vietnam
Boy	-1.671	2.426	0.071	1.730	-0.568	-3.690**	1.138	0.082
	(3.110)	(2.875)	(1.999)	(1.624)	(2.379)	(1.698)	(1.891)	(1.566)
MBP	0.132	0.148	-0.019	0.094	0.257	-0.311*	0.192	-0.129
	(0.307)	(0.237)	(0.175)	(0.122)	(0.234)	(0.183)	(0.201)	(0.121)
Urban	0.015	-0.093	0.195*	0.200	-0.089	0.057	0.116	0.046
	(0.254)	(0.212)	(0.116)	(0.123)	(0.210)	(0.161)	(0.132)	(0.114)
Boy*MBP	0.006	-0.012	-0.129	-0.270	-0.169	0.294	-0.541**	-0.047
	(0.465)	(0.346)	(0.277)	(0.187)	(0.360)	(0.254)	(0.270)	(0.185)
Boy*Urban	0.018	0.116	-0.154	-0.244	0.199	0.115	-0.121	0.147
	(0.352)	(0.292)	(0.179)	(0.189)	(0.310)	(0.239)	(0.193)	(0.186)
MBP*Urban	0.081	0.036	0.107	0.212	-0.402	0.129	-0.175	0.373
	(0.445)	(0.312)	(0.182)	(0.215)	(0.316)	(0.261)	(0.213)	(0.240)
Boy*MBP*Urban	0.007	-0.137	0.181	-0.188	0.405	-0.034	0.317	-0.097
	(0.601)	(0.430)	(0.293)	(0.313)	(0.474)	(0.353)	(0.287)	(0.345)
Observations	1247	1598	1495	1472	1174	1600	1488	1472
R-squared	0.186	0.164	0.318	0.327	0.178	0.078	0.105	0.072

Notes: Models correspond to OLS estimates of the variables shown on the nutritional indicator (length-for-age or weight-for-length). The additional controls correspond to: child characteristics (age in months, weeks of premature birth, dummies for ethnic group/caste and birth order), parental characteristics (maternal and paternal years of education, maternal height, age, proficiency in main language and religion), household characteristics (number of child's brother and sisters, wealth index, indicators for piped water, sewerage and good-quality floor), a set of variables likely to be correlated with determinants of household formation (parental age difference, an indicator for whether parents belong to the same ethnic group/caste, indicators for whether parents share the same mother tongue, and an indicator for whether parents have never migrated from the community), and their interaction by child sex. Robust standard errors are in parentheses. *Significant at 10%, **Significant at 5%, ***Significant at 1%.

It is worth noting that, although not statistically different to households in urban areas, among households in rural areas of Peru, maternal bargaining power has a significant larger effect on girls' than on boys' weight-for-length (coefficient on

*Boy*MBP*, right panel of Table 5). Similarly, although not statistically different to households in urban areas either, the negative effect of maternal bargaining power on girls' weight-for-length found in Table 4 for India, seems to be driven by households living in rural areas (right panel of Table 5). Among households in rural areas, girls show lower weight-for-length z-scores in those households where mothers have more bargaining power (coefficient on *MBP* = -0.31 standard deviations, significant at the 10% level) relative to girls living in households where women have less power.

Table 6: Effect of Maternal Bargaining Power (MBP)
by Presence of Other Young Children Aged 0-5 (MCH) in the Household

	Length-for-Age Z-Scores				Weight-for-Length Z-Scores			
	Ethiopia	India	Peru	Vietnam	Ethiopia	India	Peru	Vietnam
Boy	-1.551	2.249	0.130	1.691	-0.771	-3.587**	1.167	-0.001
	(3.064)	(2.834)	(2.004)	(1.632)	(2.376)	(1.698)	(1.869)	(1.578)
MBP	0.212	0.063	0.092	0.214*	0.296	-0.243	0.069	-0.014
	(0.329)	(0.228)	(0.148)	(0.124)	(0.233)	(0.174)	(0.161)	(0.126)
MCH	-0.184	0.126	-0.049	-0.055	-0.183	-0.069	-0.287***	-0.053
	(0.167)	(0.142)	(0.101)	(0.099)	(0.136)	(0.122)	(0.101)	(0.095)
Boy*MBP	0.157	-0.003	-0.063	-0.355*	-0.352	0.313	-0.288	-0.096
	(0.465)	(0.313)	(0.215)	(0.187)	(0.355)	(0.239)	(0.216)	(0.192)
Boy*MCH	0.042	0.031	-0.089	-0.036	0.029	0.128	0.313**	0.179
	(0.251)	(0.209)	(0.142)	(0.150)	(0.196)	(0.158)	(0.150)	(0.148)
MBP*MCH	-0.118	0.297	-0.078	-0.332*	-0.339	-0.072	0.065	-0.172
	(0.433)	(0.307)	(0.170)	(0.173)	(0.331)	(0.264)	(0.199)	(0.204)
Boy*MBP*MCH	-0.258	-0.171	0.165	0.231	0.636	-0.022	-0.124	0.110
	(0.586)	(0.457)	(0.268)	(0.256)	(0.484)	(0.371)	(0.263)	(0.285)
Observations	1247	1598	1495	1472	1174	1600	1488	1472
R-squared	0.188	0.166	0.316	0.328	0.181	0.077	0.109	0.069

Notes: Models correspond to OLS estimates of the variables shown on a nutritional indicator (length-for-age or weight-for-length). The additional controls correspond to: child characteristics (age in months, weeks of premature birth, dummies for ethnic group/caste and birth order), parental characteristics (maternal and paternal years of education, maternal height, age, proficiency in main language and religion), household characteristics (number of child's brother and sisters, wealth index, indicators for piped water, sewerage and good-quality floor), a set of variables likely to be correlated with determinants of household formation (parental age difference, an indicator for whether parents belong to the same ethnic group/caste, indicators for whether parents share the same mother tongue, and an indicator for whether parents have never migrated from the community), and their interaction by child sex. Robust standard errors are in parentheses. *Significant at 10%, **Significant at 5%, ***Significant at 1%.

Table 6 shows no evidence that differential effects of maternal bargaining power on boys and girls vary with the presence of other young children aged 0-5 in the household (coefficients on *Boy*MBP*MCH*). However, in the sample of Vietnam, among households where there are no more young children, mothers with more bargaining power allocate more resources to their daughters than to sons (coefficient on *Boy*MBP*= -0.36 standard deviations in the regression of length-for-age, significant at the 10% level). On the other hand, among households with more young children, maternal bargaining power does not have a differential effect on boys and girls.[47] In Peru, maternal bargaining power does not seem to affect boys and girls differently, whether or not living with other young children. Interestingly though, Table 6 shows some evidence of rivalry for resources in households where mothers have less bargaining power than their husbands. The presence of additional young children negatively affects girls' health (-0.29 standard deviations in the weight-for-length regression, significant at the 1% level) but does not affect boys' health (sum of coefficients on *MCH* and *Boy*MCH* (right panel of Table 6).[48] Thus, rivalry between young children for household resources seems to negatively affect girls but not boys in households where women have less power in Peru.

Overall, the differential effects of maternal bargaining power on boys' and girls' health found in the previous section (favouring girls in Peru and Vietnam and boys in India) do not seem to vary with the households' location in urban and rural or with the presence of other young children in the household. Although not statistically different from households in urban areas, there is evidence that this effect concentrates among households in rural areas. Similarly, I find evidence that among households where women have less power, rivalry with other young children in the household is associated with worse health for girls but not for boys.

[47] Sum of coefficients (*Boy*MBP*) + (*Boy*MBP*MCH*).
[48] The difference being statistically significant at the 5% level as reflected in the coefficient on *Boy*MCH*.

6. Robustness Checks

This section shows additional checks for whether the estimated effects are being driven by behavioural differences between parents. Following Thomas (1994), this section shows estimates for growth in child weight (Table 7) and length (Table 8). The first round of the YL survey includes information on weight at birth and current weight for children in the four countries, and thus, the first set of regressions of growth in weight is made for the four countries. In contrast, child length at birth is only available in the sample of Peru, and thus, a regression of growth in length for this sample is presented as an additional robustness check at the end of this section.

Assuming that there are no differences in the effect of parental characteristics on girls' and boys' weight (and length) at birth, if behavioural differences between the mother and the father exist, the estimated parameters for the child's current weight (and length) models should reflect these differences. If this is the case, the estimated parameters on the growth in weight (and growth in length) and the ones on the current weight (and length) specifications should be similar.

The last specification included in this section (right panel of Table 8) accounts for potential bias due to the omission of paternal height in all the previous models. Again, this regression uses data that is only available in the YL sample of Peru.

6.1. Growth in Weight

Growth in weight is calculated as the child's current weight minus the child's birth weight, in kilograms. Although the four YL countries questionnaires collected both variables, the latter component variable (weight at birth) shows a significant number of missing values in all the countries, particularly on the relatively less developed countries, Ethiopia and India. To compare the estimated parameters, the results shown below are restricted to a sample where values are different from missing, in the two indicators. For each country, Table 7 reports the estimations of the most complete set of covariates on child weight (columns I) and growth in weight (columns II).

The considerable reduction on sample size due to missing data on birth weight in Ethiopia and India makes it even more difficult to identify any differential impact of bargaining power on children's health. Similarly to Table 4 for weight-for-length z-scores, I find no differential effect of bargaining power on child weight in Ethiopia, India and Vietnam.

Table 7: Effect of Maternal Bargaining Power (MBP) on Child Growth in Weight (kg)

	Ethiopia		India		Peru		Vietnam	
	(I)	(II)	(I)	(II)	(I)	(II)	(I)	(II)
Boy	2.404	1.649	0.511	0.132	3.049	3.558*	1.830	-0.510
	(6.964)	(7.337)	(2.718)	(2.444)	(2.136)	(2.132)	(1.502)	(1.830)
MBP	0.694	0.955	0.099	0.158	0.094	0.082	0.044	0.013
	(0.608)	(0.641)	(0.209)	(0.223)	(0.149)	(0.150)	(0.119)	(0.127)
Boy*MBP	-0.991	-1.050	0.195	0.080	-0.404**	-0.384*	-0.187	-0.019
	(0.734)	(0.773)	(0.278)	(0.290)	(0.202)	(0.198)	(0.174)	(0.181)
Observations	215	215	632	632	1312	1312	1473	1263
R-squared	0.500	0.383	0.359	0.360	0.388	0.380	0.368	0.348

Notes: Models correspond to OLS estimates of the variables shown on child weight in Kg. (columns I) or growth in weight in Kg. (columns II). The additional controls correspond to: child characteristics (age in months, weeks of premature birth, dummies for ethnic group/caste and birth order), parental characteristics (maternal and paternal years of education, maternal height, age, proficiency in main language and religion), household characteristics (number of child's brother and sisters, wealth index, indicators for piped water, sewerage and good-quality floor), a set of variables likely to be correlated with determinants of household formation (parental age difference, an indicator for whether parents belong to the same ethnic group/caste, indicators for whether parents share the same mother tongue, and an indicator for whether parents have never migrated from the community), and their interaction by child sex. *Significant at 10%, **Significant at 5%, ***Significant at 1%.

For Peru, however, I find that maternal bargaining power has a differential effect according to the gender of the child. Both in the model of child weight and growth in weight, mothers with more bargaining power seem to allocate more resources to girls relative to boys. This evidence is similar to the one shown in Table 4 for the indicator of weight-for-length. What is more important from Table 7, is that coefficients are statistically identical between models (I) and (II), reflecting that the estimated effects respond to parental behaviour.[49]

6.2. Further Robustness Checks for the Peruvian YL Sample

Growth in Length

As mentioned above, child length and length at birth was only collected for the YL sample of Peru, and thus this section corresponds only to this country's sample. Again both regressions (left panel in Table 8) correspond to a reduced sample for which values are different from missing, in both indicators. Similar to the robustness check performed above for growth in weight, I run a regression on growth in length, defined as current length minus length at birth. Assuming that parental characteristics do not affect differently child length at birth, I expect that if current child length reflects parental behavioural differences, then the coefficients of the model on child length and child growth in length should be similar.

The left panel of Table 8 shows that, although the coefficients in the model for child length and growth in length are not precisely estimated, they are statistically similar. This suggests that behavioural differences between parents affect the current child health outcome.[50]

[49] The null hypothesis of equality of the coefficients *MBP* (maternal bargaining power) in the model of child weight (I) and growth in weight (II), is largely rejected in Ethiopia (p-value=0.30), India (p-value=0.65), Peru (p-value=0.83) and Vietnam (p-value=0.97). Similarly, the null hypothesis of equality of the coefficients *Boy*MBP* (measuring the differential effect of maternal bargaining power on boys and girls) in the model of child weight (I) and growth in weight (II), is largely rejected in Ethiopia (p-value=0.87), India (p-value=0.48), Peru (p-value=0.80) and Vietnam (p-value=0.27).

[50] The null hypothesis of equality of the coefficients *BP* (maternal bargaining power) in the model of child length (I) and growth in length (II), is largely rejected in the sample of Peru (p-value=0.75). Similarly, the null hypothesis of equality of the coefficients *Boy*BP* (measuring the differential effect of maternal bargaining power on boys and girls) in the model of child length (I) and growth in length (II), is largely rejected in Peru (p-value=0.95).

Paternal Height

All the previous specifications have included maternal height to capture the role of genetics on child nutritional indicators. However, to better examine the effect of genetics, one ideally would include paternal height as well. It is expected that maternal and paternal height are correlated due to the matching process of household formation. To check how much child health outcomes reflect parental behaviour and background characteristics and how much they are driven by omitted phenotype and genotype characteristics, I use data of paternal height only available in the YL sample of Peru (right panel of Table 8).

Table 8: Effect of Maternal Bargaining Power (MBP) on Growth in Child Length and Including Both Parents' Height, Peru

	(I)		(II)			
			Length-for-Age		Weight-for-Length	
	Length (cm.)	Growth in Length (cm.)	Only mother's height	Both parent's height	Only mother's height	Both parent's height
Boy	1.416	-5.419	0.905	-0.368	-0.997	0.809
	(6.745)	(7.940)	(2.407)	(2.822)	(2.315)	(2.861)
MBP	0.378	0.505	0.080	0.075	0.189	0.189
	(0.399)	(0.509)	(0.152)	(0.150)	(0.161)	(0.161)
Boy*MBP	-0.099	-0.138	-0.137	-0.203	-0.539**	-0.515**
	(0.566)	(0.730)	(0.223)	(0.222)	(0.219)	(0.219)
Maternal height	0.128***	0.049	0.061***	0.055***	-0.009	-0.010
	(0.028)	(0.035)	(0.009)	(0.010)	(0.010)	(0.010)
Paternal height	-	-	-	0.026***	-	0.002
	-	-	-	(0.008)	-	(0.009)
Maternal height	-0.004	0.025	-0.011	-0.014	0.013	0.016
(cm.)*Boy	(0.040)	(0.048)	(0.015)	(0.015)	(0.014)	(0.015)
Paternal height	-	-	-	0.012	-	-0.015
(cm.)*Boy	-	-	-	(0.012)	-	(0.013)
Observations	889	889	989	989	989	989
R-squared	0.676	0.528	0.336	0.355	0.112	0.114

Notes: The left panel (I) correspond to OLS estimates on child length and growth in length (in centimeters) and the right panel (II) correspond to OLS estimates on child length-for-age and weigth-for-length z-scores including only maternal or both parents' height. In addition to the variables shown, these models control for child characteristics (age in months, weeks of premature birth, dummies for ethnic group/caste and birth order), parental characteristics (maternal and paternal years of education, maternal height, age, proficiency in main language and religion), household characteristics (number of child's brother and sisters, wealth index, indicators for piped water, sewerage and good-quality floor), a set of variables likely to be correlated with determinants of household formation (parental age difference, an indicator for whether parents belong to the same ethnic group/caste, indicators for whether parents share the same mother tongue, and an indicator for whether parents have never migrated from the community), and their interaction by child sex. *Significant at 10%, **Significant at 5%, ***Significant at 1%.

Unlike the first YL wave, waves 2 and 3 of the sample from Peru include paternal height. Following the procedure mentioned in Section 3.2, biological parents' height was imputed from wave 2 and 3 for those with missing information in the first wave. Since the sample with complete information on parental height might be different to the complete sample considered for Peru,[51] the right panel of Table 8 considers only a reduced sample with complete information on parental height. The two specifications, one similar to Table 3 (length-for-age) and the other to Table 4 (weight-for-length), include regressions where only maternal height and both parents' height are considered.

The effect of parental height on child weight-for-length is small and statistically insignificant. In fact, adding paternal height does not improve the fit of the model, relative to the specification containing only maternal height.[52] In contrast, maternal and paternal heights have a larger and significantly positive effect on children's length-for-age. Parental height is found to have a similar effect on boys' and girls' length-for-age. Moreover, adding paternal height to the first specification of child length-for-age slightly reduces the estimated effect of maternal power on children's health. This is likely due to a correlation between parents' height associated with their matching in the marriage market. The effect of bargaining power is still statistically insignificant. In the specification of child length-for-age, adding paternal height significantly improves the fit of the model relative to the one only containing maternal height.[53]

[51] Households with no-missing information of paternal height seem to be poorer than those with missing information of paternal height. This may reflect correlations between the presence of the father during the interview and unemployment status. They are statistically different in the following observable characteristics: children show smaller weight-for-length z-scores, parents are less educated, mothers are less able to speak the main language in the community, poorer household wealth index, less access to sewerage and good-quality floors, parents are not migrants and are mainly located in rural areas.

[52] The chi-squared value of the likelihood ratio test is 2.27, no statistically significant.

[53] The chi-squared value of the likelihood ratio test is 29.32, statistically significant at the 1% level.

To sum up, even though the inclusion of paternal height improves the estimation of child health, its correlation with maternal height suggests that accounting for maternal height already captures a substantial effect of the genetic component on the chosen indicators of child health. Therefore, I expect that the absence of information on paternal height in the other three samples would not bias the estimates significantly. Moreover, the robustness checks suggest that the considered child health outcomes are partially affected by differences in behaviour between parents.

7. Conclusion

This study explores whether the allocation of household resources to boys' and girls' health vary with parental gender preferences. Following Ermisch (2003) and Thomas (1994), I use a model that allows me to disentangle the effect of parental characteristics on the technology of child health production from the effect of parental preferences. In particular, I explore the effect that parental relative education (a *distribution factor*) has on the household demand for child health. The empirical analysis uses a unique dataset of young children in developing countries: the younger cohort of the first wave of the Young Lives (YL) project data for Ethiopia, India (Andhra Pradesh), Peru and Vietnam. These datasets contain rich information on several health and nutritional indicators, parental, household and community characteristics, and enable the comparison across countries. The YL data also permit to focus the analysis on very young children (aged 6-18 months) living in poor households in four developing countries.

In the absence of a source of exogenous variation (as in Duflo, 2000 and 2003; Lundberg et al., 1997), an indicator of relative education has been used to capture parental bargaining power. It is reasonable to argue that since having relatively more education is highly correlated with having more opportunities in the labour market and at higher wages, then the parent with relatively more education is likely to have more bargaining power in the household's resource allocation process. More education is also associated with acquiring relevant skills that might be used in the intrahousehold discussions about resource allocation (Beegle et al., 2001). In this study, I assume that more education indicates more power in the intrahousehold resource allocation process, and thus, women who are better educated than their husbands should be able to impose their preferences and allocate more resources towards commodities they care more about.

After testing for the presence of the son-biased stopping rule in the YL sample, the empirical strategy consists of estimating two indicators of child health that reflect short-run shocks (weight-for-length) and long-run shocks (length-for-age) z-scores, controlling for characteristics of the child, parents, household and genetic components. Moreover, in contrast to most of the literature using relative education as a measure of maternal power, I control for its potential endogeneity, associated with assortative mating of the parents in the marriage market. Finally, I perform some robustness checks controlling for the presence of behavioural effects driven the estimates and for additional controls (paternal height).

The estimated effects of maternal power on child health vary across countries and the indicator chosen. In the YL sample of Peru and Vietnam (and Ethiopia but not precisely estimated), maternal bargaining power has a general positive effect on children's health but it also has a differential effect, which suggests that mothers prefer to allocate more resources to their daughters in comparison to their sons. In contrast in the Indian sample, girls living with mothers with more power in the household receive fewer resources than girls in other households.

A possible channel through which maternal power may affect child health is related to breastfeeding practices. Particularly in India, empowered mothers weaned their children sooner than other mothers. In the absence of good supplementary food and hygienic conditions, weaning has a negative effect on child health, particularly increasing the risk of infections and diarrhoeal diseases, which is primarily reflected in the short-run indicator of weight-for-length. I find that girls are weaned sooner than boys in India and boys are weaned before girls in Peru. In some extent, it reflects differences in maternal behaviour related to child gender.

I do not find evidence of heterogeneous differential effect of maternal bargaining power on girls' and boys' health by urban/rural location and by the presence of other young children in the household turn out to be informative. However, among households located in rural areas of India and Peru there seem to be differences in the allocation of resources between boys and girls. Similarly, some evidence of competition for household resources affecting girls' health is also found in the samples from Peru and Vietnam.

The robustness checks confirm that differences in parental behaviour affect child health. Under the valid assumption that there is no sex-selective abortion or discrimination before birth, an initial specification of child weight and growth in weight for the four countries and a specification of child length and growth in length only for Peru show that differences in parental behaviour affects the current child status. Finally, the results are robust when including additional information about parental genotype characteristics (paternal height in addition to maternal height).

In sum, even though the YL countries present considerable diversity in socioeconomic and cultural backgrounds, some common effects have been found. In Peru and Vietnam, mothers with more bargaining power seem to allocate more resources to their daughters. In India, the opposite effect is found and boys show better health when mothers have more power in household decisions.

The relatively small effects found for the YL sample in India are likely to be due to the relatively better situation of Andhra Pradesh respect to the rest of the country. As Sommerfelt and Arnold (1998) and Marcoux (2002) argue, the large gender discrimination against girls seems to be focused on particular populations and not around the whole country.

The differential effects of maternal bargaining power on girls' and boys' health found in this paper are however small in magnitude. As Duflo (2005) mentions, even in countries with the strongest preferences for boys, households deciding under normal circumstances do not differentiate between boys and girls, but in extreme circumstances (e.g. negative economic shocks, such as droughts or floods), households are likely to discriminate between boys and girls.

The different effects of parental characteristics on boys' and girls' outcomes might respond to diverse motivations, which are beyond the goals of this paper. On the one hand, differences coming from technology are related to the sexual division of labour; father-son and mother-daughter pairs tend to spend more time together, have closer relationships and thus, each parent might play a more important role in the development of the child (Lamb, 1976 and 1987). On the other hand, each parent may have different expectations about gender differentials in future returns to investments and then parental preferences about investment might vary according to the gender of the child. Elderly mothers are more likely to have more contact with their children if at least one is a girl (Spitze and Logan, 1990). Hess and Waring (1978) shows that daughters, relative to sons, tend to give more support to their widowed mothers after the father's death. Gender differences might also respond to systematic gender differences in prices of human-capital investments (e.g. differences in relative wages for girls' and boys' labour, scarce supply of goods and services) and in these cases preferences are independent of any within household cost-profit analysis. It might also be that parental preferences for investment in human capital are oriented to redress imbalances among children.

8. References

Apps, P.F. and R. Rees. 1997. "Collective Labor Supply and Household Production." *Journal of Political Economy* 105 (1):178–90.

Arnold, F., S. Kishor, and T.K. Roy. 2002. "Sex-Selective Abortions in India." *Population and Development Review* 28 (4):759–785.

Barcellos, S.H., L. Carvalho, and A. Lleras-Muney. 2012. "Child Gender and Parental Investments In India: Are Boys And Girls Treated Differently?" Working Paper 17781, National Bureau of Economic Research.

Barker, D.J. 1990. "The Fetal and Infant Origins of Adult Disease." *BMJ: British Medical Journal* 301 (6761):1111.

Basu, K. and R. Ray. 2002. "The Collective Model of the Household and an Unexpected Implication for Child Labor: Hypothesis and an Empirical Test." Policy Research Working Paper Series 2813, The World Bank.

Becker, G.S. 1991. *A Treatise on the Family*. Cambridge, MA: Harvard Univ. Press.

Beegle, K., E. Frankenberg, and D. Thomas. 2001. "Bargaining Power within Couples and Use of Prenatal and Delivery Care in Indonesia." *Studies in Family Planning* 32 (2):130–146.

Behrman, J.R. 1988. "Intrahousehold Allocation of Nutrients in Rural India: Are Boys Favored? Do Parents Exhibit Inequality Aversion?" *Oxford Economic Papers* 40 (1):32–54.

Behrman, J.R. and A.B. Deolalikar. 1990. "The Intrahousehold Demand for Nutrients in Rural South India: Individual Estimates, Fixed Effects, and Permanent Income." *The Journal of Human Resources* 25 (4):665–696.

Bélanger, D., T.H.O. Khuat, J. Liu, T.T. Le, and V.T. Pham. 2003. “Are Sex Ratios at Birth Increasing in Vietnam?” *Population (English Edition, 2002-)* 58 (2): 231–250.

Bloom, D.E. and P.H. Reddy. 1986. “Age Patterns of Women at Marriage, Cohabitation, and First Birth in India.” *Demography* 23 (4):509–523.

Blundell, R., P.A. Chiappori, T. Magnac, and C. Meghir. 2007. “Collective Labour Supply: Heterogeneity and Non-Participation.” *The Review of Economic Studies* 74 (2):417–445.

Blundell, R., P.A. Chiappori, and C. Meghir. 2005. “Collective Labor Supply with Children.” *Journal of Political Economy* 113 (6):1277–1306.

Browning, M., F. Bourguignon, P.A. Chiappori, and V. Lechene. 1994. “Income and Outcomes: A Structural Model of Intrahousehold Allocation.” *Journal of Political Economy* 102 (6):1067–96.

Browning, M. and P.A. Chiappori. 1998. “Efficient Intra-Household Allocations: A General Characterization and Empirical Tests.” *Econometrica* 66 (6):1241–1278.

Browning, M., P.A. Chiappori, and Y. Weiss. 2011. *Family Economics*. Cambridge University Press.

Chakravarty, A. 2012. “Gender Bias in Breastfeeding and Missing Girls in Africa: The Role of Fertility Choice.” *Mimeo*.

Chiappori, P.A. 1988. “Rational Household Labor Supply.” *Econometrica* 56 (1):63–90.

———. 1992. “Collective Labor Supply and Welfare.” *Journal of Political Economy* 100 (3):437–467.

———. 1997. “Introducing Household Production in Collective Models of Labor Supply.” *Journal of Political Economy* 105 (1):191–209.

Chiappori, P.A. and I. Ekeland. 2009. "The Microeconomics of Efficient Group Behavior: Identification." *Econometrica* 77 (3):763–799.

Dauphin, A., E. Lahga, B. Fortin, and G. Lacroix. 2011. "Are Children Decision-Makers within the Household?" *The Economic Journal* 121 (553):871–903.

Dev, MS. and C. Ravi. 2003. "Macroeconomic Performance and Policies." *Economic and Political Weekly* 38 (12-13):22–29.

Duc, L.T., N.P. Ngoc, T.M. Chau, N.V. Tien, and V.T Son. 2008. "Young Lives: Vietnam Round 2 Survey Report." *Young Lives, DFID, University of Oxford*.

Duflo, E. 2000. "Child Health and Household Resources in South Africa: Evidence from the Old Age Pension Program." *The American Economic Review* 90 (2):393–398.

———. 2003. "Grandmothers and Granddaughters: Old-Age Pensions and Intrahousehold Allocation in South Africa." *World Bank Economic Review* 17 (1):1–25.

———. 2005. "Gender Equality in Development." *BREAD Policy Paper* 11.

Ermisch, J. 2003. *An Economic Analysis of the Family*. Princeton University Press.

Escobal, J. and E. Flores. 2008. "An Assessment of the Young Lives Sampling Approach in Peru." *Young Lives Technical Note, DFID (University of Oxford)* 3.

Fafchamps, M. and A.R. Quisumbing. 2002. "Control and Ownership of Assets within Rural Ethiopian Households." *Journal of Development Studies* 38 (6):47–82.

———. 2005. "Marriage, Bequest, and Assortative Matching in Rural Ethiopia." *Economic Development and Cultural Change* 53 (2):347–380.

Frankenberg, E. and D. Thomas. 2003. "Household Decisions, Gender and Development: Measuring Power." *In A. Quisumbing., Ed., Household Decisions,* *Gender and Development: A Synthesis of Recent Research* International Food Policy Research Institute, Washington, D.C.

Gertler, P. and P. Glewwe. 1992. "The Willingness to Pay for Education for Daughters in Contrast to Sons: Evidence from Rural Peru." *The World Bank Economic Review* 6 (1):171–188.

Gitter, S.R. and B.L. Barham. 2008. "Women's Power, Conditional Cash Transfers, and Schooling in Nicaragua." *The World Bank Economic Review* 22 (2):271–290.

Guilmoto, C.Z. 2007. "Characteristics of Sex-Ratio Imbalance in India, and Future Scenarios." In *4th Asia Pacific Conference on Reproductive and Sexual Health and Rights, October 29-31, 2007, Hyderabad, India*. 1–32.

———. 2009. "The Sex Ratio Transition in Asia." *Population and Development Review* 35 (3):519–549.

Haddad, L. and J. Hoddinott. 1994. "Women's Income and Boy-Girl Anthropometric Status in the Cote d'Ivoire." *World Development* 22 (4):543–553.

Hao, L., V.J. Hotz, and G.Z. Jin. 2008. "Games Parents and Adolescents Play: Risky Behaviour, Parental Reputation and Strategic Transfers." *The Economic Journal* 118 (528):515–555.

Haughton, D. and J. Haughton. 1997. "Explaining Child Nutrition in Vietnam." *Economic Development and Cultural Change* 45 (3):541–556.

Hess, B. and J. Waring. 1978. "Parent and Child in Later Life: Rethinking the Relationships." In *Child Influences on Marital and Family Interactions: A Life Span Perspective*, edited by Richard Lerner and Graham Spanier. New York: Academic Press, 241–273.

Hoddinott, J. and L. Haddad. 1995. “Does Female Income Share Influence Household Expenditures? Evidence from Côte d’Ivoire.” *Oxford Bulletin of Economics and Statistics* 57 (1):77–96.

Jayachandran, S. and I. Kuziemko. 2011. “Why Do Mothers Breastfeed Girls Less than Boys? Evidence and Implications for Child Health in India.” *The Quarterly Journal of Economics* 126 (3):1485–1538.

Knodel, J., V.M. Loi, R. Jayakody, and V.T. Huy. 2005. “Gender Roles in the Family.” *Asian Population Studies* 1 (1):69–92.

Lamb, M.E. 1976. “The Role of the Father: An Overview.” In *The Role of the Father in Child Development*, edited by Michael Lamb. New York: Wiley, 1–36.

———. 1987. *The Father’s Role: Cross Cultural Perspectives*. London: Lawrence Erlbaum Associates.

Lundberg, S. 1988. “Labor Supply of Husbands and Wives: A Simultaneous Equations Approach.” *The Review of Economics and Statistics* 70 (2):224–235.

Lundberg, S., J. Romich, and K. Tsang. 2009. “Decision-Making by Children.” *Review of the Economics of the Household* 70 (1):1–30.

Lundberg, S.J., R.A. Pollak, and T.J. Wales. 1997. “Do Husbands and Wives Pool Their Resources? Evidence from the United Kingdom Child Benefit.” *Journal of Human Resources* 32 (3):463–480.

Marcoux, A. 2002. “Sex Differentials in Undernutrition: A Look at Survey Evidence.” *Population and Development Review* 28 (2):275–284.

Martorell, R. and J.P. Habicht. 1986. “Growth in Early Childhood in Developing Countries.” In *Human Growth: A Comprehensive Treatise*, vol. 3, edited by F. Falkner and J. Tanner. New York: Plenum Press, 2 ed., 241–262.

Miguel, E. and M. Kremer. 2004. "Worms: Identifying Impacts on Education and Health in the Presence of Treatment Externalities." *Econometrica* 72 (1):159–217.

NCTPE, (National Committee on Traditional Practices of Ethiopia). 2003. "Old Beyond Imaginings: Ethiopia, Harmful Traditional Practices." Tech. rep., Addis Ababa, Ethiopia: NCTPE.

Parsons, D.O. and C. Goldin. 1989. "Parental Altruism and Self-Interest: Child Labor among Late Nineteenth-Century American Families." *Economic Inquiry* 27 (4):637–659.

Pitt, M.M. and M.R. Rosenzweig. 1985. "Health and Nutrient Consumption Across and Within Farm Households." *The Review of Economics and Statistics* 67 (2):212–223.

Quisumbing, A.R. 2003. "Food Aid and Child Nutrition in Rural Ethiopia." *World Development* 31 (7):1309–1324.

Reggio, I. 2011. "The Influence of the Mother's Power on Her Child's Labor in Mexico." *Journal of Development Economics* 96 (1):95–105.

Schady, N. and J. Rosero. 2007. "Are Cash Transfers Made to Women Spent Like Other Sources of Income?" Policy Research Working Paper Series 4282, The World Bank.

Schultz, T.P. 2005. "Productive Benefits of Health: Evidence from Low-Income Countries." *Health and Economic Growth: Findings and Policy Implications. MIT Press, Cambridge MA*: 257–286.

Sommerfelt, A.E. and F. Arnold. 1998. "Sex Differentials in the Nutritional Status of Young Children." *Too Young to Die: Genes or Gender*: 133–153.

Spitze, G. and J. Logan. 1990. "Sons, Daughters, and Intergenerational Social Support." *Journal of Marriage and Family* 52 (2):420–430.

Srinivasan, S. and A. Bedi. 2011. "Ensuring Daughter Survival in Tamil Nadu, India." Oxford Development Studies 39 (3):253–283.

Strauss, J. and D. Thomas. 1998. "Health, Nutrition, and Economic Development." *Journal of Economic Literature* 36 (2):766–817.

Subramanian, S. and A. Deaton. 1991. "Gender Effects in Indian Consumption Patterns." *Sarvekshana* 14 (4):1–12.

Thomas, D. 1990. "Intra-Household Resource Allocation: An Inferential Approach." *The Journal of Human Resources* 25 (4):635–664.

———. 1994. "Like Father, like Son; Like Mother, like Daughter: Parental Resources and Child Height." *The Journal of Human Resources* 29 (4):950–988.

United Nations Development Programme, UNDP. 2000. "Human Development Report 2000." Tech. rep., New York: Oxford University Press.

WHO. 1986. "Use and Interpretation of Anthropometric Indicators of Nutritional Status." *Bulletin of the World Health Organization* 64 (6):929–41.

WHO Multicentre Growth Reference Study Group. 2006. *WHO Child Growth Standards: Length/height-for-age, Weight-for-age, Weight-for-length, Weight-for-height and Body Mass Index-for-age: Methods and Development*. Genève: World Health Organization.

World Bank. 2004. "World Development Indicators 2004." Tech. rep., Washington, D.C.: The World Bank.

Yamaguchi, K. 1989. "A Formal Theory for Male-Preferring Stopping Rules of Childbearing: Sex Differences in Birth Order and in the Number of Siblings." *Demography* 26 (3):451–465.

9. Appendix A

Table A.1: Selected Poverty and Development Indicators in 2000

A. General Statistics

Country	HDI (2000) rank (a)	Pop. below $1 a day (%) (b)	Life expectancy at birth (c)		% population with access to… (c)					
					Drinking water			Improved sanitation		
			Female	Male	Rural	Urban	Total	Rural	Urban	Total
Ethiopia	171	26.3	55	52	19	87	29	4	24	7
India	128	34.7	62	60	77	94	82	13	49	23
Peru	80	18.1	71	67	56	91	81	28	80	65
Vietnam	108	17.7	72	68	72	94	77	43	78	51

B. Child Health and Nutritional Status Statistics

Country	Age (years)	% below median… (d)						Infant mortality rate* [per 1000 live b.] (c)	
		weight/age		length/age		weight/length			
		3SD	2SD	3SD	2SD	3SD	2SD	Female	Male
Ethiopia	[0.00 - 0.49]	9.8	20.8	9.6	22.2	6.1	16.1	80	104
	[0.50 - 0.99]	17.2	37.1	18.0	37.4	7.6	21.5		
	[1.00 - 1.99]	19.3	44.6	34.4	60.0	6.0	18.5		
India	[0.00 - 0.49]	14.5	34.3	11.4	27.4	10.3	24.4	67	65
	[0.50 - 0.99]	16.0	38.4	16.6	36.5	7.9	22.7		
	[1.00 - 1.99]	21.7	47.0	32.7	58.4	7.2	21.0		
Peru	[0.00 - 0.49]	1.4	4.5	4.2	11.9	0.8	1.5	31	35
	[0.50 - 0.99]	2.0	4.8	5.2	16.6	0.5	2.5		
	[1.00 - 1.99]	1.3	5.6	11.4	32.4	0.4	1.2		
Vietnam	[0.00 - 0.49]	2.8	8.1	9.0	17.2	0.8	7.8	23	23
	[0.50 - 0.99]	5.8	12.7	7.9	29.0	1.5	5.9		
	[1.00 - 1.99]	6.7	27.5	21.4	47.5	1.1	7.4		

Sources: (a) Human Development Report 2000 (UNDP, 2000). (b) World Development Indicators (The World Bank, 2004). (c) WHO Statistical Information System (WHOSIS website: http://www.who.int/whosis/en/index.html). *Probability of dying between birth and age 1. (d) WHO Global Database on Child Growth and Malnutrition (website: http://www.who.int/nutgrowthdb/en/), SD = Standard Deviations. Note: The percentage of population living in poverty in Vietnam corresponds to 1998. The athropometric measures in India corresponds to 1998 - 1999.

A.2 General Description of the YL Sample by Country

The Young Lives team in Ethiopia in charge of the round 1 sample was formed by the Ethiopian Development Research Institute (with researchers from Addis Ababa University) and Save the Children UK in Ethiopia. The whole country YL sample consists of 1,999 children. However, further data quality checking and sample definitions reduced the sample size up to 75 percent of the original sample, depending on the analysis performed.[54] In addition to the five regions presented above (Addis Ababa, Amhara, Oromia, SNNP and Tigray), the sample is mainly distributed in nine ethnic groups (Agew, Amhara, Gurage, Hadiva, Kambata, Oromo, Sidama, Tigrian, Wolavta) and five religious groups (Muslim, Catholic, Protestant, Orthodox, Evangelist).

In India, Young Lives was implemented in Andhra Pradesh by two organisations, the Centre for Economic and Social Studies (CESS), Hyderabad and Save the Children UK (India). The fieldwork for round 1 was undertaken in Andhra Pradesh between September and December 2002. The sample of children is representative of the three regions of Andhra Pradesh: Rayalaseema, Coastal Andhra and Telangana. The original sample size for the 'younger cohort' contains information on 2011 households, from which further quality checks reduce the sample up to 95 percent of its original size.[55] In addition to the three regions, the YL sample in Andhra Pradesh is mainly distributed in four castes (Scheduled Castes, Scheduled Tribes, Backward Castes and Other Castes) and five religious groups (Christian, Muslim, Buddhist, Hindu, none).

[54] This rate does not take in consideration a regression model presented later on the paper on 'growth in weight' that only consider 276 observations for Ethiopia due to a high number of missing observations on the 'birth weight' variable.

[55] Andhra Pradesh suffered the lowest cut in sample size in the four YL countries. However, this rate does not take in account a regression model presented later on the paper on 'growth in weight' that only consider 842 observations for India due to missing observations on the 'birth weight' variable.

In Peru the Young Lives partners for round 1 were the Nutrition Research Institute (IIN), the Group for the Analysis of Development (GRADE) and the local office of Save the Children UK. The original YL 'young cohort' sample consists of 2,052 children, but was reduced due to further data quality checks and sample definitions by up to 81 percent of the original sample, depending on the analysis performed.[56] In addition to the three geographical regions presented above (Coast, Mountains and Jungle) the sample is mainly distributed in five ethnic groups (White, Mestizo-including Andean Indian-, native of the Amazon, Black and Asiatic) and four religious groups (Catholic, Evangelist, Mormon and none).

In Vietnam the Young Lives round 1 was implemented by three organisations: the Research and Training Centre for Community Development (RTCCD), the General Statistics Office (GSO) and Save the Children UK. The original sample size for the 'younger cohort' is 2,000 households. Further sample definitions for the objectives of this paper and quality of data checks reduced the sample size up to 83 percent of its original size.[57] The Vietnamese YL sample is distributed in four regions (Northern uplands, Red river delta, Central coastal and Mekong river delta); among seven ethnic groups (Kinh, H'mong, Ede, Ba Na, Nung, Tay and Dao), and among six religious groups (Christian, Buddhist, Protestant, Ancestor Worship, Cao Dai and none).

---

[56] This rate does not take in consideration a regression model presented later on the paper on 'growth in weight'. Peru is however, the country with less missing values on 'birth weight' variable and then the regression is performed with 1,664 households.

[57] Similar to Peru, the YL sample for Vietnam shows relatively less missing values on the variable 'birth weight' and then the regression model on 'growth in weight' could be performed without losing so many households.

Figure A.1: Weight-for-Age, by Sex and Country

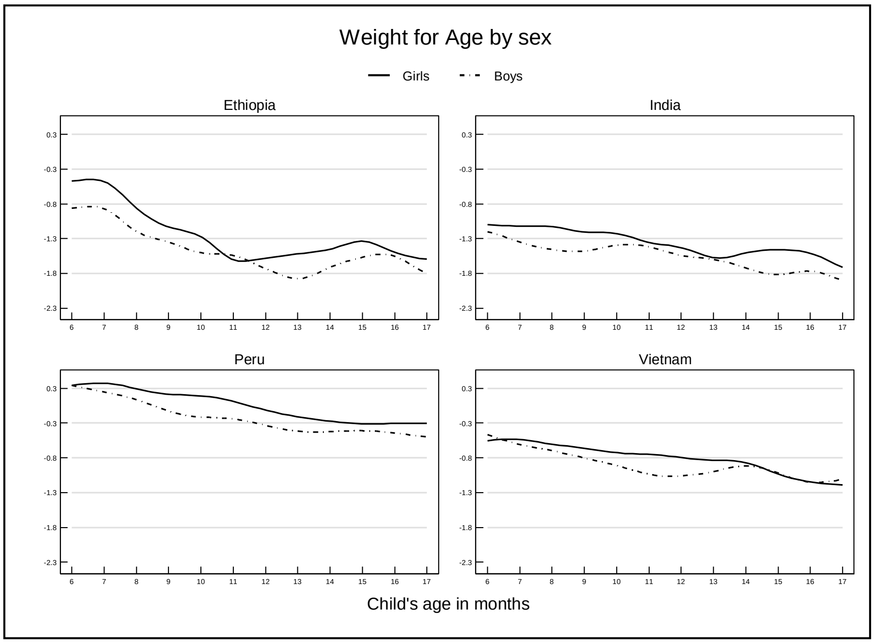
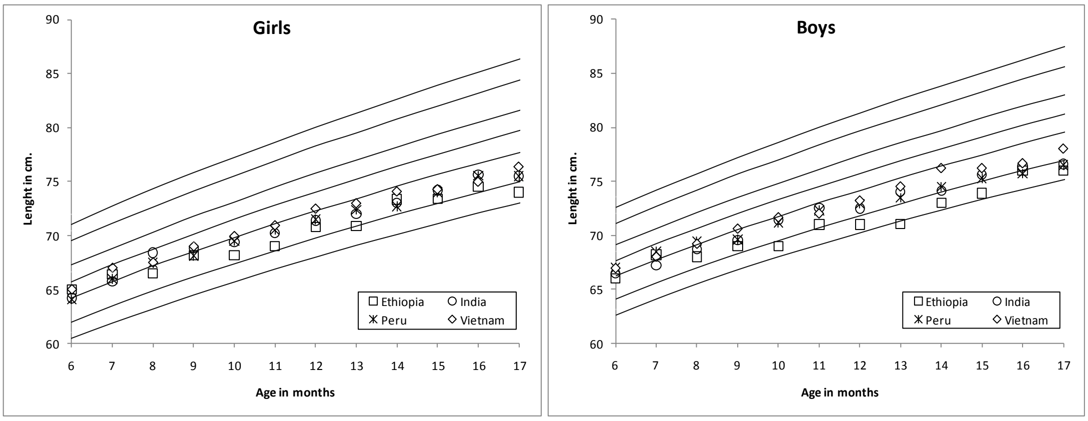


Figure A.2: YL Children's Length on WHO Growth Curves



Source: Own calculations using the YLS - round 1 datasets