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Intra-household Resource Allocation when Food Prices Soar
Impacts on Child Growth in Indonesia

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Abstract

An unanticipated spike in food prices can increase malnutrition among the poor with lasting consequences, but parents can protect the most vulnerable within the family by distributing scarce food to minimize adverse impacts. To find evidence of this strategy, we use anthropometric and consumption data from Indonesia, collected before and after the 2007/08 food price crisis. Our results indicate that soaring food prices had a significant and uneven impact on growth among children. Using household fixed effects, we find that the negative impact was significantly larger among larger children, as measured by the initial height z-score. We find that children with low height z-scores at the start of the crisis gained ground relative to their peers during the crisis, consistent with food-resource allocations in their favor. The findings remain robust when controlling for possible differential impacts by gender, family size and food producer status. We conclude that the food price crises had negative long-term impacts on children, and that parental behavior protected the most vulnerable. For Indonesian policy makers, our results indicate that safeguarding family food security should be a priority when targeting specific groups of children is difficult.

Keywords: food price crisis, child growth, human capital, siblings, Indonesia

JEL Classifications: Q11, Q18, O12, O15

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1. Introduction

Poor households devise multi-faceted livelihood strategies to manage and mitigate many risks that jeopardize their welfare. Informal insurance mechanism is a vital component that relies on help from extended family members and neighbors in time of need. However, such insurance mechanisms can fail when systemic events affect all members of the mutual scheme. In the absence of effective intervention by aid providers, families are left to cope based on their own limited resources. There are several lines of defense, as families draw down savings, liquidate assets and change the composition of their consumption; however, in desperate times, the final choice can come down to deciding how food will be shared among family members.

How parents treat children with different conditions when facing a food security crisis is not well studied. More generally, the question of how parents invest in children with different conditions in response to negative shocks is unclear in the literature, though some evidence has been provided. Griliches (1979) conjectures that parents compensate for gaps in child endowments by human capital investments and studies, for example by Pitt et al. (1990) and Li, et al. (2010), provide evidence supporting compensating behavior. Parental investments can also reinforce gaps among children under certain conditions (Becker and Tomes, 1976).

In this paper, we focus on surging food prices in Indonesia during the 2007/08 international food crises, a cardinal systemic risk that diminishes the capacity of poor households to meet minimal nutritional needs. To look for evidence of intra-family resource allocation, we examine within-sibling variations in growth, which are attributable to variation in nutrition levels during the crises. We calculate local food price indices to account for differences in the transmission of global price shocks and the effects of differing diet compositions. Further, we construct standardized child anthropometric nutrition-status measures, height-for-age and weight-for-age z-scores, to detect health outcomes (WHO, 2006). This measure is also used to identify initial differences among siblings.

The analysis uses Indonesian household panel data collected in 2007 and 2010 from rural villages. The first round was fielded during the second to third quarter of 2007, immediately before the financial crisis; and the follow-up survey was conducted in 2010 after the crisis had subsided. The timings offer us an ideal setting to assess the impact of food price spikes on child growth. The geographic coverage of the survey over seven provinces in five macro regional islands takes advantage of significant variations in village-level food price changes.

Our findings suggest that soaring prices had a significant and negative impact on child growth, but also show that negative impacts were more severe among less vulnerable children in the family, that is, young family members with initially higher height for age z-score. This suggests that parents

protect the most vulnerable of their children as food was allocated among family members. In a study to investigate whether parents make compensatory and reinforcing investments in different dimensions of human capital across children in response to health shocks, Yi et al. (2015) show that the family acts as a net equalizer in response to child early health shocks across children. The empirical results of this paper are similar to theirs in spirit that the most vulnerable children are protected by their parents from adverse impacts of the food price crisis. The evidence supports the proposition, conjectured by Griliches (1979), that parents tend to equalize growth among their children in response to negative shocks.

Our findings imply that the food price crises can have long-term impacts on child human capital formation, but the impact is not uniform across siblings. Parents intervene to protect the most vulnerable among their children to equalize their long-term human capital outcomes. Policy implications are discussed in the concluding section.

2. Background

Following decades of relative stability, global food prices spiked between the summer of 2007 and June 2008. Initially triggered by poor wheat harvests in Ukraine and Australia at a time when grain inventories were already low, the crisis spread quickly to the thinly traded rice market due, in part, to over-reactive policy decisions by key global market participants. Rice prices, which had climbed modestly during the summer, registered month-over-month increases of 21, 24 and 42 percent in February, March, and April of 2008. The average price for April 2008 of \$US 907 per ton is the highest monthly price recorded in the World Bank's series for rice, which begins in 1960. The scale and rapid onset of the price increases unleashed widespread social and political unrest (Slayton, 2009; Bellemare, 2015).

The first major food price-spike of the 21st Century was large by modern standards. In a review of international food prices, Larson (2018) notes that the crises anchored a three-year period in which real food prices increase by 54 percent, a run-up surpassed only twice between 1900 and 2015. The largest run-up occurred during by a 5-year recovery period, 1932-37, when agricultural prices returned to earlier levels following a sharp decline at the start of the Great Depression. The second period, which most resembles the 2007/08 event, occurred during the 1972-74 OPEC crisis when the World Bank's food price index shot up by 70 percent.

Figure 1: Monthly international rice prices, 2000-2017

Note: Price series is for 5 percent broken, milled white rice, f.o.b. Bangkok. Source: World Bank (2018).

The sharp change in food prices played out differently among households, creating a continuum of outcomes, even among the poor (Swinnen, and Squicciarini, 2012). Producing food for household consumption is a mainstay of rural livelihood strategies, and this provided a measure of protection for many poor rural households. Moreover, households producing a surplus of food likely benefited from higher prices, as did households with livelihoods linked to agriculture. Consequently, poor rural smallholders are thought to be less vulnerable to food price spikes than the landless and urban poor (Ruel et al, 2010).

The crisis garnered much attention in the academic literature. To date, most studies about the crisis fall into one of three categories. The first set of studies focuses on the events precipitating the crises and the implications for policy. Examples include Mitchell (2008); Demeke, Pangrazio, and Maetz (2008); Headey and Fan (2008); Meijerink, Roza, and van Berkum (2009); Slayton (2009); Webb (2009); Brinkman et al. (2010); Timmer (2010), Headey (2011); Wodon and Zaman (2010); Berazneva, Julia, and Lee (2013); Rutten, Shutes, and Meijerink (2013); and Clapp (2017). A second group centers on simulation-based estimates of the impact of the crisis on the ranks of the poor. Often cited examples include Ivanic and Martin (2008; 2014); USDA (2009); de Hoyos and Medvedev (2011) and Warr and Yusuf (2014). Another set of studies focuses on the degree to which the surge in international prices affected local prices, for example, Dawe (2009); Ortiz, Chai, and Cummins (2011); Dawe and Maltoglou (2014); Yang et al. (2015); and Jamora and von Cramon-Taubadel (2016).

Still, few papers to date focus on the chief concern of policy makers at the time - whether the crisis would have a negative impact on the nutritional status of the poor. Among these, two of the papers rely partly on micro-based simulation exercises to measure price impacts in Malawi. In the first paper, Ecker and Qaim (2011) use household data from Malawi to estimate income and price

elasticities of food demand and nutrient consumption, then use the estimated system to simulate price and income interventions. They conclude that income-related policies are better suited than price policies to improve nutrition. The second paper, by Harttgen, Klasen and Rischke (2016), also uses a simulation-model based on micro-economic survey data from Malawi to gauge the impact of high staple-food prices on calorie consumption and poverty. They conclude that price shocks have a considerable impact on food security with particularly strong effects on poor net food buyers in rural and urban areas. A third paper, by Dimova and Ghakou (2013), uses household survey data from Côte d'Ivoire to identify winners and losers during the crisis based on calculated benefit ratios. They conclude that urban consumers, who depend more on rice than rural households, were hurt most by the global crisis. A final paper worth mentioning focuses on a longer time horizon that includes the global price crises. Using data from 2001 to 2008, Grace, Brown and McNally (2014) show that birth weights in rural Kenya rise and fall with local maize prices.

To the best of our knowledge, only three peer-reviewed papers use household data to measure the effects of the crisis on food security. All find that the crisis negatively impacted the poor despite the precautions taken by households and governments.

Among them, only our companion paper (Yamauchi and Larson, 2019) examines the effects of the crisis in Indonesia. In the paper, we use the same data used in this paper to show that soaring food prices had a significant and negative impact on child growth for families, but that the worse effects of the crisis were fully mitigated among households that produced a portion of their own food.

The remaining papers find evidence of impact from the crisis in El Salvador and Bangladesh. The earliest paper, by de Brauw (2011), uses household data from El Salvador to show that height-for-age z-scores declined among young children during the crises, but that households with access to remittances from international migrants experienced lower declines. In their paper, Akter and Basher (2014) note that the direct impacts of the international price crisis are hard to extract in Bangladesh because the international price crisis occurred contemporaneously with two natural disasters, monsoon floods followed by Cyclone Sidr. That said, the authors conclude that soaring food prices heightened food insecurity in rural Bangladesh, a condition that was aggravated by subsequent income shocks.

Furthermore, we examine whether poor households reallocated diminished supplies of food to protect the most vulnerable family members. Redirecting food among family members is the last action families can take when other precautions against food insecurity fail. Although no other papers have examined household reallocations in the context of the international food price crises, the phenomenon is well studied in other settings.

3. Empirical Strategy

Our empirical strategy includes three key elements. First, we assume that the 2007/08 food price spike was not fully anticipated, thus, a component of the observed changes in food prices is exogenous. As discussed, the 2007/08 price event succeeded an extended period of relative stability for food prices, and the scale and sudden on-set of the event make it unlikely that the full extent of the eventual run-up was fully anticipated even after it had begun. Moreover, we find statistical support for this assumption in our earlier paper (Yamauchi and Larson, 2019). Second, based on evidence from other food security crises, we posit that the adverse impacts of food price spikes may be non-uniform among the siblings. And finally, although we are not able to view the full price trajectory, we expect to find evidence of forced periods of malnutrition due to the crises in the stocks of physical capital evident in our anthropometry data.

3.1 Measurement, Identification and Specification

We use child nutrition status, specifically the height-for-age and weight-for-age z-scores, to measure child human capital. To measure price changes, price indices are constructed from price and quantities, and consumption data captured in two rounds before and after the 2007/08 food price crisis. First, we compute the median price for each food commodity in each village for each period. Both value-based expenditure shares and calorie-based consumption shares are used to construct the indices that are used for our analysis. For the value-based indices, village-average consumption quantities are valued by median village-prices to calculate an aggregate consumption budget and commodity expenditure shares. For the caloric measures, observed consumption quantities are converted to calories; total calories consumed are then calculated along with commodity shares. The base year for the indices is set to 2007.

The consumption data distinguishes between household-produced food and food from other sources. We use this to calculate a discrete indicator based on the proportion of food consumption (in terms of calories) supported by own production. We then use the indicator in our estimation of the consequences of food-producing and other livelihood strategies during the crises.

We use the following first-differenced model to measure the impact of food price spikes on child nutrition status with household fixed effects:

$$\Delta g_{ij} = \beta_1 g_{ij}^0 \Delta P_v + X_{ij}^0 B' + \Delta \varepsilon_{ij} \quad (1)$$

where Δg_{ij} is change in nutrition status (weight or height) of child i , in household j and village v , ΔP_v is the change in village-level price index (using either quantities or calories consumed weights), and g_{ij}^0 is the initial nutrition status (measured by the initial height z-score). X_{ij}^0 is a vector of initial

individual characteristics, including the child's age, gender, birthweight, and g_{ij}^0 . The term $\Delta\varepsilon_{ijv}$ is an error difference term, with a distribution centered on zero. The set $[\beta_1, B]$ are parameters to estimate.

A potential correlation between ΔP_v and g_{ij}^0 is a concern. First, the household may invest in child human capital to increase their resilience to possible food price spikes if they believe that such events are likely to occur. To minimize the risk of a possibly endogenous ΔP_v we use a Laspeyres index, based on the 2007 pre-crisis consumption bundles:

$$\Delta P_v = \frac{\sum_s p_{v,s,t} q_{v,s,0}}{\sum_s p_{v,s,0} q_{v,s,0}} (100) = 100 \sum_s \frac{p_{v,s,t}}{p_{v,s,0}} w_{v,s} \quad (2)$$

where p and q are village-median prices and village-average consumption values for commodities $s = 1, 2, \dots, S$ for $t = 0 = 2007$ and $t = 1 = 2010$; the share-weights are $w_{v,s} = [p_{v,s,0} q_{v,s,0} / \sum_s p_{v,s,0} q_{v,s,0}]$. As the index uses weights determined prior to the price spike, it represents a village-level measure of the potential first-order price shock prior to any adjustments of expenditures. The measure also accounts for fixed village-level transaction costs arising from commodity-and-place specific policies and markets. Further, the index is based on pre-crises price, therefore, the 2010 index values give the percentage change in village prices, weighted by value or by calories. Moreover, we use household fixed effects to have inference from the within-sibling variations. Consequently, a variation in parental behavior toward siblings, prior to the crisis, should be accounted for.

3.2 Descriptions of the Survey and Data

This section describes the survey data from Indonesia. The data is from panel household surveys in Indonesia, conducted jointly by Indonesian Center for Agricultural Socio Economic and Policy Studies, International Food Policy Research Institute, and Japan International Cooperation Agency. The primary source of our data is the village and household surveys conducted in 2007 for 98 villages in 7 provinces (Lampung, Central Java, East Java, West Nusa Tenggara, South Sulawesi, North Sulawesi, and South Kalimantan) under the Study of Effects of Infrastructure on Millennium Development Goals in Indonesia (IMDG-1) by the Japan Bank for International Cooperation (JBIC) and the International Food Policy Research Institute. Figure 2 shows locations of surveyed villages. In 2010, the survey team revisited all the 98 sample villages to re-interview sample households and their splits. Out-migrants were also tracked through either direct or phone interviews.

Figure 2 Locations of surveyed villages

Source: IMDG 2007, 2010

The 2007 survey was designed to overlap with villages covered in the 1994/95 PATANAS survey conducted by ICASEPS to build household panel data. The 1994/95 PATANAS survey was representative of Indonesia's major agro-climatic zones. In 2007, the ICASEPS/IFPRI team visited those villages to expand the scope of research as a general household survey. Thus, the IMDG-1 sample inherited the representation of major agro-climatic zones in the country. Moreover, the 2007 round added 51 new villages in the same areas (sample districts). These new villages were selected with the following criteria. Villages that had received relatively substantial amounts of government infrastructure projects during the period of 1995 to 2005 (mainly funded by either the Japan Bank for International Cooperation or the World Bank) were listed from the sample districts. The new villages were randomly sampled from the list. In the above procedure, the expanded sample also represents the major agro-climatic zones. When revisiting the original-sample villages in 2007, 20 households from the 1994/95 sample were resampled per village, and the split households (defined below) were included. In the new villages, 24 households were sampled in each village. 2007. The total number of households in the 2007 survey was 2262. The 2010 survey followed the sample households in all 98 villages.

Between the surveys, some household members separated from the original 2007 households to start their own new households or to join other households. For survey purposes, a split household is defined as a new household in which (i) a member of the original household became the head of a new household, and (ii) the household resides in the same village. By contrast, household members

who moved out of the original village or simply joined other households within the village are not called splits in the survey. The total number of households including split cases is 2442. From 2007 to 2010, 9.25% of the 2010 households were split households under the above definition. 12.97% of the sample individuals recorded as household members in 2007 migrated out, that is, either moved out of their village or joined other households in their village. The proportion is 11.82% once the group is restricted to those aged 20-55 in 2007.¹

Anthropometry data were collected in both 2007 and 2010. Child's height, weight, and his/her mother's weight were measured in the field. In addition, birthweight was recorded from official sources. However, the age coverage differs between the two rounds. In the 2007 round, children aged 0 to 6 were covered, and the 2010 round extended the age range to cover children aged up to 12. Following a standard procedure, we calculated the height-for-age z-scores for children aged 0 to 12 and the weight-for-age z-scores for children aged 0 to 10. In the panel analysis, we use the sample of children aged 0-6 in 2007 (thus, corresponding to the cohort aged 3-9 in 2010). As we look at potential effects from intra-household food allocation, we restrict our sample to households with more than one child. All told, this leaves us with a panel of 353 children.

The surveys collected consumption and expenditure data using the following formats. For food consumption, which is of our primary interest, respondents were asked to provide item-by-item information on quantity consumed and, if purchased, the purchase price during a one-week recall period. They were also asked about the share of total consumption produced by the household and the share received as a gift. All physical measurement units were converted to kilograms, either in the field or during processing, and we use prices per kilogram as our price measure. From the above data, we also computed the proportion of calories produced by own production. Expenditure and calories shares were constructed for six food subcategories: (i) staple foods, (ii) fish, meat, tofu and tempe, (iii) beans and pulses, (iv) vegetables and fruits, (v) milk and eggs, and (vi) all other food items.²

We use both calorie-weighted and value-weighted indices in our analysis because of potentially significant differences in the way households adjust to higher prices. Sometimes, households adapt by moving to lower-quality lower-priced commodity grades, for example, by purchasing bags of rice with a higher proportion of broken grains. Especially for staple foods, these lower grades usually

¹ In the case of split households, child anthropometry data were collected since the full household questionnaire was used. However, in the case of out-migrants who joined other households in the same village or moved out from the village, child anthropometry data were compromised since due to time constraint, the survey team had to shorten the household questionnaire. The tracking rules have some implications on potential attrition bias, which turned out to be inconsequential in our analysis below.

² The following food groups are included in "Other": (vii) oil/fat, (viii) dried foods, (ix) spices, (x) sugar/sweets, and (xi) beverages.

provide the same number of calories per kilogram. Consequently, a value-weighted index can underestimate the impact of a price increase for staple goods when nutritionally equivalent lower-priced grades are available. A calorie-weighted index addresses this concern. Alternatively, an increase in food prices can lead households to shift to food items that are less nutritious, but which generate the same number of calories; for example, when consumers switch from protein rich meat to calorie-rich staples. In this case, caloric weights mask an expenditure adjustment with important consequences for childhood development.³ The two types of adjustment are not mutually exclusive, and households may adjust in both ways.

As a practical matter, the differences are small in our setting. Figure 3 compares the distribution of village Laspeyres price indices for 2010 using value-share and calorie-share weights. The calorie-weighted index is dominated by the value-weighted index, but the two distributions are quite similar.

Figure 3: Distributions for value-weighted and calorie-weighted price indices.

Table 1 shows descriptive statistics of the Laspeyres price index by province using either quantities or calories as weights. The table illustrates two points. First, the index is larger when weights are constructed from consumption quantities. This implies that households tend to consume more calories from foods that had smaller price increases in the subsequent period. Second, regardless of

³ See, for example, the discussion in D'Souza and Jolliffe (2014) and the references therein.

which index is used, the means are quite similar across provinces, which indicates that the impact of the crises on food prices were national rather than provincial. Nevertheless, we find significant within-province variations of food price changes, which we use in the child-growth regressions.

Table 1: Value and calorie weighted Laspeyres indices for 2010, by province

	Villages	Mean	Std. Dev.	Min	Max
<u>Lampung</u>					
Value-weighted	16	260.1	45.9	184.23	343.1
Calorie-weighted	16	231.6	41.2	150.4	297.3
<u>Central Java</u>					
Value-weighted	12	252.8	83.4	162.8	479.9
Calorie-weighted	12	243.2	80.1	161.5	472.2
<u>East Java</u>					
Value-weighted	11	258.0	38.4	219.4	352.5
Calorie-weighted	11	233.8	23.0	204.4	272.7
<u>NTB</u>					
Value-weighted	13	228.1	56.7	152.8	324.2
Calorie-weighted	13	231.2	71.1	163.2	389.3
<u>South Kalimantan</u>					
Value-weighted	16	272.0	30.5	242.6	336.7
Calorie-weighted	16	253.3	27.1	209.3	301.1
<u>North Sulawesi</u>					
Value-weighted	12	224.0	38.3	180.7	289.1
Calorie-weighted	12	191.0	25.9	157.7	244.6
<u>South Sulawesi</u>					
Value-weighted	18	225.1	41.4	155.9	346.9
Calorie-weighted	18	207.2	38.7	151.3	315.3

Note: By design, the indices all equal 100 in 2007, the base-year for the indices' weights.

Table 2 shows food shares in value and calorie and Laspeyres index by food category. First, we observe that the average calorie-weighted share of staples increased from 46 to 58 percent, while the average value-weighted share slightly decreased from 28 to 25 percent. In contrast, the share decreased in calories for fish/meat/tofu/tempe and vegetables/fruits. Consistently, these categories experienced the largest price increase between observation points. Thus, the table indicates that households shifted consumption from nutritionally-rich to nutritionally-poor foods, thereby increasing the risk of micronutrient deficiencies.

Table 2: Food shares and price change by food category

Food Category	Villages	2007	2010
<u>Staple</u>			
Laspeyres (value-weighted)	98	100.00	134.65
Value share	98	0.28	0.25
Calorie share	98	0.46	0.58
<u>Fish/meat/tofu/tempe</u>			
Laspeyres (value-weighted)	98	100.00	186.14
Value share	98	0.22	0.19
Calorie share	98	0.06	0.05
<u>Beans and pulses</u>			
Laspeyres (value-weighted)	98	100.00	171.86
Value share	98	0.01	0.01
Calorie share	98	0.02	0.02
<u>Vegetables/fruits</u>			
Laspeyres (value-weighted)	98	100.00	282.98
Value share	98	0.11	0.12
Calorie share	98	0.07	0.04
<u>Milk/egg</u>			
Laspeyres (value-weighted)	98	100.00	176.43
Value share	98	0.05	0.06
Calorie share	98	0.02	0.02

Note: Value shares are shares of total food expenditures, inclusive of other categories, and therefore do not sum to one; calorie shares exclude other sub-categories.

4. Empirical Results

This section summarizes the empirical results. As discussed, we use two specifications of the Laspeyres index to measure of food price changes, one based on value-shares and the other on calorie-shares. Results using both formulations are presented in the following tables.

Table 3 presents benchmark results on the height-for-age z-score. Column 1 shows the results from a biological production function that does not incorporate the economic impacts of food price changes. Height displays a regressive dynamic; changes in the height z-score are larger if the initial z-score is smaller; that is, height z-scores tend to converge as young children age. This could be caused by either biological or behavioral factors such as parents' interventions that aim to equalize growth among children. The next two columns show results from models that include the economic price component. Column 2 shows the results when a value-weighted Laspeyres price index based on local village prices. The results show that an increase in food prices augments the negative effect of the initial z-score; that is, an adverse effect of soaring food prices is greater for larger, better-endowed children. If the initial z-score is positive, the effect of rising food prices on change in the z-score is certainly negative. A qualitatively similar effect is found when using a price index weighted by calories (Column 3).

These results need to be interpreted with care. It might be that under poor nutritional conditions, children with greater height potential are unable to fully realize that potential in a situation where all siblings are treated equally. The result is also consistent with the notion that parents reallocate resources to less-robust siblings, explicitly penalizing larger children. In either case, what is clear is that intrahousehold food allocations appear to have favored the small-for-age siblings. Said differently, we find no evidence that parents failed to protect their most vulnerable children.⁴

Table 3: Impacts of price change on the height-for-age z-score

Dependent variable: change in height z-score			
Model	Biological	Price index weights	
		Value	Calorie
Initial height z-score	-0.7572 (16.59)	-0.3547 (2.35)	-0.4588 (3.45)
Initial height * price change		-0.0017 (2.96)	-0.0013 (2.62)
Initial age (months)	0.0338 (2.14)	0.0314 (2.20)	0.0318 (2.02)
Initial age squared	-0.0004 (1.68)	-0.0003 (1.52)	-0.0003 (1.49)
Male child	-0.0562 (0.38)	-0.0831 (0.58)	-0.0732 (0.50)
Log of birthweight (kg)	1.0217 (1.46)	1.2284 1.77	1.2351 1.77
Household fixed effects	Yes	Yes	Yes
Number of observations	353	353	353
R squared (within)	0.594	0.606	0.602

Note: Sample drawn from children age 6 and younger in 2007. Numbers in parentheses are absolute t-values using robust standard errors with province clusters. Price indices reflect local village prices.

We conduct a similar analysis using the weight-for-age z-score (Table 4). Though we confirm that weight also tends to converge (i.e., a negative effect of the initial z-score), the parameter on the interaction between weight and prices is statistically insignificant regardless of how the price index is constructed. Birthweight has a significantly positive effect on change in the weight z-score in all columns. We speculate that the difference between the two sets of results stems from the fact that weight can be added quickly in the short run, while height tends to be cumulative and also irreversible to a degree. In this sense, height-for-age effects are a better indicator of whether a temporary drop in nutrition intake might have long-term consequences for child development.

⁴ Attrition rate of children is 5.65% over the three years for the height-for-age z score in the full sample. Attrition comes from the following three possibilities: (i) children were not followed up in the 2010 survey under the tracking rules, (ii) anthropometry data were not collected for some reasons (e.g., children were temporarily away from their households), and (iii) outliers that become missing when age-standardized z scores were computed. Yamauchi and Larson (2019) examined whether the main explanatory variables used in the main analysis are significantly correlated with attritions and concluded that they are not, which leads to more confidence in our main results.

Table 4: Impacts of price change on the weight-for-age z-score

Dependent variable: change in weight z-score			
Model	Biological	Price index weights	
		Value	Calorie
Initial weight z-score	-0.6179 (7.06)	-0.4650 (1.66)	-0.5794 (2.36)
Initial weight * price change		-0.0007 (0.56)	-0.0002 (0.17)
Initial age (months)	-0.0016 (0.11)	-0.0031 (0.20)	-0.0020 (0.13)
Initial age squared	0.0001 (0.49)	0.0001 (0.58)	0.0001 (0.51)
Male child	-0.0919 (0.63)	-0.1040 (0.71)	-0.0938 (0.64)
Log of birthweight (kg)	1.4111 (2.02)	1.5032 (2.14)	1.4353 (2.04)
Household fixed effects	Yes	Yes	Yes
Number of observations	353	353	353
R squared (within)	0.311	0.313	0.312

Note: Sample drawn from children age 6 and younger in 2007. Numbers in parentheses are absolute t-values using robust standard errors with province clusters. Price indices reflect local village prices.

We next look for evidence on gender effects in the allocation of food resources made scarce by rising prices. We do so by adding an additional interaction term. In the presence of gender discrimination, the additional variable in which initial-height-and-price-change-and-male should capture an additional effect not measured by the initial-height-and-price-change interaction term or the initial-height-and-male interaction term. Results are reported in Table 5. The t-scores associated with the three-way-interactive term are not statistically significant at the 90 percent confidence level. We, therefore, conclude that parents in our sample do not discriminate among their children based on gender. Similarly, no significant difference was detected between younger and older siblings.

Table 5: Testing for gender discrimination

Dependent variable: change in height z-score			
Model	Biological	Price index weights	
		Value	Calorie
Initial height z-score	-0.6844 (10.06)	-0.0600 (0.34)	-0.2000 (1.16)
Initial height * Male	-0.1182 (1.30)	-0.5924 (1.73)	-0.5416 (1.75)
Initial height* price change		-0.0024 (4.33)	-0.0020 (3.47)
Initial height *price change* Male		0.0018 (1.43)	0.0017 (1.41)
Initial age (months)	0.0361 (2.29)	0.0364 (2.33)	0.0356 (2.25)
Initial age squared	-0.0004 (1.85)	-0.0004 (1.85)	-0.0004 (1.76)
Male child	-0.2622 (1.18)	-0.3895 (1.65)	-0.3746 (1.60)
Log of birthweight (kg)	0.9694 (1.39)	1.1714 (1.73)	1.1638 (1.68)
Household fixed effects	Yes	Yes	Yes
Number of observations	353	353	353
R squared (within)	0.598	0.615	0.610

Note: Sample drawn from children age 6 and younger in 2007. Numbers in parentheses are absolute t-values using robust standard errors with province clusters. Price indices reflect local village prices.

We next look for evidence that children with a larger number of siblings were affected differently during the crises than their peers. As before, we introduce an additional three-way interaction term to the base model that captures the interactions between initial height, price changes, and the number of siblings in the family. As shown in Table 6, the estimated parameter on the three-way-interaction term is quantitatively small and statistically indistinguishable from zero. We therefore do not find indications that children in large households were harmed more than their peers during the food price crises.

Table 6: Testing that children in large households were penalized during the crisis.

Dependent variable: change in height z-score		
Model	Price index weights	
	Value	Calorie
Initial height z-score	-0.3510 (2.26)	-0.4564 (3.27)
Initial height* price change	-0.0015 (2.54)	-0.0012 (2.32)
Initial height* price change* number of siblings	0.0000 (0.33)	0.0000 (0.18)
Initial age (months)	0.0318 (2.05)	0.0320 (2.04)
Initial age squared	-0.0004 (1.56)	-0.0004 (1.52)
Male child	-0.0798 (0.55)	-0.0714 (0.49)
Log of birthweight (kg)	1.2301 (1.77)	1.2373 (1.78)
Household fixed effects	Yes	Yes
Number of observations	353	353
R squared (within)	0.606	0.602

Note: Sample drawn from children age 6 and younger in 2007. Numbers in parentheses are absolute t-values using robust standard errors with province clusters. Price indices reflect local village prices.

Finally, we look for evidence that children in food nonproducer households were affected differently during the crises than children in food producer households. Yamauchi and Larson (2019) shows that children in nonproducer households suffered from soaring food prices more seriously than children in producer households, but their study did not look at inter-sibling differences in the impact. As before, an additional three-way interaction term is introduced to the base model that captures the interactions between initial height, price changes, and the indicator that takes the value of one if they are food producers and zero otherwise. As shown in Table 7, the estimated parameter on the three-way-interaction term is quantitatively small and statistically indistinguishable from zero. In both nonproducer and producer households, therefore, an adverse effect of soaring food prices is greater for larger, better-endowed children, though the average effect of soaring food prices was greater in nonproducer households.

Table 7: Testing that children in nonproducer households were penalized during the crisis.

Dependent variable: change in height z-score		
Model	Price index weights	
	Value	Calorie
Initial height z-score	-0.3655 (2.43)	-0.4559 (3.43)
Initial height* price change	-0.0017 (2.97)	-0.0014 (2.77)
Initial height* price change* Nonproducer	0.0002 (0.83)	0.0003 (1.42)
Initial age (months)	0.0316 (2.02)	0.0319 (2.02)
Initial age squared	-0.0003 (1.51)	-0.0003 (1.47)
Male child	-0.0794 (0.55)	-0.0703 (0.48)
Log of birthweight (kg)	1.2020 (1.73)	1.2079 (1.73)
Household fixed effects	Yes	Yes
Number of observations	353	353
R squared (within)	0.607	0.604

Note: Sample drawn from children age 6 and younger in 2007. Numbers in parentheses are absolute t-values using robust standard errors with province clusters. Price indices reflect local village prices.

5. Conclusions

This paper examines the impact of food price spikes experienced during the 2007/08 food price crisis on child nutrition status using rural household panel data in Indonesia. Our results show that the crisis had significant and negative, but non-uniform impacts on child growth varied by the initial conditions. We find that children with low height z-scores at the start of the crisis gained ground relative to their peers during the crisis, consistent with food resource allocations in their favor. The findings did not show any significant gender difference. Further, the key findings remain robust with respect to family size and food producer status. We conclude that the food price crises had negative long-term impacts on children, but that parental behavior protected the most vulnerable.

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