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ORIGINAL PAPER

Hungry for free trade? Food trade and extreme hunger in developing countries

Sébastien Mary [1,2]

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Abstract
One of the greatest challenges of the twenty-first century is to ensure that the world population has reliable access to adequate, affordable and nutritious food sufficient to avoid hunger. Agricultural trade liberalization is often considered a central element of economic strategies aiming at improving food security in developing countries. Many, however, argue that most developing countries may not benefit from freer agricultural trade and that liberalization may accentuate food insecurity. From an empirical perspective, little is known about the effects of trade on food security in developing countries. We estimated the effects of food trade openness on extreme hunger in developing countries using a novel two-step approach. First, we estimated the reverse causal impacts of hunger on food trade openness using rainfall anomalies as instrumental variables to generate exogenous variation in hunger. In a second step, we estimated the effect of food trade openness on hunger using the residual food trade openness that is not driven by hunger as an instrument. We found that a 10% increase in food trade openness would increase the prevalence of undernourishment by about 6%. We also found evidence that developing countries reduce food trade openness as a response to increased hunger, suggesting protectionist policies. A percentage point increase in undernourishment prevalence would decrease food trade openness by 0.9%. Our results suggest that countries may be better off adopting food self-sufficiency for some time, despite such actions clashing with World Trade Organization's regulations and current agenda.

Keywords Food trade openness · Extreme hunger · Undernourishment · Trade liberalization · Food supply · Producer prices

JEL codes O24 · Q17 · Q18

1 Introduction

One of the greatest challenges of the twenty-first century is to ensure that the world population has reliable access to adequate, affordable and nutritious food, sufficient to avoid hunger (World Food Summit 1996). Undernutrition, which currently affects approximately 815 million people (FAO et al. 2017), has substantial consequences on the cognitive and physical development and health of affected populations, and the potential economic growth of affected countries (e.g. Black et al. 2013; IFPRI 2014; World Bank 2017; McGovern et al. 2017).

Agricultural trade liberalization has arguably been accused of being one of the main causes of undernutrition in developing countries (WTO 2009; De Schutter 2011; FAO 2015a, 2015b; Gayi 2006; Singh 2014; Matthews 2014; Diaz-Bonillo 2013). Free trade opponents typically argue that most developing countries are not competitive enough to take advantage of agricultural trade liberalization, especially in a context where the existence of large farm subsidies in developed countries prevents access to export markets and potential gains for developing countries (Bureau et al. 2006; Wise 2008; Koning and Pinstrup-Andersen 2007; Tokarick 2008). This paper investigates the impacts of food trade openness on extreme hunger in developing countries.

The relationships between food trade and nutrition are complex (Diaz-Bonillo 2013; McCorriston et al. 2013; FAO 2016). First, increased food trade openness should result in higher availability of calories and nutrients and improved

✉ Sébastien Mary
smary@depaul.edu

1 Department of Economics, DePaul University, 1 East Jackson Boulevard, Chicago, IL 60604, USA

2 Department of Economics, The University of Kansas, 1460 Jayhawk Blvd, Lawrence, KS 66045, USA

nutritional diversity. The increased food supply should then result in decreased consumer prices, easing the economic access for net food buyers. However, decreased agricultural prices affect the livelihoods of many rural and farm households (Singh 2014; McCorriston et al. 2013). Also, trade may likely bring about productivity gains from higher specialization (Dithmer and Abdulai 2017). This might, nevertheless, be accompanied by higher concentration in input and output markets with uncertain effects on rural and farm incomes and structural change. More fundamentally, increased food trade may also result in changes in the food supply available to consumers. Substitution effects between domestic products and imported products, with potentially less calorific or nutritional values, may occur (Hawkes 2006). Further, the effect of greater trade openness on food stability is ambiguous as it may lessen the impact of food shortages via higher imports but also make developing countries, especially net food importers, more dependent on imports, and therefore on the (sometimes volatile) export policies of their trading partners (Bezuneh and Yeheyis 2012; FAO 2016). Overall, it is not clear whether food trade openness necessarily ensures food security.

In contrast with the literature on trade and poverty and economic growth (e.g. Winters et al. 2004; Brückner and Lederman 2015), there is little evidence on the impacts of trade on food security (Dithmer and Abdulai 2017). A few case studies support the existence of links between trade and nutrition (Madeley and Solagral 2001; FAO 2006; Thow and Hawkes 2009) but these qualitative analyses cannot disaggregate the effects of trade from other existing economic and nutritional shocks and policy changes. Recently, Dithmer and Abdulai (2017) showed that trade openness increases food supply and dietary diversity in developed and developing countries using System Generalized Method of Moments (SGMM) estimations. A systematic review by McCorriston et al. (2013) concludes there is no consensus on the impacts of agricultural trade liberalization on food security. Levine and Rothman (2006) also find that trade openness decreases child stunting using a mixed sample of developed and developing countries.

Most of the existing literature typically links trade to indicators of food availability or access, such as food quantities or prices, rather than food utilization.[1] This may be misleading because while trade openness may increase the amount of food, this does not necessarily mean that those who need it the most would benefit. Nutrition gains may be restricted to urban areas or populations that are not actually at risk.[2]

Also, previous studies have estimated the impact of trade on food security metrics lumping data from developed and developing countries together. Yet, the impacts of agricultural trade liberalization depend on multiple mitigating factors, such as the structure of trade, the vertical structure of food markets or the existence of market distortions (Sexton et al. 2007; Magrini et al. 2017). These factors are likely to be quite different in developed and developing countries. Thus, past studies may be misrepresentative of the links between trade and food security in developing countries.

From a methodological perspective, the estimation of the effects of trade openness on nutrition is complicated by the existence of potential reverse causality between hunger and trade. Dithmer and Abdulai (2017) indeed suggest that developing countries may be implementing more protectionist policies as a response to past food security shocks. Typically, estimating the impact of trade openness on hunger requires valid instruments for trade openness using two-stage least squares (2SLS) instrumental variables (IV) regressions in static models. In our context finding valid instruments may however be difficult. System Generalized Method of Moments (SGMM) estimators can also be used for dynamic models but they suffer from potential identification and weak instrument issues (Bazzi and Clemens 2013).

Instead, we used an alternative approach accounting for the reverse causal effect of hunger on trade openness that has been used in a relatively comparable context (Brückner 2013; Brückner and Lederman 2015). The estimation strategy identifies the causal impact of food trade openness on hunger by extending a two-step procedure. First, we estimate the reverse causal impacts of hunger on food trade openness using rainfall anomalies as instrumental variables to generate exogenous variations in hunger. In a second step, we estimate the effect of food trade openness on hunger using the residual food trade openness that is not driven by hunger as an instrument. A key advantage of this approach, unlike traditional 2SLS-IV or SGMM estimations, is that it allows estimating the reverse causal effects of hunger on food trade openness.

Given this background we estimated the effects of food trade openness on undernourishment prevalence using a sample of 52 developing countries between 1990 and 2013. Our paper contributes to the existing literature in two main aspects. First, this is the first study to specifically examine the relationships between food trade and hunger in developing countries. Second, the approach estimates the reverse causal effects between trade openness and hunger. This will provide an indication with respect to potentially protectionist policies implemented in developing countries following the occurrence of food security crises.

We find that a 10% increase in food trade openness would increase the prevalence of undernourishment by approximately 6% in our sample. We also find suggestive evidence that developing countries engage in protectionist policies with respect to food trade as a result of hunger crises. A percentage point increase in the prevalence of hunger would decrease food trade openness by 0.9% in our sample of developing countries.

[1] Levine and Rothman (2006) is a notable exception.

[2] This would be in line with a recent literature linking trade openness and rising overweight and obesity (Miljkovic et al. 2015, 2017).

The paper is structured as follows. Section 2 presents the empirical model. Section 3 describes the data. Section 4 presents the empirical results and discusses the findings. Section 5 concludes.

2 Empirical model and identification

To investigate the impacts of food trade openness on nutrition, we specify the following static model with fixed effects (FE):

$$H_{it} = \beta \ln(TO_{it}) + \alpha x_{it} + \mu_i + \pi_t + \varepsilon_{it} \quad (1)$$

Where countries are indexed by i and year by t; H_{it} is the prevalence of undernourishment of country i in year t; TO_{it} is food trade openness; x_{it} is a vector of independent variables, namely, the logarithms of non-food agricultural trade openness, non-agricultural trade openness and gross domestic product (GDP) per capita, and foreign direct investment (FDI) and urbanization; μ_i is a time-invariant country-specific effect and π_t are year dummies.

β is the key coefficient of interest as it captures the effects of food trade openness on hunger. We use a linear-log model that implies that food trade openness has diminishing marginal returns towards hunger reductions.[3] The inclusion of country and period fixed effects accounts for the presence of time invariant omitted variables and common shocks affecting both undernourishment and trade openness, respectively. All other independent variables, but food trade openness, are treated as exogenous. This can be relaxed (see robustness analyses where GDP per capita is treated as endogenous). Yet, we do not find evidence of endogeneity between hunger and other independent variables.

The impact of trade openness on hunger will depend on the size of the sector affected by trade and therefore we weight our trade openness measures using 'participation' effects following Headey (2013). Each trade openness measure is weighted by the share of the (non) food sector within GDP. In essence, participation effects allow accounting for the level of economic development of each country, and implicitly the size of the (non) agricultural and (non) food sector within the economy.

We also include non-food agricultural trade and non-agricultural trade openness measures as both may be affecting hunger through indirect channels. For example, trade may increase the quality and availability of medical services, or provision of social services (Levine and Rothman 2006). Non-food trade openness may affect incomes in rural areas and therefore the nutritional status of local populations.

Nutrition models should follow the theory implied by existing conceptual frameworks of undernutrition (Smith and Haddad 2015). In particular, studies examining the effect of various factors on nutrition indicators such as Smith and Haddad (2015), Soriano and Garrido (2016), or Mary et al. (2018a, b), often rely on the UNICEF's conceptual framework offering a classification of the causes of undernutrition (UNICEF 1998) to justify the inclusion/exclusion of variables in their models. This classification has often provided a theoretical background for modelling undernutrition in the empirical literature, where most studies typically focused on either the role of immediate, underlying, or basic factors (Mary et al. 2018a, b; Soriano and Garrido 2016; Smith and Haddad 2015). We follow the literature by including only basic factors that generally determine the political and economic context and the amount of human and physical capital and resources available.[4]

There is an extensive literature on the impacts of economic growth and urbanization on nutrition (e.g. Smith and Haddad 2015; Soriano and Garrido 2016; Mary et al. 2018a, b). Economic growth increases the populations' average incomes allowing them better access to and more consumption of health goods and services, social services and education. It may also increase the public supply of social and health infrastructures, as well as the provision of nutrition-relevant services. In addition, the current rate of expansion of urban agglomerations in developing countries has spawned severe challenges for the provision of basic services such as adequate housing, water and sanitation systems, as well as provision of health clinics and schools (Mary et al. 2018a, b). We last included FDI as there is evidence that it may improve food security (Mihalache-O'Keef and Li 2011). FDI may increase economic growth and household incomes, leading to higher calorific and nutritional intakes.

We did not include additional variables in the baseline model because the inclusion of additional variables might come at a risk of introducing potentially endogenous variables; for example, nutrition-relevant policies or foreign aid inflows are likely targeted at countries experiencing higher hunger rates. Thus the choice of a parsimonious model is driven by the desire to avoid the issue of bad controls caused by the inclusion of potentially endogenous regressors (Angrist

[3] Conceptually, we think a linear-log form seems more adequate. Indeed, our specification suggests that as trade openness increases, the gains in terms of hunger reductions become smaller, but remain positive. A quadratic specification would mean that past a certain level, trade openness could result in hunger increases. On a purely statistical basis, we also estimated a model using a linear-linear form and then compared the adjusted R-2 and compared the residuals, finding that the linear-log form performed best. We also tested the inclusion of a squared term in our main model and the coefficient turned out to be statistically insignificant.

[4] The exclusion of climate variables such as rainfall or temperature from the 'basic' model is consistent with the classification of climate as an immediate and underlying cause of undernutrition. Similarly, the inclusion of food aid would not respect this principle as food aid is considered to be an underlying factor (Ickes et al. 2015). On a more practical level, the coverage of data on food aid, food consumer price indexes or agricultural foreign direct investment is particularly limited for developing countries and/or the pre-2000s period. For example, the OECD discourages the use of sector aid data (and thus food aid) before 2002 because the coverage of sector disbursements is incomplete.

and Pischke 2008). As an additional robustness check, to make sure our baseline results are not affected by bad control issues, we discuss the results of the reduced form of the empirical model in which only food trade openness is included as independent variable.

While we warn about the risk of bad controls, we nonetheless provide the results of an extended model, in which we include inflation based on the consumer price index, the share of agriculture in total GDP to account for the sectoral composition of the economy, and an indicator of the socioeconomic pressures at work in society that could constrain government action or fuel social dissatisfaction, and an indicator of the level of ethnic tensions in a country. Inflation is a proxy for food CPI. The share of agriculture in GDP captures indirectly the existence of geographical inequalities via the sectoral relative importance of rural areas (versus urban areas) since most of the agriculture in developing countries is located in rural areas. The indicator of socioeconomic pressures indirectly generally captures the existence of economic inequalities (via poverty or unemployment). The ethnic tensions indicator accounts for the existence of ethnic inequalities.

2.1 Identification strategy

We take the first differences of the model in Eq. (1) to remove the fixed effects and obtain:

$$\Delta H_{it} = \lambda \Delta \ln(TO_{it}) + \phi \Delta x_{it} + \eta_t + \varpi_{it} \tag{2}$$

Equation (2) represents our main empirical model.

As explained in the introduction, the identification of the causal impact of trade openness on hunger is complicated because of the reverse causality between trade and hunger. To account for this issue, we adapted the two-step approach devised by Brückner and Lederman (2015) in a somewhat similar context. As a first step, assuming ex-ante the potential endogeneity of food trade openness, the effect that hunger has on food trade openness can be estimated using an instrumental variables (IV) approach. Equation (3) presents the estimation:

$$\Delta \ln(TO_{it}) = \zeta \Delta H_{it} + \delta \Delta x_{it} + \sigma_t + \varphi_{it} \tag{3}$$

The statistical significance of ζ provides a direct endogeneity test of food trade openness. Estimating the effect that hunger has on trade openness requires an exogenous source of variation for hunger. We used rainfall anomalies,[5] defined as the yearly deviation from the long run rainfall level observed over the full period in a country; we included a quadratic term of rainfall anomalies to capture the nonlinear effects between rainfall and hunger. Rainfall anomalies (z) are generally defined as follows:

$$z_{it} = \frac{a_{it} - \bar{a}_i}{\bar{a}_i}$$

Where $\bar{a}_i$ is the mean rainfall level, observed over the full time period, in country i; $a_{it}$is the rainfall level in year t. If z_{it} is positive (negative), this indicates that the average rainfall in year t in country i is above (below) the long run level.

We then construct an adjusted food trade openness series where the response of food trade openness to hunger is 'partialled out' using estimates from Eq. (3):

$$\Delta \ln\left(TO^{*}_{it}\right) = \Delta \ln(TO_{it}) - \zeta \Delta H_{it} - \delta \Delta x_{it} \tag{4}$$

$\ln\left(TO^{*}_{it}\right)$ is free of the endogeneity bias and can be used to instrument $\ln(TO_{it})$ in the 2SLS estimation of Eq. (2). The second step estimation is exactly identified. Instruments (i.e. rainfall anomalies) are also included as independent variables in the second-step estimation.

There are several reasons why this identification strategy, that is, using current rainfall anomalies to instrument hunger, is plausible. First, rainfall anomalies are exogenous and random. Second, rainfall anomalies are likely correlated with undernourishment prevalences via food production (e.g. Rabassa et al. 2014). The key identifying assumption is that current rainfall anomalies are unexpected food supply-sided shocks that affect hunger and hence food trade openness as a response to increased hunger.

The exclusion restrictions imply that rainfall anomalies have no large effects on trade openness, other than through nutrition channels. To provide some supporting evidence of the exclusion restrictions, in line with Brückner and Lederman (2015), we provide additional estimations where rainfall anomalies are included as right hand side variables in the IV estimations of Eq. (3). If the associated coefficients are not statistically significant, then the additional estimations confirm the validity of the exclusion restrictions.

Also, we used the Hansen's J-test of over-identification of all instruments to investigate the validity of our instruments. This tests the null hypothesis that the instruments are uncorrelated with the error term. If the *p* value is below 10%, all instruments are not exogenous and therefore invalid. The strength of instruments was further examined by reporting the Kleibergen-Paap rk Wald F statistic (or F-statistic) that we compared with the critical values from Stock and Yogo (2005) for testing weak instruments. We tested the null hypothesis that the maximum size distortion was greater than 10, 15 or 20%.

If weak instruments affected the first-step we used the Continuous Updating Estimator (CUE), which has been found to perform better than 2SLS or bias-corrected Generalized Method of Moments estimators in the case of weak

[5] We originally used temperature anomalies but found that as instruments they do not meet the exclusion restrictions in this setting.

instruments (Donald et al. 2009). Note that the many weak instrument maximal relative bias is zero for CUE estimators.

2.2 Alternative modelling specifications

In robustness analyses we tested two variants of the model in Eq. (2). First, while the differencing process removes the fixed effects from Eq. (1), Headey (2013) recommends adding fixed effects to the differenced equation to control for unobserved country-specific trend factors that influence changes in hunger (labelled 'country trend effects', ξ_i). Eq. (2) can be rewritten as:

$$\Delta H_{it} = \varrho \Delta \ln(TO_{it}) + \upsilon \Delta x_{it} + \chi_t + \xi_i + \psi_{it} \quad (5)$$

The identification strategy outlined above still applies in this case.

Second, Soriano and Garrido (2016) suggested including a lagged term of hunger, H_{it-1}, to account for unobserved time-varying factors and potential serial correlation. This model is equivalent to a dynamic model with a lagged dependent variable. This captures the existence of time-varying omitted factors (Mary et al. 2018a, b; Soriano and Garrido 2016). Eq. (2) can be marginally modified:

$$\Delta H_{it} = \alpha' H_{it-1} + \lambda' \Delta \ln(TO_{it}) + \phi' \Delta x_{it} + \eta'_t + \varpi'_{it} \quad (6)$$

Using the same identification strategy, the additional lagged term was instrumented using its first and second lags following Brückner and Lederman (2015). The model can be estimated using GMM (or 2SLS-IV).

3 Data

The sample included 52 developing countries for a total of 998 observations. Table 1 displays descriptive statistics for the dataset. In this paper, we collected data on the prevalence of undernourishment and other determinants for low income and middle income countries between 1990 and 2013. Given the limitations in existing data sources, the selection of countries and years was mainly determined by available data on food trade. We found no outlier using the Hadi procedure.[6] Appendix Table 8 provides a full list of countries.

3.1 Dependent variable

Hunger is measured by the prevalence of undernourishment or the percentage of the population whose food intake is below the minimum level of dietary energy consumption and insufficient to meet dietary energy requirements continuously. The methodology behind this indicator, and its strengths and weaknesses, have been discussed elsewhere (e.g. Cafiero 2014; Soriano and Garrido 2016). For example, despite food availability, food insecurity may still affect many people in a country due to inadequate access to food by poor households. In a related manner, the average food available to each person, even corrected for possible effects of low income, is not a strong predictor of food insecurity among the population (Mary et al. 2018a, b).

Despite these limitations, this indicator has been used to monitor the progress towards the MDG1c target, which aimed to halve world hunger between 1990 and 2015 and is therefore a particularly appropriate dependent variable for our empirical analysis. Let's note that data on undernourishment prevalences are reported as three-year averages, so we used the three-year averages of all variables included in the empirical model in Eq. (2) to match the undernourishment prevalence data.

3.2 Trade openness

Trade openness is defined as the sum of exports and imports divided by GDP. Building food trade openness first requires agricultural exports and imports, as well as agricultural GDP. Agricultural exports and imports were built following Sanjuan-Lopez and Dawson (2010). First, we collected data on total GDP, total exports and imports of goods and services in constant local currency units from the World Bank's World Development Indicators (WDI) database.

Data on total agricultural exports and imports, all crop and livestock products, and total merchandise exports and imports were obtained from the Food and Agricultural Organization (FAO) database. We then calculated the ratios of agricultural exports (imports) to the total amount of exports (imports) and multiplied these ratios by the total amount of exports (imports) of goods and services in constant local currency units, to obtain agricultural imports and exports. Agricultural GDP was calculated by multiplying the share of agriculture of total value added from the WDI database and total GDP in constant local currency units. We could then build (non) agricultural trade openness, as the sum of agricultural exports and imports divided by (non) agricultural GDP.

Lastly, we can decompose further agricultural trade openness between food trade openness and non-food trade openness. The share of food exports (imports) within agriculture as well as the share of food GDP within agricultural GDP can be calculated using data from the FAO. We can then construct food trade openness and non-food trade openness as subcomponents of agricultural trade openness.

3.3 Other variables

GDP per capita is in constant US dollars 2011 based on purchasing power parity (PPP). An international dollar has the

[6] This is implemented using the BACON command in Stata.

Table 1 Descriptive statistics of the sample

Variables	Source	(1) N	(2) Mean	(3) Standard dev.	(4) Min.	(5) Max.
Undernourishment prevalence (%)	WDI	998	18.13	11.58	5	58.80
Foreign direct investment, % of GDP	WDI	998	3.55	4.31	−1.98	40.53
Urbanization rate (%)	WDI	998	48.40	18.59	14.33	87.49
Rainfall anomaly	CRU	998	1.19	8.96	−31.20	35.0
Rainfall anomaly, squared	CRU	998	0.81	1.39	0	12.3
GDP per capita, in 2011 USD PPP	WDI	998	6499	4645	456.0	22,431
Non-food trade openness (%)	FAO	998	5089	13,025	28.72	103,851
Food trade openness (%)	FAO	998	76.82	107.0	2.332	986.1
Non-agricultural trade openness (%)	FAO	998	70.72	35.75	8.865	248.3

Notes: the non-food agricultural trade openness ratio may seem exceptionally high. This is because the non-food agricultural sector represents a very small share of agricultural GDP in developing countries

same purchasing power over GDP as the U.S. dollar has in the United States. Foreign direct investments[7] (as per cent of GDP) are the net inflows of investment (new investment inflows less disinvestment) to acquire a lasting management interest (10% or more of voting stock) in an enterprise operating in an economy other than that of the investor. Urbanization refers to the proportion of total population living in urban areas. Data for GDP, FDI, urbanization and hunger were taken from the WDI database.

Rainfall (and temperature) data is from the Climate Research Unit (CRU) and the Tyndall Centre for Climate Change Research (TYN) of the University of East Anglia, from the dataset CRU TS v. 3.23. The gridded climate dataset (referred to as CRU TS3.10) is built from monthly observations at meteorological stations across the world's land areas, interpolated into 0.5° latitude/longitude grid cells covering the global land surface (excluding Antarctica), and combined with an existing climatology to obtain absolute monthly values.

4 Results

The aim of this paper is to estimate the impacts of food trade openness on the prevalence of undernourishment in developing countries. We include year dummies in all regressions. Robust country-clustered standard errors were used. As explained earlier, the econometric approach relies on two steps, first, estimating the reverse causal effects of food trade on hunger, and second, estimating the effects of food trade on hunger once the reverse causal effects have been accounted for.

4.1 First step: The effect of hunger on food trade openness

Table 2 focuses on the first-step of the approach and displays the CUE estimates of the effect that undernourishment has on food trade openness. We report the CUE estimates because of the presence of weak instruments in the 2SLS-IV estimates (available in Appendix 4). Column 1 presents the first-stage estimates that link rainfall anomalies to undernourishment for the food openness regression.

The coefficients associated with rainfall instruments were all statistically significant and indeed confirm that increased rainfalls from the long run level in a country are associated on average with decreased hunger prevalence, though large rainfall anomalies may have negative consequences. These estimates are in line with what one would expect about the relationships between rainfall and undernourishment via food production shocks.

More importantly, the coefficient for undernourishment prevalence in column 2 is negative (−0.009) and significant at 5%; a percentage point increase in undernourishment prevalence would decrease food trade openness by approximately 0.9%.[8] This suggests that developing countries reduce food trade openness as a response to increased hunger prevalence. This may be interpreted as the implementation of protectionist policies in the food sector and confirms the conjecture of Dithmer and Abdulai (2017) that food security shocks may lead governments to implement protectionist policy responses.

Policymakers in developing countries may benevolently adopt protectionist measures resulting in decreased food trade openness because they believe that such policies will help restore food security (Mgeni et al. 2018). First, policymakers in developing countries might feel that the large program of agricultural subsidies in developed

[7] FDI data do not give a complete picture of international investment in an economy. Balance of payments data on FDI do not include capital raised locally, an important source of investment financing in some developing countries (World Bank 2018).

[8] This is calculated as $(e^{-0.009} - 1) \approx 0.89\%$.

Table 2 The effect of undernourishment prevalence on food trade openness

Dependent variable	(1) First stage Hunger	(2) CUE Food Trade Openness
Hunger prevalence		−0.009**
		[0.03]
Rainfall anomaly	−1.238***	
	(−2.73)	
Rainfall anomaly, squared	10.126***	
	(3.14)	
Urbanization	−0.485	0.004
	(−0.93)	(0.20)
Non-food agricultural trade openness	8.143	0.853**
	(0.81)	(2.45)
Non-agricultural trade openness	0.419	−0.030
	(1.23)	(−1.15)
FDI	−0.096***	−0.001
	(−3.43)	(−0.55)
Food trade openness		
Logged GDP per capita	−8.303***	−0.250*
	(−3.43)	(−1.73)
Observations	998	988
Number of countries	52	52
First-stage, F-stat	n.a.	10.32
Hansen J, p value	n.a.	0.94

Robust z-statistics in parentheses.*** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$, # $p < 0.15$. Stock-Wright LM p values between square brackets. n.a..: non-applicable. Our specification used an instrument set with rainfall anomaly and a quadratic term for rainfall anomaly

economies provides unfair advantages to their exports (Devadoss 2006). Within this line of thought, the reduction in openness will likely reduce external competition for domestic producers. Second, there is some evidence that support to agricultural producers may be beneficial to several dimensions of food security in developing countries (e.g. Magrini et al. 2017). However, Magrini et al. (2017) also show that too much protection may be counterproductive. Protectionist measures may not also be the most efficient strategy for supporting local agriculture (Mgeni et al. 2018).

Further, governments' responses to hunger crises may also be motivated by politicians' personal preferences, electoral incentives or by lobbying (Fulgitini and Shogren 1992; Bellemare and Carnes 2015). For example, interest groups in farming may influence the provision of farm subsidies by supporting incumbent governments or the campaign of candidates who support domestic agriculture or have more protectionist views (Karnik and Lalvani 1996).

From a methodological standpoint, the presence of negative reverse causality implies that the Ordinary Least Squares (OLS) estimate of the impact of trade openness on hunger prevalence is biased downwards. If this is not controlled for, reverse causality may result in unreliable estimates.

The p value for the Hansen J test on the over-identifying restrictions in column 2 Table 2 (0.94) cannot reject that the instruments are uncorrelated with the second-stage error. We can reject the null hypothesis that the maximum size distortion is greater than 10% as the Kleibergen-Paap statistic is 10.32, well above the associated Stock-Yogo value (8.68).

Also, Appendix Table 9 provides intuitive tests for the exclusion restrictions for trade openness. The validity in the IV estimation of Eq. (3) relies on assuming that rainfall anomalies affect trade openness only through undernourishment. To examine whether there are significant direct effects of rainfall anomalies on trade openness beyond hunger, we present the results of regressions in which rainfall anomalies are introduced as right hand side variables in Eq. (3), respectively in Table B1. The coefficients associated with rainfall anomalies are all statistically insignificant in these additional regressions (columns 2 and 3), therefore providing reassuring evidence for the exclusion restrictions. The coefficients for hunger in Table B1 are in line of that of column 2 in Table 2.

4.2 Second step: The effect of food trade openness on hunger

Table 3 column 2 presents the estimation results of the second step. First, column 1 shows the OLS estimates. The coefficient for food trade openness is positive (1.341) but statistically insignificant, implying that food trade openness has no impact on undernourishment. As explained earlier, we suspect that the OLS estimate may be biased downwards because of the negative reverse causality between trade openness and hunger.

Once we account for the negative reverse causal impacts of undernourishment on trade openness in column 2, we find that food trade openness increases the prevalence of undernourishment. In line with the endogeneity bias, the 2SLS estimate is higher than the OLS estimate.[9] According to column 2 (using the residuals based on the CUE estimates in Table 2), a 10% increase in food trade openness would increase undernourishment prevalence by about 6%.[10] This result contrasts with the literature that has found that trade openness increased food security metrics (e.g. Brandao and Martin 1993; Tokarick 2008; Anderson 2016, 2017; Dithmer and Abdulai 2017). However, these studies do not typically look at food utilization. Levine and Rothman (2006) found that trade openness reduced stunting, an indicator of food utilization, but also acknowledge that their identification strategy was not adapted for child stunting.

Increased food trade openness will logically increase the amount of calories and nutrients, but our results suggest that those who need it the most (living in rural areas for the most part) may not benefit from increased food trade. Perhaps, most of these gains may be restricted to urban areas because of weak physical and economic linkages between urban and rural areas. In addition, the increased food supply likely affects food producer prices downwards and thus farmers' and rural households' incomes. In a related manner, gains from trade may be mainly captured by intermediates along the food chain and may not accrue to the producers and rural areas. This is especially likely in developing countries where food markets are riddled with significant imperfections (Sexton et al. 2007). Overall, our results fit within the literature that recognises the heterogeneous effects of trade openness (Singh 2014; Anderson 2016) and suggest that a holistic assessment on how trade affects food security, through exploring potential effects on supply-sided (externally and domestically) and demand-sided channels is needed.

Furthermore, the coefficient for non-food trade openness is insignificant but the coefficient for non-agricultural trade openness is positive and significant suggesting that higher non-agricultural trade openness may increase hunger. The effects are economically small as a 10% increase in non-agricultural trade openness would result in a 0.4% increase in the prevalence of undernourishment. This may be driven by the effects of increased competition by non-agricultural imports.

Also, the coefficients for GDP per capita and FDI are negative and statistically significant, suggesting that economic growth and higher FDI reduce extreme hunger. For example, a 10% increase in GDP per capita would decrease hunger prevalence by 3.5%. Estimates are in line with existing studies (e.g. Soriano and Garrido 2016; Mary et al. 2018a, b). Similarly, a 10% increase in FDI would result in 0.2% decrease in hunger prevalence. Last, the Kleibergen-Paap statistic (2375) is well above the Stok-Yogo value (16.4), clearly suggesting that weak instruments are not an issue. Similarly, the Ramsey test suggests omitted variables are not a problem (p-val.: 0.103).

Next, we tested whether our main finding with respect to food trade openness is biased by bad controls by reporting the results of a reduced form model without any control variables, except food trade openness, in column 3. The coefficient is very close to that of column 2 and is statistically significant, providing reassuring evidence that our main finding is not biased by bad controls (Angrist and Pischke 2008). According to column 3, a 10% increase in food trade openness would increase undernourishment prevalence by about 6.4%.

We also provide the results of an extended model in the Appendix (Table 11), despite the risk that additional controls may bias the estimates and affect inference. Interestingly, the estimate for food trade openness remains positive and statistically significant, though much lower than in column 2 Table 3 (3.55 versus 10.86).

Further, our modelling specification controls for time invariant heterogeneity and common shocks. One might be concerned that time varying heterogeneity affects our estimates. To check the robustness of our results, we provide additional estimations in the Appendix for the dynamic model developed in Eq. (6). The dynamic model captures the existence of time varying omitted variables (Mary et al. 2018a, b; Soriano and Garrido 2016). The food trade openness estimate in column 1 Table D1 is positive, 8.587, and statistically significant at 1%. We also provide the results of the model with country trend effects developed in Eq. (5). Again, the estimate for food trade openness is positive and statistically significant. These additional results lessen any concerns that we may have about potential bias due to time invariant and time varying heterogeneity.

Last, we investigate the robustness of our findings to two modelling choices: (1) model in which GDP per capita is endogenous[11]; (2) the use of 2SLS-IV estimations for the first

[9] While we follow Brückner and Lederman (2015) who included the instruments in the second step (here shown in column 2), it is important to note that the models in columns 1 and 2 are not exactly the same. We provide a full comparison of models including and excluding the instruments in Appendix 4 Table 12. The differences are minimal.

[10] This is calculated as: $\frac{10.86}{18.13} * \frac{10}{100} \approx 5.99\%$.

[11] An extension of the approach to two potentially endogenous variables can be found in Mary et al. (2017).

Table 3 The effect of food trade openness on hunger prevalence

Hunger prevalence	(1) OLS-FE	(2) 2SLS-IV	(3) 2SLS-IV
Food trade openness, logged	1.341	10.862***	11.640***
	(0.69)	(4.66)	(4.16)
Non-food trade openness, logged	6.291	−0.207	
	(0.66)	(−0.02)	
Non-agricultural trade openness, logged	0.469	0.793**	
	(1.32)	(2.42)	
GDP per capita, logged	−8.090***	−6.450***	
	(−3.65)	(−3.45)	
Urbanization	−0.581	−0.578	
	(−1.18)	(−0.99)	
FDI	−0.094***	−0.100***	
	(−3.20)	(−3.45)	
Observations	998	988	998
Number of countries	52	52	52
Ramsey RESET, p value	0.09	0.10	0.33
First-stage, F-stat	n.a.	2375	1807

Robust z-statistics in parentheses.*** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$, # $p < 0.15$. n.a..: non-applicable. Regressions in columns 2 and 3 control for rainfall and rainfall squared (see Appendix Table 12 for full tables)

step of the approach. The estimation results can be found in Appendix 4. The main pattern of results remains the same in these additional checks.

4.3 Causal pathways between food trade openness and hunger

We now examine what drives the effect of food trade openness on the prevalence of undernourishment. Table 4 examines the impact of food trade openness on several potential causal mechanisms: food supply (in kilocalories – kcal – per day per capita), agricultural producer prices, and GDP per capita in the food sector. All variables are expressed in logarithms. Data for food supply and agricultural producer prices are from the FAO database. The FAO Agricultural Producer Price Indices measure annual changes in the selling prices received by farmers (prices at the farm-gate or at the first point of sale). The indices are constructed using price data in Standardized Local Currency. Food GDP per capita (in 2011 international dollars) was calculated using the share of food in value added combining data from the FAO and WDI databases.

We use the same methodology presented in Section 3 to account for endogeneity, though we use temperature anomalies as instruments[12] when rainfall anomalies are not relevant instruments, and find that the abovementioned variables are all endogenous to food trade openness. Appendix Table 10 summarizes the first-step estimation results. Models in Table 4 are reduced-form estimates to avoid the problem of bad controls as explained above, but control for the instruments in the second step.

Table 4 reports the second-step estimates. In short, we find that increased food trade openness increases food supply but decreases food GDP per capita and agricultural producer prices. The food supply coefficient in column 1 (0.028) is positive and significant at 1%. It implies that a 10 per increase in food trade openness would result in 0.3% increase in food supply, or a paltry gain of approximately 8 kcal per day per capita. This is much lower than the estimates found in Dithmer and Abdulai (2017) though they used data from both developed and developing countries.

It is important to realize that this final estimated impact results from various opposite effects. For example, increased food imports automatically increase food supply, while food exports reduce it. Also, the increased food imports may compete with local farm products. The extent to which competitive pressures affect local production will determine whether food supply is increased largely, moderately or even decreased.

Moreover, the coefficient for food GDP per capita in column 2 (−0.247) is negative and significant at 1%, implying that a 10% increase in food trade openness would result in a 2.4% decrease in food GDP per capita. This might be the result of structural change as small scale farmers may not be able to deal with increased competitive pressures and leave the agricultural sector (and move to the cities to find manufacturing or services jobs). In a somewhat related manner, we find increased food trade openness significantly decreases agricultural producer prices. The effect is large as a 10% increase in food trade openness would reduce producer prices by about

[12] We do so for food supply and producer prices.

Table 4 Transmission channels between food trade openness and hunger

Dependent variable	(1) 2SLS-IV Food Supply	(2) 2SLS-IV Food GDP per capita	(3) 2SLS-IV Producer prices
Food trade openness	0.028***	−0.247***	−0.469*
	(4.56)	(−3.26)	(−1.83)
Observations	995	995	983
Ramsey RESET test, p value	0.58	0.86	0.13
First-stage, F-stat	1104	127.4	65.75

Robust z-statistics in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$, # $p < 0.15$. n.a..: non-applicable. Regressions control for instruments. Estimates are from reduced form models

4.7%. This definitely explains why food GDP per capita is negatively affected.

Overall, these estimates suggest that the negative effects on producer prices and GDP per capita in the food sector dominate the positive, but rather small, impacts on food supply. These findings somewhat resonate with a key conclusion from the literature review by McCorriston et al. (2013) who found that prices play a central role in determining the effects of trade openness on food security.

4.4 Alternative trade indicators

We repeated the analysis using alternative trade measurements, namely, import penetration and export-GDP ratios (Table 5). When we used import penetration ratios, the coefficient in column 1 that summarises the first-step estimation indicates that developing countries reduce import penetration as a response to increased hunger. A percentage point increase in hunger prevalence would reduce food import penetration by 1.5%. This is in line with the result from Table 2. Further, we find that the coefficient in column 2 is significant and positive (13.251), suggesting that a 10% increase in import penetration would increase hunger by 7.3%. This provides additional evidence that trade, especially food imports, can hurt developing countries from a food utilization standpoint.

On the contrary, we find that a 10% increase in the export-GDP food ratio would decrease hunger by 5.4% (column 4). While this suggests that access to export markets and potential higher market prices is beneficial for developing countries in our sample, our findings on the effect of food trade openness suggest that the gains from accessing export markets are largely outweighed by the effects of import competition. This may partly be because of the existence of trade-distorting farm subsidies in developed economies. In the meantime, developing countries may be better off by adopting food self-sufficiency until they reinforce and develop comparative advantages in the food sector and/or developed economies drastically reform and eliminate their subsidy programs.

4.5 Heterogeneity

In this section, we examine whether the average effect of food trade openness is conditional on various dimensions, such as the quality of institutions, the level of ethnic tensions, or the overall net food trade position, that is, whether the country is a net importer or exporter in a given year. Unfortunately, we cannot offer estimates for different sub-regions of the world (e.g. Sub-Saharan Africa, Latin America) because the approach requires a minimum number of countries and observations to be implemented. Table 6 displays the conditional coefficients for the impact of food trade openness on hunger at different percentiles of various institutional indicators.[13]

Data for the quality of institutions are from the International Country Risk Guide (ICRG)'s Political Risk Services database. In particular, we examine the conditional effect of food trade openness through the following dimensions: *control of corruption* within the political system, including bribery, nepotism, or secret party funding; *law and order*, accounting for the strength and impartiality of the legal system and popular observance of the law; *bureaucratic quality*, accounting for the instructional strength and quality of the bureaucracy; *democratic accountability*, accounting for the responsiveness of the government to its people's demands. We also examine the mediating role of ethnic tensions in a country. The *degree of ethnic tensions* captures tensions in a country due to racial, nationality or language divisions, and overall the extent to which opposing groups are willing to compromise and be tolerant. This indicator is also taken from the ICRG.

We also tested for the presence of interactions in the first step, that is, in the models examining the effect of hunger on trade openness; we did not find that this effect is conditional on the quality of institutions or the level of ethnic tensions. More importantly, the main message across Table 6 is that the negative effect of food trade

[13] Full tables of the interaction models can be requested from the author.

Table 5 The effect of (hunger) food trade on (food trade) undernourishment prevalence

Dependent variable	(1) CUE Food import penetration	(2) 2SLS-IV Hunger prevalence	(3) CUE Food export-GDP	(4) 2SLS-IV Hunger prevalence
Food import penetration		13.251*** (4.41)		
Food export-GDP				−9.759*** (−5.01)
Hunger prevalence	−0.015*** [0.00]		0.012** [0.02]	
Observations	998	998	997	997
Number of countries	52	52	52	52
Ramsey RESET, *p* value	0.60	0.15	0.40	0.05
First-stage, F-stat	10.41	658	9.78	2188

Robust z-statistics in parentheses.*** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$, # $p < 0.15$. Stock-Wright LM p values between square brackets. Other coefficients are omitted for the sake of space

openness on hunger becomes much larger as the control of corruption, law and order, bureaucratic quality, or democratic accountability, worsen. For example, according to column 1 Table 6, a 10% increase in food trade openness would result in an increase in hunger prevalence by 15.6, 14.1 and 13.2%, respectively at the 50th, 50th and 95th percentiles of corruption control. The mediating effect of governance on the effect of trade on undernutrition has long been recognized (FAO 2016). For example, improved institutions may allow farmers to capture more value along the food chain. Similarly, as governance improves, the voice of food insecure populations is more likely to count and frame the policymaking so that governments respond equitably to their needs. Table 6 especially suggests institutional change that gives more attention to corruption and bureaucratic quality could somewhat mitigate the effects of food trade openness.

Last, we show how the effect of conflict on hunger worsens as ethnic tensions are increased in column 5 Table 6. For example, a 10% increase in the level of ethnic tensions would result in an increase in hunger prevalence by approximately 21% at the 5th percentile of ethnic tensions, that is, when tensions between ethnic groups are the most acute. Arguably, pre-existing tensions between ethnic groups can indeed accelerate the inequities associated with food trade. This is suggestive that strategies aiming at preventing the negative effects of food

Table 6 The effect of food trade openness on hunger at different percentiles of institution

Dependent variable: hunger	Institution				
Percentile of institution	Control of corruption (1)	Law and order (2)	Bureaucratic quality (3)	Democratic accountability (4)	Ethnic tensions (5)
Food trade openness, 5th	28.219*** (4.26)	38.762*** (5.33)	26.883*** (3.31)	32.753*** (5.38)	37.921*** (4.42)
Food trade openness, 10th	27.580*** (4.27)	37.609*** (5.34)	24.647*** (3.38)	32.065*** (5.37)	36.997*** (4.42)
Food trade openness, 50th	25.701*** (4.29)	36.017*** (5.34)	24.647*** (3.38)	30.205*** (5.35)	35.720*** (4.42)
Food trade openness, 90th	24.761*** (4.30)	33.272*** (5.34)	23.515*** (3.41)	28.197*** (5.31)	33.519*** (4.42)
Food trade openness, 95th	23.822*** (4.31)	31.433*** (5.33)	21.952*** (3.45)	26.802*** (5.28)	32.419*** (4.41)

Robust z-statistics in parentheses.*** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$, # $p < 0.15$

Table 7 The effect of food trade openness on hunger prevalence: net importers versus net exporters

Hunger prevalence	(1) Food exporter	(2) Food importer
Food trade openness, logged	−2.338	12.175***
	(−1.08)	(3.83)
Non-food trade openness, logged	15.588	−4.520
	(1.40)	(−0.34)
Non-agricultural trade openness, logged	0.185	1.096**
	(0.40)	(2.33)
GDP per capita, logged	−10.365***	−4.716**
	(−2.97)	(−2.21)
Urbanization	−0.089	−1.252
	(−0.16)	(−1.13)
FDI	0.001	−0.112***
	(0.02)	(−4.23)
Observations	509	489
Ramsey RESET, *p* value	0.13	0.41
First-stage, F-stat	20,135	3336

Robust z-statistics in parentheses.*** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$, # $p < 0.15$. n.a..: non-applicable. Regressions control for rainfall and rainfall squared

trade should pay particular attention to sustained social cohesion and nation-building.[14]

Another interesting result across Table 6 is that no matter what the quality of institutions and the degree of ethnic tensions, increased food trade openness is consistently associated with increased hunger. This suggests that improving the quality of institutions is not enough to prevent the negative consequences of food trade in developing countries.

Last, we examine whether the effect of food trade openness on hunger is conditional on the food trade position of a country, that is, whether a country is a net food importer or exporter. Table 7 provides the additional estimates. The coefficient for food trade openness in column 1 Table 7 is negative (−2.338) suggesting that food trade openness may decrease hunger in net food exporting countries. The estimate is however statistically insignificant.

On the contrary, the coefficient for food importers is positive (12.175) and statistically significant indicating that food trade openness increases hunger for net food importers. A 10% increase in food trade openness would increase hunger by about 8% in net food importers. The impacts of food trade openness estimated in Table 7 seem to confirm that developing countries that have a net food exporting position are less affected by openness. This could be because these countries enjoy comparative advantages and that further access to export markets result in additional productivity gains offsetting the decreased producer prices. This might also suggest that policymakers may be better off financing productivity-enhancing measures in the agricultural sector to eliminate the negative effects of distorted trade (Mgeni et al. 2018).

[14] It is important to note we cannot directly compare the estimates in Table 6 with those of Table 3 because the interaction models include more variables than in Table 3.

5 Conclusions

The Sustainable Development Goals (SDG) adopted in 2015 call for an end to malnutrition in all its forms and for all people by 2030. Free trade has often been accused of causing undernutrition in low income and middle income countries. This paper has estimated the impact of food trade openness on the prevalence of undernourishment in a sample of 52 developing countries between 1990 and 2013.

We found evidence that food trade openness causes undernutrition in developing countries. In particular, a 10% increase in food trade openness would increase undernourishment prevalence by about 6%. We show the negative effect of trade on extreme hunger is robust to alternative model specifications. Furthermore, we show that these effects on hunger are driven by decreased GDP per capita in the food sector and decreased agricultural producer prices despite gains in the food supply as a result of increased food trade openness. We also find that developing countries typically restrict food trade openness as a response to hunger crises since a percentage point increase in hunger prevalence would result in a 0.9% decrease in food trade openness. This may be interpreted as evidence of ‘food protectionism’.

Further we found heterogeneous effects of food trade across several dimensions of governance, but found little support that improving the quality of institutions would make food trade more beneficial for developing countries. A more important finding relates to the net food trade position of a country. In particular, we found that food exporters do not suffer from increased food trade openness, perhaps because they already enjoy comparative advantages. On the contrary, food net importers seem to be affected by food trade. This may be because the negative effects of food trade openness, due to agricultural distortions or lack of efficiency in the domestic food sector, do not offset the positive effects of having access to a larger food supply. This warns against examining the effect of trade with a purely supply-sided lens.

While the SDG offers a great opportunity and platform to provide a better environment to correct for the effects of free trade, our results suggest that developing countries may be better off adopting food self-sufficiency and focus on improving the efficiency of domestic producers for some time despite such actions clashing with World Trade Organization’s regulations and current agenda. Combined trade and health and food interventions must be re-designed and implemented to fight undernourishment.

Last, we have examined one side of the double burden of malnutrition, that is, undernutrition. Further research could make use of the methodology developed in this paper to examine the effects of trade openness on overweight and obesity in developing countries.

Compliance with ethical standards

Conflict of interest The author declares no conflict of interest.

Appendix 1

Table 8 List of 52 countries

Algeria	Honduras	Nicaragua
Armenia	India	Nigeria
Azerbaijan	Indonesia	Niger
Bangladesh	Iran	Pakistan
Bolivia	Jamaica	Panama
Brazil	Jordan	Peru
Burkina Faso	Kazakhstan	Philippines
Cameroon	Kenya	Senegal
Colombia	Lebanon	South Africa
Costa Rica	Madagascar	Sri Lanka
Cote d'Ivoire	Malawi	Tanzania
Dominican Republic	Malaysia	Thailand
Ecuador	Mali	Togo
Egypt	Mexico	Tunisia
El Salvador	Mongolia	Turkey
Ethiopia	Morocco	Vietnam
Ghana	Mozambique	Zimbabwe
	Namibia	

Appendix 2. Test of exclusion restrictions in the first step

Table 9 Test of exclusion restrictions (first step)

	(1) Size	(2) Test 1 2SLS-IV	(3) Test 2 CUE
Dependent variable	Food trade openness	Food Trade openness	Food Trade openness
Rainfall anomaly	0.030	−0.001	−0.001
	(1.30)	(−0.08)	(−0.07)
Rainfall anomaly, squared	−0.316***		
	(−2.93)		
Hunger prevalence		−0.009	−0.010
		[0.11]	[0.20]
Observations	998	998	998
First-stage, F-stat	n.a.	9.84	5.61
Stock-Yogo 15% size	n.a.	8.96	5.33
Hansen J, p value	n.a.	n.a.	0.46

Robust z-statistics in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$, # $p < 0.15$. Stock-Wright LM p values between square brackets. Test 1 only uses the quadratic term of rainfall anomalies in the instrument set. The equation is exactly identified and is estimated via 2SLS-IV. Hence the p value for the Hansen test is not reported. Test 2 adds temperature anomalies in the instrument set. The model is over-identified and estimated via CUE

Appendix 3. First-stage estimates for causal pathways

Table 10 Transmission channels: first-step estimates

Food Trade Openness	(1) CUE	(2) 2SLS-IV	(3) 2SLS-IV
Food GDP per capita	1.593***		
	[0.00]		
Food supply		−5.512**	
		[0.02]	
Producer prices			0.606*
			[0.07]
Observations	995	995	983
First-stage, F-stat	6.010	26.23	532.7
Hansen J, p value	0.53	0.35	0.39

Robust z-statistics in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$, # $p < 0.15$. Stock-Wright LM p values between square brackets. Models in columns 1, 2 and 3 do not include other independent variables. Models 2 and 3 use temperature anomalies in the instrument set

Appendix 4. Robustness analyses

Table 11 The effect of food trade openness on hunger prevalence

Hunger prevalence	(1) Dynamic model GMM	(2) Country trend effects 2SLS-IV	(3) Endogenous GDP 2SLS-IV	(4) 2SLS/first step 2SLS-IV	(5) Extended model 2SLS-IV
Food trade openness, logged	8.587***	4.263**	10.862***	10.482***	3.555*
	(4.03)	(2.14)	(4.64)	(4.59)	(1.66)
GDP per capita, logged	−6.903***	−7.907***	−6.186***	−6.515***	−8.698***
	(−4.26)	(−3.49)	(−3.32)	(−3.48)	(−4.56)
Urbanization	−0.459	−0.464	−0.591	−0.575	−0.316
	(−0.82)	(−0.86)	(−1.01)	(−0.99)	(−0.57)
Non-food trade openness, logged	−3.153	12.846	−0.183	0.086	22.605**
	(−0.31)	(1.08)	(−0.02)	(0.01)	(2.53)
Non-agricultural trade openness, logged	0.441	0.668**	0.780**	0.780**	0.052
	(1.43)	(2.06)	(2.37)	(2.39)	(0.19)
FDI	−0.054#	−0.084**	−0.100***	−0.100***	−0.117***
	(−1.56)	(−2.32)	(−3.47)	(−3.45)	(−3.31)
Lagged hunger	−0.036***				
	(−6.48)				
Inflation (CPI)					−0.001
					(−1.17)
Share of agriculture in GDP					−0.154***
					(−3.50)
Socio-economic conditions					0.059
					(0.35)
Ethnic tensions					0.197
					(1.10)
Observations	916	998	998	998	896
Ramsey RESET, p value	n.a.	0.02	0.11	0.10	0.10
First-stage, F-stat	1319	49,636	1191	2648	84,799

Robust z-statistics in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$, # $p < 0.15$. Regressions control for rainfall and rainfall squared

Table 12 The effect of food trade openness on hunger prevalence

Hunger prevalence	(1) OLS-FE Baseline	(2) OLS-FE	(3) 2SLS-IV	(2) 2SLS-IV Baseline
Food trade openness, logged	1.341	1.576	10.873***	10.862***
	(0.69)	(0.78)	(4.65)	(4.66)
Non-food trade openness, logged	6.291	6.932	−1.107	−0.207
	(0.66)	(0.73)	(−0.11)	(−0.02)
Non-agricultural trade openness, logged	0.469	0.473	0.798**	0.793**
	(1.32)	(1.39)	(2.42)	(2.42)
GDP per capita, logged	−8.090***	−8.034***	−6.465***	−6.450***
	(−3.65)	(−3.65)	(−3.38)	(−3.45)
Urbanization	−0.581	−0.499	−0.669	−0.578
	(−1.18)	(−0.96)	(−1.19)	(−0.99)
FDI	−0.094***	−0.097***	−0.097***	−0.100***
	(−3.20)	(−3.50)	(−3.15)	(−3.45)
Rainfall anomaly		−1.254***		−1.351***
		(−2.78)		(−2.66)
Rainfall anomaly, squared		10.276***		11.164***
		(3.36)		(3.63)
Observations	998	998	998	988
Number of countries	52	52	52	52
Ramsey RESET, p value	0.09	0.06	0.40	0.48
First-stage, F-stat	n.a.	n.a.	2358	2375

Robust z-statistics in parentheses.*** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$, # $p < 0.15$. n.a..: non-applicable

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Dr. Sebastien Mary is a Lecturer of Economics at the University of Kansas and a research affiliate at DePaul University who researches the interactions between economics, agriculture, and nutrition. He has previously held positions at the Joint Research Centre of the European Commission, IFREMER and the University of Aberdeen. He has published on the impact assessment of agricultural reforms and rural development, the impact of nutrition-sensitive foreign aid on extreme hunger and global food security.