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RESEARCH ARTICLE

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The effects of agricultural trade openness on food price transmission in Latin American countries

Insa Flachsbarth* and Alberto Garrido

CEIGRAM, Universidad Politécnica de Madrid. Department of Agricultural Economics and Social Sciences. Madrid, Spain

Abstract

Trade of agricultural commodities has grown significantly in most Latin American countries (LAC) over the last two decades. However, after the international food price surges in 2006-08 and 2011-12 concerns about food access of the poor arose. Within a panel framework containing six LAC (Argentina, Brazil, Chile, Colombia, Mexico and Peru), we used a single equation error correction model to identify possible cointegrating relationships between the food consumer price index (CPI) and a set of trade related and domestic variables. The main focus of the study was to examine how different levels of trade openness impact international food price transmission to domestic markets. Our results confirm that deeper market integration increases global price transmission elasticities. In other words, more agricultural trade openness proves to elevate food CPIs during global price spikes. Thus, for poor consumers world price shocks can be deteriorating in the short-run and domestic food prices will slowly converge to a higher long-run equilibrium. Especially in increasingly integrated economies, effective policies to buffer food price shocks should be put in place, but must be carefully planned with the required budget readily available. We also found that exchange rate appreciations can buffer price shocks to a certain extent and that monetary policies seem to be an appropriate means for stabilizing food prices to safeguard food access of the poor population.

Additional key words: international food trade; market liberalization; global price shock; food security; food access; error correction model; consumer food prices.

Introduction

Since the mid-1980s, reducing agricultural and non-agricultural trade distortions has led to structural changes in Latin American countries (LAC), affecting food prices as well as economic development (Anderson *et al.*, 2011). According to the law of one price, trade liberalization leads to a price increase of exported food products and a price decrease of imported food, because domestic prices adjust towards global price levels (Goodwin *et al.*, 1990; Miljkovic, 1999). However, the recent experience during the international food price crisis of 2006-08 and 2011-12 was a painful lesson to some of the poorer importing countries (Headey & Fan, 2008; Hoyos & Medvedev, 2011; Attanasio *et al.*, 2013; Rodriguez-Takeuchi & Imai, 2013). ECLAC (2008) states that food exporters that increasingly sell into international markets have experienced accelerated food price inflation. In this context, the transmission of high international prices into domestic prices has received attention (Benson *et al.*, 2008; Dawe, 2008; Alemu & Ogundeji, 2010; Cudjoe *et al.*, 2010; Jalil & Tamayo, 2011; Minot, 2011; Baquedano & Liefert, 2014). All these studies show country and crop specific results,

* Corresponding author: insa.niemeyer@gmail.com
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Abbreviations used: ADF (Augmented Dickey-Fuller); ADL (autoregressive distributed lag); CPI (Consumer Price Indices); FGLS (feasible generalized least squares); GMM (generalized method of moments); iid (identical independently distributed); IMF (International Monetary Fund): LAC (Latin American countries); M2 (money supply); PP (Phillip Perron); SEECM (single equation error correction model).

suggesting different price transmission rates from international agricultural commodity markets to developing countries' domestic markets. Since food prices to a large extent determine poor consumers' ability to access sufficient food, these findings are very relevant in the global and regional food security discussion. Especially in those countries where food purchases comprise a large share of households' total expenditure, high transmission rates in times of rising international prices can push people into deeper poverty causing malnutrition and hunger. Many authors argue that countries that are more integrated into world markets might show higher world price transmission rates. However, the impact of trade openness on transmission rates has not been empirically investigated. From a national and sectoral policy perspective though, it would be crucial to be able to directly relate world price transmission to trade liberalization tendencies in the agricultural sector. Deeper knowledge about these interdependencies would help to design effective national food security programs (Dorward, 2012; Dawe & Maltsoglou, 2014).

The aim of this study is to examine the degree of cointegration between world food commodity prices and domestic food consumer prices. Specifically, we evaluate the impacts of liberalized agricultural trade regimes on price transmission rates for a panel of large LAC's trading nations (Argentina, Brazil, Chile, Colombia, Mexico and Peru). The choice of these six countries is based on a combination of factors. First, the sample includes the two largest emerging world food exporters (Argentina, Brazil), one of the most opened and export-oriented in the world for two decades (Chile), one of the most import-dependent for staple goods and with growing market integration in a trade block (Mexico and North American Free Trade Agreement, NAFTA), and two mid-size countries which recently changed its trade regimes (Colombia and Peru). Secondly, the six countries show remarkable growth of agricultural exports and imports (Willaarts *et al.,* 2014). And thirdly, except Argentina and Chile, the remaining four still face significant, though decreasing rates of food insecurity among the poorest. We use our estimation results to calculate actual transmission rates in crises years. Complementary to international market forces, we follow Durevall *et al.* (2013) and include some domestic macroeconomic causes possibly influencing food price movements.

Thus, our study complements and expands existing literature in several ways. Although the error correction model has been widely used with respect to price transmission analyses (see *e.g.* (Cudjoe *et al.*, 2010; Minot, 2011; Baquedano & Liefert, 2014), the effect of trade openness is still unsettled in the literature. Moreover, most previous studies on price transmission do not control for domestic food price determinants or the effects from movements in exchange rates. To our knowledge, only Dawe & Slayton (2010), Baek & Koo (2014) and Baquedano & Liefert (2014) take into account exchange rate effects, and Durevall *et al.* (2013) also consider agricultural supply and demand shifters. In contrast to many other papers, our analysis studies changes in food consumer price indices (CPI), instead of specific food items. We choose this approach, first because idiosyncratic food habits and the composition of the food basket vary significantly across the selected sample of countries. Secondly, the recent literature on food price transmission both for developed, mid-income and emerging countries has emphasized the role of macro-economic aspects (Dorward, 2012). And thirdly, using the food CPI allows for more general conclusions with regard to food security issues, because price changes of different products might lead to substitution effects within the food consumption basket.

Altogether, the results allow for drawing some conclusions about interactions of global market integration and urban food price changes in LAC, considering current global market trends and different trade regimes. Although our results do not allow us to directly make conclusions about the effects of trade on food access of the poor, certainly changing urban food prices is one major driver for improving or exacerbating malnutrition.

Material and methods

Methods

Our empirical analysis is motivated by a composition of price determinants of tradable and non-tradable food items. We define the data generating process of the general price level as a function of the prices of tradable (P_{agT}) and non-tradable (P_{agNT}) goods:

$$P_A = f(P_{agT}, P_{agNT}) \qquad [1]$$

Following the «law of one price», the price of tradable goods in a small open economy is determined in the world market. Depending on policy circumstances, such as tariffs, trade quotas or export taxes, the price can deviate from the world price level. This implies that the price of tradable goods (P_{agT}) is determined by the world market (PW_{ag}), the exchange rate (xrt) as well as marketing margins ($margin$) and tax/subsidy wedges (ta) (Diaz-Bonilla & Robinson, 2010):

$$P_{agT} = xrt \times (1 + ta) \times (1 + margin) \quad [2]$$

Non-tradable products can neither be imported nor exported, and thus follow the market clearing condition (Gros & Hefeker, 2002). This means that their price is determined endogenously by the interaction between domestic demand (Q^d_{agNT}) and domestic supply (Q^s_{agNT}):

$$Q^d_{agNT} = Q^s_{agNT} \quad [3]$$

In developing and emerging countries where a large share of the total household expenditure is on food, the demand for agricultural products is expected to be highly influenced by aggregate demand (Ahsan *et al.*, 2011). Thus, we used money supply ($M2$) or alternatively the national unemployment rate ($Unemp$) as a proxy for total demand. Aggregate supply was proxied by world oil prices (Pw_{oil}) because an increase in oil prices is followed by an increase in input costs which in turn affects agricultural supply (see Hanson, 1993; Nazlioglu & Soytas, 2011; Durevall *et al.* 2013) and also has been considered in the literature as shifter in the formation process of marketing margins in food chains (Leibtag, 2009; Davidson *et al.,* 2011). If we subsume marketing margins ($margin$) and trade wedges (ta) under a trade openness indicator (top), domestic agricultural consumer food prices can be expressed by the following function:

$$P_A = f(Pw_{ag}, xrt, top, M2^1, Pw_{oil}) \quad [4]$$

Empirical model

In general when dealing with macroeconomic time series data, one has to test whether the variables contain unit roots (non-stationarity), and if so whether they are cointegrated (Österholm, 2004). Our data showed these features, so we formulated an error correction model which permitted us to describe both the long-run equilibrium relationship and the short-run dynamics between some independent variables and the dependent variable that were derived in Eq. [4]. More specifically, we could estimate the extent to which consumer prices reacted to changes in world prices, exchange rate or money supply movements and the time it takes to adjust domestic consumer food prices to the new long-run equilibrium after a shock of one of the three variables (Baquedano *et al.,* 2011). In addition to these three variables an interaction term between the world price index and trade openness was introduced to obtain insights about the effect of trade liberalization tendencies on long and short-term price transmission rates.

According to De Boef & Keele (2008), in time series analysis an explanatory variable may have only short term causal effects on the dependent variable or both short and long-run causal effects as described above. Short-run effects may occur at any lag, but the effect does not persist into the future. Thus, apart from the described variables with possibly long-run effects, we included an agricultural supply shifter, namely world prices of crude oil as derived above, where we assume only short-run effects.

To formulate the error correction model, we depart from an autoregressive distributed lag (ADL) model as described in Eq. [5]. This general form of the model is reported as an ADL(1,1) process which means that one lag of the dependent and one lag of the possibly cointegrated independent variables are considered as regressors. However, we make no a priori assumptions about appropriate lag length in the model, but in our estimation procedure we follow a general to specific approach and eliminate insignificant lags to obtain a more parsimonious model.[2]

$$\begin{gathered}P_{A_{it}} = \alpha_0 + \alpha_1 P_{A_{it-1}} + \psi_0 Pw_{AG_{it}} + \psi_1 Pw_{AG_{it-1}} + \phi_0 xrt_{it} + \\ \phi_1 xrt_{it-1} + \zeta_0 M2_{it} + \zeta_1 M2_{it-1} + \beta_0 top_{it} + \beta_1 top_{it-1} + \\ \omega_0 (Pw_{AG_{it}} * top_{it}) + \omega_1 (Pw_{AG_{it-1}} * top_{it-1}) + \\ \eta_0 Pw_{oil_{it}} + \tau yd_t + \nu_i + \varepsilon_{it}\end{gathered} \quad [5]$$

where i represents the cross-section (country), t the different years of the panel and yd are year dummies, ν_i are country fixed effects and ε represents the iid error term.

The standard modeling approach when dealing with non-stationary and cointegrated variables has been the two step (Engle & Granger, 1987) error correction

[1] Alternatively for M2 (in local currency units or as a share of GDP) the unemployment rate is used to proxy demand.
[2] We begin with an ADL(8,8) which implies eight lags for dependent and independent variables. Taking the AIC and BIC criterion we eliminate insignificant lags.

model. In our study, however, we rely on a single equation error correction model (SEECM), because it has the advantage that not all time series variables need to have unit roots (Lütkepohl, 2005; Banerjee *et al.,* 1998). For the panel data we applied unit root tests following Levin *et al.* (2002) and Im *et al.* (2003). For the world price data which repeat in each cross-section, we used the regular Augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP) stationarity tests.

To derive the unrestricted SEECM, we add and subtract lags of the variables in Eq. [5] which yields in Eq. [6]:

$$\begin{aligned} \Delta P_{A_{it}} &= \alpha_0 + (\alpha_1 - 1) P_{A_{it-1}} + \psi_0 \Delta Pw_{AG_{it}} + (\psi_0 + \psi_1) Pw_{AG_{it-1}} + \\ &\phi_0 \Delta xrt_{it} + (\phi_0 + \phi_1) xrt_{it-1} + \zeta_0 \Delta M2_{it} + (\zeta_0 + \zeta_1) M2_{it-1} + \\ &\beta_0 \Delta top_{it} + (\beta_0 + \beta_1) top_{it-1} + \omega_0 (\Delta Pw_{AG_{it}} * \Delta top_{it}) + \\ &(\omega_0 + \omega_1)(Pw_{AG_{it-1}} * top_{it-1}) + \eta_0 \Delta Pw_{oil_{it}} + \tau yd_t + \nu_i + \varepsilon_{it} \end{aligned} \quad [6]$$

After substituting and factoring common parameters, we arrive at:

$$\begin{aligned} \Delta P_{A_{it}} &= \alpha_0 + \delta P_{A_{it-1}} + \lambda_0 \Delta Pw_{AG_{it}} + \lambda_1 Pw_{AG_{it-1}} + \\ &\theta_0 \Delta xrt_{it} + \theta_1 xrt_{it-1} + \kappa_0 \Delta M2_{it} + \kappa_1 M2_{it-1} + \\ &\mu_0 \Delta top_{it} + \mu_1 top_{it-1} + \pi_0 (\Delta Pw_{AG_{it}} * \Delta top_{it}) + \\ &\pi_1 (Pw_{AG_{it-1}} * top_{it-1}) + \eta_0 \Delta Pw_{oil_{it}} + \tau yd_t + \nu_i + \varepsilon_{it} \end{aligned} \quad [7]$$

where $\delta = (\alpha_1 - 1)$; $\lambda_0 = \psi_0$; $\lambda_1 = (\psi_0 + \psi_1)$; $\theta_0 = \phi_0$; $\theta_1 = (\phi_0 + \phi_1)$; $\kappa_0 = \zeta_0$; $\kappa_1 = (\zeta_0 + \zeta_1)$; $\mu_0 = \beta_0$; $\mu_1 = (\beta_0 + \beta_1)$; $\pi_0 = \omega_0$; $\pi_1 = (\omega_0 + \omega_1)$. Collecting common terms after rearranging Eq. [7], we obtain the following SEECM equation:

$$\begin{aligned} \Delta P_{A_{it}} &= \alpha_0 + \vartheta \Delta Pw_{AG_{it}} + \rho \Delta xrt_{it} + \kappa \Delta M2_{it} + \\ &\mu \Delta top_{it} + \pi (\Delta Pw_{AG_{it}} * \Delta top_{it}) + \eta \Delta Pw_{oil_{it}} + \\ &\delta \big(P_{A_{it}-1} - \gamma Pw_{AG_{it-1}} - \varphi xrt_{it-1} - \xi M2_{it-1} - \\ &\varepsilon top_{it-1} - \chi (Pw_{AG_{it-1}} * top_{it}) \big) + \tau yd_t + \nu_i + \varepsilon_{it} \end{aligned} \quad [8]$$

where $\vartheta = \lambda_0$; $\rho = \theta_0$; $\kappa = \kappa_0$; $\mu = \mu_0$ and $\eta = \eta_0$ are the short-run elasticities. The long-run elasticities are given by $\gamma = (1 - \lambda_1/\delta)$; $\varphi = (1 - \theta_1/\delta)$; $\xi = (1 - \kappa_1/\delta)$; $\varepsilon = (1 - \mu_1/\delta)$; $\chi = (1 - \pi_1/\delta)$; δ represents the error correction term which indicates the speed of adjustment to the new long-run equilibrium. To obtain consistent standard errors of the long-run coefficients we perform the Bewley's transformation (for details see De Boef & Keele (2008) and footnotes in Baquedano & Liefert, 2014).

A cointegration relationship assumes that and that ($\gamma \neq 0$; $\varphi \neq 0$; $\xi \neq 0$; $\varepsilon \neq 0$; $\chi \neq 0$) and ε_{it} is stationary I(0). If none of the long-run coefficients is significantly different from zero the variables are not cointegrated, while a coefficient of one would mean complete transmission. Coefficients can also take on values larger than one, which would mean that that the dependent variable over-shoots when the independent variable experiences a shock.

Eq. [8] was estimated using a panel approach and controlling for country fixed effects to limit omitted variable bias. After testing for serial correlation in the idiosyncratic errors, following Wooldridge (2002), we assumed the error term of this model follows a first-order autoregressive process. Further, we needed to correct for heteroscedasticity. Thus, the model was estimated using STATA's xtpcse and xtgls commands which execute a Praise Winsten estimator and a Feasible Generalized Least Squares estimator, respectively.[3] These models assume weak exogeneity of all independent variables, meaning that causality runs from world prices, exchange rate and money supply to domestic consumer food prices. Since this might be a strong assumption for the exchange rate, money supply, and maybe even world prices[4], we treated those variables as endogenous using a system GMM estimator (Blundell-Bond)[5] to check whether the simultaneous equation bias is severe. To guarantee the robustness of the estimator, we tested the moment conditions for no serial correlation in the idiosyncratic errors. Further we tested whether the moment conditions used are valid by implementing the Sargan test (for details see Arellano & Bond, 1991).

Applying the relatively new approach of using panel data in an error correction model[6] instead of single

[3] For details on methods see Greene (2003).

[4] World prices and domestic prices would only be endogenously determined if the country was a large importing or exporting nation with enough market power to change the world price. Since our sample includes very large agricultural trading nations like Brazil and Argentina, simultanous endogeneity might be problematic. However, we are using food price indices containing many different products, so it more likely that all countries are price takers.

[5] For details see Blundell & Bond (1998).

[6] Panel data methods for non-stationary data were first developed in the early 1990s and only recently applied more frequently. For details see *e.g.* Kao (1999), Madala & Wu (1999), Levin *et al.* (2002).

country estimates has some advantages and disadvantages. Since trade data are only available annually, this approach allowed us to estimate a possible long-run relationship between variables, even though the time dimension would have been too short for making reliable inference for any single country estimation. A weakness of the panel data estimation might have to do with sample selection bias in the estimates. To check whether our estimates suffered from this bias, we compared model results from the full six country panel with results from only including five countries in the regression (omitting either Argentina or Brazil or Chile or Colombia or Mexico or Peru).

Our study is meant to discuss food price transmission after recent food price spikes with a specific focus of evaluating the effects of more trade openness in a sample of LAC. First, we applied actual degrees of agricultural trade openness within each country in the year of the price spike to compare food price transmission across countries. Secondly, we constructed a set of counterfactual degrees of trade openness in the year of the price shock to measure by how much food price transmission depends on the degree of trade liberalization. SEECM provides the long-run coefficients γ, χ and the speed of adjustment $|\delta|$ as well as the short run coefficients ϑ and π. Employing such a multiplicative interaction model allowed us to calculate the marginal effect of world price changes on domestic food CPI, depending on the level of trade openness, which is given as by:

$$\frac{\partial P_A}{\partial Pw_{AG}} = \gamma + \chi * top$$

(respectively for short-run world price elasticities). Note that the marginal effect can still be significantly different from zero even if the coefficient of the interaction term were insignificant, because the standard error of the marginal effect is not a direct output of the regression result, but must be calculated as follows (Brambor, 2005):

$$\hat{\sigma}_{\frac{\partial P_A}{\partial Pw_{AG}}} = \sqrt{var(\hat{\gamma}) + (top)^2 var(\hat{\chi}) + 2(top)cov(\hat{\gamma}\hat{\chi})} \quad [9]$$

Furthermore, we calculated the median lag-length of the 2008 shock's effect on domestic food prices. We followed Baquedano & Liefert (2014) and defined the median lag-length as the number of periods at which at least half of the new equilibrium value of the domestic food price from the world price shock is reached. Mathematically, this can be expressed as follows (De Boef & Keele, 2008):

$$m = \frac{\sum_{t=0}^{T} \nu_t}{\sum_{t=0}^{\infty} \nu_t} \quad [10]$$

where $\Sigma_{t=0}^{\infty}\nu_t$ is the long run transmission elasticity of the world price. $\Sigma_{t=0}^{T}\nu_t$ is the summation across the number of periods T of the adjustment process towards the new food price equilibrium. Thus, the median lag-length is reached when $m \geq \gamma + \chi * \text{top}$.

Table 1 summarizes some hypotheses of the effects of each variable on domestic consumer food prices. Note that we limit ourselves to interpreting the marginal effect of the interaction term, instead of the constitutive terms "world price index" and "trade openness". The reason is that in the applied multiplicative interaction model, γ captures the marginal effect of a 1% increase in world prices when trade openness is zero which is an unrealistic assumption.

Data

Our dataset contains six cross-sections (Argentina, Brazil, Chile, Colombia, Mexico and Peru) and includes time series data from 1995 until 2013 which adds up to 114 observations. Suppl. Table S1 [pdf online] gives an overview of all data with sources and Suppl. Table S2 [pdf online] shows the corresponding descriptive statistics of each variable.

Fig. 1 illustrates the dynamics in our six LAC, with respect to global and domestic food price movements, the general CPI trend as well as changes money supply and the exchange rate between 1995 and 2008. There was a certain co-movement of world food prices and domestic food prices, especially in Argentina, Chile and Peru. In all countries, except Argentina, food prices rose faster than the general price level after 2007 which coincides with the recent high price trend in international food commodity markets. Suppl. Fig. S1 [pdf online] visualizes developments of trade openness in the agricultural sector between 1995 and 2013. In the early 2000's, agricultural trade as a share of agricultural GDP began to rise sharply, particularly in Brazil, Chile and Mexico.

Results

Before discussing the long-term and short-term determinants of domestic consumer food prices in light of agricultural trade openness, we give a brief overview

Table 1. Hypotheses of the relevant variables' effect on domestic food price levels

Variable	Hypothesis	Ground
Interaction term between world agricultural prices and agricultural trade openness indicator (Trade openness * World price index)	Food prices increase	World food price transmission is expected to increase with higher degrees of trade openness, because countries would be more affected by international price fluctuations than rather closed economies.
Exchange rate	Food prices increase	In the short-run, a currency depreciation makes imports more expensive, and thus prices of tradables rise. A depreciation also makes exports more competitive in the world market, thereby increasing demand for LAC's food products and hence prices.
		In the long-run, under the assumption of perfect arbitrage, higher prices will lead to an appreciation of the exchange rate which in turn leads to a downward price adjustment.
Money supply	Food prices increase	If money supply increases, aggregate demand increases which leads to a higher price level in the economy, including food prices.
Unemployment rate	Food prices decrease	If unemployment increases, aggregate demand decreases which leads to a lower price level in the economy, including food prices.
World oil prices	Food prices increase	If input costs in agricultural production increase, aggregate supply decreases which leads to increasing food prices.

of the stationary properties of our data. Table 2 shows that most time series in levels contain unit roots, but are stationary in first differences. Only the unemployment rate, and maybe the exchange rate and trade openness (depending on the unit root test applied) are also stationary in levels.[7] This indicates that the SEECM is the appropriate method, because the model is not restricted to non-stationary data. An alternative way to treat non-stationary data would be to estimate a model with variables transformed to first differences. However, this approach does not capture the long-run properties of cointegrated variables. We therefore only performed SEECM estimations.

Our estimates of domestic food price determinants are reported in Table 3, showing the results of the Praise Winsten regressions. Other model specifications and estimations using the FGLS estimator or the system GMM estimator are omitted because they yield very similar results to the Praise Winsten regression. This shows that endogeneity of world price indices, the exchange rate and money supply is not problematic. Table 3 is structured so that the first column shows results of estimations using the world price index of the International Monetary Fund (IMF), while the second column shows the same results using instead the world price index provided by FAO. The last two columns are reported to demonstrate robustness of the results with respect to sample selection bias. Since Brazil represents a very large exporting nation and Mexico a very large importing nation, we report five-country panel estimates without these two countries. We also ran regressions without Argentina, Chile, Colombia or Peru, but the main coefficients remain stable, so results are not reported.[8] Even though selection bias is not severe, depending on the countries included, estimates of price transmission deviate from each other to a certain extent. However, the main trends remain unchanged, so we ignore these slight differences and interpret only the six-country panel estimate that uses world prices provided by the IMF. Post-estimation diagnostic tests are reported for all models demonstrating validity of the estimation.

All coefficients of the significant variables have the expected sign, independent of the model specification

[7] All tests were conducted with and without a deterministic trend. Because of space constraints, we present only the results without trend, as they do not change our conclusions about the stationarity properties of our data series. The tests were initially conducted with a maximum of 8 lags. However, the tests results were no different when using a more parsimonious lag structure. The results of the non reported tests are available from the authors upon request.

[8] They are available from the authors upon request.

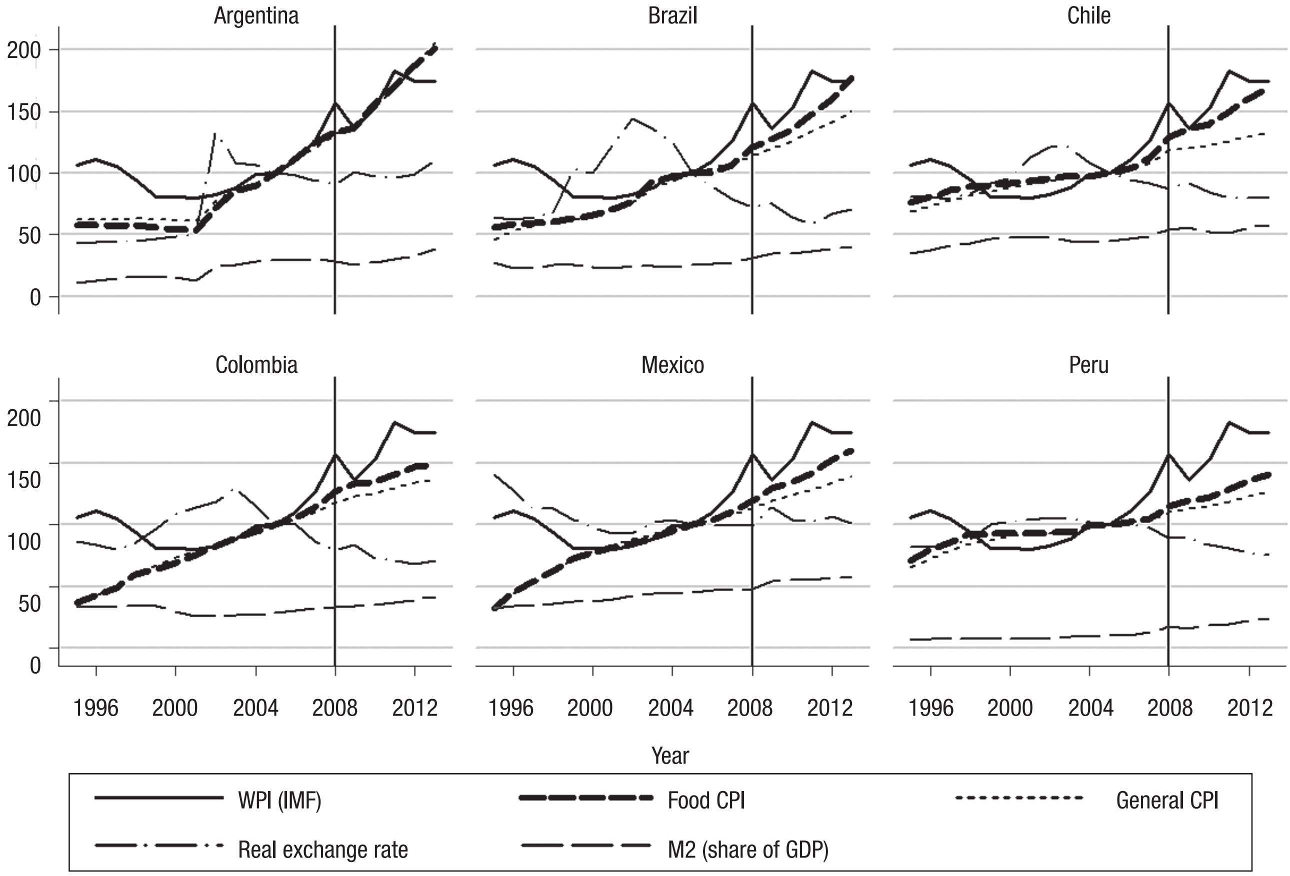

Figure 1. Evolution of the world price index (food and beverages IMF), domestic general consumer price indices (CPIs), domestic food CPIs (base year 2005), and money supply (expressed as a percentage of GDP) and real exchange rate movements (expressed as an indicator). *Source:* Data obtained from FAO (2014), ECLAC (2014) and Inter-American Development Bank (2014).

or the estimator. We are primarily interested in the marginal effect of the world food price index on the domestic food price index with varying levels of trade openness. Hence, Table 4 shows the calculated marginal effects and the corresponding standard errors according to Eq. [9]. Note that although the short-run coefficient of the interaction term is insignificant (see Table 3), the marginal effects of the parameters become statistically significant until trade openness reaches 140%. Since all countries of interest (except for Chile after 2003) are within this range, insignificance of higher degrees of trade openness does not affect our interpretations severely. Table 4 illustrates that increasing levels of trade openness especially impact the degree of price transmission in the short-run. In other words, if there is a global food price shock, the instantaneous reaction of the domestic food CPI highly depends on the level of market integration. Long-run transmission rates are also positively influenced by higher degrees of trade-openness and significant at all levels. However, the effect of trading activity is more moderate in the long-run than in the short-run.

We applied the discussed results to six LAC and estimated price transmission rates after the international food price spike of 2008. All countries showed different degrees of agricultural market integration during the price shock: Argentina 107%, Brazil 45%, Chile 220%, Colombia 45%, Mexico 80% and Peru 70% (see Suppl. Fig. S1 [pdf online]). This has led to different world food price transmission rates, varying between 100% in Brazil and Colombia, 110% in Peru, 120% in Mexico, 130% in Argentina and 170% in Chile[9] (short-run plus long-run transmission) in our six studied countries.

[9] Chile's long-run coefficient of the marginal effect is statistically significant. But the short-run coefficient is not, so there are higher uncertainties in the interpretation of the effects of trade openness on total price transmission than for the other coutries.

Table 2. Unit root tests

	In levels		In differences	
Panel unit root tests	**LLC**	**IPS**	**LLC**	**IPS**
Food CPI (log)	6.86	–0.20	**–1.20****	**–3.82****
Trade openness	2.94	**–1.46***	**–2.53*****	**–9.32*****
World price index (IMF) (log)* Trade openness	3.28	–1.00	**–1.87****	**–8.13*****
World price index (FAO) (log)* Trade openness	3.10	–0.98	**–2.71*****	**–8.10*****
Exchange rate (log)	**–12.18****	0.04	**–3.25*****	**–4.09*****
M2 (in local currency) (log)	3.81	–0.25	**–2.05****	**–4.70*****
M2 (as share of GDP) (log)	3.62	1.02	**–2.76*****	**–6.43*****
Unemployment rate (log)	**–1.67****	**–3.24*****	**–7.17*****	**–3.39*****
Unit root tests for world prices	**ADF**	**PP**	**ADF**	**PP**
World food price index (IMF) (log)	–0.25	–0.14	**–2.82***	**–3.32****
World food price index (FAO) (log)	0.00	–0.29	**–3.02****	**–3.50*****
World prices crude oil (log)	–0.12	–0.43	**–3.52*****	**–4.60*****

Note: LLC = Levin, Lin, Chu (2002), IPS = Im, Pesaran, Shin (2003). The statistics are asymptotically distributed as standard normal with a left hand side rejection area. Total number of observations (N * T) are 114. ADF = Augmented-Dickey-Fuller and PP = Phillips-Perron test applied for all world price variables 19 years being the number of observations. For all tests *, **, *** denote rejection of the null hypotheses of non-stationarity at 0.1, 0.05, 0.01 significance levels. All tests were performed in levels and first differences. Maximal lag length is selected by Schwert's rule of thumb, optimal lag length selection according to the Akaike information criterion (AIC). Method used to estimate the long-run variance of each panel's series according to llc (2002). Estimated with STATA's xtunitroot and dfuller/pperron commands.

In 2008 world food prices (International Monetary Fund) rose by 24% compared to year 2007. This means that, ceteris paribus, food CPIs adjusted 25% in Brazil and Colombia, 27% in Peru, 29% in Mexico, 31% in Argentina and 41% in Chile as a reaction to the 2008 price shock. Fig. 2 depicts this nexus, showing that for example Argentina's food CPI strongly reacted in the short-run (13% CPI adjustment), but that the long-run trend was similar to the one of the other countries. In contrast, Colombia's short-run food CPI adjusted only by 8% due to its much lower agricultural market integration. As mentioned, the long-run adjustment rates varied less among countries. A notable effect on long-run price transmission rates only shows at very high degrees of trade openness, like in the case of Chile. Here, the long-run price transmission elasticity is about 11% higher than in Brazil, being the country showing less market integration. The estimated long-run transmission elasticities translated into a long-run food CPI adjustment of 17% in Colombia or Brazil and 20% in Chile after the world food price shock of 2008. The other countries were in between 17 and 20%. Fig. 3a illustrates these long-run adjustment pathways in our six countries. Fig. 3b shows the speed of the long-run adjustment beginning in 2008. The median lag-length was reached within the first three years after the price shock. Hence, the bulk of food price adjustment took place in the short-run.

To get a very clear picture of how transmission rates change according to different levels of trade openness, we constructed three different counterfactual scenarios of trade openness. If the level of agricultural trading activity in LAC had only been around 20%, total price transmission rates would have been reduced to 94%. Thus, as a reaction to the 2008 price spike, domestic food CPIs would only have increased by 23%. In contrast, if a country had had degrees of trade openness of 100% or 200%, the estimated price transmission rates would have been at 125% or 163%, respectively. This would have caused total food CPI adjustments of 30% or 39% (short plus long-run effects).

Apart from world market integration and volatile global food prices, we investigated the role of some macroeconomic factors that influence food prices (see Fig. 1). These variables can be relevant from a policy perspective. Table 3 shows that exchange rate movements have quite strong positive short-run and long-run effects on domestic food prices. This means that imports get more expensive if a currency depreciates, making domestic food more expensive. Simultaneous-

Table 3. Long-run and short-run elasticities of food price determinants

	Panel of six countries		**Panel of five countries (without Brazil)**	**Panel of five countries (without Mexico)**
	World price index IMF	**World price index FAO**	**World price index IMF**	**World price index IMF**
	Praise Winsten	**Praise Winsten**	**Praise Winsten**	**Praise Winsten**
Long run elasticities				
EC	**–0.2323*****	**–0.2282*****	**–0.2316*****	**–0.1914*****
Standard error	(0.028)	(0.028)	(0.025)	(0.032)
World price index	**0.6801*****	**0.6271*****	**0.5368*****	**0.8050*****
Standard error	(0.023)	(0.018)	(0.016)	(0.055)
Trade openness	**–0.0030*****	**–0.0026*****	**–0.0082*****	**–0.0026*****
Standard error	(0.000)	(0.000)	(0.000)	0.0000
Trade openness* World price index	**0.0006*****	**0.0005*****	**0.0018*****	**0.0005*****
Standard error	(0.000)	(0.000)	(0.000)	(0.000)
Exchange rate	**0.5834*****	**0.5921*****	**0.5662*****	**0.6100*****
Standard error	(0.002)	(0.002)	(0.002)	(0.004)
M2 (local currency)	**0.0530*****	**0.0454*****	**0.0247*****	**0.0197*****
Standard error	(0.003)	(0.003)	(0.002)	(0.005)
Short run elasticities				
World price index	**0.1850*****	**0.1577*****	**0.1598*****	**0.2238*****
Standard error	(0.042)	(0.033)	(0.039)	(0.050)
Trade openness	**–0.0010*****	**–0.0009****	**–0.0008****	**–0.0009****
Standard error	(0.000)	(0.000)	(0.000)	(0.000)
Trade openness* World price index	0.0032	0.0018	0.0035	0.0033
Standard error	(0.003)	(0.002)	(0.003)	(0.003)
Exchange rate	**0.3067*****	**0.3171*****	**0.3194*****	**0.2841*****
Standard error	(0.034)	(0.034)	(0.034)	(0.037)
M2 (local currency)	**0.1259*****	**0.1280*****	**0.1372*****	**0.0896****
Standard error	(0.039)	(0.039)	(0.036)	(0.045)
Crude oil prices	–0.0044	–0.0039	–0.0053	–0.0116
Standard error	(0.014)	(0.014)	(0.013)	(0.016)
Post-estimation				
Observations	108	108	90	90
R^2	0.760	0.76	0.85	0.70
Wooldrige test for autocorrelation (*p* value)	(0.113)	(0.183)	(0.452)	(0.211)
Levin-Lin-Chu unit-root test of residuals (*p* value)	(0.000)***	(0.000)***	(0.000)***	(0.000)***
IPS unit-root test of residuals (*p* value)	(0.037)**	(0.020)**	(0.003)***	(0.066)*

Notes: *** $p \leq 0.01$, ** $p \leq 0.05$, * $p \leq 0.1$. All estimations were performed using the natural log of each variable and with country fixed effects, except "Trade openness" which was performed in levels, because it is already a relative measure in %. EC = Error correction term or speed of adjustment. Unit root test: Ha = residuals are stationary confirming cointegration. Tests for autocorrelation: Ha = no serial correlation of the error term.

ly, a depreciation makes export markets more competitive which in turn increases foreign demand for domestic food products, and hence food prices. Related to year 2008, exchange rate movements were rather moderate, so appreciations or depreciations did not have a strong effect during that time. Argentina, Chile and

Table 4. Short and long-run marginal effects of world price shocks under different degrees of trade openness

Degrees of trade openness	Panel of six countries		Panel of five countries (without Brazil)		Panel of five countries (without Mexico)	
	Marginal effect short-run	Marginal effect long-run	Marginal effect short-run	Marginal effect long-run	Marginal effect short-run	Marginal effect long-run
0	**0.185***** (0.041)	**0.680***** (0.033)	**0.160***** (0.039)	**0.537***** (0.027)	**0.224***** (0.050)	**0.805***** (0.048)
20	**0.250***** (0.064)	**0.693***** (0.030)	**0.230***** (0.060)	**0.574***** (0.024)	**0.289***** (0.070)	**0.817***** (0.045)
40	**0.315***** (0.111)	**0.706***** (0.027)	**0.301***** (0.106)	**0.611***** (0.022)	**0.355***** (0.119)	**0.828***** (0.042)
60	**0.380**** (0.164)	**0.719***** (0.025)	**0.371**** (0.156)	**0.648***** (0.021)	**0.420***** (0.174)	**0.840***** (0.040)
80	**0.445**** (0.219)	**0.732***** (0.025)	**0.441**** (0.207)	**0.685***** (0.021)	**0.486**** (0.231)	**0.851***** (0.038)
100	**0.510*** (0.273)	**0.745***** (0.025)	**0.512*** (0.259)	**0.721***** (0.021)	**0.551**** (0.289)	**0.863***** (0.037)
120	**0.575*** (0.328)	**0.758***** (0.027)	**0.583*** (0.311)	**0.758***** (0.023)	**0.617***** (0.345)	**0.874***** (0.037)
140	**0.640*** (0.383)	**0.771***** (0.027)	**0.653*** (0.362)	**0.795***** (0.025)	**0.683*** (0.405)	**0.886***** (0.038)
160	0.704 (0.438)	**0.784***** (0.033)	**0.724*** (0.414)	**0.832***** (0.028)	0.748 (0.463)	**0.897***** (0.040)
180	0.769 (0.493)	**0.797***** (0.036)	**0.794*** (0.466)	**0.869***** (0.032)	0.814 (0.522)	**0.909***** (0.042)
200	0.834 (0.548)	**0.811***** (0.040)	**0.865*** (0.518)	**0.906***** (0.035)	0.879 (0.581)	**0.921***** (0.045)
220	0.899 (0.603)	**0.824***** (0.045)	0.935 (0.571)	**0.943***** (0.039)	0.945 (0.639)	**0.932***** (0.048)

Note: *** $p \leq 0.01$, ** $p \leq 0.05$, * $p \leq 0.1$. Standard errors are given in parentheses below the parameter estimates. The marginal effects measure the reaction of world food price changes on domestic food prices, considering different degrees of trade openness. Trade openness is measured as the sum of export and import values over agricultural GDP. Marginal effects were obtained from estimates using the World price index (IMF).

Mexico showed almost no movement in exchange rates. Only Brazil, Colombia and Peru saw their domestic currencies appreciate against the US dollar by 5 to 6%. These currency appreciations had a food price depressing effect between 1.6 and 2% in the short-run, and around 3.5% in the long-run starting in 2008. Opposed to the exchange rate, real money supply increased between 2007 and 2008 at rates between 10% in Mexico and almost 60% in Peru. According to Table 3, a 100% increase in money supply increases food prices by 13% in the short-run and an additional 5% in the long-run.

Discussion

Globally, hunger still affects 868 million people, of which 49 million are located in LAC (FAO, 2012). However, over the last decade Latin America stands among the regions that achieved larger improvements in fighting hunger. FAO (2012) states that food security is not only about sufficient disposable food supplies in quantitative terms, but primarily about the challenge of food access in economic terms. Since food makes up a large share of poor consumers' consumption basket, price changes are one factor affecting their purchasing power. Thus, food security is partially linked to food price developments, especially in urban areas. Note that Latin America is a more urbanized society than other developing and emerging countries (Poelhekke, 2011). We investigated the dynamics in six LAC, namely Argentina, Brazil, Colombia, Chile, Mexico and Peru. These countries showed large improvements in food-security indicators over the last decade, but still suffer from incidences of food insecurity. At the same time, all six countries are increasingly integrated into global food markets, being both large importers and exporters.

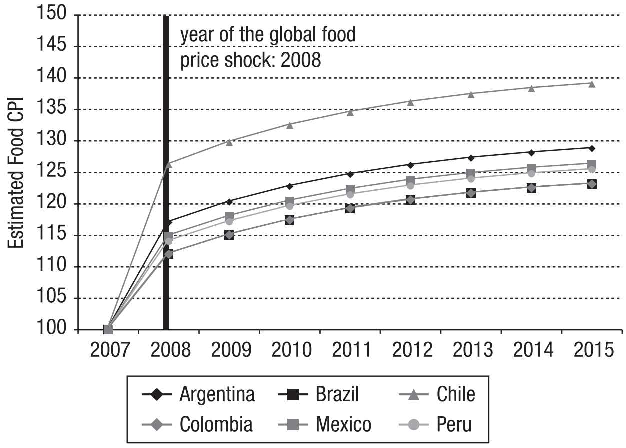

Figure 2. Short-term and long-term estimated food CPI (measured as an index) in each country after a 24% world price shock in year 2008 (ceteris paribus), taking into account different trade openness levels in the different countries. Trade openness levels are as follows: Argentina: 107%, Brazil: 45%, Chile: 220%, Colombia: 45%, Mexico: 80% and Peru: 70%. Short-run price transmission rates vary between 35% and 147%, long-run price transmission elasticities vary between 66% and 89%, depending on the level of trade openness. In 2015 about 90% of total price transmission is reached. Food CPI base year 2007. *Source:* Authors' calculations from regression results.

Especially in light of the global food price crises in 2007/08 and 2011/12, questions of the impacts of agricultural trade and more market integration on food security in developing and emerging countries arose. In order to analyze the relationship between global food price shocks, trade openness and domestic food price movements, we used an error correction framework to estimate how different degrees of agricultural market integration influenced world price transmission rates.

In line with other authors, we find that increasing world prices will transmit into domestic prices. Although some authors claim that in many countries price transmission is not high (Benson *et al.*, 2008; Minot, 2011; Baquedano & Liefert, 2014), we estimated quite elevated long-term price transmission rates between 0.71 and 0.82 in LAC. Our findings show that international trade and market integration has led to different degrees of price transmission rates in the studied countries. Especially in the short-run, world price shocks affect countries' food prices differently, depending on the degree of trade openness. Argentina and Chile are very dependent on food imports and exports which resulted in high transmission rates. In these two countries, the 2008 world price shock of 24% led to an overproportionate instantaneous increase in domestic food prices, and the long-run price equilibrium was an additional 18 to 20% higher than in 2007. On the contrary, a country like Brazil trades very large volumes of agricultural commodities, however, their agricultural sector still produces even larger amounts for their domestic market. Our results show that this lower degree of agricultural market integration has also led to a lower degree of food price transmission, especially in the short-run. These findings are particularly relevant with regard to the expected growing global food demand in the future. It is projected that global food commodity prices will be higher and more volatile than in the past. So, countries that are very dependent on food imports and exports will strongly participate in these global developments. Especially the drastic short-run price transmission rates can be devastating for countries with high degrees of trade openness. Hence,

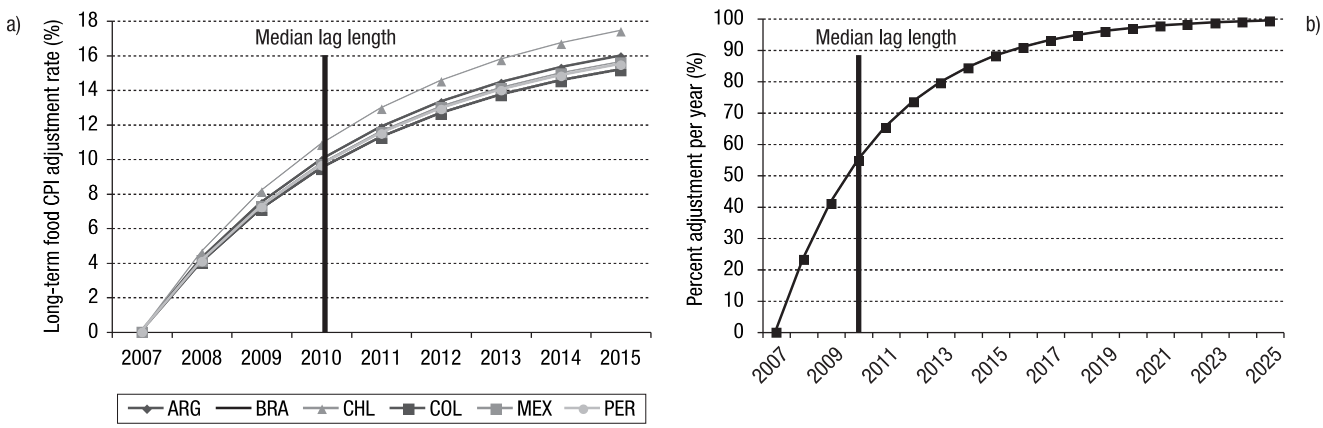

Figure 3. a) Long-term food CPI response in each country due to a 24% world price increase between year 2007 and 2008, taking into account different levels of trade openness in the six countries (in %). b) Yearly percent adjustment rates of total long-term CPI adjustment. Median lag length is reached in period 3 (year 2010). *Source:* Authors' calculations from regression results.

domestic policies should be in place to buffer price shocks. This is not only crucial because higher prices hurt urban food consumers, but also because food price transmission also leads to more domestic food price volatility, leading to higher price uncertainty. It has been argued that stocks holding could be one appropriate policy to be able to react counter-cyclical in times of price shocks (Trostle, 2008; Serra & Gil, 2013). However, this would further drive total global demand in order to build up stocks which in turn further increases world prices (Headey, 2011). According to FAO (2012), the six countries of investigation have implemented a few policies to guarantee food access of the poor as a reaction to increasing world market prices. To mention a few: besides price intervention by establishing maximum food prices for certain commodities, policies were put in place for redistributing food and providing food in elementary schools, and cash tranfer programmes like Bolsa Familia in Brazil or Progresa in Mexico.

Although our results confirm higher world price transmission rates with increasing degrees of trade openness, it does not mean that trade only harms food consumers. First, trading nations can buffer domestic supply shocks by substituting lower food production by imports. Secondly, trading nations turn to the production of those products for which they have comparative advantages, and thus produce at lower costs which should lower consumer prices (Vousden, 1990). Therefore, policies should not necessarily aim at returning to protectionism, but rather focus on establishing effective safety nets to stabilize food prices in times of global shocks.

Our results also illustrate that currency appreciations can in parts buffer world price transmission. For example the Brazilian, Colombian and Peruvian currencies appreciated between 5 and 6% between 2007 and 2008 which made food imports less expensive and exports less competitive and thereby decreased the Peruvian food CPI. Apart from international market forces, our results show that food prices are also affected by domestic macroeconomic factors. A policy relevant variable is money supply. By managing money supply through effective monetary policies, a country can also regulate food price (inflation) to a certain extent. Inflation targeting regimes have been adopted in Brazil, Chile, Mexico and Peru. All countries show relatively strong financial systems, however Brazil and Peru show rather weak fiscal systems (Garcia-Solanes & Torrejón-Flores, 2009). Gonçalves & Salles (2008) confirm that developing countries adopting an inflation targeting regime did not only experience greater drops in inflation, but also in volatility of CPIs. Thus, promoting macroeconomic stability and well-functioning institutions seem to be a crucial factor in stabilizing food prices, safeguarding sufficient food access to the poor.

A possible weakness of our analysis could result from the fact that we modeled six LAC within a panel framework. On the one hand, this allows for drawing more general conclusions on trade and food price transmission than conducting country-specific estimates. On the other hand, we cannot differentiate between effects from trade openness in net importing and net exporting nations. Due to the fact that we have to deal with yearly trade data, country observations would not be sufficiently reliable over a time period of 19 years, though. Another constraint of this study might be caused by the fact that we only look at the total food CPI and do not look at different subgroups of the food CPI. Specific grains, meat and fruits and vegetables CPIs might have been impacted differently by world price shocks and trade openness. Unfortunately, in most of the six countries, the subgroups were not available for the entire period of investigation. Too short time spans make the econometric analysis unreliable. We therefor, sticked to the overall food consumption basket. So, future research should provide more detailed results on different food items and allow for different outcomes depending on a countries net trade position.

As a reaction to the international commodity price spikes, the recent scientific literature has been abundant with analyses on price transmission, impacts of price shocks and policy responses to manage and cope with them (Attanasio *et al.*, 2013; Rodriguez-Takeuchi & Imai, 2013; Baek & Koo, 2014; Baquedano & Liefert, 2014). However, little attention has been paid to the empirical analysis of the interaction between price transmission rates and the level of trade openness. Therefore, the main novelty of this study lies in the examination of the interdependencies between food price transmission rates and varying degrees of agricultural market integration. We also consider different macroeconomic variables that can be relevant for policy advice. We found that increasing levels of trade openness elevate food price transmission rates after price shocks, especially in the short-run. Short-run price transmission elasticities vary between 33 and 89% in the six countries of investigation, depending on the degree of market integration. Long-run transmission elasticities are more independent of trade openness, being at rates between 71 and 81%. Hence, more trade openness brings with it more price instability in the short-term under world price shocks and the resulting persistence in the long-term.

Clearly, immediate effects require different policy approaches to a world price shock than long-term effects. To reduce households' income shocks caused by a sudden and large increase of the price of food, some degree of price management for basic staples may be warranted, coupled with income support or cash transfers. Besides price interventions by establishing maximum food prices for certain commodities temporarily, policies for redistributing food as well as providing food in elementary schools and the poorest households can be appropriate, but must be carefully planned with the required budget readily available.

The study shows that with increasing global market integration, a large proportion of consumer food prices are determined by global forces. But a significantly large proportion too is also due to other macroeconomic factors (exchange rate and money supply). The exchange rate shows an elasticity of 0.31 in the short-run and 0.58 in the long-run. Thus, currency appreciations can buffer shocks from world prices. The elasticity of money supply is 0.13 in the short-run and 0.05 in the long-run. Thus, monetary policies that promote macroeconomic stability seem to be an appropriate means for stabilizing food prices, safeguarding sufficient food access of the urban poor in LAC.

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