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Determinants of Food Price Transmission from International to Domestic Markets

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Abstract

Food price transmission determines the speed of and extent to which international food price shocks affect domestic markets and consumers, an issue of particular importance in low-income countries. Incomplete food price transmission is also implicated in exacerbating international food price volatility because it makes global supply and demand less elastic. This paper explores the factors that influence the degree of price transmission from international maize, rice, and wheat markets to domestic food markets. Using 147,000 observations of monthly food prices, we first estimate food price transmission in 711 markets in 81 countries. Then, the degree of food price transmission across these 711 markets is estimated as a function of country and commodity characteristics.

The results from the first stage indicate that about half (48%) of the domestic food prices were cointegrated with their corresponding international grain price. Among the vector error correction (VEC) models, the median price transmission is 52% after 6 months and 69% after 12 months. Among the vector autoregressive (VAR) models, price transmission was markedly lower, with the median leveling off below 7% after four months. The second-stage results suggest that both market frictions and policy play a role in incomplete transmission of food prices. Food price transmission is significantly lower in landlocked countries, where grain imports or exports are a small share of domestic use, and in Africa and Asia relative to Europe and Latin America. Price transmission is also lower where grain tariffs are high and variable over time, suggesting that trade policy plays a role. These results appear relatively robust, with only minor changes when we estimate the model with different dependent variables, other combinations of independent variables, different thresholds for the stationarity tests, and alternative cointegration tests. The results have implications for the measurement of the welfare impact of international food price shocks and for strategies to reduce the volatility of international grain prices.

Keywords: food prices; price transmission; developing countries; international trade; trade policy; price volatility; error correction model

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1. Introduction

International markets for grains such as maize, rice, and wheat are characterized by periods of relative stability interrupted by sharp increases in prices lasting 6–12 months, often precipitated by conflict, poor harvests, or policy changes. Recent examples of spikes in grain prices include the food price crises of 2007/08 and 2010/11 and the Russian invasion of Ukraine in 2022, each of which caused staple grain prices to more than double over the course of a few months (FAO, 2025a). International shocks induce increases in staple food prices throughout the world, but the adverse impact on poverty and food security is greatest in low-income countries where staple grains are a large share of households budgets and where safety net programs are weak or nonexistent (Ivanic and Martin, 2008).

The severity of impact depends, in large measure, on the speed and degree of price transmission from international grain markets to domestic food markets. Some countries are partially insulated from international food price shocks because they are largely self-sufficient. Others attempt to insulate their food markets from international price shocks with trade policy or public grain stocks, but their success in doing so is mixed (Martin et al., 2026).

Food price transmission varies considerably across countries and commodities, and this heterogeneity has several implications for food security and trade policy. First, studies of the poverty and welfare impact of international grain price shocks depend on estimates of food price transmission. In the absence of relevant estimates of price transmission, many studies use simplifying assumptions, which may under or over-estimate the local impact. Second, although empirical studies frequently show slow or incomplete food price transmission, the causes are not well understood. For example, is it the result of normal market frictions or due to the use of trade policy to insulate local markets from international shocks? Third, although efforts to insulate local markets from global shocks are understandable, the combined effect is to make global grain supply and demand more inelastic, thus exacerbating the international price volatility. If incomplete food price transmission is due to grain policies, international coordination could reduce global food price volatility, but if it is due to market frictions, the remedies are different and perhaps more elusive.

In this paper, we examine price transmission from international grain markets to domestic food markets and explore the determinants of the degree of price transmission. In the first stage of the analysis, we estimate food price transmission using vector autoregressive (VAR) models and vector error correction (VEC) models along with 147,000 monthly price observations for 711 domestic–international price pairs covering 81 countries. In the second stage, we estimate the degree of food price transmission as a function of country and commodity characteristics.

The results from the first stage indicate that about half (48%) of the domestic food prices were cointegrated with their corresponding international grain price. Among the VEC models, the median price transmission is 52% after 6 months and 69% after 12 months. Among the VAR models, price transmission was markedly lower, with the median leveling off below 7% after 4 months.

The second-stage analysis suggests that food price transmission is significantly higher for countries with a high net trade ratio in the grain (net trade as a proportion of domestic use). It is also higher for Europe and Latin America than for other regions. Conversely, it is lower for landlocked countries, processed food products, and retail prices. Finally, food price transmission is significantly reduced by country product-specific tariffs and tariff variability. Thus, we find evidence that partial food price transmission is caused by both the natural friction of markets and policy efforts to insulate domestic markets from international price shocks. The results have implications for the welfare impact of international grain price shocks and for strategies to reduce international price volatility.

Dozens of studies have examined the transmission of international food price shocks to domestic food markets, and many of these have speculated about reasons for differences in price transmission. However, this study is one of a handful that have examined the determinants of grain price transmission econometrically, and to our knowledge, it is the first to incorporate tariff policy as a determinant, which turns out to be highly predictive.

The paper is organized as follows: Section 2 reviews the literature on food price transmission, including the economic theory behind price transmission and empirical studies that quantify the degree of price transmission. In Section 3, we describe the data and methods used in this study. Section 4 provides the results of the analysis, including tests of the statistical properties of food prices, the results of many estimates of food price transmission, and a meta-analysis of the country and product characteristics that "predict" the degree of food price transmission. The conclusions are given in Section 5.

2. Background on incomplete price transmission

2.1. Conceptual framework

The strong form of the "law of one price" says that identical products will have the same price in two markets thanks to spatial arbitrage. However, this is valid only under the unusual circumstance where there are perfectly competitive markets, instant transactions, and no cost of transportation. In real life, of course, these assumptions rarely if ever hold, so price transmission from one market to another is lagged and incomplete (Isard, 1977; Kindelberger, 1989). Price transmission may be partial for at least six reasons, as described below.

Imperfect substitutes. Imported and local grains may not be "identical." Differences between local and imported commodities make them imperfect substitutes and create price differences. For example, yellow maize accounts for the bulk of international maize trade, but white maize is strongly preferred in sub-Saharan Africa (Conforti, 2004).

Imperfect competition. Market power in agricultural markets is another proposed explanation for incomplete price transmission. The idea that cartels of traders could conspire to transmit price decreases more slowly than price increases has an intuitive appeal, although it is challenging to develop a rigorous theoretical model to support this idea (Weldegebriel, 2004).

Lagged response. It often takes several months between the time a trader contracts to import grain from overseas until its delivery in domestic markets, particularly in landlocked countries. Because of this, the process of spatial arbitrage can be slow and price differences may persist over time (Hsu and Goodwin, 1995).

Exchange rates: In the case of price transmission between markets that use different currencies, a shift in the real exchange rate will matter. For example, a sharp increase in export revenue will strengthen the currency, making grain imports from global markets less expensive and depressing the local currency price of grains even without a change in the US dollar price on world markets (Mundlak and Larson, 1992; Baquedano and Liefert, 2014).

Transportation costs. Transportation costs are a major cause of deviation from the law of one price, particularly for food grains, which have low value–bulk ratios. For imported grain crops in sub-Saharan Africa, the cost of sea freight and overland transportation may represent more than half the final price. This creates a large gap between the import parity price, above which imports are profitable, and the export parity price, below which exports become profitable. When domestic prices fall between these two, international trade is not profitable and spatial arbitrage does not operate, disconnecting domestic grain prices from international prices. Autarky and disconnected domestic prices may occur seasonally and/or regionally within the country (Takayama and Judge, 1974; Fackler and Goodwin, 2001).

Trade taxes. Finally, taxes and other policy restrictions on international trade are common. A fixed tariff will increase the cost of importing grain but will not reduce price transmission unless it (at least occasionally) chokes off imports of the commodity. This is more likely if imports represent a small share of domestic use. In addition, many countries try to stabilize domestic food prices by raising (lowering) tariffs in response to lower (higher) world food prices. Similarly, food exporters restrict (liberalize) exports in response to higher (lower) international grain prices. And quantitative barriers, if binding, will break the transmission of prices from one market to another. If government licenses are required to trade or if there are obstacles to purchasing foreign exchange, trader response to changes in international prices may be delayed or blocked entirely by administrative procedures, resulting in imperfect price transmission (Bale and Lutz, 1979; Martin and Anderson, 2012; Martin and Minot, 2022). Frequent changes in trade policy can increase commercial risk in international trade, adding a risk premium associated with trade (Jayne et al., 2006; Martin et al., 2026).

If we assume that transportation costs and taxes are fixed in per ton terms (rather than proportional), the rules of spatial arbitrage can be concisely expressed as a mixed complementarity problem:

$$P_w - t \leq P_d \leq P_w + t \perp T \tag{1}$$

where P_w is the world price, P_d is the domestic price, t is the per unit cost of transportation, and T is net exports.[1] The ⊥ symbol denotes a mixed complementary relationship in which at least one of the following holds:

$$P_w - t = P_d \text{ and } T \geq 0 \tag{2}$$

$$P_d = P_w + t \text{ and } T \leq 0 \tag{3}$$

$$P_w - t < P_d < P_w + t \text{ and } T = 0 \tag{4}$$

Equations 2–4 imply that either a) the domestic price is equal to the export parity price, allowing exports, b) the domestic price is equal to the import parity price, allowing imports, or c) the domestic price lies between import and export parity prices and there is no trade (Rutherford, 1995). The implication for price transmission is that international and domestic prices are no longer related to each other at times or in places where the local market becomes autarkic.

If some portion of transportation costs is fixed, then price transmission elasticities will be greater than or less than one. For importing countries, domestic prices are higher than international prices, so if marketing margins are fixed in value terms, domestic prices may rise by the same dollar amount but a smaller proportion compared to international prices. Thus, a 1% increase in the international price would result in domestic price increase of less than 1% even in the absence of market friction or trade barriers. Likewise, if the domestic price for an exporter is less than the international reference price, the elasticity of price transmission may be greater than one even without market friction or trade barriers (Sharma, 2003).

[1] The above framework can be expanded to incorporate *ad valorem* tariffs and proportional transportation costs, but it makes the relationships less transparent and does not contribute any new insights.

2.2. Empirical research on food price transmission from international to domestic markets

A vast body of research has looked into price transmission across space, time, and product type (see reviews by Fackler and Goodwin (2001) and von Cramon-Taubel and Goodwin (2021)). A search of the online database AgEcon for publications with "price transmission" yielded 492 publications 10 years ago (Kouyate and von Cramon-Taubadel, 2016). A similar analysis carried out recently by the author resulted in over 900 publications (University of Minnesota, 2026).

Research on price transmission has been motivated largely by the idea that co-movement of prices in different markets can be interpreted as a sign of efficient, competitive markets, while lack of co-movement is an indication of market failures. However, as noted above, even temporary autarky caused by transportation costs can break the co-movement of domestic and international prices even if markets are efficient and there are no policy barriers to trade (Harriss, 1979; Barrett, 1996; Baulch, 1997). Other motivations for price transmission research include studying the impact of international food price shocks on consumers and producers in developing countries (Minot, 2011), measuring the role of price insulation in exacerbating international price volatility (Martin and Anderson, 2012), and testing the market power of traders (Weldegebriel, 2004).

Early studies of price transmission used simple correlation coefficients of contemporaneous prices. Later research used regression analysis with lagged world prices, reflecting the fact that international prices may take more than a month to be reflected in domestic prices (Ravallion 1986). The seminal work by Granger and Newbold (1974) showed that many time-series variables (including prices) are nonstationary and that standard regression analysis will yield "significant" coefficients even when no relationship exists. Engle and Granger (1987) developed methods of analyzing nonstationary data, namely cointegration analysis and the error correction model. Johansen (1988) proposed a widely used test of cointegration.

Three variants of the standard methods for analyzing price transmission have emerged. Asymmetric price transmission allows the speed and magnitude of price transmission to vary depending on whether the initial shock is positive or negative (Enders and Granger, 1998; von Cramon Taubadel, 1998). Second, threshold integration allows an impact only when the impulse reaches a certain size, such as a minimum gap between two prices (Balke and Fomby, 1997). Finally, GARCH models have been used to examine the transmission of both mean prices and price volatility from one market to another (Engle and Kroner, 1995; Ceballos et al., 2017).

Research on price transmission from international to domestic grain markets shows mixed results (see Table 1). Mundlak and Larson (1992) found price transmission elasticities averaging 0.95, but these results are probably due to the fact that they regressed annual domestic price levels on international price levels, ignoring the effect of non-stationarity. Quiroz and Soto (1995) repeated the analysis with similar data using an error correction model and found no relationship between domestic and international prices for 30 of the 78 countries examined. Even in countries with a relationship, the convergence was very slow.

Later studies generally used monthly data and tested cointegration at the country-commodity level rather than the country level. In general, these studies tend to find cointegration in one-third to two-thirds of the commodities examined (see Sharma, 2003; Conforti, 2004; Ghoshray, 2011; Greb et al., 2016; Iqbal and Babcock, 2016). On the other hand, Baquedano and Liefert (2014) found 84% of domestic prices were cointegrated with world markets, while Minot (2011) and Ceballos et al. (2017) found less than a quarter were. Some of these differences undoubtedly reflect the choice of commodities and countries.

Several studies identified patterns in the price transmission. In an analysis of 67 cereal prices in sub-Saharan Africa, Minot (2011) noted that wheat and rice prices were more integrated with world markets than maize prices, which was attributed to the fact that most African countries are heavily dependent on rice and wheat imports but close to self-sufficient in maize. Baquedano and Liefert (2014) found a similar pattern across commodities. Two studies of maize markets in Africa found that linkages to prices in neighboring countries were stronger than links to international maize markets (Hatzenbuehler et al., 2017; Pierre and Kaminski, 2019).

Table 1. Selected studies on food price transmission from international to domestic markets

Source	Data and methods	Results
Mundlak and Larson, 1992	Static regression analysis of annual export prices and domestic prices in 58 countries.	High rates of price transmission with average elasticity of 0.95.
Quiroz and Soto, 1995	ECM analysis of similar data as Mundlak and Larson for 78 countries.	No cointegration in 30 countries. Very slow transmission in almost all others.
Sharma, 2003	ECM analysis for domestic rice prices in 16 Asian countries	No cointegration in 8 of 16 countries. Average long-run elasticity was 0.65.
Baffes and Gardner, 2003	Estimates price transmission for 8 countries and 10 commodities for 31 prices.	Cointegrated prices for 3 countries and for selected commodities in the other 5 countries.
Conforti, 2004	ECM analysis of monthly prices from 7 countries and annual prices for 9 countries.	Among the monthly cereal prices, 5 of 10 were integrated. No cointegration for rice in Indonesia or wheat in Egypt due to policy.
Ghoshray, 2011	Threshold ECM analysis of 18 commodity prices from 6 Asian countries.	Three of the 5 cereal prices were cointegrated with world prices.
Minot, 2011	ECM analysis of 67 were cereals prices in 12 African countries.	Cointegration in 13 of the 67 cereal prices. More often with wheat and rice than maize.
Baquedano and Liefert, 2014	ECM analysis of 61 cereal prices in 28 countries.	Long-run relationships found in 51 of 61 pairs. Wheat and rice were more cointegrated than maize and sorghum.
Greb et al., 2016	Analysis of 678 price transmission estimates from literature and ECM analysis of 499 FAO prices. Analysis of factors in transmission.	79% of prices in the literature were cointegrated, but just 55% in data.

Iqbal and Babcock, 2016	Asymmetric ECM analysis of 317 prices of wheat, rice, maize, and soybeans.	40% of price pairs were cointegrated. Mean elasticity was 0.36 for maize and 0.52 for soy.
Kouyate and von Cramon Taubadel, 2016	Meta-analysis of ECM estimates of 708 cereal prices from previous studies.	Cointegration was found in 60% of all studies and 81% of all published studies.
Bekkers et al., 2017	AR model in first differences of link between world food price index and food price indices in 147 countries.	One-third have significant long-run pass through from international to domestic food price indices.
Ceballos et al., 2017	GARCH analysis of price and price volatility transmission to 41 cereal prices in 27 countries.	10 of the 41 prices had significant price transmission. Transmission stronger for rice and wheat than maize and sorghum.
Hatzenbuehler et al., 2017	ECM analysis of price transmission for maize, rice, and other staples in 6 Nigerian markets.	All maize and rice prices are linked to international prices. Maize prices linked more strongly to prices in neighboring countries.
Arnade et al, 2017	ECM analysis of price transmission to 3 cereals and 6 other food items in China.	Chicken and soybeans are the most linked to world markets, and rice is the least.
Pierre and Kaminski, 2019	Global VAR model of maize prices in 27 African countries, considering impact of regional and global prices.	Half of countries were cointegrated with other markets. Regional linkages were stronger than links to global maize markets.
Guo and Tanaka, 2019	GARCH analysis of volatility transmission to wheat price in 10 countries.	Significant transmission of wheat prices in 9 of the 10 countries.
Guo and Tanaka, 2020	GARCH analysis of volatility transmission to wheat prices in 10 countries.	Significant volatility transmission in all 10 countries.
Emediegwu and Rogna, 2024	Panel VAR analysis of quasi-synthetic price data on four cereals in 23 countries.	Significant and relatively quick transmission for rice, wheat, and maize but not sorghum.

A few studies have estimated the determinants of food price transmission econometrically. Goletti et al. (1995) showed that rice price integration among pairs of the 65 districts of Bangladesh was significantly related to distance, road density, and rail density. Kouyate and von Cramon-Taubadel (2016) carried out a meta-analysis of over 700 cointegration analyses of cereal price transmission. They found that cointegration was more common and the speed of adjustment was faster when the two markets were in the same country and when the distance between them was shorter. Guo and Tanaka (2019 and 2020) studied the determinants of transmission of price volatility, finding that transmission is weaker if the country is more self-sufficient in the grain and if it consumes competing grains.

Greb et al. (2016) analyzed 678 estimates of international-to-domestic price transmission drawn from the literature. Cointegration with world markets was more likely in wheat markets than maize or rice markets, less likely for West African markets compared to Asian markets, and less likely for retail compared to wholesale prices. Surprisingly, a high net import ratio reduced the likelihood of cointegration, while a state trading enterprise increased it. A separate analysis in the same report examined international–domestic price transmission using data for 499 cereal price pairs. Maize and wheat prices were less likely to be cointegrated than rice prices, and East and West African markets were *more* likely to be cointegrated than Asian markets. As expected, retail prices were less cointegrated than wholesale prices, and cointegration was less common in countries with a state grain-trading enterprise. Being landlocked, trade openness, and a measure of ease of trade were not significant factors.

Finally, Bekkers et al. (2017) examined the determinants of the long-run coefficient of price transmission from the FAO global food price index to food price indices in 147 countries. They find price transmission of food indexes was negatively associated with per capita GDP and average agricultural tariff rates and positively associated with the import ratio, at least in some specifications.

3. Data and methods

This study examines the factors influencing the transmission of international food prices to domestic markets. In the first stage, we estimate price transmission parameters using monthly data for a large number of pairs of prices of wheat, rice, and maize products. For each pair, the international price is a standard reference price for one of the three commodities in a major exporting country, while the domestic price is a wholesale or retail price of grain, flour, or wheat bread in one market in a country. In the second stage, we estimate the degree of price transmission as a function of country- and country-commodity-specific factors.

3.1. Data

This section describes the price data used in the cointegration analysis (the first stage) and the variables used to estimate the determinants of food price transmission (the second stage).

International food prices

The international food prices are monthly reference prices of the three grains, maize, rice, and wheat, from major exporting countries of each commodity. Specifically, we used the wholesale price of No. 2 yellow maize in the United States (FOB Gulf of Mexico), the wholesale price of Super A1 white rice in Thailand (FOB Bangkok), and the wholesale price of No. 2 hard red winter wheat in the United States (FOB Gulf of Mexico). All three prices were obtained from the Food Monitoring and Price Analysis (FPMA) Tool of the Food and Agriculture Organization of the United Nations (FAO) (FAO, 2025a). These prices are expressed in nominal US dollars per ton.

Domestic food prices

The domestic food prices were also obtained from the FMPA Tool (FAO, 2025a). In order to work in a common currency, we used price series that were converted to US dollars by the FAO. The analysis starts with all 1,956 domestic prices of maize, maize flour, rice, wheat, wheat flour, and wheat bread in the FPMA database, representing 128 countries. We then impose three requirements: the price series must have at least 120 monthly observations, it should have no more than 5% missing observations, and the ratio of highest to lowest price should not be more than 20. About one-third of the prices (659) were removed for not having enough observations, another 208 were dropped for having too many missing observations, and 33 were excluded because the range was implausible, leaving 1,056 domestic prices. Later, 336 price were dropped because they did not pass stationarity tests and 9 were removed because they were from countries for which we do not have explanatory variables for the second stage. In the end, the analysis of the determinants of food price transmission used 147,000 monthly price observations in 711 domestic price series from 81 countries.

Explanatory variables

We hypothesize that various factors influence the degree of price transmission from international to domestic food markets (see Table 2). In general, we expect any factor that increases the cost of trading grain to or from the country will reduce price transmission. This is because higher costs create a wider band between the import parity price and the export parity price, which implies that it is more likely at any given time that the country is autarkic. While trade is interrupted, there is no spatial arbitrage to link domestic and international prices of the commodity. Thus, we expect landlocked countries, countries more distant from the main exporter, and countries with high tariffs to have lower transmission of prices from international to domestic markets. Similarly, a country with a small trade ratio in a commodity (such as one importing just 5% of its maize requirements) is more likely to be occasionally autarkic, thus reducing the average price transmission over time. We use separate trade ratios for importers and exporters to allow for asymmetric effects. And if there are economies of scale in international shipments, a large trade volume in the commodity will reduce costs and improve price transmission.

Table 2. Explanatory variables used in second stage

Variable	Description	Source	Expected impact on price transmission
Ln GPD pc	Natural logarithm of per capita gross domestic product in US dollars in 2015 using 2021 purchasing price parity	World Bank Development Indicators (World Bank, 2025)	Positive
Landlocked	Binary variable indicating that the country is landlocked	UN classification (UN, 2025)	Negative
Ln distance	Natural logarithm of the sea distance from the exporting country to the country in question in km	CERDI Sea Distance Database. (Bertoli et al., 2016)	Negative
Avg tariff (%)	Average tariff rate for the commodity in the country over 2015–2020	WTO Tariff and Trade Data (WTO, 2025)	Negative
SD tariff	Standard deviation of the tariff rate for the commodity in the country over 2015–2020	WTO Tariff and Trade Data (WTO, 2025)	Negative

Ln trade	Natural logarithm of the sum of the quantity of imports and exports of the commodity in the country in 1000 tons over 2011–2015	FAO Food Balance Sheet (FAO, 2025b)	Positive
Import ratio (%)	For importers, the ratio of net imports to domestic utilization of the commodity in the country	FAO Food Balance Sheet (FAO, 2025b)	Positive
Export ratio (%)	For exporters, the ratio of net exports to domestic utilization of the commodity in the country	FAO Food Balance Sheet (FAO, 2025b)	Positive
Flour	Binary variable indicating that the domestic price is of maize or wheat flour	Description of price data. (FAO, 2025a)	Negative
Bread	Binary variable indicating that the domestic price is of wheat bread	Description of price data. (FAO, 2025a)	Negative
Retail	Binary variable indicating that the domestic price is a retail price rather than a wholesale price	Description of price data. (FAO, 2025a)	Negative
Calorie pct	The average percentage share of calories in the national diet coming from the commodity over 2011–2015	FAO Food Balance Sheet (FAO, 2025b)	Negative
Maize	Binary variable for maize prices (wheat is the omitted category)	Description of price data. (FAO, 2025a)	Unknown
Rice	Binary variable for rice prices (wheat is the omitted category)	Description of price data. (FAO, 2025a)	Unknown
LAC	Binary variable for countries in Latin America and the Caribbean (Africa is the omitted category)	UN classification. (UN, 2025)	Positive
Europe	Binary variable for countries in Europe (Africa is the omitted category)	UN classification. (UN, 2025)	Positive
Asia	Binary variable for countries in Asia and the Pacific (Africa is the omitted category)	UN classification. (UN, 2025)	Positive

Given that high-income countries tend to have better transportation infrastructure, we expect them to have stronger price transmission. For similar reasons, we expect more developed regions to have stronger price transmission than other regions. Since the regional binary variables and per capita GDP operate through similar causal channels, collinearity between the two may reduce the measured effects of each.

We expect the prices of flour and bread to show lower price transmission from international grain prices because the process of milling flour and baking bread introduce other costs which may vary over time, thus weakening the linkage between international grain prices and the food price. For the same reason, we expect retail prices to be more weakly connected to international markets than wholesale prices.

If a staple grain accounts for a large share of the caloric intake of the country, the prices of the commodity are likely to be politically sensitive, motivating the government to insulate domestic prices from international price shocks. We hypothesize that price transmission will be weaker (stronger) for country–commodity pairs where the grain represents a large (small) share of the caloric intake.

If the government attempts to insulate domestic food markets from international price volatility, this will be reflected in a higher standard deviation of the commodity-specific tariff. Thus, we expect the standard deviation of the tariff rate to be negative related to price transmission.

3.2. Methods

In the first stage of the analysis, we check the statistical properties of the international and domestic prices to determine the appropriate type of econometric model and estimate the relationship between the international and domestic prices econometrically. In the second stage of the analysis, we estimate the degree of price transmission from the first stage as a function of commodity and country characteristics. These steps are described below.

Testing the price series

After excluding domestic price series that were too short, had too many missing values, or too wide a range, we carry out three tests on the price data. The first test is for unit roots (non-stationarity) in the prices levels and the first differences in prices. We use the generalized least squares version of the Dickey-Fuller test of stationarity. Elliott, Rothenberg, and Stock (1996) show that this test is more powerful than the augmented Dickey-Fuller test. Since cointegration analysis is based on the assumption that the levels are nonstationary but the first differences are stationary, these tests determine whether or not the price series will be included in the analysis.

The second test is to determine the optimal lag length for the analysis. Typically, this is done by selecting the lag length that generates the lowest information criterion. Lutkepohl (2005) shows that the Hannan-Quinn information criterion (HQIC) and the Schwarz-Bayesian information criterion (SBIC) provide consistent estimates of the true lag order.

The third test is to determine the number of cointegrating equations in the system. In our case, there are just two variables in each model (the international and domestic prices), so there can be either 0 or 1 cointegrating equations, implying no cointegration or cointegration, respectively. We use the Johansen trace test to determine the number of cointegrating equations and to identify the econometric model to use.

Econometric analysis of the impact of international prices on domestic prices

For each pair of domestic and international prices, we apply one of two econometric models. If the Johansen trace test indicates that the domestic and international prices are cointegrated, we use a VEC model. The VEC model for the analysis of prices can be expressed as follows:

$$\boldsymbol{\Delta p_t} = \boldsymbol{\alpha} + \boldsymbol{\Pi p_{t-1}} + \sum_{i=1}^{k} \boldsymbol{\Gamma_i \Delta p_{t-i}} + \boldsymbol{\varepsilon_t} \tag{5}$$

where

$\mathbf{p}_t$ is a vector of the log of prices at time t, and $\mathbf{\Delta p}_t$ is a vector of the first difference of the log of prices at time t;

$\boldsymbol{\alpha}$ is a vector of estimated parameters that describe the trend component;

$\mathbf{\Pi}$ is a matrix of estimated parameters that describe the long-term relationship and the speed of adjustment toward the long-term relationship;

$\mathbf{\Gamma}_i$ is a matrix of estimated parameters that describe the short-run relationship between prices for time t-i; and

$\boldsymbol{\varepsilon}_t$ is a vector of error terms.

In this study, we use a two-variable version of the VEC model, which generates estimates of the effect of international prices on domestic prices and vice versa. Since most markets can be considered "small countries" in the staple grain markets, we focus on the parameters describing the effects of world prices on domestic prices. Applying the Johansen normalization to the parameters in $\mathbf{\Pi}$ in equation (5), the determinants of domestic prices can be expressed as follows:

$$\Delta p_t^d = \alpha + \theta\left(p_{t-1}^d + \beta_0 + \beta_1 p_{t-1}^w\right) + \sum_{i=1}^{k} \delta_i \, \Delta p_{t-i}^w + \sum_{j=1}^{k} \varphi_j \, \Delta p_{t-j}^d + \varepsilon_t \tag{6}$$

where

Δp_t^d is the first difference of the log of domestic prices at time t;

Δp_t^w is the first difference of the log of international prices at time t;

α, θ, β_0, β_1, δ_i, and φ_j are estimated parameters derived from the parameters in equation (1); and

ε_t is the error term.

The coefficients in the error correction model can be interpreted as follows:

α is the constant, reflecting any long-term trend in p^d not explained by p^w;

θ reflects the speed of adjustment toward the long-run equilibrium;

β_0 is the constant in the long-run equilibrium;

β_1 represents the negative of the long-run elasticity of the domestic price with respect to the international price;

δ_i is the short-run impact of changes in the international price on changes in domestic prices with a lag of i months; and

φ_j is the autoregressive coefficient, reflecting serial correlation in the changes in the domestic price with a lag of j months.

Alternatively, if the Johansen trace test suggests that there is no cointegration between the two variables, we use a VAR model to represent the short-run relationship between first differences in domestic and international prices. We work in first differences because testing confirms that prices are generally nonstationary in levels but stationary in first differences. The generalized VAR of first differences in prices can be expressed as follows:

$$\Delta \boldsymbol{p_t} = \boldsymbol{\alpha} + \sum_{i=1}^{k} \boldsymbol{\Gamma_i} \Delta \boldsymbol{p_{t-i}} + \boldsymbol{\varepsilon_t} \tag{7}$$

where the variables and parameters are defined as in equation (5). Because our interest is in the effect of international prices on domestic prices, we focus on the portion of this equation that has domestic prices as the dependent variable:

$$\Delta p_t^d = \alpha + \sum_{i=1}^{k} \delta_i \, \Delta p_{t-i}^w + \sum_{j=1}^{k} \varphi_j \, \Delta p_{t-j}^d + \varepsilon_t \tag{8}$$

where variables and parameters are defined as in equation (6). In this model, changes in domestic prices are a linear function of k lagged values of changes in domestic prices and k lagged values of changes in international prices.

Both the VEC model and the VAR model generate a large number of parameters, all of which affect the relationship between international prices and domestic prices. In order to collapse this information into an index of the degree of price transmission, we use the impulse response function (IRF), which summarizes the effect of a one unit "impulse" of the independent variable (the international price) on the dependent variable (the domestic price) after a given number of periods (months). Specifically, the IRF values after 6 months is used as the main measure of price transmission, though we test the sensitivity using 3 and 12 months as well.

Assessing the determinants of price transmission

In the second stage of our analysis, the degree of price transmission for the price pairs is estimated as a function of variables described in Section 3.1. The degree of price transmission is measured by the IRF value at six months. This analysis includes diagnostic tests to check if the assumptions behind ordinary least squares regression are met, as well as test of the sensitivity of the results to the three approaches to addressing heteroskedasticity, three alternative dependent variables, different sets of independent variables, different levels of significance for the stationarity test, and different types of cointegration test.

4. Results

The results are divided into three sections: the testing of the price data, the estimation of food price transmission using VEC and VAR models, and the analysis of the determinants of price transmission.

4.1. Tests of price data

Testing the stationarity of the price data

We apply the GLS Dickey-Fuller test of non-stationarity to the three international grain prices and 1,056 domestic food prices, each in log levels and in log first differences. We cannot reject a unit root (non-stationarity) in the levels of all three international prices at either 1% or 5% level of significance. However, we reject non-stationarity in the first differences of all three prices at either level of significance (see Table A1).

With regard to the 1,056 domestic food prices, non-stationarity in price levels is rejected for about one-quarter (27%) of the prices at the 5% confidence level. With regard to the first difference of domestic prices, non-stationarity is rejected for the vast majority of price series (94%) at the 5% level. In summary, the conditions for using cointegration analysis are met in all three of the international grain prices and about two-thirds of the domestic grain prices (see Table A1). The main results are based on a 5% confidence level, but we test the sensitivity of the results to a) adopting a 1% level of confidence and b) accepting all price series regardless of the outcome of the stationarity tests.

Testing the optimal lag length

The lag length for the econometric models (k) is set by selecting the lag length that minimizes an information criterion.[2] Two widely used measures are employed: the Hannan-Quinn information criterion (HQIC) and the Schwarz-Bayesian information criterion (SBIC). The optimal lag length is two periods (months) for about two-thirds of the 1,056 domestic prices according to both tests (see Table A2). Where they disagree, the HQIC always results in a longer optimal lag length than the SBIC. We adopt the HQIC to select the lag length for modeling each price series, though we also examine the sensitivity of the results to adopting the SBIC instead.

[2] The test determines the optimal lag length for the underlying VAR model in levels. For the VEC model (equation 1) and the VAR model in first differences (equation 2), the lag length (k) will be the optimal length minus one.

Testing for cointegration of domestic and international prices

The cointegration test determines whether there is a stable long-term equilibrium relationship between each domestic price and the corresponding international price. Three cointegration tests are applied to the 1,056 pairs of domestic and international prices. According to the Johansen trace statistic, 48% of the domestic and international price pairs are cointegrated (see Table A3). The SBIC test results in a much smaller share of cointegrated relationships (27%), while the HQIC yields a higher share (61%). In all three tests, there are five price pairs (0.7%) for which the cointegration test failed, suggesting that the price levels are stationary. We adopt the results of the Johansen trace statistics for our main results but explore the sensitivity of the results to using the other two tests.

4.2. Econometric analysis of the relationship between domestic and international prices

Of the 1,056 price pairs examined in the previous section, we reject 336 pairs because either the domestic price level is stationary or because the first difference in domestic prices is nonstationary. This leaves 720 domestic prices from 84 countries that meet our requirements for cointegration analysis. Of these, 342 pairs are cointegrated according to the Johansen trace statistic, so we apply the VEC model from equation (1). For the other 378 pairs, cointegration is rejected, so we use the VAR model in first differences, shown in equation (3). For each price pair, the lag length (k) is determined by the HQIC test.

Vector error correction models

In the first iteration of the two models, we included constants: α and β_0 in equation (2) and α in equation (4). However, the estimated values of these coefficients were close to zero and almost never statistically significant. In the VEC model, the first constant (α) was statistically significant at the 5% level in less than 1% of the models, while the constant in the cointegrating equation (β_0) was never statistically significant. Similarly, the constant in the VAR model (α) was significant at the 5% level in just less than 1% of the models. In light of these results, we reran the VEC and VAR models without these constants, setting α=0 and β_0=0. Since a non-zero α would imply a trend in p^d even if p^w remained unchanged, restricting α to zero is consistent with intuition and spatial arbitrage. A nonzero β_0 in the VEC model is neither suggested nor prohibited by theory, but the data indicate that β_0 is not statistically significant.

Table 3 shows the summary statistics of the results of the second iteration of the 342 VEC models. All estimated values of the speed of adjustment (θ) are negative, as expected, since a positive deviation from the long-term relationship should lead to a negative adjustment in the domestic price. A large majority (86%) of the θ coefficients are statistically significant at the 5% level. Among those that are statistically significant, the average value is −0.064, meaning that domestic prices move about 6% of the way toward the equilibrium relationship with the international price in each period (month).

The second coefficient in Table 3 is β_1, the coefficient measuring the long-term relationship between the domestic price and the international prices. All but one of the 342 VEC models have a β_1 that is statistically significant at the 5% level. All the estimated values are negative, as expected, indicating a positive long-term relationship between domestic and international prices.[3] Since the prices are in logarithms, $-\beta_1$ is the long-term elasticity of the domestic price with respect to international prices. The estimated elasticities range from 0.9 to 1.9, with a mean value of 1.1.

The next set of coefficients (δ) are estimates of the short-term impact of changes in international prices on changes in domestic prices. The number of estimated coefficients declines with longer lags because the optimal lag length varies across models (see Table A2). Less than half of the δ coefficients are statistically significant, and those that are significant have mean values that are positive but less than 0.2.

The last set of coefficients (φ) are the autoregressive coefficients, describing serial correlation in changes in domestic prices. Most of the estimated values of φ are statistically significant, particularly for one lag. It is difficult to interpret the individual values of δ and φ because there are multiple lags and they interact with each other. It is the combined effect that matters, as discussed later.

Table 3. Descriptive statistics of coefficients in VEC model

Coefficient	Among all coefficients		% significant at 5% level	Among significant coefficients			
	N	Mean		N	Mean	Minimum	Maximum
θ	342	-0.057	85.7	293	-0.064	-0.280	-0.010
β_1	342	-1.137	99.7	341	-1.140	-1.922	-0.948
δ_1	326	0.075	37.4	122	0.157	-0.258	0.363
δ_2	47	0.010	27.7	13	0.017	-0.367	0.263
δ_3	13	0.117	76.9	10	0.180	-0.119	0.346
φ_1	47	-0.054	56.4	184	0.142	-0.458	0.568
φ_2	13	-0.002	53.2	25	-0.097	-0.293	0.262
φ_3	0	0.000	15.4	2	-0.018	-0.169	0.133

Source: Analysis of data from FAO (2025a).

Vector autoregressive model

Table 4 summarizes the results of 378 VAR models, where the Johansen test rejected a long-term relationship between domestic and international prices. The δ coefficients, representing the short-run effect of changes in international prices on domestic prices, are generally positive (as expected), though often not statistically significant.

The φ coefficients indicate the auto-regressive component of changes in domestic prices. More than half of these are statistically significant, and the average value among significant coefficients was negative for

[33] As shown in equation (2), the long-term relationship is expressed as ($p^d_{t-1} + \beta_0 + \beta_1 p^w_{t-1}$). Since this expression is the residual from regressing p^d on p^w, β_1 will be negative if the two variables are positively related.

two of the three lags. Again, it is difficult to interpret these coefficients individually since it is the combined effect that matters, as described in the next section.

Table 4. Descriptive statistics of coefficients in VAR model

Coefficient	Among all coefficients		%	Among significant coefficients			
	N	Mean	significant at 5% level	N	Mean	Minimum	Maximum
δ_1	378	0.056	13.0	49	0.187	-0.131	0.388
δ_2	81	0.094	35.8	29	0.203	0.077	0.580
δ_3	13	0.077	46.2	6	0.105	-0.234	0.425
φ_1	378	0.046	51.9	196	0.082	-0.606	0.654
φ_2	81	-0.143	66.7	54	-0.201	-0.372	0.268
φ_3	13	-0.076	76.9	10	-0.122	-0.386	0.278

Source: Analysis of data from FAO (2025a).

Summarizing the degree of price transmission

The IRF describes the effect of a one-unit change in the independent variable on the dependent variable, combining the effects of short- and long-term impact after a given number of periods. In our case, the IRF gives the effect of a unit increase in the international grain price (p^w) on the domestic price (p^d) after a specified number of months, expressed as a percentage of the initial impulse. Figure 1 summarizes the distribution of IRF values as a function of the number of months after the shock for the domestic prices that are cointegrated with the corresponding international price. Each line represents a percentile of the IRF values across the 342 VEC model results. The median IRF rises from 13% after 1 month to 69% after 12 months.

Figure 2 summarizes the IRF values for the 378 VAR model results. The impact of international prices on domestic prices is more muted because the VAR models do not include a long-run relationship pulling the domestic prices to an equilibrium level. Since the optimal lag is usually one, two, or three months, the impact of international prices on domestic prices ceases after a few months. The median IRF is less than 7% after 12 months.

Figure 3 shows the IRF values of all 720 models, combining VEC and VAR models. The results are a weighted average of the previous two graphs, with the median IRF rising from 14% after 1 month to 33% after 12 months.

Figure 1. Summary of impulse response functions in VEC models (N=342)

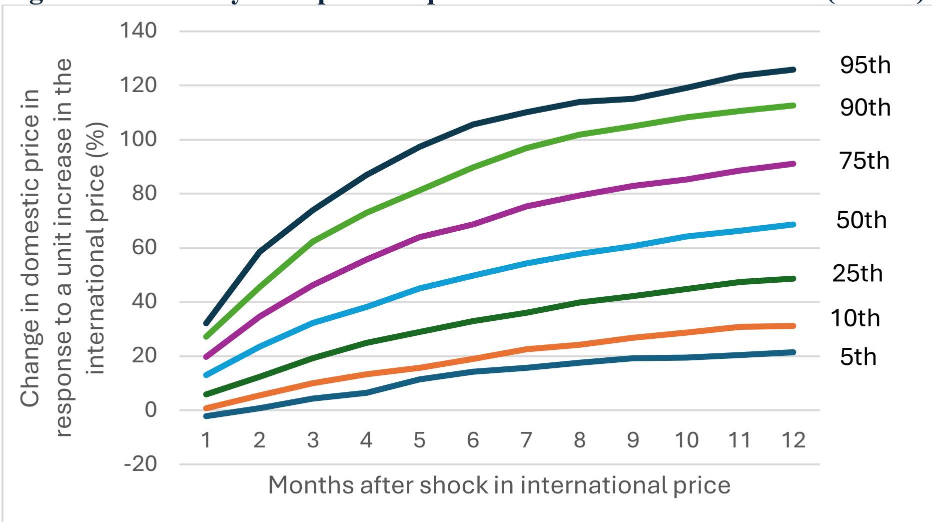


Figure 2. Summary of impulse response functions in VAR models (N=378)

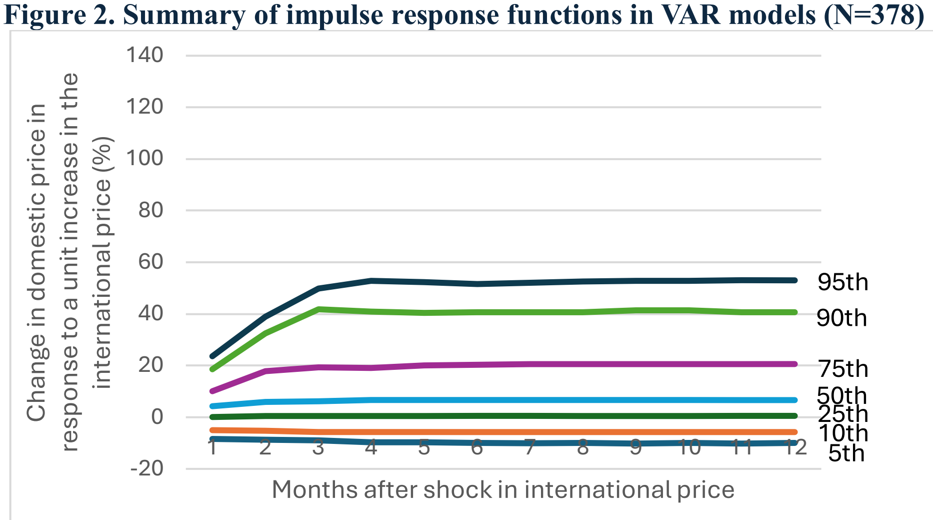
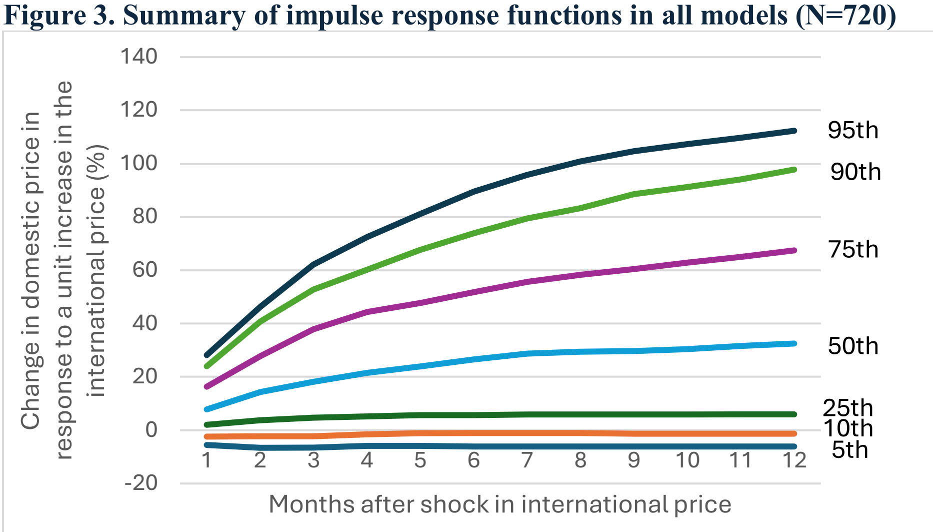


Figure 3. Summary of impulse response functions in all models (N=720)



4.3. Determinants of food price transmission

This section describes the results of regression analysis to estimate the degree of food price transmission as a function of country and country–commodity specific explanatory variables, listed in Table 1. Although we have estimates of the degree of price transmission for 720 domestic food prices from 84 countries, the trade and tariff explanatory variables are not available for Djibouti, South Sudan, and Somalia. As a result, the analysis of the determinants of price transmission focuses on 711 price pairs from 81 countries.

Descriptive statistics

The descriptive statistics for the three dependent variables and the 19 explanatory variables are shown in Table 5. The first three rows give the statistics for the three dependent variables, corresponding to the impulse response function (IRF) after 3, 6, and 12 months. The remaining rows show the summary statistics for the independent variables.

Table 5. Descriptive statistics of variables

Variable	N	Mean	Std. dev.	Minimum	Maximum
IRF 3 (pct)	711	22.22	22.30	-37.29	101.91
IRF 6 (pct)	711	31.11	30.57	-47.05	131.65
IRF 12 (pct)	711	39.76	38.78	-37.52	172.35
Ln GDP pc (USD)	711	9.00	0.81	7.27	11.02
Landlocked	711	0.25	0.43	0.00	1.00
Ln distance (km)	711	9.30	0.89	0.00	10.03
Avg tariff (pct)	711	26.09	36.53	0.00	513.00
SD tariff (pct)	711	5.33	14.31	0.00	54.77
Ln trade (1000 t)	711	6.61	1.93	0.47	10.16
Import ratio (%)	711	38.22	39.46	0.00	133.33
Export ratio (%)	711	9.75	50.39	0.00	1156.55
Flour	711	0.27	0.44	0.00	1.00
Bread	711	0.11	0.31	0.00	1.00
Retail	711	0.62	0.49	0.00	1.00
Caloric pct (%)	711	21.46	15.19	0.12	62.13
Maize	711	0.12	0.33	0.00	1.00
Rice	711	0.43	0.50	0.00	1.00
Wheat	711	0.44	0.50	0.00	1.00
LAC	711	0.22	0.41	0.00	1.00
Africa	711	0.29	0.45	0.00	1.00
Europe	711	0.05	0.21	0.00	1.00
Asia	711	0.45	0.50	0.00	1.00

Source: Author's analysis. See Table 2 for data sources.

Main results

The first column of Table 6 gives the results of an ordinary least squares (OLS) regression of IRF6 as a function of the full set of 17 independent variables (maize is omitted from the commodity variables and Africa from the region variables to avoid perfect collinearity). The degree of food price transmission after six months is significantly associated (at the 5% level of confidence) with 10 of the independent variables. The model explains about one-quarter of the variance in the dependent variable.

Before interpreting the results, we consider three diagnostic tests of the OLS results. The first is the Ramsey regression specification-error test (RESET) for omitted variables using powers of the fitted values of the dependent variable. The result is that $F(24, 669) = 7.25$, which rejects the null hypothesis of no omitted variables at the 1% confidence level, suggesting that there are omitted variables. Later, we examine the results after adding quadratic terms to the explanatory variables (see Table 8), but even with quadratic terms we are not able to exclude the possibility of missing variables.

The second is the calculation of variance inflation factors (VIFs) to check for multicollinearity. Although close correlation among independent variables does not violate the assumptions behind OLS, it would increase the standard error of the correlated variables, thus affecting the interpretation of the result. None of the VIF values for the independent variable is greater than 5.0, suggesting there is no issue of multicollinearity in the independent variables.

Third, we apply the Breusch–Pagan/Cook–Weisberg test for heteroskedasticity, which tests for differences in the variance of the residuals as a function of the independent variables. The test statistic is $\chi2(17)=50.27$, which rejects the null hypothesis of homoskedasticity at the 1% confidence level. With heteroskedasticity, estimates of the coefficients remain unbiased, but they are not efficient. More seriously, the standard errors and tests of significance are not reliable.

One approach to addressing heteroskedasticity is to calculate Huber-White sandwich standard errors, also known as robust standard errors, as shown in the second column of Table 9. The coefficients are the same as the OLS coefficients in the first column, but the standard errors are valid even with some types of misspecification. None of the coefficients lose or gain statistical significance compared to the OLS results in the first column.

Table 6. Determinants of food price transmission: Main results

	OLS		Robust		GLS	
Ln GDP pc	-6.537	**	-6.537	**	-6.646	**
	(1.615)		(1.699)		(1.313)	
Landlocked	-5.891		-5.891		-6.326	*
	(3.018)		(3.110)		(2.850)	
Ln distance	1.064		1.064		3.988	**
	(1.267)		(1.589)		(1.545)	
Avg tariff (%)	-0.169	**	-0.169	*	-0.026	**
	(0.035)		(0.067)		(0.008)	
SD tariff	-0.391	**	-0.391	**	-0.331	**
	(0.085)		(0.066)		(0.054)	
Ln trade qty	2.019	**	2.019	*	-0.302	
	(0.766)		(0.825)		(0.730)	
Import ratio (%)	0.108	**	0.108	*	0.208	**
	(0.039)		(0.045)		(0.036)	
Export ratio (%)	0.023		0.023		0.354	**
	(0.022)		(0.053)		(0.057)	
Flour	-2.633		-2.633		-4.357	*
	(3.408)		(3.063)		(2.032)	
Bread	-11.665	*	-11.665	**	-14.234	**
	(4.635)		(3.993)		(3.517)	
Retail	-7.202	**	-7.202	*	-5.502	**
	(2.646)		(2.844)		(1.929)	
Caloric pct	0.301	**	0.301	**	0.623	**
	(0.096)		(0.108)		(0.096)	
Grain						
Rice	3.552		3.552		-2.654	
	(3.581)		(3.542)		(3.475)	
Wheat	-2.933		-2.933		-6.821	
	(4.490)		(4.208)		(4.228)	
Region						
LAC	15.350	**	15.350	**	20.106	**
	(3.316)		(3.374)		(3.125)	
Europe	41.991	**	41.991	**	42.161	**
	(6.989)		(8.422)		(6.885)	
Asia	6.000		6.000		7.112	
	(3.866)		(4.168)		(3.734)	
Intercept	62.103	**	62.103	**	35.452	
	(20.277)		(23.332)		(20.295)	
N	711		711		711	
R-squared	0.26		0.26			
Adjusted R-squared	0.25		0.25			
F statistic	14.57		28.43			

** p<.01, * p<.05

A more advanced method of dealing with heteroskedasticity is to use feasible generalized least squares (GLS) in which the variance of the error terms is estimated as a function of exogenous variables, and observations are weighted by the inverse of the estimated variance. In the first iteration, the full set of 17 independent variables is used to estimate the variance, and in the final version, the six independent variables that were statistically significant are retained. The third column of Table 6 shows the results using GLS to adjust for heteroskedasticity (the variance model is found in the second column of Table A4 in the Annex). Thirteen of the coefficients are statistically significant at the 5% level. Of these, 10 have the expected sign.

The coefficient on the landlocked binary variable is significant and negative, as expected. It suggests that price transmission after six months is 6.6 percentage points lower in landlocked countries than in countries with a coastline, other factors held constant.

The coefficient on average tariff is negative and significant, indicating that higher tariffs are associated with lower price transmission between international and domestic markets. This is not surprising given that tariffs raise the cost of importing and create a wider band of international prices over which international trade is not profitable, thus dampening spatial arbitrage that would otherwise link international and domestic prices. The magnitude is small: a 10 percentage point increase in the average tariff is associated with a 0.3 percentage point decrease in price transmission after six months.

The variability of tariffs over time is also significantly and negatively associated with food price transmission. As described in Martin et al. (2026), many countries change tariff levels over time, raising (lowering) tariffs when international prices are low (high), in an effort to insulate domestic food prices from international market volatility. These results confirm that countries with more variable tariff rates have lower food price transmission.

The coefficients on the import ratio and export ratio are significant and positive, as expected. A country that relies more heavily on imports or exports is less likely to be occasionally autarkic, and theory suggests that continuous trade means that domestic prices are linked to international prices without interruption. A 10 percentage point increase in the import (export) ratio is associated with a 2.1 (3.5) percentage point increase in price transmission after six months.

Both flour and bread have negative and significant coefficients, as expected. Six months after a shock in international prices, price transmission is 4 percentage points lower for flour than grain, and it is 14 percentage points lower for bread than grain. Similarly, domestic retail food prices show price transmission that is 5.5 percentage points lower than domestic wholesale prices. These patterns are expected because the margin between international and domestic prices includes other costs (labor, rent, transport, and processing) that dilute the connection between international and domestic prices.

Finally, domestic food prices in Europe and Latin America and the Caribbean (LAC) have higher levels of price transmission than those in Africa, the omitted category. This may reflect better transportation infrastructure and/or more competitive markets. The magnitude of these effects is large; for example, other things being equal, food price transmission in Europe is 42 percentage points higher than in Africa, even after controlling for other factors.

Three coefficients are significant but have the “wrong” sign. First, higher per capita GDP appears to weaken price transmission, other things being equal. This is contrary to expectations because higher-income countries generally have better port, communication, and road infrastructure. However, the inclusion of the regional variables complicates the interpretation because average per capita GDP varies across regions. Later, we show that excluding the three regional variables causes the per capita GDP coefficient to lose statistical significance (see Table 8). Second, long distances between the major exporter and the country appear to strengthen price transmission, contrary to expectations. This relationship holds even when the regional variables are removed. However, the magnitude is modest: a 10% increase in distance is associated with price transmission rising by 0.4 percentage points. Finally, the caloric share coefficient is positive, implying that if a country relies heavily on the commodity in its diet, price transmission is stronger. This undermines our hypothesis that the price of these staple grains would be politically sensitive, leading to government efforts to insulate domestic prices from international shocks.

It is worth noting that the total quantity of trade in the grain (imports plus exports) is not significantly related to price transmission, unlike the previous two models in this table. This suggests that economies of scale in grain trade may be less important than expected. Furthermore, there is no statistically significant difference in six-month price transmission between rice, wheat, and maize after controlling for other factors.

Alternative dependent variables

In Table 7, we use alternative dependent variables, namely the degree of food price transmission after different periods of time: 3 months, 6 months, and 12 months. For all three, we use the GLS model, estimating the variance of the error term using the same six independent variables as above. The second column in this table repeats the results from the GLS column of Table 9 (the variance models for all three are in Table A4 in the Annex).

The first two columns of Table 7 are broadly comparable. One notable difference is that the average tariff does not significantly affect food price transmission after three months, but it does after six months. Similarly, price transmission in Asia is significantly higher than in Africa in the IRF 3 model but not in the IRF 6 model.

Comparing the second and third column, the results are again quite similar. The only differences in significance are that the landlocked variable becomes statistically insignificant in the 12-month price transmission and the wheat variable becomes significant, suggesting weaker price transmission for wheat than maize after 12 months. Overall, the results in this table indicate that the statistically significant determinants of food price transmission do not vary much across 3-, 6-, and 12-month time frames.

Table 7. Determinants of food price transmission: Alternative dependent variables

	IRF_3		IRF_6		IRF_12	
Ln GDP pc	-3.604	**	-6.646	**	-8.770	**
	(1.037)		(1.313)		(1.605)	
Landlocked	-6.303	**	-6.326	*	-4.774	
	(2.151)		(2.850)		(3.637)	
Ln distance	2.066	*	3.988	**	4.047	*
	(0.975)		(1.545)		(2.043)	
Avg tariff (%)	-0.063		-0.026	**	-0.028	**
	(0.045)		(0.008)		(0.008)	
SD tariff	-0.305	**	-0.331	**	-0.348	**
	(0.045)		(0.054)		(0.064)	
Ln trade qty	-0.651		-0.302		-0.185	
	(0.589)		(0.730)		(0.899)	
Import ratio (%)	0.089	**	0.208	**	0.298	**
	(0.030)		(0.036)		(0.045)	
Export ratio (%)	0.251	**	0.354	**	0.409	**
	(0.046)		(0.057)		(0.070)	
Flour	-3.452	*	-4.357	*	-4.945	*
	(1.757)		(2.032)		(2.383)	
Bread	-10.487	**	-14.234	**	-16.802	**
	(2.786)		(3.517)		(4.363)	
Retail	-6.035	**	-5.502	**	-5.244	*
	(1.721)		(1.929)		(2.198)	
Caloric pct	0.284	**	0.623	**	0.986	**
	(0.077)		(0.096)		(0.121)	
Grain						
Rice	-1.207		-2.654		-4.503	
	(2.639)		(3.475)		(4.331)	
Wheat	-1.125		-6.821		-10.932	*
	(3.153)		(4.228)		(5.281)	
Region						
LAC	10.712	**	20.106	**	29.288	**
	(2.326)		(3.125)		(3.951)	
Europe	29.418	**	42.161	**	50.178	**
	(5.319)		(6.885)		(8.331)	
Asia	6.243	*	7.112		8.951	
	(2.750)		(3.734)		(4.750)	
Intercept	33.325	*	35.452		48.771	
	(13.955)		(20.295)		(26.394)	
N	711		711		711	

** p<.01, * p<.05

Alternative independent variables

Some of the independent variables "share" a channel of influence on the level of price transmission. For example, the main explanation for regional differences in price transmission is that the level of per capita GDP (and associated infrastructure development) varies across regions. Thus, income and region may "share" some influence. In this section, we examine the sensitivity of the results to using different combinations of independent variables as well as quadratic terms.

The first column of Table 8 shows the base results from Table 6 (GLS column) and Table 7 (IRF 6 column). In the second column, the commodity and regional variables are excluded from the analysis. The third model shows a version with quadratic terms included (the variance models for all three are in Table A5 in the Annex).

In the second column, the main effect of excluding regional and commodity variables is that per capita GDP is no longer statistically significant. Although it is not positive, as expected, it is no longer negative and significant as it is in the base version. This confirms that the unexpected negative association between per capita GDP and price transmission in the base model is an artifact of including regional dummy variables.

The third column of Table 8 shows the base model with additional quadratic terms for three variables: distance, trade quantity, and caloric share (these are the only three that were significant when running a version with quadratic terms for all eight continuous variables). The average marginal effect of the logarithm of distance on price transmission is 4.4, meaning that a 10% increase in distance is associated with a 0.4 percentage point increase in price transmission after six months. Across the values of distance, the marginal effect remains positive, ranging from 2 to 5 (see Table A6).

In the case of the volume of grain trade, the quadratic relationship follows an inverted U-shape, with the maximum price transmission at 478,000 tons of trade (exports plus imports). The positive relationship at lower volumes can be explained by economies of scale in sea freight. It is understandable that price transmission would plateau at some point, but it is less clear why price transmission would decline with larger volumes.

The quadratic relationship effect of caloric share on price transmission has a U-shape, with the minimum price transmission when the grain accounts for just 7% of total calories consumed. For most of the range of caloric share, the price transmission is rising (see Table A6). This indicates that the importance of the staple grain in the diet does not consistently result in political sensitivity and policies to insulate domestic prices from international shocks.

Table 8. Determinants of food price transmission: Alternative independent variables

	Base model		Without region & product variables		With quadratic terms	
Ln GDP pc	-6.646	**	-1.837		-6.348	**
	(1.313)		(1.237)		(1.355)	
Landlocked	-6.326	*	-8.181	**	-7.650	**
	(2.850)		(2.842)		(2.933)	
Ln distance	3.988	**	4.830	**	-9.392	*
	(1.545)		(1.584)		(3.977)	
Avg tariff (%)	-0.026	**	-0.056	**	-0.026	**
	(0.008)		(0.008)		(0.007)	
SD tariff	-0.331	**	-0.471	**	-0.295	**
	(0.054)		(0.046)		(0.054)	
Ln trade qty	-0.302		-0.790		14.042	**
	(0.730)		(0.671)		(2.974)	
Import ratio (%)	0.208	**	0.107	**	0.159	**
	(0.036)		(0.033)		(0.038)	
Export ratio (%)	0.354	**	0.362	**	0.368	**
	(0.057)		(0.059)		(0.057)	
Flour	-4.357	*	-4.986	**	-3.956	*
	(2.032)		(1.785)		(2.010)	
Bread	-14.234	**	-19.546	**	-15.353	**
	(3.517)		(3.223)		(3.595)	
Retail	-5.502	**	-7.521	**	-4.485	*
	(1.929)		(2.068)		(1.815)	
Caloric pct	0.623	**	0.575	**	-0.216	
	(0.096)		(0.091)		(0.274)	
Grain						
Rice	-2.654				-1.829	
	(3.475)				(3.403)	
Wheat	-6.821				-5.956	
	(4.228)				(4.219)	
Region						
LAC	20.106	**			17.102	**
	(3.125)				(3.193)	
Europe	42.161	**			45.182	**
	(6.885)				(7.010)	
Asia	7.112				3.704	
	(3.734)				(4.012)	
(Ln distance)^2					0.741	**
					(0.272)	
(Ln trade)^2					-1.138	**
					(0.234)	
(Cal pct)^2					0.014	**
					(0.005)	
Intercept	35.452		1.371		63.031	**
	(20.295)		(21.064)		(20.854)	
Number of observations	711		711		711	

** p<.01, * p<.05

Alternative levels of confidence for tests of stationarity

In the main results shown in Table 6, we limit the analysis to pairs of prices where GLS Dickey-Fuller test failed to rejected non-stationarity for prices in levels and rejected non-stationarity for the first difference in prices at the conventional 5% confidence level. We test the sensitivity of these results to two alternatives: a) we do not impose any stationarity tests on the price series and b) we test stationarity with the GLS Dickey-Fuller test at the 1% confidence level. The absence of a test for stationarity gives us a larger sample (1,011 price pairs) and results in three changes (see Tables A7 and A8 in the Annex). The distance variable is not statistically significant without the stationarity test, which is more plausible than our base model which has a significant but unexpected positive sign on the distance variable. In addition, rice and wheat have significant coefficients suggesting that price transmission for these commodities is 7 and 10 percentage points less than price transmission for maize.

Considering the case where we use a 1% level of confidence for stationarity, the sample is somewhat larger (833 price pairs), and the only difference in significance is that the average tariff is no longer statistically significant. None of the other variables change in significance nor (among those that are significant) in sign.

Alternative tests for cointegration

In the main results, we use the Johansen trace statistic to determine whether each pair of international and domestic prices is cointegrated. If the prices are cointegrated, the VEC model is used, but if they are not, the VAR model is used. Would the results change if we tested cointegration using the Hannan-Quinn information criterion (HQIC) or the Schwarz-Bayesian information criterion (SBIC)? We find that the results are relatively insensitive to the cointegration test (see Tables A9 and A10 in the Annex). If we use the HQIC test of cointegration, the landlocked and flour variables become statistically insignificant, while the wheat variable turns significant and negative. Alternatively, using the SBIC test of cointegration there is just one change: the average tariff loses its statistical significance. The signs of the significant coefficients are the same across all three models, and even the magnitudes are fairly similar.

5. Conclusions

This paper explores the factors that influence the degree of food price transmission from international maize, rice, and wheat markets to domestic food markets. The analysis makes use of all available FAO price data on maize, rice, and wheat products that meet our criteria, covering 147,000 domestic price observations, 711 domestic price series, and 81 countries. The first-stage analysis finds that about half (48%) of the domestic food prices were cointegrated with their corresponding international grain price. Among the VEC models, the average long-run elasticity of the domestic price with respect to the international price is 1.1 and the median price transmission is 52% after 6 months and 69% after 12 months. Among the VAR models, price transmission was markedly lower, with the median leveling off below 7% after four months.

In the second stage, we find that food price transmission after six months is, as expected, statistically greater in countries with a coastline, where the grain tariff is lower, where tariffs are less variable, where grain imports or exports are a large share of domestic use, for grains compared to flour or bread, for wholesale prices compared to retail prices, and in Europe and Latin America compared to Africa. Unexpectedly, price transmission is also greater among low-income countries, though this finding disappears if we remove regional variables. Distance and caloric share in the diet also increase food price transmission, against expectations, though the effects are small.

These results appear relatively robust, with only minor changes when we estimate the model with different dependent variables (3- and 12-month price transmission), other combinations of independent variables, different thresholds for the stationarity tests, and alternative cointegration tests.

An important issue is whether incomplete food price transmission is caused by natural friction of markets such as occasional autarky due to high transportation costs or caused by protective policies, particularly efforts to insulate domestic markets from international price shocks. This study finds evidence for both. Regarding market friction, being landlocked and trading only a small share of domestic utilization are consistently associated with lower food price transmission. This effect has been discussed in various studies of price transmission (Barrett and Li, 2002; Conforti 2004; Minot, 2011) and found to be statistically significant in a few (Guo and Tanaka, 2019 and 2020). This supports the view that occasional autarky in a commodity breaks the link between domestic and international prices, thus reducing price transmission. On the policy side, both the average level of grain-specific tariffs and the variability in those tariffs appear to dampen food price transmission. Although the use of tariffs to insulate domestic markets from international shocks has been discussed (Bale and Lutz, 1979; Martin and Anderson, 2012; and Martin and Minot, 2022), to our knowledge, this is the first study to demonstrate econometrically that both average tariffs and the variability in tariff rates have a significant effect on price transmission in food grain markets.

The finding that retail prices, flour prices, and bread prices show weaker linkages to international grain prices is not surprising. Previous research has found that price transmission from international markets is greater for wholesale than retail prices (Greb et al., 2016), though the finding of differences between grains and processed grain products appears to be new. Like transport costs, processing costs increase the margin and allow other cost factors (such as fuel costs and labor costs) to introduce “noise” into the link between international and domestic prices.

The strong differences in food price transmission across regions presumably reflects differences in road, port, and communication infrastructure, though, as we have seen, per capita GDP does not seem to capture this effect well. Future research may be able to better identify the primary factors behind these regional differences.

These results have two implications for research on food security and trade policy. Studies of the local poverty and welfare impact of international food price shocks often rely on assumptions about the degree of transmission to domestic food markets. For country–commodity combinations that are part of this study, the estimated IRF values from the first stage can be used to improve assessments of the welfare impact. For country–commodity combinations not included in this study, the second-stage regression model can be used to infer the rates of food price transmission over 3, 6, and 12 months.

The results also have implications for strategies to reduce international food price volatility. The study confirms that high and variable grain tariff rates reduce food price transmission and thus exacerbate international food price volatility. Thus, there is some scope for multilateral trade agreements to reduce international food price volatility. At the same time, various market frictions associated with temporary autarky also play an important role in limiting food price transmission. For this reason, food price transmission would be slow and incomplete even in a free-trade regime, and efforts to reduce these frictions (where possible) will also contribute to reducing international food price volatility.

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Annex

Table A1. GLS Dickey-Fuller tests of non-stationarity

International prices	N	Tau	Critical value 5%	Critical value 1%	Whether H_o (non-stationarity) is rejected 5%	1%
		Descriptive statistics for test statistic				
Level (ln P^w)						
US No. 2 yellow maize	1	-2.36	-2.90	-3.48	No	No
Thai Super A1 white rice	1	-2.72	-2.90	-3.48	No	No
US hard red winter wheat	1	-2.03	-2.90	-3.48	No	No
First difference (Δln P^w)						
US No. 2 yellow maize	1	-10.71	-2.90	-3.48	Yes	Yes
Thai Super A1 white rice	1	-9.45	-2.90	-3.48	Yes	Yes
US hard red winter wheat	1	-12.14	-2.90	-3.48	Yes	Yes
Domestic prices	**N**	**Mean**	**Min**	**Max**	**Share of cases where H_0 rejected (%) 5%**	**1%**
Level (ln P^d)						
Maize & maize flour	223	-3.127	-6.344	-1.066	58.7	31.8
Rice	412	-2.204	-6.841	-0.252	18.9	8.5
Wheat, flour, bread	421	-2.344	-9.770	-0.334	19	6.2
Total	1056	-2.455	-9.770	-0.252	27.4	12.5
First difference (Δln P^d)						
Wheat, flour, bread	223	-8.46	-14.760	-1.610	94.6	91.5
Rice	412	-8.81	-17.490	-1.890	94.7	93.4
Maize & maize flour	421	-8.39	-15.490	-1.680	94.3	92.4
Total	1056	-8.57	-17.490	-1.610	94.5	92.6

Source: Analysis of data from FAO (2025a).

Note: The null hypothesis (H_o) is that the variable is non-stationary.

Table A2. Optimal lag length according to Hannan-Quinn and Schwarz-Bayesian information criteria

	Hannan-Quinn information criterion		Schwarz-Bayesian information criterion	
Optimal number of lags	Number of prices	Pct of prices	Number of prices	Pct of prices
1	156	14.8	328	31.1
2	737	69.8	702	66.5
3	124	11.7	22	2.1
4	39	3.7	4	0.4
Total	1,056	100.0	1,056	100.0

Source: Analysis of data from FAO (2025a)

Table A3. Tests of cointegration between international and domestic prices

	Johansen trace statistic		Schwarz-Bayesian information criterion		Hannan-Quinn information criterion	
Cointegrating vectors	Number of prices	Pct of Prices	Number of prices	Pct of Prices	Number of prices	Pct of Prices
0	546	51.7	764	72.4	405	38.4
1	504	47.7	286	27.1	645	61.1
.	6	0.6	6	0.6	6	0.6
Total	1,056	100	1,056	100.0	1,056	100.0

Source: Analysis of data from FAO (2025a).

Table A4. Models of variance for Table 7

	Variance in IRF3 model		Variance in IRF 6 model		Variance in IRF 12 model	
Ln GDP pc	-0.072		-0.188	**	-0.287	**
	(0.069)		(0.067)		(0.067)	
Ln distance	-0.026		-0.144	*	-0.168	**
	(0.069)		(0.061)		(0.059)	
Avg tariff	-0.000		-0.011	**	-0.015	**
	(0.003)		(0.002)		(0.002)	
SD tariff	-0.034	**	-0.034	**	-0.035	**
	(0.004)		(0.004)		(0.004)	
Export ratio (%)	0.005	**	0.005	**	0.005	**
	(0.001)		(0.001)		(0.001)	
Retail	-0.454	**	-0.699	**	-0.703	**
	(0.141)		(0.130)		(0.129)	
Intercept	7.094	**	10.191	**	11.869	**
	(0.929)		(0.874)		(0.865)	
Number of observations	711		711		711	

** p<.01, * p<.05

Source: Analysis of data from FAO (2025a).

Table A5. Models of variance for Table 8

	Base model		Without region & product variables		With quadratic terms	
Ln GDP pc	-0.188	**	-0.094		-0.166	*
	(0.067)		(0.069)		(0.067)	
Ln distance	-0.144	*	-0.147	*	-0.020	
	(0.061)		(0.065)		(0.072)	
Avg tariff	-0.011	**	-0.009	**	-0.014	**
	(0.002)		(0.003)		(0.002)	
SD tariff	-0.034	**	-0.036	**	-0.033	**
	(0.004)		(0.004)		(0.004)	
Export ratio (%)	0.005	**	0.005	**	0.005	**
	(0.001)		(0.001)		(0.001)	
Retail	-0.699	**	-0.699	**	-0.545	**
	(0.130)		(0.138)		(0.134)	
Intercept	10.191	**	9.443	**	8.776	**
	(0.874)		(0.922)		(0.936)	
Number of observations	711		711		711	

** p<.01, * p<.05

Source: Analysis of data from FAO (2025a).

Table A6. Marginal effects for quadratic variables in Table 11

Independent variable	Percentile	Value of x	Marginal effect (dy/dln(x))
Ln(distance)	5th	2,150	1.97
	10th	3,262	2.59
	25th	8,275	3.97
	50th	13,778	4.73
	75th	19,496	5.24
	90th	20,071	5.28
	95th	20,290	5.30
Ln(trade)	5th	23	6.87
	10th	51	5.09
	25th	166	2.41
	50th	1,041	-1.77
	75th	3,854	-4.75
	90th	7,455	-6.25
	95th	14,651	-7.79
Caloric pct	5th	2.5	-0.15
	10th	4.9	-0.08
	25th	9.5	0.05
	50th	18.0	0.28
	75th	29.7	0.60
	90th	44.7	1.02
	95th	48.9	1.13

Source: Analysis of data from FAO (2025a).

Table A7. Determinants of food price transmission: Alternative levels of confidence for the stationarity tests

	No DF test		DF at 5% level		DF at 1% level	
Ln GDP pc	-7.396	**	-6.646	**	-7.412	**
	(1.274)		(1.313)		(1.265)	
Landlocked	-6.456	*	-6.326	*	-6.813	*
	(2.649)		(2.850)		(2.707)	
Ln distance	2.255		3.988	**	3.489	*
	(1.495)		(1.545)		(1.520)	
Avg tariff	-0.022	*	-0.026	**	-0.018	
	(0.009)		(0.008)		(0.009)	
SD tariff	-0.423	**	-0.331	**	-0.428	**
	(0.055)		(0.054)		(0.054)	
Ln trade qty	-0.240		-0.302		-0.638	
	(0.693)		(0.730)		(0.690)	
Import ratio (%)	0.114	**	0.208	**	0.152	**
	(0.035)		(0.036)		(0.035)	
Export ratio (%)	0.288	**	0.354	**	0.274	**
	(0.052)		(0.057)		(0.052)	
Flour	-4.463	*	-4.357	*	-5.241	**
	(1.746)		(2.032)		(1.844)	
Bread	-15.263	**	-14.234	**	-13.200	**
	(3.529)		(3.517)		(3.440)	
Retail	-6.743	**	-5.502	**	-5.423	**
	(1.776)		(1.929)		(1.887)	
Caloric pct	0.509	**	0.623	**	0.432	**
	(0.090)		(0.096)		(0.091)	
Grain						
Rice	-6.569	*	-2.654		-5.150	
	(2.917)		(3.475)		(3.185)	
Wheat	-9.950	**	-6.821		-7.095	
	(3.819)		(4.228)		(3.953)	
Region						
LAC	9.231	**	20.106	**	15.599	**
	(2.797)		(3.125)		(2.932)	
Europe	31.839	**	42.161	**	46.295	**
	(6.168)		(6.885)		(6.161)	
Asia	-0.860		7.112		6.805	
	(3.583)		(3.734)		(3.502)	
Intercept	76.946	**	35.452		59.627	**
	(19.161)		(20.295)		(19.550)	
Number of observations	1011		711		833	

** p<.01, * p<.05

Source: Analysis of data from FAO (2025a).

Table A8. Model of variance for Table A7

	Model of variance for no DF test		Model of variance for DF at 5% level		Model of variance for DF at 1% level	
Ln GDP pc	-0.129	*	-0.188	**	-0.175	**
	(0.057)		(0.067)		(0.063)	
Ln distance	-0.131	*	-0.144	*	-0.160	**
	(0.059)		(0.061)		(0.059)	
Avg tariff	-0.007	**	-0.011	**	-0.007	**
	(0.001)		(0.002)		(0.001)	
SD tariff	-0.043	**	-0.034	**	-0.039	**
	(0.003)		(0.004)		(0.004)	
Export ratio (%)	0.003	**	0.005	**	0.004	**
	(0.001)		(0.001)		(0.001)	
Retail	-0.507	**	-0.699	**	-0.661	**
	(0.104)		(0.130)		(0.116)	
Intercept	9.643	**	10.191	**	10.228	**
	(0.789)		(0.874)		(0.819)	
Number of observations	1011		711		833	

** p<.01, * p<.05

Source: Analysis of data from FAO (2025a).

Table A9. Determinants of food price transmission: Alternative tests of cointegration

	Johansen		HQIC		SBIC	
Ln GDP pc	-6.646	**	-6.681	**	-5.943	**
	(1.313)		(1.321)		(1.273)	
Landlocked	-6.326	*	-4.763		-5.957	*
	(2.850)		(2.841)		(2.745)	
Ln distance	3.988	**	4.635	**	4.393	**
	(1.545)		(1.601)		(1.569)	
Avg tariff	-0.026	**	-0.032	**	-0.012	
	(0.008)		(0.009)		(0.007)	
SD tariff	-0.331	**	-0.409	**	-0.285	**
	(0.054)		(0.055)		(0.051)	
Ln trade qty	-0.302		-0.157		-0.189	
	(0.730)		(0.735)		(0.709)	
Import ratio (%)	0.208	**	0.155	**	0.191	**
	(0.036)		(0.036)		(0.035)	
Export ratio (%)	0.354	**	0.299	**	0.298	**
	(0.057)		(0.056)		(0.054)	
Flour	-4.357	*	-3.809		-3.806	*
	(2.032)		(2.102)		(1.852)	
Bread	-14.234	**	-15.646	**	-15.445	**
	(3.517)		(3.522)		(3.363)	
Retail	-5.502	**	-4.201	*	-4.070	*
	(1.929)		(1.940)		(1.783)	
Caloric pct	0.623	**	0.485	**	0.506	**
	(0.096)		(0.097)		(0.093)	
Grain						
Rice	-2.654		-3.584		-2.542	
	(3.475)		(3.486)		(3.368)	
Wheat	-6.821		-8.758	*	-5.028	
	(4.228)		(4.242)		(4.102)	
Region						
LAC	20.106	**	17.058	**	16.826	**
	(3.125)		(3.118)		(3.008)	
Europe	42.161	**	48.567	**	38.801	**
	(6.885)		(6.739)		(6.744)	
Asia	7.112		4.978		3.822	
	(3.734)		(3.713)		(3.612)	
Intercept	35.452		39.087		22.969	
	(20.295)		(20.925)		(20.271)	
Number of observations	711		711		711	

** p<.01, * p<.05

Source: Analysis of data from FAO (2025a).

Table A10. Model of variance for Table A9

	Model of variance for Johansen test		Model of variance for HQIC test		Model of variance for SBIC test	
Ln GDP pc	-0.188	**	-0.226	**	-0.134	
	(0.067)		(0.069)		(0.069)	
Ln distance	-0.144	*	-0.190	**	-0.160	**
	(0.061)		(0.064)		(0.056)	
Avg tariff	-0.011	**	-0.010	**	-0.011	**
	(0.002)		(0.002)		(0.002)	
SD tariff	-0.034	**	-0.033	**	-0.037	**
	(0.004)		(0.004)		(0.004)	
Export ratio (%)	0.005	**	0.005	**	0.004	**
	(0.001)		(0.001)		(0.001)	
Retail	-0.699	**	-0.687	**	-0.660	**
	(0.130)		(0.130)		(0.133)	
Intercept	10.191	**	10.930	**	9.786	**
	(0.874)		(0.932)		(0.861)	
Number of observations	711		711		711	

** p<.01, * p<.05

Source: Analysis of data from FAO (2025a).

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