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Food Trade Policy and Food Price Volatility

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Notices

ABSTRACT

Food trade barriers in many countries are systematically adjusted to insulate domestic markets from world price changes—a response not predicted by traditional political economy models. In this study, policymakers are assumed to minimize the political costs associated with changing domestic prices and deviating from longer-run political-economy equilibria. Error correction techniques applied to domestic and world price data for rice and wheat collected to measure trade policy distortions allow estimation of policy response parameters. The results suggest that systematic short-run price insulation reduces shocks to domestic prices but sharply increases world price volatility and the costs of trade distortions. However, idiosyncratic domestic price shocks resulting from inefficient policy instruments such as quantitative restrictions increase domestic price volatility relative to the magnified volatility of world prices—frequently outweighing the stabilizing impacts of price insulation. This fundamentally changes our understanding of the impacts of price-insulation—from a zero-sum game where some countries reduce the volatility of their prices using beggar-thy-neighbor policies that raise price volatility elsewhere, into one where price volatility rises in most countries. National policy reforms to move away from discretionary, destabilizing policies could lower costs, reduce volatility in domestic and world prices, and facilitate reform of international trade rules.

Keywords: trade policy; food prices; behavioral economics; econometric models; food price volatility.

JEL Codes: F13; F14; Q17; Q18; C51.

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1. Introduction

Food price volatility is an important concern for policymakers, especially in developing countries (Timmer 2000). High food prices put at risk the food security of vulnerable net buyers of food, while low prices may impoverish farmers reliant on food sales for their incomes. Surges in world food prices have occurred with alarming frequency in recent years, including in 2008, 2011, and 2022. Many countries seek to stabilize domestic food prices using trade policies such as export restrictions or changes in import tariffs that can have major impacts on the volatility of world prices (Giordani, Rocha, and Ruta 2016; Martin and Anderson 2011). Key questions addressed in this paper are the extent to which these interventions affect the volatility of international and domestic prices, and what reforms of current policies might lead to better outcomes.

An important role for international trade in food is to help diversify food supplies. Such diversification can greatly reduce the volatility of supply for staple foods—relative to relying only on local food production—and reduce the vulnerability of countries' populations to food supply shocks (Burgess and Donaldson 2010). Once countries open their markets to food trade, they may be able to further reduce domestic price volatility by partially insulating their markets against changes in world prices. When world prices of food staples such as wheat rise, exporters frequently introduce export restrictions, while importers often reduce import barriers to avoid having their domestic prices rise in line with world prices (Giordani, Rocha, and Ruta 2016). When world prices fall, countries sometimes raise import tariffs or use export subsidies to avoid declines in their domestic prices (Martin and Anderson 2011). These actions may, in turn, magnify the impacts of primary shocks—such as those due to weather or conflict—on world food prices, both upward and downward (Martin and Minot 2022).

The important paper by Giordani, Rocha, and Ruta (2016) provides a new theoretical rationale for the observed approach of countries using price insulation in markets for staple foods—a response not predicted by most political-economy models. Drawing on applications of the theory of loss aversion to trade policy by Freund and Özden (2008) and by Tovar (2009), they argue that the losers from food price changes are more strongly motivated to press for policies that compensate for their losses than the gainers are to press for policies that preserve their newfound gains. Their model complements the risk-aversion based explanation for price insulating policies proposed by Pieters and Swinnen (2016). Both approaches are consistent with a longstanding, informal view that policymakers in many developing countries feel an imperative to stabilize the domestic prices of staple foods (Timmer 2010). And it is not only in developing countries that these pressures arise. Anderson and Nelgen (2012) show that policymakers in rich countries lowered protection to rice and wheat during the 1972–74 price surge and raised protection during the 1984–86 period of depressed prices.

With rare exceptions that span the two literatures, such as Swinnen (1994), two bodies of work on food trade policies exist in parallel: one focuses on the political-economy forces determining equilibrium levels of protection (Grossman and Helpman 1994; De Gorter and Swinnen 2002;) the other (Martin and Anderson 2011; Giordani, Rocha, and Ruta 2016) focuses on changes in protection due to disturbances in food markets. But these two policy responses are interrelated. If protection rates change when world prices change, this disturbs the political-economy equilibria that underly the observed tendency for protection to be systematically positive in some countries and negative in others (Anderson 2009). For parameter estimation and policy analysis, what seems to be needed is a model encompassing efforts to both stabilize domestic prices and to return to the level of protection that reflects the political-economy equilibrium.

The purpose of this paper is to improve our understanding of the policies followed by countries in response to changes in world food prices. We develop and estimate a policy model that encompasses both short- and long-term policies for trade in two key staple foods—rice and wheat—using newly available long data series on prices of staple foods collected to measure protection rates. Importantly, these data eliminate nonpolicy influences—such as the additive margins emphasized by Hoffman et al (2024)—that cause conventional measures of world and domestic prices to respond differently to shocks, leaving only the effects of trade policy. The resulting parameter estimates help to understand trade policies for these staple foods and provide a potential basis for analysis of reforms to improve their performance.

An important innovation of our analytical framework is that it allows changes in domestic prices of staple foods to be decomposed into <u>systematic</u> changes in response to changes in world prices and <u>idiosyncratic</u> or random price shocks. Systematic policy changes follow our policy model, responding both to changes in world prices and to deviations from the political-economy equilibrium. While it is possible for insulating policies to stabilize domestic prices at the expense of destabilizing world prices (Sampson and Snape 1980), many policy approaches used to insulate domestic policies do not do this successfully. Instead, they introduce idiosyncratic shocks to domestic prices from a range of sources, such as:

- Supply shocks when quantitative restrictions such as export bans are in effect,
- The timing and magnitude of adjustments to administered domestic prices, and
- Changes in the success of interest groups' application of political pressure.

As noted by Martin and Anderson (2011) and Giordani, Rocha, and Ruta (2016), systematic responses to changes in world prices are likely to be correlated across countries and hence to magnify the effects of shocks to world prices. By contrast, most of the impacts of idiosyncratic national policy shocks on world prices are country-specific so that their effects on world price volatility will largely be diversified away, like idiosyncratic shocks to individual assets in investment portfolios (Elton and Gruber 1997). However, these idiosyncratic national price shocks directly affect domestic prices, and hence the welfare of incompletely diversified producers and consumers in these economies.

Notably, a key finding is that idiosyncratic shocks and the magnification of world prices associated with systematic policy shocks outweigh the efforts of most policymakers to stabilize their domestic prices relative to their volatility in the absence of price insulation. In many cases, the price volatility caused by idiosyncratic shocks is enough to destabilize domestic prices even relative to world market prices that have been destabilized by the magnification effects of systematic policy responses. While earlier studies such as Anderson, Martin and Ivanic (2017) suggested that use of price-insulating policies was a zero-sum policy game—with those willing to insulate more than the average able to reduce their domestic price volatility relative to a global free trade environment—our analysis suggests the game is likely negative-sum for almost all countries.

The next section of this paper examines the motivations of policymakers and develops a modeling approach to represent policy responses to world price changes, while also allowing for idiosyncratic policy shocks. The third section considers the data used. The fourth section outlines the approach to estimation. The fifth section presents results for world rice and wheat markets, while the sixth section interprets the results and discusses their policy implications. The final section summarizes and draws conclusions.

2. Modeling Policy Responses

Whenever a country trades a homogenous commodity, its domestic price can be linked to the world price using a tariff equivalent $(1+t)$ that summarizes the protective effect of trade measures such as tariffs and/or quotas. In levels, this may be written:

$$P = (1+t)\cdot P^w \tag{1}$$

where P is the domestic price; P^w is the external price for the same commodity at the same point in the marketing chain; and t is the tariff or export tax equivalent of the border measures applying to this commodity.

Following the logic of Grossman and Helpman's seminal paper (1994, p. 842), political economy bargaining between interest groups and the government determines an equilibrium proportional tariff equivalent, t^*, that depends on generally stable parameters, such as the elasticity of import demand/export supply, the share of domestic production in total consumption, import/export status, and the extent to which producers and consumers of the commodity are organized. The terms-of-trade explanation for trade policy proposed by Bagwell and Staiger (2002) postulates that trade barriers are determined by similarly stable parameters such as the elasticities of foreign demand for exports and foreign supply of imports. In this case, trade policy is motivated by efforts to take advantage of market power, that is, the ability of large importers and exporters to influence world prices.

In logs, equation (1) yields:

$$p^* = \tau^* + p^w \tag{2}$$

where p^* is the log of the desired level of domestic prices; τ^* is the log of $(1+t^*)$, capturing the effects of influences such as the constant price elasticities and political-economy weights considered by Grossman and Helpman and/or Bagwell and Staiger; and p^w is the log of the world price.[1] That the coefficient on p^w is one in this relationship, as implied by the Grossman-Helpman and Bagwell-Staiger models, is a potentially testable hypothesis.

If models in which policymakers maintain a stable optimal tariff were a complete representation of trade policy, then the log of the domestic price, p, would equal p^* at all times. Then the ratios of domestic and world prices would be constant and changes in the log of the domestic price would be given by:

(3) $$\Delta p_t = 1 \cdot \Delta p_t^w$$

In contrast, models in which policymakers seek to stabilize domestic prices relative to world prices imply price transmission coefficients of less than one. This study uses a model in which domestic prices are adjusted by a coefficient δ, where $0 \leq \delta \leq 1$, encompassing: the full insulation case where $\delta=0$; intermediate values of δ involving partial insulation; and the case of full price transmission (and constant levels of protection) where $\delta=1$.

The formulation used in this paper is consistent with a situation where policymakers have two distinct goals: (i) avoiding sharp changes in prices, which can generate intense political reactions if they reduce the welfare of an interest group like net food buyers or sellers below an anticipated reference level; and (ii) avoiding large deviations from the political-economy equilibrium, which can also generate intense political reactions. These goals are frequently in tension as, for example, when policymakers seek to avoid sharp increases in domestic prices in response to higher world prices. This requires reductions in protection rates, frequently to below their political-economy equilibrium levels. This in turn creates pressure from the interest groups supporting the political-economy equilibrium to reduce the gap between the applied rate of protection and the political-economy equilibrium rate. If policymakers have only one instrument for these two goals, the best they can do is to minimize a weighted average of these two sources of political costs.

The Error Correction Model and Trade Policy

Much of the extensive literature on price transmission in agricultural markets focuses on situations where competition can be expected to result in equilibrium price differentials that

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[1] For small values of the tariff, $\tau=\ln(1+t^*)$ is approximately t^*.

reflect both policy and nonpolicy influences, such as costs of transport or product transformation (von Cramon-Taubadel and Goodwin 2021). Many of these studies use the Error Correction Model (ECM) to capture the dynamics of adjustment, including any policy responses, and to deal with the statistical properties of the data series (Engle and Granger 1987). We build on this approach but focus purely on policy by using data on domestic and external prices that have been adjusted to remove any differences other than those due to policies.

Nickell (1985) provides an interpretation of the ECM as representing a dynamic optimization problem where decision-makers minimize a weighted sum of the costs associated with adjusting control variables and those associated with deviations from targets. The objective function for our trade policy problem is:

(4) $$C_t = \sum_{s=0}^{\infty} \alpha^s \left[(p_{t+s} - p_{t+s-1})^2 + \lambda(p_{t+s} - p^*_{t+s})^2\right]$$

where α is a discount factor (0≤λ≤ 1); $(p_{t+s} - p_{t+s-1})^2$ represents the costs of changing domestic prices; and $\lambda(p_{t+s} - p^*_{t+s})^2$ represents the costs of deviating from the political-economy equilibrium price, where λ is a weight representing the costs to policymakers of deviations from the political equilibrium price relative to those of changing prices.

To make (4) operational, it is necessary to specify a stochastic process for $p*$ and for future world prices. A simple approach to specifying $p*$ follows the Grossman-Helpman specification in equation (2). Nickell shows that if world prices follow a second-order autoregressive scheme with a unit root—which Nickell sees as ubiquitous for macroeconomic aggregates—then equation (4) yields an ECM that is for our problem:[2]

(5) $$\Delta p_t = \delta \Delta p^w_t + \theta(p_{t-1} - \tau^* - p^w_{t-1})$$

where δ is a short-run adjustment coefficient ranging between zero and one, showing the extent to which the domestic price is adjusted in response to changes in the world price; the expression in parentheses is the deviation from the long term political-economy equilibrium tariff level in the previous period; and θ is an error correction (EC) coefficient indicating the speed of adjustment toward this equilibrium. The augmented Dickey-Fuller-Generalized Least Squares (ADF-GLS) tests (Elliott, Rothenberg, and Stock 1996) for stationarity of the world price

[2] We simplify Nickell's model by omitting the constant terms that, while important for modeling growing macroeconomic aggregates, are almost invariably insignificant in our application.

variables used in our analysis found that this model was typically the best—and almost always an acceptable—univariate representation of the data-generating process for our world price data.[3] Nickell's section III shows that the same ECM, with some additional difference terms, arises if the data-generating process for world prices involves moving-average processes or additional autoregressive terms. He also shows that, if world prices follow a pure random walk, then $\delta = -\theta$ and it simplifies to a partial adjustment model.

While equation (5) focuses on the relationship between domestic and world prices, it can easily be transformed into a relationship between world prices and the rate of protection. If we define the rate of protection in logarithms as $\tau = (p - p^w)$ and subtract Δp_t^w from both sides, equation (5) can be rewritten as:

(6) $$\Delta \tau_t = (\delta - 1)\Delta p_t^w + \theta(p_{t-1} - \tau^* - p_{t-1}^w)$$

When δ is less than one, this model has the intuitive implication that an increase in the world price causes the rate of protection to decline. In this equation, $(\delta - 1)$ is an elasticity of price insulation, as opposed to the short-run elasticity of price transmission, δ, in equation (5).

Equations (5) and (6) provide important insights into policymakers' relative weights on aversion to sharp price changes and aversion to deviation from the politically optimal relationship between domestic and world prices. The lower the price adjustment coefficient, δ, the greater the political costs of adjusting domestic prices, and the greater the extent to which price instability is exported to countries less willing or able to insulate their markets. The higher the absolute value of θ, the more rapidly policymakers return protection to its political equilibrium level following a shock to world prices. The equilibrium level of protection, τ^*, can be estimated as part of the model.

This specification for price-insulating policies is closely related to normative approaches using a welfare function incorporating risk aversion (Gouel and Jean 2015; Pieters and Swinnen 2016). The approach used here does not require knowledge of the mean value of the price series, which Wright and Williams (1990) argue is typically unknown. Further, matching observed rates of price insulation frequently requires very high values of the coefficient of relative risk aversion.

[3] This model, which can be transformed to $p_t^w = p_{t-1}^w + \gamma \Delta p_{t-1}^w + \varepsilon_t$, was the null hypothesis in ADF tests for nonstationarity of the world price variables.

Gouel, Gautam, and Martin (2016, p. 840) needed an estimate of close to 6 to match policy responses in India's wheat market, while Gandelman and Hernández-Murillo (2015) estimate this parameter at 0.98.

Defining the Equilibrium Rate of Protection

In principle, it would seem attractive to model the equilibrium rate of protection using national political economy models as in Anderson (1995) or Gawande and Hoekman (2006). However, this seems unlikely to work well given the limited sample size for each country, and the fact that agricultural protection in some economies seems to have been strongly influenced by international trade negotiations as well as by changing domestic political pressures. Thus, we decided to represent the evolution of τ^* using simple quadratic functions of time—proceeding from general to specific models (Campos et al. 2005) to avoid including unnecessary trend and trend squared terms—with structural breaks identified using the Gregory-Hansen procedure (1996) allowed where necessary.

Implications of the Loss Aversion Model for Price Insulation

Freund and Özden (2008) provide an important and nonintuitive insight into the implications of the loss aversion model: price insulation is complete over some range of price changes or, in our formulation, δ =0. This full-insulation result would have enormous implications for market stability in our context. If it applied to all countries and a primary shock caused the log of world prices to rise by Δp^w, then equation (6) shows that each country would lower its agricultural protection rate (in logs) by Δp, negating the increases in domestic prices needed to offset the primary shock to world markets, and requiring a second increase of Δp^w in world prices. This would set off another round of reductions in protection. With a price insulation coefficient of unity in all countries, the world market would clearly be unstable.

The internal validity of the theory is clear so it is important to understand the circumstances under which it might—or might not—apply to our estimated models. The Appendix identifies two possible reasons for less than complete price insulation: (i) changes in prices that typically extend beyond the range of full price insulation; and (ii) differences between initial domestic price levels and the reference price levels that, once crossed, generate intense feelings of loss.

3. Price Data for Estimation

Ideally, estimation of the policy model outlined in Section 2 would use high-quality data on domestic and external prices adjusted to the same point in the marketing chain so that changes in their relative prices reflect trade policies alone, rather than being conflated with nonpolicy influences on price transmission, such as additive margins, changes in the direction of trade, differences in product quality, and/or lags in price adjustment. The data would also cover a long period so that multiple cases of unusually high- and low-price periods are included in the analysis. Fortunately, reasonably long time series of annual data with these attributes are now available from the World Bank's Distortions to Agricultural Incentives (DAI) project (Anderson 2009) up to 2004 and the AgIncentives Consortium (Laborde et al. 2024) for subsequent years.

For our objective of making inferences about trade policies, these data are much better than standard food price series, such as those from FAO's GIEWS[4] used by Martin and Minot (2022) or FAOSTAT data on producer price indexes. To obtain reliable estimates of the trade barriers generated by any applicable trade measures, the analysts who compile these price series take many quality control steps, such as: identifying domestic and foreign products that are as similar as possible; making adjustments for any remaining quality differences; identifying the direction of trade (moving, for example, from Free on Board (FOB) prices for exports to Cost, Insurance, and Freight (CIF) prices for imports as the direction of trade changes); and adjusting for internal transport and marketing margins between the farmgate and the border (OECD 2016).

One potential nonpolicy source of shocks to domestic markets arises when there is no trade and domestic prices are between the FOB price of exports and the CIF price of imports. To assess the potential importance of this source of deviations from our models' predictions, we examined the share of country periods during which countries were classified by the analysts measuring protection as neither importers nor exporters. For rice, only 16 of 1,498 country-periods—10 in India and 6 in Uganda—were classified as nontraded. For wheat, only 46 of 1,577—24 in Ethiopia and 22 in India—were nontraded. The overall importance of nontraded

[4] The United Nations Food and Agriculture Organization's Global Information and Early Warning System on Food and Agriculture.

status in influencing our results appears to be very small and concentrated in only a few countries.

Price data at the same point in their marketing chains are available from 1955 or 1961 to 2021 in many cases, although only shorter time series were available for many transition economies and some other developing countries. These datasets provided data on rice prices for 29 economies, with the European Union (EU) treated as one trade bloc with evolving membership and the United Kingdom excluded because of structural changes associated with its accession to, and exit from, the EU. For wheat, we obtained similar data for a slightly different set of 29 economies. A small number of missing observations were replaced by linear interpolation of the logged values, as in Martin and Minot (2022), because some of the algorithms used, such as those for ADF tests, are unable to handle missing values.

Two key questions were whether to (i) conduct the analysis using prices in domestic currency or a common currency, and (ii) use nominal or real price series. We chose US dollars because the series are provided in that currency and because the relative price represented by τ is invariant to the choice of the common currency in which these prices are presented. Nominal price series have the advantage of transparency and of avoiding introduction of irrelevant variation. If the US consumer price index (CPI) is used as a deflator, irrelevant variations from the point of view of other countries—such as changes in the relationship between US traded and nontraded goods prices—are introduced whenever the US real exchange rate changes. Deflating by national CPIs introduces real exchange rate changes as apparent sources of changes in both domestic and international prices, and often sharply reduces the length of the available series. To allay potential concerns that our findings might be different had we used deflated data, we calculated the volatility of the log first-differenced series following deflation by the US CPI and found that the results were essentially the same as with nominal prices.

Rice Price Data

Figure 2 presents data on rice prices received by producers and the reference prices that producers would have received in the absence of trade policy interventions. These price data highlight sharp differences between domestic and external prices in many countries, while showing the close correspondence between the two series in other countries. Even in countries where the two series diverge for extended periods, however, the domestic series appear to respond to international prices to some degree with, for example, the producer and reference prices in India beginning and ending in similar proportion to each other. A striking feature of the graphs is that it is far from clear that the domestic prices are any less volatile than the external prices, despite the strong apparent focus on price stabilization in many discussions of trade policy for food staples (Timmer 2010).

Table 1 provides key summary statistics for the domestic and external price series for rice. The first two columns provide the volatility of domestic and world prices, defined as the standard deviation of first differences in logged prices. We use first differences of these variables rather than their levels because, as we will see, nominal prices generally appear to be nonstationary, making their variances unstable. An important insight that does not rely on parameter estimation is that domestic price volatility exceeds external price volatility in 8 of the 29 cases (or 28 percent), with a further one indistinguishable. Many cases where the volatility of the domestic price exceeds that of the external reference price are in countries such as India, where policy narratives appear to place a great deal of emphasis on achieving price stability (Timmer 2010). Pooling across all sample countries, the average within-country standard deviation (0.26) is slightly above the average standard deviation for external reference prices (0.25). These two results are striking given that price-insulating policies such as variable levies can potentially completely stabilize domestic prices (Sampson and Snape 1980) and many countries appear to be incurring substantial economic costs through policies of price insulation and storage designed to stabilize their domestic prices relative to world prices.

Table 1. Volatility of prices and protection rates, and tariff averages for rice

	Volatility of domestic price $SD(p_t-p_{t-1})$	Volatility of external reference price $SD(p^w_t - p^w_{t-1})$	Standard deviation of the tariff equivalent $SD(p_t-p^w_t)$	Mean of tariff equivalent $\mu(p_t-p^w_t)$
Australia	0.24	0.26	0.09	0.08
Bangladesh	0.26	0.19	0.23	0.01
Brazil	0.30	0.36	0.21	0.06
China	0.14	0.18	0.37	-0.15
Colombia	0.18	0.21	0.37	0.39
Dominican Rep.	0.22	0.21	0.33	0.40
Ecuador	0.19	0.22	0.28	0.11
European Union	0.15	0.24	0.42	0.13
Ghana	0.41	0.24	0.46	-0.15
Indonesia	0.19	0.20	0.25	0.12
India	0.33	0.27	0.29	-0.35
Japan	0.13	0.15	0.51	1.27
Kazakhstan	0.18	0.20	0.41	-0.39
Kenya	0.27	0.21	0.28	0.45
Korea, Rep.	0.16	0.18	0.60	0.69
Sri Lanka	0.23	0.26	0.23	0.05
Mexico	0.17	0.24	0.23	0.04
Mozambique	0.59	0.37	0.75	-0.54
Nigeria	0.26	0.26	0.31	0.21
Nicaragua	0.14	0.22	0.28	0.34
Pakistan	0.23	0.27	0.35	-0.35
Philippines	0.13	0.24	0.43	0.24
Senegal	0.18	0.23	0.33	0.16
Türkiye	0.28	0.30	0.60	0.64
Tanzania	0.28	0.27	0.46	-0.33
Uganda	0.35	0.36	0.30	0.25
United States	0.22	0.30	0.16	0.08
Viet Nam	0.23	0.24	0.19	-0.03
Zambia	0.39	0.25	0.48	-0.27
Total	0.26	0.25	0.38	0.12/0.37

Source: Authors' estimates. Notes: Countries are listed in order of ISO3 code. Variable p is the natural log of the domestic price; p^w is the log of the external price. Standard deviations for prices were computed country by country. The Total estimates are for variation within countries, obtained using regressions for the log price data against a set of 28 country dummy variables. For comparability with the econometric analysis to follow, the tariff equivalents are defined as $(p - p_w)$, which is equivalent to $\ln(1+t)$ where t is a conventional proportional tariff. For small tariffs, this is approximately equal to t. Two global mean tariffs are provided. The first, 0.14, is the simple average of the national means. The second, 0.39, is the average of the absolute values of the tariff equivalents.

The third column of Table 1 shows the standard deviation of the tariff equivalents, while the final column shows average rates of protection by country. As shown by Francois and Martin

(2004), the cost of protection in an individual country can be estimated by multiplying the slope of the import demand or export supply curve by the sum of the variance of the tariff equivalent and the square of the average tariff rate. If changes in protection rates lower domestic price volatility, the costs of protection variability can be weighed against the value of the reduction in price volatility. If changes in protection do not reduce the volatility of domestic prices, it seems difficult to justify the costs of protection volatility.

The overall average tariff is 12 percent, with some economies, such as Japan, having high average rates of protection (127 percent), and others, such as Pakistan, having substantially negative average protection rates (-35 percent). The within-country standard deviation of tariff equivalents, at 38 percent, is more than three times the average protection rate (12 percent). Since what matters for economic costs is the square of the absolute value of the protection rate, we calculated the average of the absolute values of protection rates, which is 37 percent.

Simple calculations following the formula in Francois and Martin (2004, equation (6)) with conservative supply and demand elasticities of 0.3 and -0.3 respectively suggest that the total cost of trade distortions to rice in the countries imposing them comes to $33 billion per year at average 2010 to 2021 prices—more than 10 percent of the value of rice production at world prices. Of this total, $13 billion comes from average protection rates while $20 billion comes from their volatility.

Wheat Price Data

Figure 3 presents data for wheat prices on the same basis as those for rice in Figure 2. These graphs provide some important insights into the nature of the relationships between producer and reference prices. In some countries, such as Argentina, the difference between the two prices is highly variable, being sizeable in some periods and close to zero in others. In other countries, such as Australia and Canada, the differences are quite small in most periods, and essentially zero in recent years. In a few countries, such as India, the producer price is much more stable than the reference price, but this pattern is surprisingly rare. In many other countries, such as Brazil, Kenya and Zambia, the domestic price appears to be at least as volatile as—or perhaps more volatile than—the external reference price.

Figure 2. Producer and reference prices for rice at farmgate

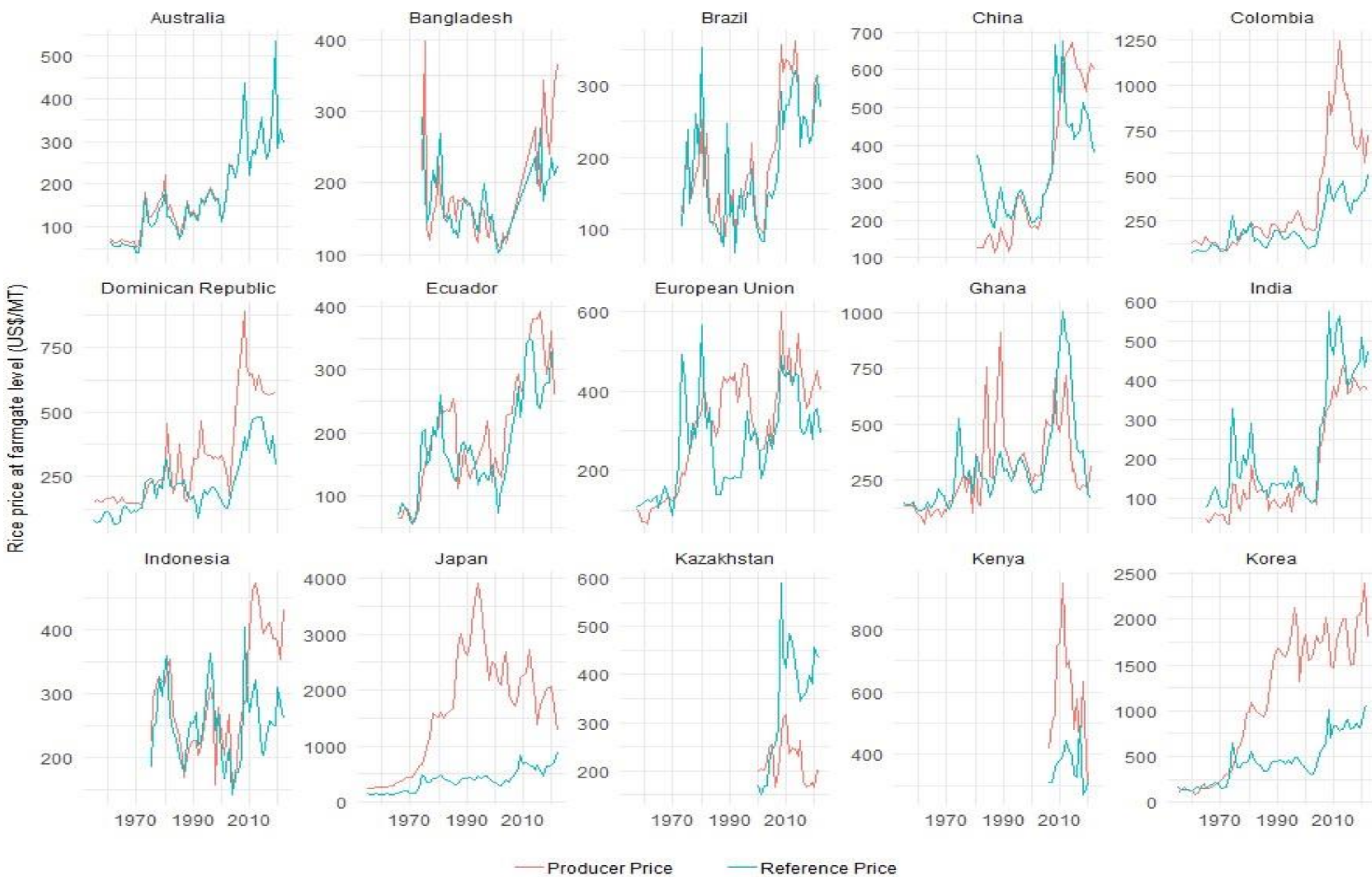


Figure 2. Producer and reference prices for rice at farmgate, continued

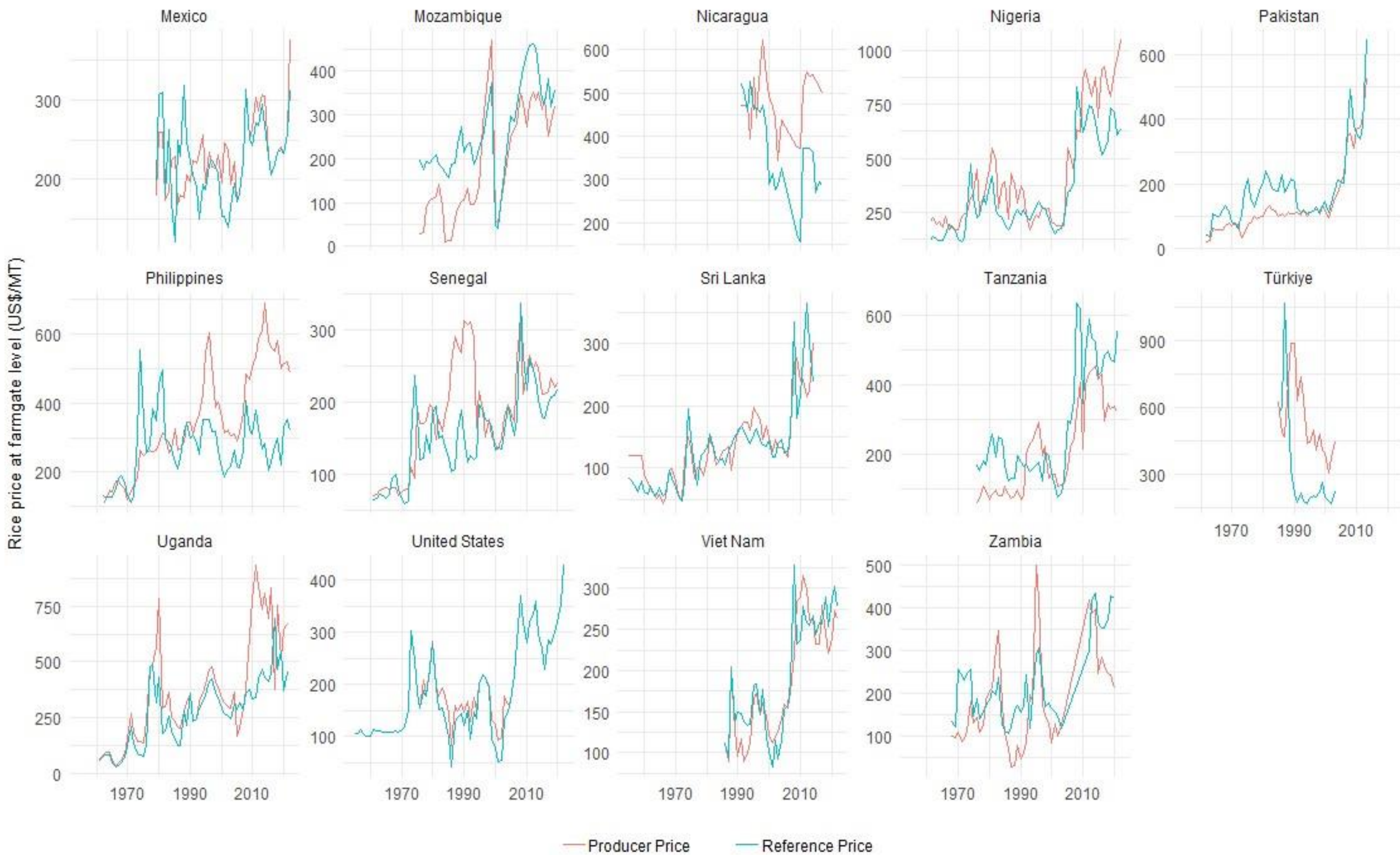


Figure 3. Producer and reference prices for wheat at farmgate

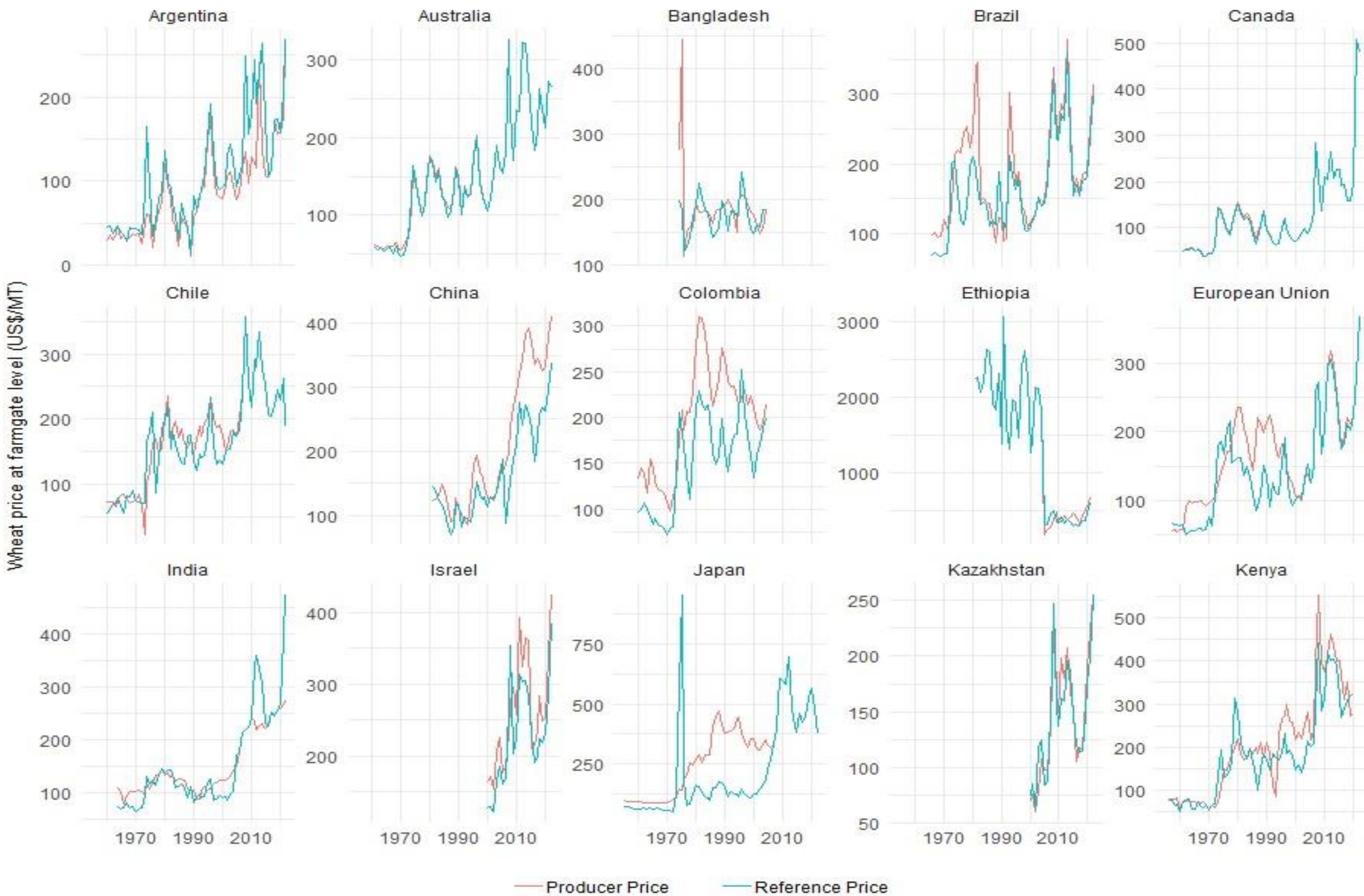
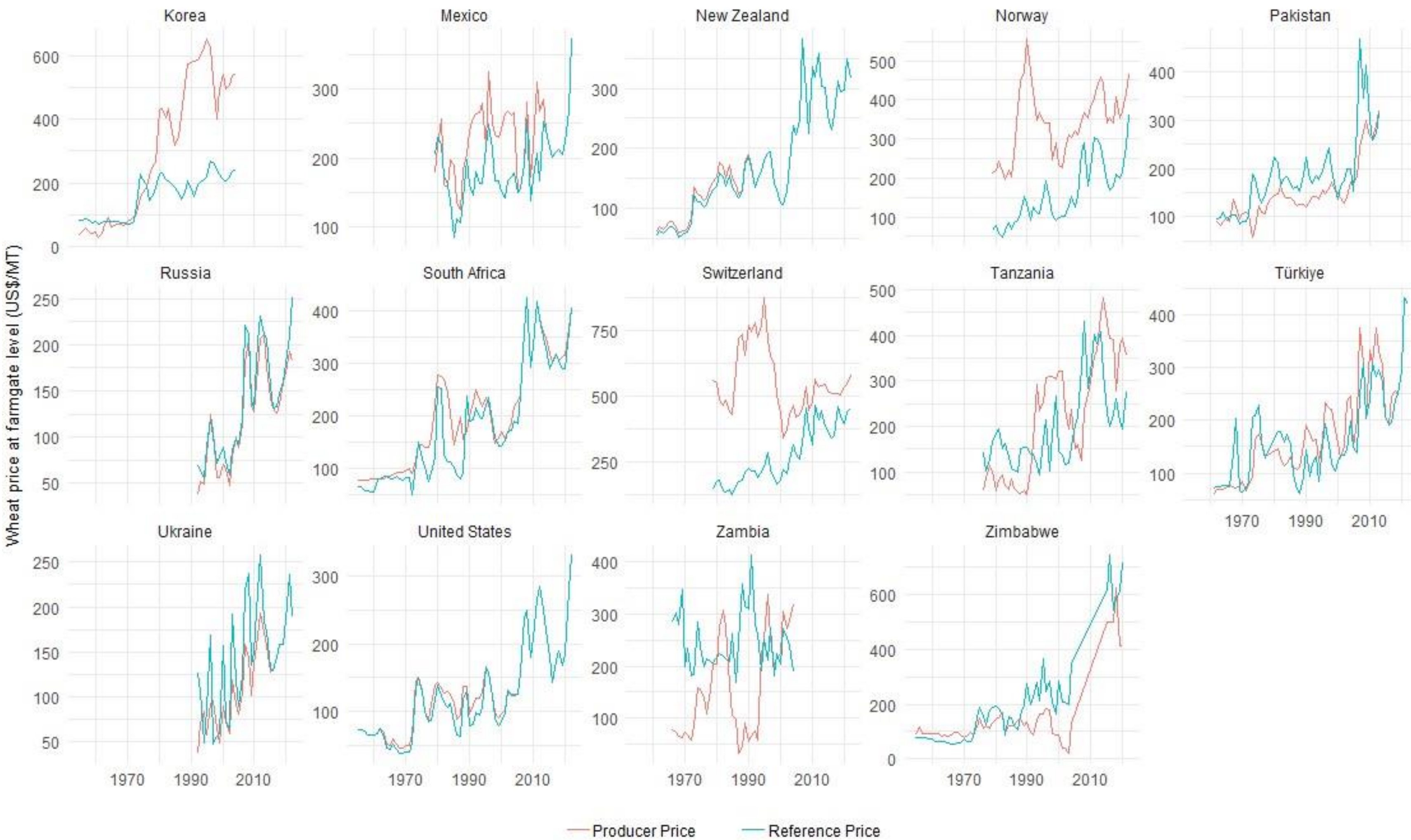


Figure 3. Producer and reference prices for wheat, continued



Summary statistics for wheat price data are presented in Table 2, following the format of Table 1. While the average volatility of domestic prices is slightly below that for external prices, domestic prices are more volatile in 9 of our 29 sample economies (or almost one-third) and the same in another 3.

Table 2. Volatility of prices and tariff rates, and tariff averages for wheat

	Volatility of domestic price $SD(p_t - p_{t-1})$	Volatility of international reference price $SD(p^w_t - p^w_{t-1})$	Standard deviation of the tariff equivalent $SD(p_t - p^w_t)$	Mean of tariff equivalent $\mu(p_t - p^w_t)$
Argentina	0.45	0.45	0.21	-0.22
Australia	0.21	0.22	0.05	0.02
Bangladesh	0.29	0.15	0.20	0.06
Brazil	0.28	0.23	0.24	0.12
Canada	0.24	0.24	0.03	0.02
Switzerland	0.12	0.16	0.47	0.78
Chile	0.28	0.24	0.24	0.05
China	0.14	0.21	0.20	0.21
Colombia	0.11	0.17	0.17	0.28
Ethiopia	0.48	0.41	0.18	0.00
European Union	0.16	0.21	0.29	0.24
India	0.09	0.16	0.21	0.07
Israel	0.21	0.22	0.14	0.14
Japan	0.12	0.36	0.56	0.45
Kazakhstan	0.23	0.29	0.14	0.00
Kenya	0.21	0.22	0.26	0.09
Korea, Rep.	0.19	0.15	0.60	0.34
Mexico	0.22	0.21	0.23	0.21
Norway	0.14	0.23	0.36	0.87
New Zealand	0.17	0.17	0.05	0.05
Pakistan	0.16	0.20	0.26	-0.20
Russia	0.28	0.25	0.13	-0.11
Türkiye	0.17	0.31	0.35	0.11
Tanzania	0.27	0.29	0.64	-0.03
Ukraine	0.31	0.52	0.35	-0.17
United States	0.20	0.23	0.10	0.08
South Africa	0.14	0.26	0.24	0.17
Zambia	0.40	0.24	0.77	-0.60
Zimbabwe	0.33	0.23	0.55	-0.28
Total	0.24	0.26	0.34	0.10/0.28

Sources and Notes: As for Table 1.

Just as for rice, the average rate of protection (10 percent) is relatively low because it covers countries with both positive and negative support. The average absolute rate of protection is much higher at 28 percent. Even this rate is lower than the standard deviation of the tariff equivalent, implying that the costs from variability of protection may be much higher than those associated with the average protection rate—the rate typically used in model-based assessments of the benefits from trade reform.

Simple calculations using equation (6) from Francois and Martin (2024) suggest that the cost of trade distortions to wheat are around $12 billion per year for the countries imposing these barriers. Of this, $3.9 billion per year arise from average rates of protection, while $8.3 billion come from the volatility of protection rates. The total cost is 6.8 percent of the value of wheat production at world prices. The cost of trade distortions to wheat are substantially lower than for rice because wheat markets are more open and both average trade distortions and their intertemporal volatility are lower.

4. The Approach to Estimation

The approach used for this analysis was to begin by examining whether the price variables being considered are nonstationary. If they are—which is generally the case for variables of this type—then a key question is whether their first differences are stationary. If so, the next question is whether domestic and world prices are cointegrated, meaning a stable relationship exists between them. If both conditions are satisfied, then an ECM like equation (5) potentially provides both reliable parameter estimates and tests of their significance. As noted by Wickens and Breusch (1988), the ECM is a transformation of an autoregressive distributed lag model and may also be used when the data in levels are stationary. If cointegration is not found or the model results are inconsistent with theory, then one possibility is that policymakers do not follow a cost-minimizing approach like equation (4). They may instead focus on self-sufficiency, while allowing domestic prices to disconnect from world prices.

Engle and Granger (1987) suggest estimating the ECM by first estimating the long-run relationship between the variables using Ordinary Least Squares (OLS). Whether the long-run relationship between these variables is stationary—and hence the series are cointegrated—can be tested using a modified Dickey-Fuller test, such as egranger in STATA, where the significance tests are adjusted because the residuals rely on estimated parameters (Schaffer 2022). Schaffer identifies this test as involving fewer distributional assumptions than, and providing a robust alternative to, the popular Johansen (1991) tests for cointegration.

Once a cointegrating relationship has been found, Engle and Granger suggest using the lagged residuals from the cointegrating equation together with the price change variables in equation (5) to estimate the dynamic adjustment parameters of the model, δ and θ. A key disadvantage of this approach is that it does not provide reliable significance tests for the long-run coefficients that are also of interest here. Addressing this problem, Inder (1993, p. 68) advocated estimating ECMs with a single dependent variable by nonlinear least squares (NLS), which he found to yield precise estimates and valid significance tests even in the presence of endogeneity in the explanatory variables. The simple Engle-Granger approach was, however, useful for checking and confirming that constant terms were not needed in our estimating equations based on equation (5).

As noted by Inder, the Bewley (1979) transformation approach provides a linear-in-parameters instrumental variable estimator for the long-run coefficients that we used as a complement and cross check. This approach, however, does not provide direct estimates of the adjustment parameters of central interest to our analysis. We also considered maximum likelihood approaches based on Johansen (1991) but concluded that a parsimonious NLS approach was more likely to generate useful results given our need to estimate with single-country samples.

The specific steps undertaken for estimation were to:

(i) Use STATA dfgls to test whether levels and first differences of prices are nonstationary,

(ii) Perform egranger cointegration tests (Schaffer 2022) for long-run relationships,

(iii) Perform Gregory-Hansen (1996) tests for structural breaks, and

(iv) Use NLS, incorporating structural breaks where necessary, to estimate the short- and long-run coefficients and their significance.

5. Results

The results for rice are presented first, followed by those for wheat. In each case, the diagnostic tests for nonstationarity, and for the potential existence of a stable relationship between domestic and world prices, are considered first, followed by results for estimation of the ECM.

Results for Rice

Table 3 presents the results from preliminary testing of the rice price series for stationarity, and of tests for cointegration between domestic and world prices. The key result in this table is that, for the price data in levels, it was rarely possible to reject the null hypothesis of a unit root, therefore suggesting that these are integrated series. By contrast, the overwhelming majority of the first-differenced variables were found to be stationary, with the few exceptions in countries like Kazakhstan and Türkiye with relatively small samples. The cointegration test results in the final column of the table suggest that domestic and world prices are cointegrated in all but a few cases. These results are consistent with an ECM specification like equation (5).

Table 3. Dickey-Fuller tests for integration and cointegration of rice prices

	Test of the null hypothesis that the series is nonstationary				
	Log of domestic price (p)	Log of intl. price (p^w)	1st diff. log domestic price (Δp)	1st diff. log intl. price (Δp^w)	Cointegration: log of domestic & intl. prices
Australia	-4.18***	-5.25***	-5.92***	-6.24***	-4.36**
Bangladesh	-1.58	-1.38	-4.31***	-3.63**	-6.14***
Brazil	-2.35	-2.15	-5.49***	-4.59***	-5.39***
China	-2.96	-1.57	-3.80***	-3.79***	-4.04**
Colombia	-2.09	-3.11*	-4.95***	-5.46***	-2.86
Dominican Rep.	-3.98***	-2.92*	-7.07***	-5.35***	-4.75***
Ecuador	-1.85	-1.86	-6.18***	-4.55***	-3.79*
European Union	-1.80	-3.09*	-4.61***	-5.88***	-3.41*
Ghana	-2.76	-3.37**	-7.33***	-5.21***	-4.77**
Indonesia	-2.23	-3.22	-5.56***	-5.21***	-4.51*[a]
India	-2.37	-2.26	-6.04***	-4.55***	-5.40***
Japan	-0.82	-3.09	-5.02***	-5.56***	-5.46**[a]
Kazakhstan	-1.57	-2.73	-2.94	-3.03*	-3.84*
Kenya	-1.37	-1.96	-1.67*	-3.49**	-3.81
Korea, Rep.	-1.14	-3.17	-5.86***	-5.32***	-6.72***[a]
Sri Lanka	-3.58**	-4.20***	-7.14***	-6.81***	-4.25***
Mexico	-3.06*	-2.87	-4.11***	-3.25*	-4.62**
Mozambique	-3.69**	-2.88	-4.88***	4.38***	-3.66*
Nigeria	-2.22	-3.08*	-6.75***	-5.99***	-4.65***
Nicaragua	-1.84	-1.84	-3.06**	-3.41**	-4.45**
Pakistan	-2.21	-1.93	-7.56***	-5.45***	-5.66***

	Test of the null hypothesis that the series is nonstationary				
	Log of domestic price (p)	Log of intl. price (p^w)	1st diff. log domestic price (Δp)	1st diff. log intl. price (Δp^w)	Cointegration: log of domestic & intl. prices
Philippines	-2.16	-2.61	-4.95***	-7.07***	-3.48
Senegal	-1.62	-4.45***	-5.92***	-7.16***	-3.13*
Türkiye	-3.02*	-0.90	-2.73	-2.54	-5.16**[a]
Tanzania	-2.20	-2.08	-4.47***	-4.92***	-4.65**[a]
Uganda	-2.46	-3.10*	-4.53***	-4.99***	-4.46***
United States	-2.90*	-2.88*	-5.91***	-5.34***	-5.86***[a]
Viet Nam	-2.25	-1.69	-3.61**	-4.10***	-4.64**[a]
Zambia	-2.6	-2.26	-4.19***	-5.45***	-3.76**

Note: p refers to the domestic price and p^w to the border price, both adjusted for transport and marketing costs to the farmgate (Anderson 2009). ADF-GLS tests were performed using the dfgls command in STATA. Where structural breaks were not detected, modified ADF tests for cointegration were performed using egranger in STATA (Schaffer 2022). Cointegration test results with the superscript [a] were obtained using the Gregory-Hansen procedure to identify potential structural breaks and test for cointegration.

The ECM was first estimated allowing a nonunitary coefficient (β_1) on p_{t-1}^{w} in equation (5), to permit testing of the Grossman-Helpman hypothesis that this coefficient is unity. Highly significant coefficients close to unity were typically obtained when no other variables were included in the EC term within parentheses in equation (5). This parallels the Mundlak and Larson (1992) finding—from OLS regressions in levels that should have provided super-consistent estimates—of coefficients on world prices that are close to one. However, when equilibrium protection rates, and/or variables for changes in these rates, were introduced in our equations, the estimated values of β_1 frequently deviated from unity. Table A.1 presents preliminary estimates of this model for reference.

It seems likely that the problems in estimating coefficients on variables other than world price in the long-run relationships result from difficulties in distinguishing reliably between intercepts and slopes in these relationships. To help avoid these difficulties, the models were re-run with the β_1 coefficient restricted to unity, as implied by theory (see Table 4 for results). These estimates include trend and trend squared terms within the EC term to allow for potential structural changes in equilibrium levels of protection as the political-economy equilibria evolve. Most importantly, this analysis confirmed the robustness of estimates of the δ and θ coefficients to the imposition of this restriction. The mean absolute difference between estimates of δ from the restricted and unrestricted models was only 0.024.

Table 4. Results for nonlinear restricted estimates of ECM models for rice, $\beta_1 \equiv 1$

	δ	θ	β_0	β_2	β_3	β_4	R^2	Sample	RMSE
Australia	0.94	-0.42	0.20	-0.004			0.97	1961-2021	0.05
	39.3	*-3.7*	*7.0*	*-4.9*					
Bangladesh	0.37	-0.80	-0.10	0.005			0.47	1974-2019	0.19
	2.3	*-5.9*	*-1.3*	*1.7*					
Brazil	0.74	-0.71	-0.20	0.02	0.0004		0.70	1973-2019	0.17
	9.9	*-5.1*	*-1.9*	*2.2*	*-1.8*				
China	0.42	-0.31	-0.56	0.025			0.37	1981-2021	0.12
	3.8	*-3.3*	*-4.1*	*4.6*					
Colombia	0.52	-0.19		0.015			0.41	1960-2020	0.14
	5.7	*-2.9*		*5.2*					
Dominican Rep.	0.39	-0.20	0.46				0.19	1955-2019	0.20
	3.2	*-2.6*	*3.6*						
Ecuador	0.26	-0.41		0.006			0.38	1966-2016	0.15
	2.6	*-4.8*		*3.6*					
European Union	0.75	-0.16	-0.60	0.06	0.001		0.67	1957-2021	0.20
	10.8	*-2.2*	*-1.3*	*1.7*	*-1.6*				
Ghana	0.93	-0.42	-0.74	0.06	-0.001		0.43	1955-2018	0.32
	5.4	*-4.0*	*-2.4*	*2.7*	*-3.0*				
Indonesia	0.67	-0.55	0.46			-0.45	0.54	1975-2021	0.14
	6.2	*-4.7*	*6.3*			*-5.3*			
India	0.73	-0.67	-0.83	0.03	-0.0003		0.71	1965-2021	0.18
	7.5	*-5.6*	*-7.3*	*3.0*	*-1.8*				
Japan	0.19	-0.22	1.04	0.07	-0.001	-0.57	0.24	1955-2021	0.12
	1.9	*-3.2*	*2.4*	*3.6*	*-4.6*	*-2.0*			
Kazakhstan	0.30	-0.66	0.51	-0.13	0.003		0.51	2000-2021	0.14
	1.8	*-3.9*	*3.2*	*-3.8*	*2.1*				
Kenya	-0.17	-0.70	0.71	0.003	-0.004		0.73	2000-2021	0.17
	-0.5	*-2.5*	*2.5*	*0.0*	*-0.9*				
Korea, Rep.	0.07	-0.23		0.07	-0.001	-0.30	0.24	1955-2021	0.15
	0.6	*-3.3*		*6.3*	*-4.3*	*1.64*			
Sri Lanka	0.58	-0.40	0.06				0.46	1955-2014	0.17
	6.4	*-3.9*	*1.0*						
Mexico	0.33	-0.35	0.07				0.28	1979-2021	0.15
	3.1	*-3.3*	*1.0*						
Mozambique	0.93	-0.32	-1.02	0.03			0.47	1976-2019	0.44
	5.1	*-2.8*	*-2.2*	*1.5*					
Nigeria	0.26	-0.46	0.37	-0.01			0.27	1961-2015	0.23
	2.1	*-4.2*	*2.7*	*-1.3*					
Nicaragua	0.42	-0.52	-0.22	0.07	-0.002		0.43	1991-2017	0.12
	3.5	*-3.2*	*-1.5*	*3.0*	*-1.9*				
Pakistan	0.43	-0.36	-0.25	-0.02	0.001		0.44	1961-2013	0.18
	4.3	*-3.9*	*-1.0*	*-1.2*	*1.8*				

	δ	θ	β_0	β_2	β_3	β_4	R^2	Sample	RMSE
Philippines	0.31	-0.23	0.18	-0.02	0.001		0.39	1962-2021	0.10
	5.3	*-3.6*	*0.9*	*-1.3*	*2.6*				
Senegal	0.18	-0.18	0.25				0.11	1961-2020	0.17
	1.7	*-2.5*	*1.9*						
Türkiye	0.12	-0.35	-2.31	0.43	-0.01	1.25	0.27	1985-2003	0.28
	0.4	*-1.8*	*-1.7*	*2.2*	*-1.7*	*1.5*			
Tanzania	0.47	-0.34	-0.52	-0.02		-1.04	0.36	1976-2021	0.24
	3.5	*-3.5*	*-1.3*	*2.8*		*-2.8*			
Uganda	0.70	-0.45	0.29				0.52	1961-2018	0.25
	7.3	*-3.9*	*3.9*						
United States	0.73	-0.10	0.13				0.92	1955-2021	0.06
	27.5	*-2.1*	*1.7*						
Viet Nam	0.71	-0.57	-0.33	0.04	0.001		0.63	1986-2021	0.15
	6.7	*-3.8*	*-2.4*	*2.4*	*-2.3*				
Zambia	0.50	-0.30	-0.25				0.23	1968-2020	0.35
	2.6	*-3.0*	*-1.6*						

Notes: For conciseness, t-statistics for the null hypothesis that the coefficient equals zero are presented in italics below the estimates, rather than providing standard errors and asterisks for significance levels. The one-tailed critical values for these equations are roughly 1.3 at the 10 percent level; 1.7 at the 5 percent level; and 2.4 at the 1 percent level. These equations were estimated subject to the restriction that the long-run coefficient on world price, β_1, is unity, as implied by the Grossman-Helpman model. The δ and θ coefficients correspond to equation (5), while estimates of τ^* involve a constant β_0, a time trend coefficient β_2, a time trend squared coefficient β_3, and a pre-regime-change dummy for structural changes identified using the Gregory-Hansen procedure, β_4. The Indonesia break was in 2010; Japan in 1984; Korea in 1985; Türkiye in 1994; and Tanzania in 1996. RMSE is the Root Mean Squared Error of the residuals.

Another feature of the results in Table 4 is the consistency of most of the estimates with the political-economy interpretation provided in Section 2. In every case except Kenya, the price adjustment coefficient, δ, is positive and lies between zero and one, while the EC term, θ, is invariably negative and between zero and one, implying that deviations from the political-economy equilibrium are reduced in subsequent periods. For results consistent with the model, δ ranges from 0.94 in Australia through 0.75 in Europe, 0.73 in the United States, 0.42 in China, 0.37 in Bangladesh, 0.15 in Japan, and 0.07 in Korea. Several countries that might seem strongly averse to transmitting changes in world prices—such as India (0.73) and Indonesia (0.67)—have quite high initial passthrough for rice. In most cases, these impact effects are highly significant. In only a few cases are these coefficients not significant at the 5 percent level using a one-tailed test.

The EC coefficients (θ) range from -0.8 in Bangladesh to -0.1 in the United States. The absolute value of this parameter generally appears to be smaller than for the δ coefficient, and the absolute values of the two coefficients appear to be positively correlated. The large differences in their magnitudes and their strong significance tests in most cases, suggest that they are not equal, so that this model does not collapse to a simpler partial adjustment model (Nickell 1985).

The implied equilibrium rates of protection are also generally plausible. Where the constant term is individually significant and there are no trend terms, this coefficient is generally comparable with the average tariff rate in Table 1. Where there are trend terms, the implied average rate of protection is similar to the mean in Table 1, and the trend provides valuable information about the evolution of the equilibrium rate over time. In Australia, for example, the mean tariff equivalent in Table 1 is the same as the equilibrium rate at the midpoint of the sample.

The Gregory-Hansen structural change coefficients (β_4) show some sharp changes in rates of protection, with a negative coefficient indicating lower protection prior to the structural break. The most notable break is perhaps the increase in protection to rice in Indonesia from 2010, a change clearly visible in Figure 2.

Results for Wheat

Table 5 shows the tests for stationarity of the logs of domestic and border prices of wheat, and of cointegration between the two series.

Table 5. Dickey-Fuller tests for integration and cointegration of wheat price data

	Test of the null hypothesis that the series is nonstationary				
	Log of domestic price (p)	Log of intl. price (p^w)	1st diff. log domestic price (Δp)	1st diff. log intl. price (Δp^w)	Cointegration: log domestic & intl. prices
Argentina	3.91***	-4.10***	-6.40***	-6.62***	-5.38**
Australia	-3.97***	-4.21***	-6.44***	-6.21***	-5.62***
Bangladesh	-2.84	-3.37*	-3.24*	-4.15***	-8.06***
Brazil	-2.82	-2.91*	-6.36***	-5.76***	-5.24**
Canada	-3.25**	-3.39**	-5.95***	-6.02***	-3.62*
Switzerland	-2.69*	-3.07*	-4.05***	-5.56***	-5.15**[a]
Chile	-3.03	-2.92*	-6.77***	-6.25***	-5.15***
China	-2.71	-2.06	-3.89***	-5.13***	-4.35***
Colombia	-2.10	-3.26*	-3.78***	-4.77***	-3.25*
Ethiopia	-2.35	-2.31	-4.64***	-5.32***	-5.50***
European Union	-2.26	-3.05*	-7.03***	-6.83***	-3.57*
India	-1.08	-1.66	-4.31***	4.65***	-4.35**
Israel	-1.48	-2.55	-2.82	-3.98***	-4.11**
Japan	-1.78	-4.70***	4.39***	-7.42***	-5.59***[a]

	Test of the null hypothesis that the series is nonstationary				
	Log of domestic price (p)	Log of intl. price (p^w)	1st diff. log domestic price (Δp)	1st diff. log intl. price (Δp^w)	Cointegration: log domestic & intl. prices
Kazakhstan	-2.05	-3.07*	-1.88	-2.73	-3.55*
Kenya	-3.17**	-2.94*	-6.61***	-6.81***	-4.83***
Korea, Rep.	-1.10	-3.08*	-4.51***	-4.85***	-3.62**
Mexico	-2.94	-3.97***	-5.01***	-6.96***	-4.42**
Norway	-2.09	-3.28*	-3.57**	-6.19***	-5.79***[a]
New Zealand	-3.15*	-3.31**	-5.72***	-5.72***	-7.52***[a]
Pakistan	-2.54	-3.70**	-8.58***	-6.57***	-4.07**
Russia	-2.33	-2.74	-5.52***	-5.40***	-5.91***
Türkiye	-3.37**	-2.80	-5.65***	-6.08***	-4.96***
Tanzania	-1.86	-2.82	-3.45**	-5.90***	-3.66
Ukraine	-3.03	-3.37*	-6.01***	-6.73***	-5.90***
United States	-4.12***	-4.82***	-5.16***	-5.19***	-4.04*
South Africa	-2.33	-4.17***	-4.98***	-5.80***	-3.89**
Zambia	-1.88	-2.22	-3.87***	-4.48***	-3.07
Zimbabwe	-2.02	-2.83	-4.96***	-6.52***	-7.50***[a]

Notes: As for Table 3.

The results in Table 5 show that the ADF-GLS tests for most of the price variables in levels (p and p^w) do not reject the null hypothesis of a unit root. Much more important for our analysis is that these tests reject the null hypothesis of a unit root for the first-differenced variables in almost all cases, implying that these variables are stationary. Most of the few cases where it does not reject have extremely small samples, such as the 22 years (2000–2021) for Kazakhstan, likely accounting for the inability to reject the null hypothesis of a unit root. The results of the cointegration tests appear to confirm the existence of a cointegrating relationship in all but two cases—Tanzania and Zambia. As with rice, the failure to reject the null hypothesis of no cointegration may also reflect the low power of this type of test. Both integration and cointegration tests support the use of ECMs.

Table 6. Results for nonlinear restricted estimates of ECM models for wheat, $\beta_1 \equiv 1$

	δ	θ	β_0	β_2	β_3	β_4	β_5	R^2	Sample	RMSE
Argentina	0.91	-0.57	-0.21					0.83	60-21	0.19
	16.6	*-4.8*	*-4.9*							
Australia	0.93	-0.53	0.07	-0.001				0.96	61-21	0.04
	38.2	*-4.8*	*3.3*	*-2.4*						
Bangladesh	0.75	-0.87	0.12	-0.01				0.60	74-04	0.19
	3.0	*-4.5*	*1.5*	*-1.1*						
Brazil	0.75	-0.52	0.45	-0.02	0.0003			0.61	66-21	0.18
	7.1	*-4.5*	*3.1*	*-1.9*	*1.4*					
Canada	0.98	-0.24	0.05	-0.001				0.99	61-21	0.02
	90.6	*-2.9*	*2.5*	*-1.6*						
Switzerland	0.48	-0.39	0.75	-0.13		0.58		0.28	79-21	0.11
	3.6	*-2.5*	*3.1*	*-1.8*		*3.3*				
Chile	0.58	-0.59	0.06					0.53	60-21	0.20
	5.6	*-5.7*	*1.3*							
China	0.39	-0.29		0.01				0.33	81-21	0.12
	3.8	*-2.4*		*4.1*						
Colombia	0.33	-0.32	0.43	-0.01				0.42	60-24	0.09
	4.1	*-4.0*	*5.0*	*-1.7*						
Ethiopia	1.07	-0.69	0.12	-0.02	0.0007			0.90	81-19	0.16
	16.9	*-4.2*	*1.0*	*-1.71*	*2.0*					
European Union	0.86	-0.21	0.47	-0.01				0.80	57-21	0.17
	15.5	*-2.7*	*2.1*	*-1.4*						
India	0.13	-0.09		0.01				0.13	64-21	0.08
	1.7	*-1.8*		*1.2*						
Israel	0.71	-0.78	0.14					0.59	00-21	0.14
	5.0	*-3.6*	*3.7*							
Japan	0.94	-0.46				1.08	-0.92	0.54	55-21	0.09
	8.0	*-2.0*				*3.0*	*-7.6*			
Kazakhstan	0.73	-0.58		0.00				0.81	00-21	0.11
	8.6	*-3.1*		*1.1*						
Kenya	0.38	-0.38	0.12					0.32	56-20	0.18
	3.7	*-4.4*	*2.1*							
Korea, Rep.	0.49	-0.16	0.02					0.28	55-04	0.17
	2.9	*-2.7*	*4.4*							
Mexico	0.65	-0.57	0.11	0.026	-0.001			0.49	79-21	0.17
	5.1	*-4.0*	*0.7*	*1.7*	*-2.1*					
Norway	0.29	-0.28	0.65			0.69		0.31	79-21	0.12
	3.4	*-3.3*	*8.0*			*4.9*				
New Zealand	1.00	-0.34				0.10		0.99	61-21	0.01
	176.1	*-3.1*				*20.7*				
Pakistan	0.01	-0.38	0.12	-0.03	0.001			0.35	62-13	0.13
	0.13	*-4.7*	*0.8*	*-2.2*	*2.3*					

	δ	θ	β_0	β_2	β_3	β_4	β_5	R^2	Sample	RMSE
Russia	0.97	-0.73	-0.09					0.90	92-21	0.09
	13.9	*-5.7*	*-3.6*							
Türkiye	0.27	-0.19	0.01					0.25	61-21	0.15
	4.1	*-2.7*	*2.4*							
Tanzania	-0.05	-0.16		0.01				0.16	76-21	0.26
	-0.4	*-2.6*		*1.2*						
Ukraine	0.51	-0.67	0.12	-0.05	0.002			0.75	92-21	0.17
	7.5	*-6.6*	*0.8*	*-1.9*	*2.1*					
United States	0.83	-0.13	0.11					0.92	55-21	0.06
	25.8	*-1.8*	*1.9*							
South Africa	0.34	-0.18	0.41	-0.005				0.36	55-21	0.11
	5.9	*-2.8*	*2.5*	*-1.12*						
Zambia	-0.17	-0.14	-0.38					0.09	66-04	0.39
	-0.6	*-1.6*	*-0.8*							
Zimbabwe	0.70	-0.22	-0.45			0.84		0.33	55-20	0.27
	4.7	*-3.0*	*2.5*			*2.1*				

Notes: As for Table 4, plus a slope shifter β_5 designed to capture changes in δ associated with reforms. The Gregory-Hansen structural change dummies refer to: β_4 – Switzerland 2001, Japan 2007; β_5 – New Zealand 1988, Norway 1995, and Zimbabwe 1969.

As for rice, the model was first run with β_1 unrestricted but it proved impossible to include the equilibrium rate of protection or other variables without the estimated β_1 values deviating from unity, as is evident in the estimates presented in Table A.2. For this reason, the results in Table 6 incorporate the restriction that β_1=1. This restriction had generally small impacts on the adjustment coefficients of primary interest, while allowing estimation of the equilibrium rates and, where necessary, their changes over time.

Probably the most important feature of the restricted results in Table 6 is the sharp differences between countries in the values of the dynamic adjustment coefficients for changes in world prices, δ, and the EC term, θ. For traditional exporters such as Argentina, Australia, Canada, and the United States, the short-run adjustment coefficients are close to one. Major exporters Europe and Russia also have high short-run rates of price transmission. For these countries, the explanatory power of the simple model used is quite high, with R^2 statistics ranging from 0.79 in Europe to 0.99 in Canada. The result for Europe—an aggregate with membership expanding over time—is somewhat surprising given the high levels of support and apparent price insulation (with, for instance, negative support during the price boom of the mid-1970s). It seems more consistent with the low levels of price insulation prevailing since 2000. Japan is a particularly interesting case, where a structural change dummy for the price coefficient results in a dramatic increase in the price transmission coefficient after 2007, when price insulation for wheat practically disappeared.

Another set of countries has short-term price transmission rates below 0.9 but above 0.3, including Bangladesh (0.75), Brazil (0.75), Israel (0.71), Kazakhstan (0.73), Korea (0.49), Mexico (0.65), Ukraine (0.51), Switzerland (0.33) China (0.39), Kenya (0.38), Colombia (0.33), and South Africa (0.34). A third set of countries, including Türkiye (0.27), Norway (0.29), and India (0.13), has small but generally still statistically significant δ values. Unsurprisingly, the second and third groups of countries include many with relatively low incomes, where staples are likely to make up large shares of consumer expenditures and sharp price increases are likely to generate hostile responses. But these groups also include higher-income countries such as Norway. The low value of δ in India may well reflect the fact that roughly one-third of the sample years were classified as nontraded.

In a small group of countries, including Pakistan (0.01), Tanzania (-0.05), and Zambia (-0.17), the short-run price transmission coefficients were close to zero and not statistically significant. Insulating domestic prices from world price shocks seems to be a particularly strong goal for policymakers in these countries. In Pakistan, this result may reflect a strong aversion to adjusting prices for this important staple. For Tanzania and Zambia, wheat may simply not be important enough for policymakers to take the political heat associated with adjusting prices when world prices change. Tanzania and Zambia have very volatile idiosyncratic policy shock terms, measured by the Root Mean Square Error (RMSE) entries in the table.

Another major difference between economies is in the explanatory power of these simple models and the volatility of their residuals. Where the explanatory power of the model is high, as in Australia with an R^2 of 0.97, the RMSE of the residuals tends to be low (0.04 in the case of Australia). At the other extreme, the R^2 of 0.10 in Zambia is associated with an RMSE of 0.39.

The political-economy equilibrium rates of protection implied by the β_0, β_2, and β_3 coefficients make an important contribution to the model. When there is only a β_0 coefficient, this is generally in the same range as the mean tariff equivalent. Where present, the β_2 and β_3 coefficients capture important apparent changes in the equilibrium protection rate, such as the sharp decline in this rate for Switzerland and for Europe, with the intercept and trend coefficients at midpoint broadly consistent with the average tariff equivalent in Table 2.

6. Interpretation and Policy Implications

A key question for policymakers concerned about the effectiveness of the current trade regime and the desirability of reforms to international trade rules is the extent to which current policies magnify the volatility of world prices. A second question for national policymakers is how effective their own policies are in stabilizing domestic prices, given both price insulation and the idiosyncratic shocks to domestic prices. Each of these questions is addressed in turn, as a prelude to discussing potential policy options for dealing with these challenges.

Magnification of World Price Shocks

The short-run adjustment coefficients estimated by country provide a basis for estimating the short-run impacts of price-insulating policies on world prices. This uses a very simple global model with: (i) a global demand curve that is the horizontal sum of the individual country demand curves, and (ii) a short-run supply curve determined by countries' past planting decisions and random output shocks. In this situation, the impact of an output shock on the world price will depend on the elasticity of demand relative to the world price, with lower elasticities of demand requiring larger adjustments in world prices to restore equilibrium. Since changes in world prices must pass through the price transmission process into domestic prices before they can influence demand, this elasticity will depend on both the demand elasticity and the elasticity of price transmission. Using the notation of our model, the elasticity of demand with respect to the world price in each market, j, is:

$$\eta_j^w = \delta_j . \eta_j^d \tag{7}$$

where η_j^w is the elasticity of demand with respect to world prices; δ_j is the elasticity of price transmission; and η_j^d is the elasticity of demand with respect to domestic prices.

Unfortunately, we know little about the relevant elasticities of demand (η_i) for rice or wheat in individual markets, which depend on the weighted elasticities of demand for final use and for stocks. If we are willing, in the absence of better information, to assume this elasticity is roughly equal across markets, then we can follow the modeling approach of Anderson, Ivanic, and Martin (2014) and use a demand-weighted average of price transmission elasticities to assess the extent to which demand elasticities with respect to world prices are reduced and hence world price changes in response to shocks are magnified by imperfect price transmission. Tentative support for the equal elasticity approach is provided by Jensen and Anderson (2017), who used the information about national elasticities of demand in the GTAP (Global Trade Analysis Project) model and found magnification results for the 2008 price crisis comparable with those assuming equal elasticities of demand in all markets.

Using data on final demand for each grain and products from FAOSTAT food balance sheet data for 2020, we calculated the short-run weighted average elasticities of price transmission for rice and for wheat (Table 7). Because what matters for the price effect of any quantity shock is the inverse of the elasticity of demand, we also present the inverse of this price transmission elasticity, which is the price magnification effect of imperfect price transmission.

Table 7. Weighted average elasticities of price transmission and magnification for rice and wheat.

	Rice	Wheat
Elasticity of price transmission from international to domestic markets	0.53	0.53
Magnification of international price shocks due to incomplete price transmission	1.9	1.9

The price magnification factor for wheat is very close to the magnification of world price shocks estimated by Martin and Minot (2022) for the 2022 wheat price shock from the Ukraine crisis. The price magnification effect for rice is somewhat larger than the 1.52 estimated by Anderson, Ivanic, and Martin (2014) for the 2008 rice price crisis. Some differences between the results of these studies would be expected given their very different methodologies, with the earlier studies comparing actual changes in domestic prices relative to changes in world prices during specific episodes. It is reassuring that the broad orders of magnitude are similar.

The ECM methodology also suggests a potentially important additional channel for magnification of shocks to world prices through changes in incentives for stockholding. To see this, first consider an upward shock, such as when a harvest failure causes world prices to rise, and countries insulate to hold down their domestic prices. This lowers the protection rate relative to its political-economy equilibrium level. If no subsequent shock to world prices occurs, the EC coefficient shows how much policymakers will raise domestic prices in the next period. Stockholders will likely understand this aspect of their countries' policies and respond to the resulting incentive to increase storage, putting additional upward pressure on world markets by increasing demand for storage in periods of shortage. Given the much higher elasticity of demand for storage relative to final demand, this channel of effect may be an important—and to date unexplored—additional source of magnification of shocks on international prices. This effect, and the dynamics of agricultural trade policy identified in this paper, may contribute to the challenges faced by Gouel and Legrand (2022) in explaining correlations between output and prices for staple foods.

The econometric analysis also provides an opportunity to assess the relative importance of systematic and idiosyncratic impacts on the volatility of domestic prices. This analysis is relevant to the decisions of policymakers in individual countries, who must generally take the volatility of world prices as determined primarily by the actions of other countries. Because the variances of the systematic and idiosyncratic shocks are uncorrelated, the variance of domestic prices may be written as:

$$\sigma_p^2 = \sigma_s^2 + \sigma_i^2 \quad (8)$$

where σ_p^2 is the overall variance of domestic prices; σ_s^2 is the systematic component of the variance of domestic prices; and σ_i^2 is the idiosyncratic component of this variance. The systematic component of price volatility is determined by the volatility of world prices and the extent to which policymakers insulate their markets from this volatility.

The results from this decomposition for rice and wheat prices are presented in Figures 4 and 5 for each of the 29 economies for which we have estimates and, in the Average entry, for the simple average of these impacts. These decompositions rely, like the estimates of agricultural distortions on which they are based, on the assumption that domestic and foreign prices of the products used to generate the price comparisons would move identically in the absence of policy intervention. Given this, the variance of domestic prices in the absence of policy intervention is the variance of external prices. The variance of domestic prices if only the systematic elements of policy applied is obtained by dividing the Model Sum of Squares for the ECM by n-1, where n is the number of observations used in the model. The actual variance of domestic prices is obtained using the Total Sum of Squares from the regression model, divided by n-1.

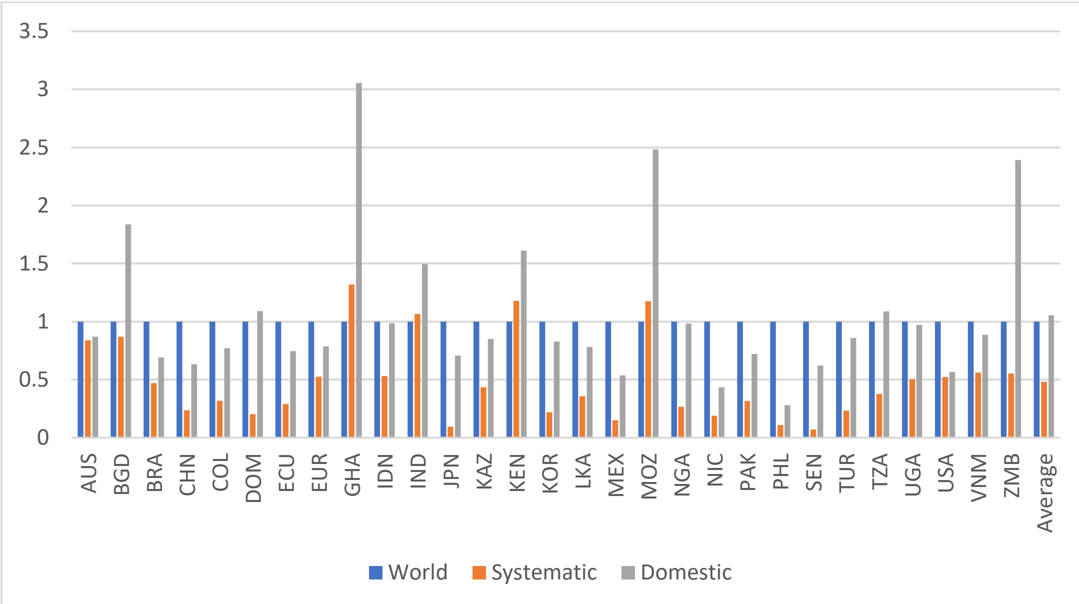
Because the variance of external prices differs substantially between countries—depending on influences like the specific products they trade and the frequency of changes in their direction of trade—all the variance measures for each country are deflated by the variance of its external reference price. This shows the price variance in each country as an index of one in the absence of policy intervention, generally falling to a level below one once systematic policy responses are considered, and then rising as the idiosyncratic elements are incorporated.

The variance decomposition results for rice reveal enormous differences in the impact and effectiveness of policies across countries (Figure 4). The simple average impacts in the final bars of the graph show a general pattern under which the systematic elements of policy reduce the volatility of domestic prices—by an average of 56 percent—while the idiosyncratic elements of policy undo this stabilizing impact and restore domestic price volatility to roughly its level in the absence of policy intervention. However, very large differences arise across countries in both the systematic and random effects of policy.

In most cases with low short-run price transmission for rice—such as China, Japan, and Korea—the systematic element of policy greatly reduces the volatility of domestic prices. Some countries with intermediate to high levels of short-run price transmission, such as India (0.73) and Mozambique (0.91), see modest increases in price volatility as their systematic policy adjustments interact with the dynamics of external prices, while other countries with high levels of price transmission, such as Australia, see modest reductions in price volatility.

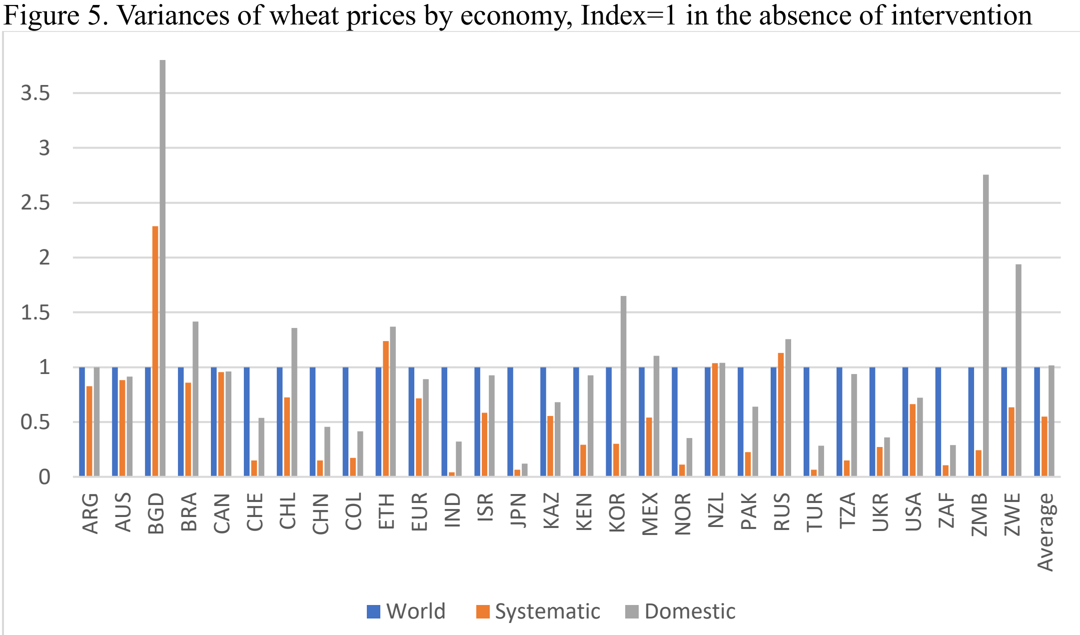
In most countries, idiosyncratic policy shocks eliminate most of the reduction in domestic rice price variability generated by stabilizing policies. Only two countries—the Philippines and Nicaragua— have combinations of systematic and idiosyncratic responses that leave the variance of their domestic prices less than one-half that of the external prices they face. However, many countries that are generally concerned about food price volatility, like Indonesia, India, Japan, and Korea, see large increases in rice price volatility from idiosyncratic shocks. In one set of countries—including Bangladesh, Ghana, Kenya, Mozambique, and Zambia—idiosyncratic policy shocks result in domestic price volatility much higher than international price volatility. Since the volatility of world prices has been greatly increased by the collective effects of price insulation, the Philippines might be the only country with lower domestic price volatility than in the global absence of price insulating policy. This result is much less satisfactory than the situation depicted by Anderson, Martin, and Ivanic (2017), where all countries with above-average insulation experience lower volatility than in the absence of insulation.

Figure 4. Variances of rice prices by economy, Index =1 in absence of intervention



Sources: Authors' estimates. Notes: Variances of the changes, rather than levels, of log prices are used because the level series are nonstationary. Averages are simple averages across countries. Countries are identified by ISO3 codes to save space.

Figure 5. Variances of wheat prices by economy, Index=1 in the absence of intervention



Sources and Notes: As for Figure 4.

Wheat price volatility reveals a broadly similar pattern (Figure 5). The simple average impacts are very similar, with systematic policy responses lowering the variance of domestic prices to 55 percent of its original level, while idiosyncratic policy shocks more than completely undo this, increasing domestic price volatility to 1.02 times the volatility of external prices. As with rice, enormous variation exists in the effectiveness of policies in reducing price volatility. Many countries, such as China, Switzerland, Colombia, Norway, and Türkiye, use systematic policy responses to greatly reduce wheat price volatility. For others, such as Argentina, Australia, and Zimbabwe, this policy approach is much less effective in reducing price volatility. For just a few countries, such as Bangladesh, Ethiopia and Russia, systematic policy responses slightly increase wheat price volatility because of interactions between the ECM parameters and the time series properties of world prices. A very important influence on the volatility of domestic prices, however, is idiosyncratic policy volatility, which, in many countries, overwhelms the stabilizing effects of systematic price insulating policies.

These results confirm the finding of Pieters and Swinnen (2016) that many countries experience much greater domestic price volatility than would be needed given their level of trade-distorting policies. The findings of this paper have important implications for policy reform at both national and global level, and for almost all the countries studied. Because the nature of these policy considerations differs so much, we deal with each domain separately.

Implications for National Policy

At the national level, an important implication of the findings of this paper is that there seems to be a strong case for reform in many countries. It seems difficult to justify policies that involve costly interventions in trade—and frequently also in storage—that generate greater price volatility than policies of nonintervention.

Potentially useful guides to policy reform might come from the literature on rules versus discretion in monetary policy. It is widely agreed that fully informed monetary policymakers can stabilize the domestic price level better by using discretionary policy interventions tailored to each specific situation than by using simple policy rules such as the Taylor rule (Taylor 2017). The critical issue is whether real-world policymakers are sufficiently well-informed about the path of the economy and the impacts of policy, and sufficiently motivated to pursue superior long-term policies, to be able to outperform simple policy rules (Kocherlakota 2016). Using an approach motivated by the rules versus discretion approach, Gouel, Gautam, and Martin (2016) concluded that simple rules-based policies for India's wheat market—one of the cases where our analysis suggests current policies have stabilized domestic relative to world markets—could achieve similar outcomes at lower cost.

Improving current policies depends a great deal on their effects. A first step is to examine whether past policies have resulted in the—all too familiar in this study—situation where, despite substantial effort to stabilize domestic prices, the volatility of domestic prices has ended up greater than that of external prices. If this is the case, then it would be possible to improve on the current situation using a less intrusive set of policies, such as a completely open trade regime, a regime with stable *ad valorem* tariffs, or a regime with modest price insulation. Even if this is not the case, it may be possible to design a set of policies that achieve current goals at lower cost. Gouel and Jean (2015) provide guidelines for reforms to trade and storage policies in small, open economies. Equation (6) in this paper provides a potential family of simple policy rules that could eliminate idiosyncratic price volatility.

If higher domestic than external price volatility is a problem, then a key question is the source of the excessive volatility. If, as seems frequently to have been the case in our study, it is high idiosyncratic volatility resulting from trade policy choices, then it is important to understand the source of that trade policy volatility. Is it simply a consequence of using rigid trade policies such as export bans that force domestic prices to adjust to accommodate all shocks to domestic output or demand? Or is it a consequence of policy settings that change for arbitrary political reasons? Or does it result from attempts to maintain unsustainable price policies that end in the collapse of the policy regime (Bardsley 1994)? Previous studies have documented the frequency with which interventionist trade regimes for African staple foods end up increasing price volatility (Jayne 2012; Minot 2014). This study suggests that policy failures of this type are much more widespread in world markets for rice and wheat.

Once the source of the problem has been identified, improved policies can be formulated. While the best policy package for each country will depend on the preferences of policymakers, some general principles might help in formulating policies. A first is that policies where tariff equivalents of policy vary over time are much more costly than those with stable protection regimes (Francois and Martin 2004). A second is that policies that require complete price stabilization—such as the use of an administered price supported by a variable import or export levy/subsidy—are likely to be very costly. A third is that price band-type policies that involve no intervention until a price trigger is reached and then strong enough intervention to support a rigid price have no clear welfare basis and are likely to be costly (Gouel and Jean 2015). Both these types of policy intervention also suffer from uncertainty about the sustainable setting of administered prices or price bands. Frequently, unsustainable settings for these prices result in the collapse of the regime, with considerable associated price instability and cost.

Implications for Global Policy

Policy formulation at the global level needs to take into consideration many of the factors influencing national policies. However, at the global level, it may be possible to more comprehensively address the collective action problems resulting from price-insulating policies. Current World Trade Organization (WTO) rules provide a strong basis for disciplining these measures. The general ban on quantitative restrictions in Article XI of the General Agreement on Tariffs and Trade (GATT) 1994 discourages measures such as export bans or import quotas, although with exceptions to prevent critical shortages of foodstuffs. Similarly, the WTO rules on market price support discourage the use of administered price supports that require policies such as variable levies to sustain them (World Trade Organization 1995, p. 63, para. 8).

A key question is whether WTO disciplines on food trade policies can be made more effective in reducing the adverse impacts of countries' policy decisions on their trading partners. Clearly, this would be much easier if countries examine their own policies and reform them along the lines suggested in the preceding subsection. Since export and import bans transform shocks to domestic output into shocks to domestic prices, they have no place in an efficient trade policy regime. If countries reformed their trade policy regimes to avoid quantitative restrictions and the type of trigger price policies that are not consistent with optimal stabilization policies, it should be much easier to secure agreement on strengthening rules to discourage those policies (Gouel, Gautam, and Martin 2016; Gouel and Jean 2015). Further analysis that examines how to improve on the current situation, where policies destabilize both domestic and world prices, seems likely to reveal more scope for improving both national and global trade policies.

7. Summary and Conclusions

The current literature provides models that explain the long-run level of protection to staple foods (for example, Grossman and Helpman 1994) and other models that explain why countries seek to resist the transmission of shocks to world food prices into their domestic markets (Giordani, Roche, and Ruta. 2016). The key goal of this paper was to develop a synthesis of these approaches that allows us to better represent countries' trade policies for staple foods. This approach allowed us to divide trade policy responses into (i) a systematic component designed to insulate domestic markets from food price shocks, and (ii) an idiosyncratic or random component resulting from policy shocks such as those created by changes in trade policy goals or output shocks when a country's market is fully insulated from world prices.

The data used for the analysis are prices of standardized products, at the same point in the marketing chain, to provide estimates of the rates of protection applying to rice and wheat. Because these price series were generally found to be nonstationary, we used ECMs to avoid spurious regression problems. Building on theoretical work by Nickell (1985), we showed that this approach represents the behavior of policymakers seeking to minimize the weighted sum of the political costs of sharply adjusting prices and of deviations from political-economy equilibria of the type considered by Grossman and Helpman (1994).

We concluded that systematic trade policy responses for rice and wheat substantially reduce the transmission of world price shocks into domestic markets—policies that would, on their own, substantially reduce the volatility of domestic prices relative to world prices. Unfortunately, a consequence of these policies is that they—by reducing the extent to which consumers and producers adjust to changed conditions— roughly double the standard deviation of world market prices, and hence quadruple their variances.

Perhaps more seriously, the idiosyncratic element of trade policy responses substantially reduces the stabilizing impacts of many countries' policies of price insulation. Improving domestic policies—perhaps by moving to simpler, rules-based approaches—could substantially reduce the volatility of domestic prices in many countries and their need for beggar-thy-neighbor policies of price insulation.

These conclusions suggest that much more effort is needed to design domestic policies that avoid the costly measures—such as quantitative restrictions and unsustainable changes in administered domestic prices— that result in high costs to the countries imposing them, increased domestic price volatility, and increased volatility in world market prices. National policy reforms that move away from these costly policies toward more resilient ones would help to both improve domestic market outcomes and reduce the obstacles to securing complementary reforms to global trade rules.

References

Anderson, K. 1995. "Lobbying Incentives and the Pattern of Protection in Rich and Poor Countries." *Economic Development and Cultural Change 43*(2): 401–423.

Anderson, K. 2009. *Distortions to Agricultural Incentives: A Global Perspective, 1955 to 2007*. Palgrave Macmillan and the World Bank.

Anderson, K., and S. Nelgen. 2012. "Trade Barrier Volatility and Agricultural Price Stabilization." *World Development 40*(1): 36–48.

Anderson, K., M. Ivanic, and W. J. Martin. 2014. "Food Price Spikes, Price Insulation, and Poverty." *The Economics of Food Price Volatility*: 311–314.

Anderson, K., W. J. Martin and M. Ivanic. 2017. "Food price changes, domestic price insulation, and poverty (when all policymakers want to be above average)" in Pingali, P. and G. Feder eds. *Agriculture and Rural Development in a Globalizing World*, Routledge, London.

Bagwell, K., and R. Staiger. 2002. *The Economics of the World Trading System*. MIT Press.

Bardsley, P. 1994. "The Collapse of the Australian Wool Reserve Price Scheme." *The Economic Journal* 104(426):1087-1105.

Bewley, R. 1979. "The Direct Estimation of the Equilibrium Response in a Linear Dynamic Model." *Economics Letters 3*(4): 357–361.

Burgess, R., and D. Donaldson. 2010. "Can Openness Mitigate the Effects of Weather Shocks? Evidence from India's Famine Era." *American Economic Review 100*: 449–453.

Campos, J., N. R. Ericsson, and D. Hendry. 2005. "General-to-specific Modeling: An Overview and Selected Bibliography." *International Finance Discussion Paper* 2005(838). Board of Governors of the Federal Reserve System. https://doi.org/10.17016/ifdp.2005.838

De Gorter, H., and J. Swinnen. 2002. "Political Economy of Agricultural Policy." In *Handbook of Agricultural Economics, Vol. 2,* edited by B. Gardner and G. Rausser, 1893–1943. Elsevier.

Elliott, G., T. J. Rothenberg, and J. H. Stock. 1996. "Efficient Tests for an Autoregressive Unit Root." *Econometrica 64*(4): 813–836.

Elton, E. J., and M. J. Gruber. 1997. "Modern Portfolio Theory, 1950 to Date." *Journal of Banking and Finance 21*: 1743–1759.

Engle, R., and C. Granger. 1987. "Co-integration and Error-correction: Representation, Estimation, and Testing." *Econometrica 55*(2): 251–276.

Francois, J. F., and W. Martin. 2004. "Commercial Policy Variability, Bindings, and Market Access." *European Economic Review 48*: 665–679.

Freund, C., and C. Özden. 2008. “Trade Policy and Loss Aversion.” *American Economic Review 98*(4): 1675–1691.

Gandelman, N., and R. Hernández-Murillo. 2015. “Risk Aversion at the Country Level.” *Federal Reserve Bank of St. Louis Review 97*(1): 53–66.

Gawande, K., and B. Hoekman. 2006. “Lobbying and Agricultural Trade Policy in the United States.” *International Organization* 60(3): 527–561.

Giordani, P. E., N. Rocha, and M. Ruta. 2016. “Food Prices and the Multiplier Effect of Trade Policy.” *Journal of International Economics 101*: 102–122.

Gouel, C., M. Gautam, and W. J. Martin. 2016. “Managing Food Price Volatility in a Large Open Country: The Case of Wheat in India.” *Oxford Economic Papers 68*(3): 811–835. https://doi.org/10.1093/oep/gpv089

Gouel, C., and S. Jean. 2015. “Optimal Food Price Stabilization in a Small Open Developing Country.” *World Bank Economic Review 29*(1): 72–101. https://doi.org/10.1093/wber/lht018

Gouel, C. and N. Legrand. 2022. *The Role of Storage in Commodity Markets: Indirect Inference*. CEPII Working Paper 2022-04. www.cepii.fr

Gregory, A. W., and B. E. Hansen. 1996. “Residual-based Tests for Cointegration in Models with Regime Shifts.” *Oxford Bulletin of Economics and Statistics 58*(3): 555–560.

Grossman, G. M., and E. Helpman. 1994. “Protection for Sale.” *American Economic Review 84*(4): 833–850.

Hoffman, C., L. Kastens, A. Portugal-Perez and S. von Cramon-Taubadel. (2024), ‘Trade Policies and the Transmission of International to Domestic Prices’ Mimeo, University of Gottingen and the World Bank.

Inder, B. 1993. “Estimating Long Run Relationships in Economics.” *Journal of Econometrics 57*: 53–68.

Jayne, T. S. 2012. “Managing Food Price Instability in East and Southern Africa.” *Global Food Security* 1(2): 143–149. https://doi.org/10.1016/j.gfs.2012.10.002

Jensen, H. G., and K. Anderson. 2017. “Grain Price Spikes and Beggar-Thy-Neighbor Policy Responses: A Global Economywide Analysis.” *World Bank Economic Review 31*(1): 158–175. https://doi.org/10.1093/wber/lhv047

Johansen, S. 1991. “Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models.” *Econometrica 59*(6): 1551–1580.

Kocherlakota, N. 2016. “Rules Versus Discretion: A Reconsideration.” *Brookings Papers on Economic Activity* 2016(2): 1–55.

Laborde, T. Lallemant, F. Majeed, A. Mamun, W. Martin and S. Tokgoz. 2024. *Introducing the Updated Ag-Incentives Consortium and Database*, www.agincentives.org

Martin, W., and K. Anderson. 2011. "Export Restrictions and Price Insulation during Commodity Price Booms." *American Journal of Agricultural Economics 94*(2): 422427.

Martin, W., and N. Minot. 2022. "The Impacts of Price Insulation on World Wheat Markets During the 2022 Food Price Crisis." *Australian Journal of Agricultural and Resource Economics 66*(4): 753–774.

Minot, N. 2014. "Food Price Volatility in Sub-Saharan Africa: Has It Really Increased?" *Food Policy 45*: 45–56. https://doi.org/10.1016/j.foodpol.2013.12.008

Mundlak, Y., and D. F. Larson. 1992. "On the Transmission of World Agricultural Prices." *World Bank Economic Review 6*(3): 399–422.

Nickell, S. 1985. "Error Correction, Partial Adjustment and All That: An Expository Note." *Oxford Bulletin of Economics and Statistics 47*(2): 119–130.

OECD. 2016. "The PSE Manual." Mimeo, *Trade and Agriculture Directorate,* OECD, Paris.

Pieters, H., and J. Swinnen. 2016. "Trading-off Volatility and Distortions? Food Policy During Price Spikes." *Food Policy 61*: 27–39. https://doi.org/10.1016/j.foodpol.2016.01.004

Sampson, G. P., and R. H. Snape. 1980. "Effects of the EEC's Variable Import Levies." *Journal of Political Economy 88*(5): 2036–1040.

Schaffer, M. 2022. *EGRANGER: Stata Module to Perform Engle-Granger Cointegration Tests and 2-step ECM Estimation* (Revised). Statistical Software Components S457210, Boston College Department of Economics.

Swinnen, J. F. M. 1994. "A Positive Theory of Agricultural Protection." *American Journal of Agricultural Economics 76*(1): 1–14. https://doi.org/10.2307/1243915

Taylor, J. B. 2017. *Rules versus Discretion: Assessing the Debate over the Conduct of Monetary Policy*, Working Paper 24149, National Bureau of Economic Research, Cambridge MA.

Timmer, C. P. 2000. "The Macro Dimensions of Food Security: Economic Growth, Equitable Distribution, and Food Price Stability." *Food Policy 25*: 283–295.

Timmer, C. P. 2010. "Reflections on Food Crises Past." *Food Policy 35*(1): 1–11. https://doi.org/10.1016/j.foodpol.2009.09.002

Tovar, P. 2009. "The Effects of Loss Aversion on Trade Policy: Theory and Evidence." *Journal of International Economics 78*(1): 154–167. https://doi.org/10.1016/j.jinteco.2009.01.012

Tversky, A., and D. Kahneman. 1991. "Loss Aversion in Riskless Choice: A Reference-Dependent Model." *The Quarterly Journal of Economics* 106(4).

von Cramon-Taubadel, S., and B. K. Goodwin. 2021. "Price Transmission in Agricultural Markets." *Annual Review of Resource Economics 13*: 15.1–15.20.

Wickens, M. R., and T. S. Breusch. 1988. “Dynamic Specification, the Long-Run and the Estimation of Transformed Regression Models.” *The Economic Journal 98*(390): 189–205.

World Trade Organization. 1995. *The Results of the Uruguay Round of Multilateral Trade Negotiations: The Legal Texts*. World Trade Organization.

Wright, B., and J. Williams. 1990. “The Behavior of Markets for Storable Commodities.” In *Paper presented at the 34th Annual Conference of the Australian Agricultural Economics Society*. AgEcon Search.

Appendix Tables

Table A.1 Results for nonlinear unrestricted estimates of ECM models for rice

	δ	θ	β_0	β_1	β_2	R^2	Sample	RMSE
AUS	0.92	-0.12		1.01		0.96	61-21	0.05
	37.1	*-1.8*		*98.2*				
AUS-2	0.92	-0.41	0.64	0.89		0.97	61-21	0.05
	40.1	*-3.9*	*5.6*	*38.8*				
BGD	0.01	-0.41	5.62	-0.10		0.23	74-19	0.23
	0.2	*-3.3*	*14.7*	*-1.3*				
BRA	0.71	-0.56		1.01		0.66	73-19	0.18
	9.3	*-4.3*		*113.8*				
BRA-2	0.67	-0.73	0.76	0.83	0.0069	0.70	73-19	0.17
	8.6	-5.1	1.6	8.7	2.6			
CHN	0.39	-0.10		1.04		0.31	81-21	0.1267
	3.4	*-1.8*		*21.0*				
CHN-2	0.40	-0.34		0.89	0.027	0.43	81-21	0.12
	3.7	*-3.3*		*37.7*	5.2			
COL	0.52	-0.19		0.98	0.017	0.41	60-20	0.14
	5.7	*-2.9*		*21.7*	*2.6*			
DOM	0.40	-0.18		1.09		0.19	55-19	0.20
	3.2	*-2.5*		*40.6*				
DOM-2	0.34	-0.36	3.34	0.32	0.019	0.27	55-19	0.19
	2.8	*-3.7*	*3.6*	*1.6*	3.2			
ECU	0.26	-0.41		1.00	0.007	0.38	66-16	0.16
	2.5	*-4.7*		*40.1*	*1.6*			
EUR	0.73	-0.10		1.04		0.66	57-21	0.20
	10.8	-1.7		24.2				
GHA	0.93	-0.28		0.96		0.36	55-18	0.33
	5.2	*-3.0*		*36.0*				
IDN	0.60	-0.33		0.98	0.011	0.43	75-21	0.15
	5.1	*-3.1*		*37.8*	*2.1*			
IND	0.80	-0.30		0.94		0.61	65-21	0.21
	7.3	*-3.1*		*52.8*				
IND-2	0.63	-0.78		0.85	0.016	0.71	65-21	0.13
	6.1	*-5.6*		*59.6*	7.1			
JPN	0.14	-0.04		1.65	-0.051	0.16	55-21	0.13
	1.4	*-1.1*		*3.3*	-0.8			
KAZ	0.28	-0.51		1.06	-0.061	0.36	00-21	0.16
	1.5	*-3.0*		*35.8*	-4.8			
KAZ-2	0.16	-0.88	2.57	0.57	-0.04	0.53	00-21	0.14
	0.9	*-4.1*	*3.4*	*3.9*	*-4.3*			
KEN	-0.19	-0.62		1.16	-0.061	0.72	00-21	0.16
	-0.6	*-2.5*		*27.8*	*-2.1*			

	δ	θ	β_0	β_1	β_2	R^2	Sample	RMSE
KOR	0.02	-0.09		1.21		0.17	55-21	0.16
	0.2	*-2.7*		*24.9*				
LKA	0.58	-0.39		1.01		0.43	55-14	0.17
	6.4	*-3.9*		*83.8*				
MEX	0.33	-0.35		1.01		0.28	79-21	0.15
	3.1	*-3.2*		*80.9*				
MOZ	0.94	-0.23		0.94		0.46	76-19	0.45
	5.1	*-2.5*		*16.5*				
NGA	0.27	-0.41		1.04		0.24	61-15	0.23
	2.1	*-3.9*		*72.7*				
NIC	0.32	-0.14		1.07		0.23	91-17	0.13
	2.5	*-1.5*		*32.4*				
NIC-2	0.30	-0.88	3.06	0.50	0.014	0.56	91-17	0.10
	2.9	*-4.3*	*5.0*	*5.1*	*3.4*			
PAK	0.45	-0.21		0.97		0.39	62-13	0.19
	4.5	*-2.7*		*33.4*				
PAK-2	0.45	-0.39		0.86	0.017	0.46	62-13	0.18
	4.8	*-3.7*		*27.3*	*3.4*			
PHL	0.26	-0.07		1.11		0.32	62-21	0.32
	4.4	*-1.9*		*20.9*				
SEN	0.18	-0.17		1.05		*0.12*	61-20	0.17
	1.7	*-2.4*		*37.8*				
TUR	-0.06	-0.12		1.07		0.10	85-03	0.28
	-0.3	*-1.1*		*10.3*				
TZA	0.43	-0.21		0.96		0.27	76-21	0.25
	3.1	*-2.6*		*28.9*				
UGA	0.71	-0.44		1.05		0.52	61-18	0.25
	7.4	*-3.8*		*74.1*				
USA	0.73	-0.10		1.03		0.92	55-21	0.06
	27.4	*-2.0*		*61.1*				
VNM	0.69	-0.42		1.00		0.58	86-21	0.15
	6.3	*-3.0*		*83.8*				
ZMB	0.50	-0.30		0.95		0.23	68-20	0.35
	2.6	*-3.0*		*31.4*				

Notes: The δ, θ, β_0, and $\beta 1$ coefficients correspond to equation (5). Coefficient β_2 is for a time trend on the long-run equilibrium tariff rate. t-statistics for the null hypothesis that the coefficient equals zero are presented in italics below the coefficient estimates. Models with coefficients other than δ, θ, and β_1 were chosen on the basis of their overall explanatory power and the significance of individual coefficients.

Table A.2. Results for nonlinear estimation of ECM models for wheat

	α	δ	θ	β_0	β_1	β_2	β_3	R^2	Sample	RMSE
ARG		0.90	-0.59		0.95			0.83	60-21	0.19
		16.7	*-5.0*		*105.7*					
AUS	0.21	0.91	-0.67		0.94			0.97	61-21	0.04
	3.1	*37.1*	*-5.3*		*63.5*					
BGD		0.76	-0.80		1.01			0.58	74-04	0.19
		3.0	*-4.4*		*117.5*					
BRA		0.79	-0.39		1.02			0.57	66-21	0.19
		7.3	*-3.8*		*80.3*					
CAN		0.98	-0.18		1.00			0.99	61-21	0.02
		89.6	*-2.4*		*323.2*					
CHE	1.39	0.42	-0.51		0.59		0.12	0.29	79-21	0.11
	2.2	*3.3*	*-2.8*		*4.1*		*5.5*			
CHL		0.59	-0.59		1.01			0.54	60-21	0.20
		5.6	*-5.7*		*117.6*					
CHN		0.39	-0.23		1.06			0.31	81-21	0.12
		3.8	*-2.1*		*55.8*					
COL		0.33	-0.26		1.06			0.36	60-24	0.09
		4.0	*-3.4*		*99.0*					
ETH		1.08	-0.55		1.00			0.89	81-19	0.16
		16.5	*-3.5*		*145.0*					
EUR		0.86	-0.18		1.04			0.79	57-21	0.18
		15.3	*-2.3*		*45.7*					
IND		0.14	-0.11		1.03			0.13	64-21	0.08
		1.9	*-2.0*		*42.8*					
ISR		0.72	-0.75		1.03			0.66	00-21	0.14
		5.0	*-3.5*		*133.2*					
JPN		0.08	-0.04		1.19			0.10	55-21	0.12
		2.0	*-1.5*		*11.1*					
KAZ		0.73	-0.54		1.01			0.81	00-21	0.11
		8.3	*-2.9*		*101.8*					
KEN		0.39	-0.38		1.02			0.32	56-20	0.18
		3.8	*-4.4*		*87.5*					
KOR		0.48	-0.09		1.18			0.25	55-04	0.17
		2.8	*-2.0*		*15.1*					
MEX		0.48	-0.09		1.18			0.25	79-21	0.17
		2.8	*-2.0*		*15.1*					
NOR		0.20	-0.25	4.07	0.36			0.23	79-21	0.13
		2.3	*-2.4*	*4.8*	*2.0*					
NZL		1.00	-0.02		0.99			0.99	61-21	0.01
		165.2	*-0.9*		*37.6*					
PAK		0.00	-0.33		0.97			0.30	62-13	0.14
		-0.04	*-4.2*		*81.0*					

	α	δ	θ	β_0	β_1	β_2	β_3	R^2	Sample	RMSE
RUS		0.96	-0.70		0.98			0.91	92-21	0.09
		13.7	*-5.6*		*186.8*					
TUR		0.26	-0.39	2.16	0.48	0.02		0.32	61-21	0.14
		3.9	*-3.7*	*3.4*	*3.5*	*5.0*				
TZA		-0.05	-0.14		1.05			0.15	76-21	0.26
		-0.4	*-2.3*		*18.2*					
UKR		0.49	-0.66		0.98			0.72	92-21	0.18
		7.0	*-6.2*		*91.6*					
USA		0.83	-0.11		1.02			0.93	55-21	0.06
		25.6	*-1.6*		*73.9*					
ZAF		0.33	-0.20	1.11	0.83			0.37	55-21	0.11
		5.8	*-2.9*	*1.9*	*7.1*					
ZMB		-0.18	-0.14		0.93			0.10	66-04	0.39
		-0.7	*-1.6*		*10.7*					
ZWE		0.66	-0.16		0.94			0.33	55-20	0.28
		4.5	*-2.4*		*22.4*					

Notes: The α, δ, θ, β_0, and β_1 coefficients correspond to equation (15). Coefficient β_2 is for a deterministic trend in the equilibrium protection rate and β_3 in these models is a slope-shifter for the coefficient on world prices associated with policy reforms in Switzerland. The t-statistics for the null hypothesis that the coefficient equals zero are presented in italics below the coefficient estimates.

Appendix

Implications of the Loss Aversion Model for Price Insulation Coefficients

To understand the striking Freund-Özden result of full price insulation, it is helpful to consider a case with two interest groups—consumers and producers—where world prices rise from an initial political-economy equilibrium in which the world price is the reservation price for consumers. Because consumers feel the resulting loss much more strongly than producers feel their gain, net consumers apply much more pressure for relief from the price rise than producers do to retain their gains.[5]

Since the political-economy welfare function is at a maximum in the initial equilibrium, the marginal political cost of a small deviation from the political-economy equilibrium is zero and policymakers will respond to a small rise in the world price by reducing protection enough to completely offset the impact of the world price rise on the domestic price. However, because the political-economy welfare function is convex in protection rates, the marginal cost of deviating from the equilibrium rate of protection rises as this deviation increases. At some point, the marginal cost of deviating from the political-economy equilibrium exceeds the weight on consumer losses relative to gains, so policymakers cease reducing the rate of protection and allow domestic prices to rise.

Tversky and Kahneman (1991) do not specify how reference prices are determined in the loss aversion model, and we cannot be sure how they might be determined in this case. If world prices for wheat and rice follow a random walk, then their price in the last period is the best predictor of their price this period, making it a plausible candidate for the reference price. If they are characterized by the second-order autoregressive process described by Nickell (1985, p. 124), then this would be the case when the price was the same in the past two periods. However, studies such as Giordani, Roche, and Ruta (2016) based on behavioral models often use reservation prices, such as three- or five-year moving averages, rather than the price in the previous year. Because world markets for rice and wheat do not appear to be explosive, what matters for both theory and market behavior is the value of the δ coefficient in equation (5).

[5] Tovar (2009) estimates the relevant parameters and finds the weight on losses relative to gains to be very large.

Figure A.1. Relationship between product prices and protection

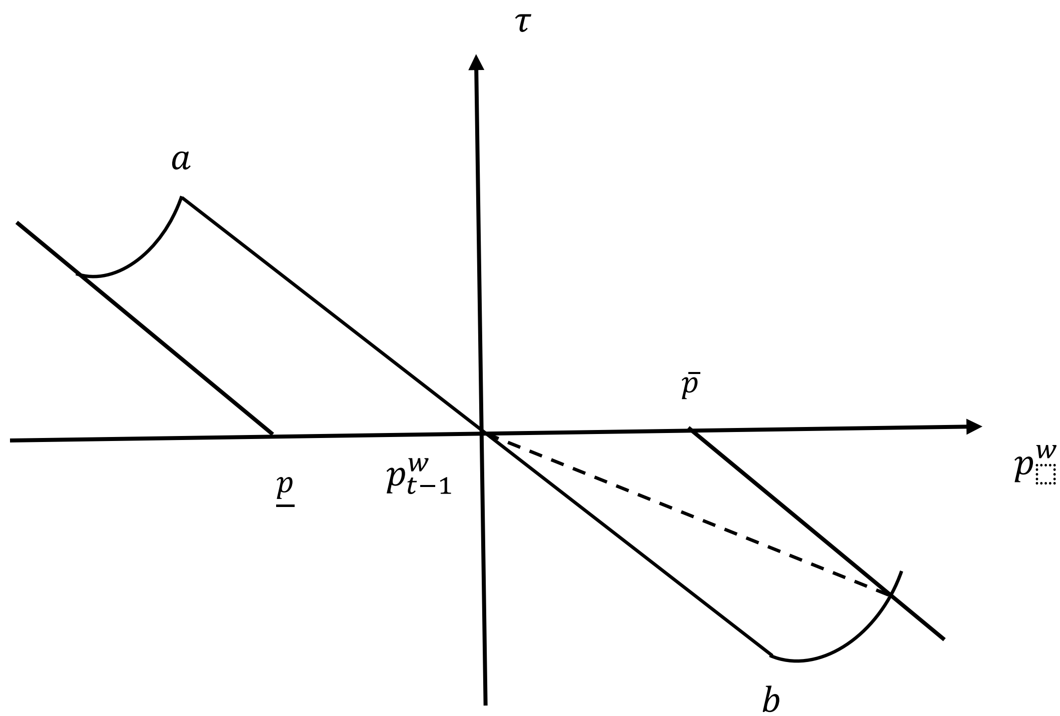


Figure A.1, drawing on Figure 4 of Freund and Özden (2008) and Figure 3 of Giordani, Roche, and Ruta (2016), is useful in identifying at least two ways in which a finding that $\delta > 0$, and hence insulation is incomplete, might be consistent with behavioral theory. The first involves shocks that change prices beyond the range of full price insulation. The second involves reference prices that differ from the world price in the previous period.

In Figure A.1, the world price is shown on the horizontal axis and the rate of protection on the vertical axis. The rate of protection begins at zero when the world price is p_{t-1}^{w}. If the reference price equals p_{t-1}^{w} and the world price rises to the level consistent with point *b*, then the rate of protection is expected to fall one for one with the world price increase, because the elasticity of price insulation, $(\delta - 1)$ in equation (6), equals -1 or, equivalently, the elasticity of price transmission (δ) is zero. If the world price rises further, Freund and Özden show that the protection rate declines in absolute value because the marginal cost of protection exceeds the marginal gain from loss aversion. In Figure A.1, this range of declining protection is shown by the upward sloping curve beginning at point *b*. If typical shocks to world prices are large enough to take the tariff rate into this range, then estimated price transmission coefficients might lie in the range $0<\delta<1$. Given the repeated findings that price-insulating behavior increases the volatility of world prices (Martin and Anderson 2011; Giordani, Roche, and Ruta 2016; this paper), policy itself is likely an important contributor to world price shocks being large enough to exceed the range of full price insulation.

Another theory-consistent potential explanation for less than full price insulation would be reference prices that differ from p_{t-1}^{w}. If, for instance, the reservation price for consumers is $\overline{p}$, then the compensating reduction in protection associated with a world price increase does not begin until the world price reaches $\overline{p}$. As shown by the dashed line in Figure A.1, either of these differences could explain less than full compensation of consumers following a rise in the world price, even given the strong internal validity of the theory. The same logic would apply where the world price falls and the reference price for producers is $\underline{p}$.

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