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SPECIAL ISSUE
Understanding International Commodity Price Fluctuations

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Implications of domestic price insulation for global food price behavior

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A B S T R A C T

Rapid changes in global food prices in recent years are widely viewed as a serious threat to global development. While various sources of price instability in agriculture have been identified, little attention appears to have been given to the importance of changes in trade policies that insulate domestic prices from world markets as a source of volatility in world prices. A contribution of this paper is to show that these interventions are dynamically more complex than simple proportional insulation. Insulation against an initial price increase in world prices increases the magnitude of that increase, while subsequent adjustments to the level of protection change the fundamental nature of price volatility. We find such policies are widespread and increase the volatility of world prices while not reducing the volatility of domestic prices because of the collective action problem involved in this form of policy intervention.

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Sudden and major changes in global prices of key food commodities in the recent years have raised serious concerns among global policy makers. Food price spikes represent a direct threat to global development by adversely affecting poor households with large food expenditure shares. The International Monetary Fund and the World Bank (World Bank, 2011) argue that sharp increases in food prices raise poverty, lower the level and quality of nutrition, and reduce the consumption other essential non-food services such as healthcare and education, all of which negatively affect future growth. Because of the important developmental consequences of food price volatility, it is important

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to understand its sources and, even more importantly, identify feasible policy options for dealing with its adverse consequences.

Much of the attention on global price volatility in recent years has focussed on the effects of weather shocks on output, and the interaction between these shocks and commodity stock levels (Cafiero et al., 2011). While there is evidence that the frequency of many weather shocks, such as extreme temperatures and intense precipitation events, has increased over time, these changes have been very gradual (Easterling et al., 2000) making them an unlikely cause of the recent increase in volatility. Further, there is evidence that, despite the apparent increase in the volatility of climate outcomes, the volatility of yields for key staples has been decreasing (Gollin, 2006). Other sources of volatility that have been emphasized include government programs such as biofuel mandates, which link the price of grains to those of volatile fuel prices; however the evidence regarding the impact of such mandates on global price volatility remain inconclusive (Zhang et al., 2010). Another frequently-cited potential source of volatility is the financialization of commodity markets and the potential for instability resulting from this linkage, but careful analysis raises serious doubts about whether this factor is an important source of increased price volatility Irwin et al. (2009).

Another important feature of commodity price behavior is strong positive autocorrelation. This was highlighted in the original formulation of the storage model in Deaton and Laroque (1992), and the apparent inability of this model to generate the observed level of autocorrelation was a major cause of its rejection in Deaton and Laroque (1996, 1995). However, Cafiero, et al. (2011) resuscitated this model after finding that it could generate both price volatility and strong patterns of positive autocorrelation for seven key commodities—coffee, copper, jute, maize, palm oil, sugar, and tin—but not for wheat or rice.

Whatever the fundamental sources of shocks to global food prices, there is considerable evidence that insulating trade policy can play a major role in increasing the volatility of these prices. Martin and Anderson (2012) conclude that almost half of the increase in world rice prices between 2005 and 2008 could be attributed to the effects of the policies that countries used to insulate themselves from increases in world prices. When there is an adverse shock that raises world prices, the use of insulating trade policies raises world prices by restricting export supplies from exporting countries, and by increasing import demand from importing countries. If all countries respond in the same way, this policy response is completely ineffective—the reductions in domestic prices resulting from insulation are exactly offset by increases in the world price. While countries do not, in fact, respond in exactly the same way to price rises, Anderson, Ivanic and Martin (2013) found that this behavior was collectively ineffective if its goal was to protect poor people from the adverse impacts of higher food prices. If changes in trade policies involve intertemporal dynamics, rather than simply attempting to reduce the volatility of domestic prices relative to world prices, they may also influence the intertemporal correlations of commodity prices, as well as changing their volatility.

Anderson and Nelgen (2011) show that protection changes reduce the extent to which changes in world prices result in changes in domestic prices. They also show that the relationship between domestic and international prices is not a consistent one, with insulation being greater when prices first increase, and diminishing when price increases are sustained. This pattern of insulating behavior has potentially important implications not just for the volatility of world prices but also for their intertemporal dynamics of these prices.

In this paper, we examine the role of trade protection policies in destabilizing global food markets in more detail. Taking advantage of a large dataset of historical agricultural protection (Anderson, 2009) in a number of regional markets and for a number of important food commodities, we estimate the price insulation parameters that represent the policies countries use to shield their markets from changes in global prices. Then, in the second part of the analysis, we use a global general equilibrium model to calculate the impact of the observed insulation policies on the volatility of global prices driven by the observed volatility of regional supply and draw conclusions.

1. Methodology

To address the questions raised in this work, we use two distinct methodologies. We first use an econometric approach to uncover key parameters describing the relationship between agricultural

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protection and global prices. In the second part of the analysis, we employ a global model to simulate stochastically the implications of this insulating behavior for the volatility of international and domestic prices.

1.1. Estimation of insulation parameters

In our econometric analysis, we explore the relationship between agricultural prices at the border, P^b, for key agricultural commodities and protection (T) defined as the ratio between domestic prices (P) and border prices as $T = P/P^b$. We follow Nickell (1985) in assuming that policy makers face quadratic costs of deviating from desired levels of protection and of changing the domestic price from its previous level. Such a model can be represented by the Error Correction Model (ECM) widely used to represent imperfect transmission of prices into domestic markets (see, for example, Greb, et al., 2012). The ECM also has the desirable feature of providing valid statistical tests for coefficient estimates even in cases where the data series are nonstationary.

Our objective is somewhat different from most standard applications of the ECM to imperfect transmission of border price changes into domestic prices. Much of the focus of studies of price transmission is on the short term responses, where prices do not respond immediately to changes in world price. This tends to lead to an emphasis on use of the highest-available frequency data, such as monthly price series, in these analyses. Our focus is on adjustments in policies such as tariffs or the use of export restrictions, in response to changes in world prices and to deviations of initial protection from a target level of protection perhaps determined by a model such as that proposed by Grossman and Helpman (1994). Many adjustments in protection rates, such as those resulting from imposition of export bans or export quotas, or temporary reductions in tariff rates, are frequently introduced for extended periods, such as a season at a time, with a view to ensuring food availability over that period. For this reason, and because the data we need on changes in border and domestic prices are frequently available only on an annual basis, we use annual data in this analysis.

We begin with a standard ECM for the response of domestic prices to changes in world prices.

$$\Delta p_t = (\alpha + 1)\cdot\Delta p_t^b + \beta\left[p_{t-1} - \gamma p_{t-1}^b - \theta\right] + \varepsilon \quad (1)$$

where p is the log of the domestic price, P; and p^b is the log of the border price. The coefficient $(\alpha + 1)$ is the price transmission coefficient, which indicates the proportion of any given change in border prices that is transmitted into domestic prices. The error-correction coefficient β reflects the extent to which domestic prices are adjusted in response to past differences between their actual and desired levels. As noted above, the partial adjustment coefficient takes into account the costs (assumed to be quadratic) resulting from changes in domestic prices and the error-correction coefficient takes into account the underlying quadratic costs of being away from the tariff rate that maximizes policy makers' political-economy objectives.

For our purposes, it is convenient to rearrange this model by subtracting $(p_t^b - p_{t-1}^b)$ from both sides, and rearranging the term in square brackets by adding and subtracting p_{t-1}^b to obtain:

$$\Delta \tau_t = \alpha\Delta p_t^b + \beta\left[\tau_{t-1} - (\gamma - 1)p_{t-1}^b - \theta\right] + \varepsilon \quad (2)$$

where τ is the log of the protection rate (T).

If we follow the logic of standard models of protection policy, such as Grossman and Helpman (1994), the long run rate of protection does not depend on the world price, but rather on factors such as the elasticity of demand for the product and the share of production supplied domestically. Even when augmented for the myriad political-economy factors (see Anderson, Rausser and Swinnen, 2013) that determine rates of assistance to agriculture, it is not clear that the long-run level of the world price of the particular commodity should play a role in its long-run level of assistance. We therefore proceeded by testing the performance of the model with and without this variable. Given disappointing results in the model including the border price, particularly in terms of testing for cointegration, we excluded this variable from the final analysis.

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Given the marked changes in rates of assistance to agriculture over time—which appear to reflect factors such as changing views on agricultural trade policy, the success of international trade negotiations and other broad factors, we added a variable for the five-year average global average rate of assistance to agricultural commodities other than the one under review, τ^*, to the estimating equation. In addition, we add a time trend to each equation to allow for the diverse forces that cause long-run changes in protection rates to agriculture within each country. The resulting equation is:

$$\Delta\tau_t = \alpha\Delta p_t^b + \beta\left[\tau_{t-1} - \gamma\tau_{t-1}^* - \delta t - \theta\right] + \varepsilon \quad (3)$$

where τ is the log of the power of taiff, p^b is the log of the border price, τ^* is a five-year average of the log of the global level of trade-weighted protection for crops other than the one being considered, t the time variable (in years).

As shown above, this equation is very similar to that in error correction models widely used in the literature on price transmission (see Greb et al. (2012)) except that our focus is on the determination of the rate of protection, rather than on the domestic price level at a given level of protection. We also omit the equations for determination of border prices that are part of the Greb et al. Vector Error Correction Model. Although there is clearly feedback from protection rates in a few large countries to the world price, this will not be important for most countries in our sample and we are concerned that estimating the coefficients on this relationship simultaneously with those on the determination of protection may bias our estimates of the coefficients of primary interest.

Before proceeding with the estimation of Equation (3), we test for stationarity of the τ, p^b and τ^* variables, and for cointegration of the long-run relationship in square brackets. The first step in this analysis uses an Augmented Dickey Fuller (ADF) test to check for integration of the variables in the levels. We then test to see whether these series are stationary after differencing. We then use Ordinary Least Squares to estimate the relationship between the hypothesized long-run explanatory variables and use an ADF test on the residuals from this series to test for the existence of a cointegrating relationship. Having confirmed stationarity of the differenced data and the existence of long-run relationships between the variables of interest, we then turn to estimating Equation (3) using non-linear least squares.

The interpretations and expected values of the coefficients are the following: coefficient α is a coefficient of price insulation, which represents the attempts by governments to adjust the level of protection to mitigate the impact of changes in world prices on domestic prices; it is likely to range from zero for those countries that do not insulate to negative one among the countries that attempt to insulate their domestic prices from global prices completely. Coefficient β represents the speed with which the government reverts to its target level of protection after having moved away from this level in previous periods; we expect β to be between negative one and zero. Coefficient γ determines the long run relationship between the country's protection and the global level of agricultural protection and we would generally expect it to be positive. Coefficient θ is the intercept in the long-run tariff equation. Finally, coefficient δ determines the time trend in protection for the particular product and may assume any value.

1.2. Measuring the implications of insulation for global price volatility

Our second methodology involves the use of a general equilibrium model to generate data for use in a Monte Carlo simulation to explore the role of the estimated coefficients in affecting global and domestic price volatility. We first define a baseline scenario in which all volatility in global prices is attributed to the observed volatility in yields from their trends. Then, we introduce an alternative scenario with protection being affected by change in global prices and also dependent on the previous change. By comparing the volatility of global and domestic prices in those two scenarios, we are able to attribute changes in the volatility of these series to the observed pattern of protection adjustment.

2. Data

We use several data sources in the analysis presented in this paper. The first was the database on Distortions to Agricultural Incentives which provides details on protection for a number of countries

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and a number of agricultural commodities (Anderson et al., 2008). This database provides estimates of border and domestic price levels, which jointly determine the level of protection, adjusted for costs of transportation and marketing to allow comparability. An unusual feature of this database is that some European countries are listed as members of the European Union for some products while other products and countries are listed separately. By necessity, we have followed the approach used in this database.

In the second stage of the analysis, we use the FAO's FAOSTAT database on yields and production of crops and primary animal products to measure annual changes in yields between 1993 and 2007 and to calculate historical levels of yield variations around their linear trends for a set of ten global regions. We consider eight important food commodities which are traded globally, are important parts of human diet, and for which we have sufficient data. They include three cereals (maize, rice and wheat), one oilseed crop (soybeans), one processed crop (sugar) and three types of meat (beef, pork and poultry).

We use the GTAP database (Version 8) described by Dimaranan (2006) to support our global computable general equilibrium model analysis. This database provides us with the necessary information on the structure of the global economy, particularly the agricultural production and trade by region that were essential for this analysis. To reduce the size of the model while focusing on the eight key food commodities, we aggregate the database into the ten regions and eleven commodities listed in Table 1. Because maize, soybeans, poultry and pork are not separate commodities in the GTAP database, we use MSplitCom and JoinCom utilities to make appropriate changes in the aggregation.

3. Results

Before turning to estimation results, we look at three examples presented in Fig. 1 through 3. Fig. 1 depicts annual price changes for border and domestic prices for maize in the United States. Clearly, changes in world prices for this commodity translate nearly one for one into changes in domestic prices—there appears to be almost no price insulation. When insulation is somewhat greater and there is a low propensity to return to the target tariff level, changes in domestic prices are often dampened without any significant autocorrelation, as with rice in Zambia in Fig. 2. A particularly strong example of insulation is shown in Fig. 3 for the price of wheat in India.

3.1. Testing for integration and cointegration

The results of testing for integration of the explanatory variables and for the existence of cointegrating relationships are presented in Table 2 through Table 9. In all but a few cases, we are unable to reject the null hypothesis of a unit root for the series in the levels. When we examine the first-differenced series, however, the null-hypothesis of a unit root is convincingly rejected in almost all cases, which suggests that the long-run run series are of the same order of integration (one) and that the differenced series are stationary.

Key cointegration statistics for the long run relationship in square brackets in Equation (3) serve as a test of whether a long run relationship exists between the tariff rate for each commodity and its

Table 1
Commodities considered and the regional aggregation used in modeling.

Commodities	Regions
Rice	Oceania
Wheat	East Asia
Sugar	Southeast Asia
Beef	South Asia
Maize	North America
Soybeans	Latin America
Pork	EU
Poultry	Middle East and North Africa
Other food	Sub-Saharan Africa
Manufactures	Rest of the World
Services	

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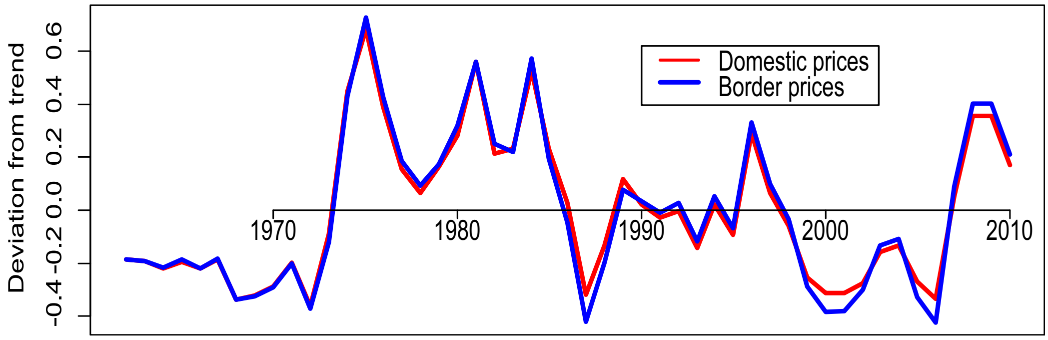

Fig. 1. Annual change in domestic and border prices, USA, maize.

explanatory variables. These tests are performed using augmented Dickey–Fuller tests on the residuals from OLS estimates of this relationship. The results are presented with a significance level of 10 percent indicated by a *, a significance level of 5 percent indicated by ** and *** indicating a significance level of 1 percent. These results suggest that the residuals of these regressions are much more frequently stationary than was the case in Greb et al. While they found only 55 percent of their price transmission relationships to be cointegrating, we find this in 64 percent of cases.

We proceed to estimate Equation (3) on the available protection data for eight food commodities, which are both important consumption items and highly traded, as identified in the previous section. The estimated coefficients are reported in Tables 10–17. The vast majority of the key coefficients (insulation coefficient α, and error-correction coefficient β) are estimated with significance levels of 10 percent or better. In each case where these coefficients are estimated with the lowest significance level (10 percent), they are of the expected sign. At a higher level of significance (5 percent) we observe one exception for the insulation coefficient, which is estimated to be positive for pork in New Zealand.

There is considerable diversity in the estimated coefficients, with cases where commodity prices are known to be strongly linked to world prices yielding low estimated price insulation coefficients. In the case of the USA, maize, poultry, soybeans and wheat have small estimated insulation coefficients, while sugar and dairy products have much higher insulation coefficients. Low insulation coefficients are observed for many product-country pairs where markets are known to be closely linked to world markets, such as Australia, Canada, Argentina and Thailand for wheat and maize. Higher degrees of price insulation are very common in European countries for many products, including the grains for which the EU sets variable tariffs based on the gap between external reference prices and an administered domestic price (European Commission, 2011). For wheat, for instance, the estimated insulation coefficient is −0.6, and the error correction term is very low at −0.16 suggesting a low rate of return to the political-economy target level of protection. Relatively high rates of price insulation are observed in cases such as rice (−0.82) and wheat (−0.76) in Japan, wheat in India (−0.86) and wheat in Norway (−0.76).

In addition to the coefficient estimates, we test for first-order autocorrelation of the residuals which show that the residuals in our estimation are generally uncorrelated. This is important for the validity of the significance tests in the regression, and in suggesting that the fairly parsimonious lag structure that we have used satisfactorily captures the dynamic behavior of the data.

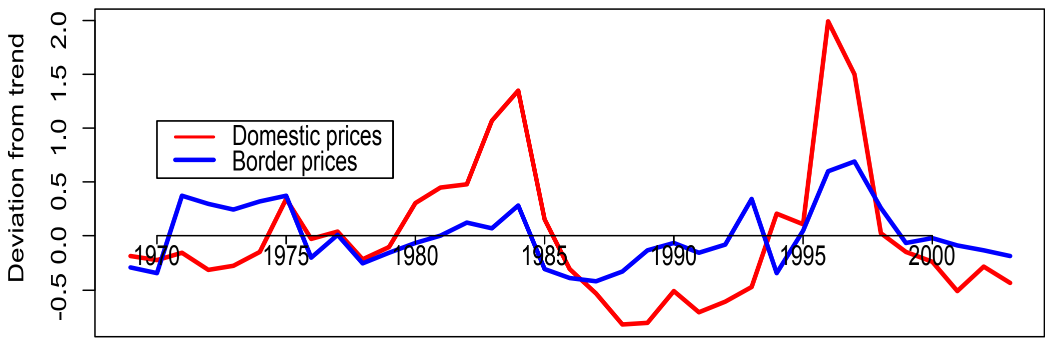

Fig. 2. Annual change in domestic and border prices, Zambia, rice.

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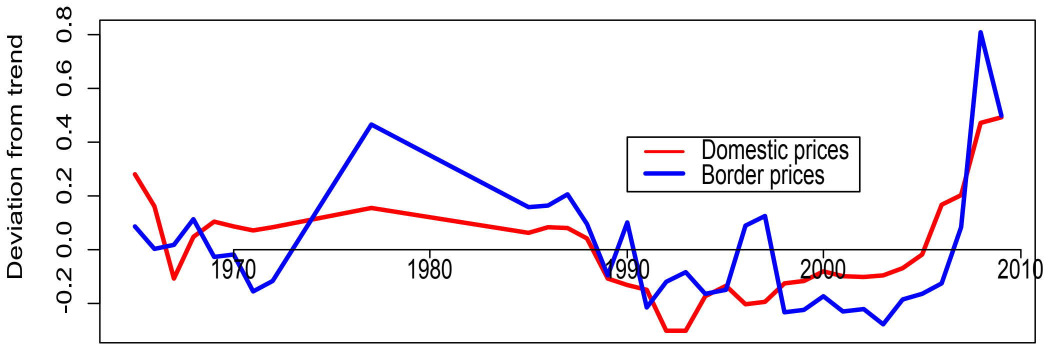

Fig. 3. Annual change in domestic and border prices, India, wheat.

We use FAO trade weights to aggregate coefficient estimates with a significance level of 1 percent or better into global estimates which we report in Table 18. The aggregate values suggest noticeable differences in the levels of price insulation and tariff correction for the eight commodities considered. We estimate that the level of tariff correction for all commodities is within a narrow range between −0.58 and −0.31 which represents the overall desire of governments to keep their tariffs around a given level. We also find insulation parameters to vary across commodities with the absolute value of the insulation coefficient being highest for sugar (−0.76) and pork (−0.59), and lowest for soybeans and maize (−0.2).

3.2. *Implications for the volatility of global prices*

In order to study the implications of the observed price insulation through protection and the subsequent tariff correction on the volatility of global and domestic prices, we use a global general equilibrium model (GTAP) to run a set of stochastic Monte Carlo simulations. This model is used primarily to weight the different countries and region, taking into account both the size of each market and the relevant elasticities of demand in that market. In each stochastic simulation, we run the model

Table 2
Integration and co-integration statistics for beef.[a]

Region	ADF test for tariff (log)	ADF test for change in tariff (log)	ADF test for global tariff (log)	ADF test for change in global tariff (log)	ADF test for border price (log)	ADF test for change in border price (log)	ADF test for cointegration factor
Argentina	−1.24	−5.12**	−3.02	−3.79**	−2.39	−5.06**	−1.42
Australia	−2.42	−5.72**	−2.86	−3.65**	−4.16**	−5.46**	−3.13
Brazil	−3.49*	−5.8**	−1.13	−3.18	−2.08	−3.71**	−3.94**
Canada	−1.89	−7.89**	−2.86	−3.65**	−2.51	−5.37**	−1.92
Chile	−2.5	−5.46**	−2.95	−3.98**	−3.77**	−5.87**	−2.58
Colombia	−3.47*	−5.74**	−3.02	−3.79**	−2.09	−5**	−3.42*
Ecuador	−3	−5.21**	−3.16	−3.82**	−3.61**	−4.94**	−4.68**
Egypt	−1.78	−4.34**	−3.02	−3.79**	−1.95	−4.34**	−1.44
EU	−3.08	−3.98**	−2.86	−3.65**	−3.02	−3.66**	−3.36*
Finland	0.63	−3.58**	−2.99	−3.21	−3	−3.75**	1.1
Iceland	−3.4*	−4.03**	−1.08	−3.4*	−2.88	−2.97	−3.42*
Italy	−2.94	−4.18**	−3.01	−3.66**	−2.45	−3.56**	−3.09
Japan	−1.61	−6.28**	−2.86	−3.65**	−2.14	−6.73**	−1.23
New Zealand	−2.06	−5.89**	−2.86	−3.65**	−2.51	−4.77**	−2.68
Philippines	−2.21	−5.21**	−3.05	−3.76**	−1.88	−4.74**	−3.22*
South Africa	−3.83**	−5.79**	−2.95	−3.98**	−3.11	−6.63**	−4.8**
South Korea	−3.64**	−4.9**	−2.9	−3.91**	−1.94	−5.41**	−3.83**
Sweden	−3.42*	−4.43**	−2.4	−3.37*	−2.58	−3.99**	−3.54*
Switzerland	−3.7**	−5.77**	−1.08	−3.4*	−1.97	−2.83	−3.73**
Taiwan	−3.31*	−7.52**	−2.85	−3.83**	−1.44	−5.71**	−3.53*
USA	−3.41*	−7.77**	−2.86	−3.65**	−2.64	−5.37**	−3.68**

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

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Table 3
Integration and co-integration statistics for maize.[a]

Region	ADF test for tariff (log)	ADF test for change in tariff (log)	ADF test for global tariff (log)	ADF test for change in global tariff (log)	ADF test for border price (log)	ADF test for change in border price (log)	ADF test for cointegration factor
Argentina	−2.97	−6.96**	−2.74	−3.4*	−2.78	−5.04**	−3.54**
Australia	−4.23**	−7.43**	−2.95	−3.7**	−2.88	−4.91**	−4.44**
Austria	−3.66**	−6.51**	−2.74	−3.4*	−1.91	−5.75**	−4.65**
Belgium	−3.62**	−7.79**	−2.74	−3.44*	−2.67	−6.29**	−3.65**
Brazil	−2.77	−7.53**	−2.46	−3.04	−2.04	−5.68**	−2.75
Canada	−4.36**	−7.5**	−2.46	−3.2*	−3.33*	−5.57**	−4.42**
Chile	−4.97**	−7.33**	−2.86	−3.65**	−2.84	−5.66**	−5.51**
Colombia	−3.77**	−6.42**	−2.74	−3.4*	−2.3	−6.15**	−3.8**
Ecuador	−3.23*	−7.25**	−3.15	−3.48*	−2.8	−5.32**	−5.42**
EU	−3.57**	−7.47**	−2.51	−3.17	−3.17	−5.93**	−3.72**
India	−3.59**	−4.98**	−1.71	−2.37	−2.07	−4.75**	−3.89**
Indonesia	−2.8	−7.38**	−2.85	−3.02	−3.39*	−5.4**	−2.75
Netherlands	−2.91	−5.81**	−3.04	−2.7	−2.57	−5.53**	−2.99
New Zealand	−1.66	−4.16**	−0.24	−3.18	−3.36*	−4.65**	−2.56
Pakistan	−3.05	−9.03**	−2.74	−3.4*	−2.56	−5.61**	−3.54**
Philippines	−3.49*	−6.56**	−2.81	−3.37*	−2.18	−4.5**	−4.02**
Portugal	−4.27**	−6.43**	−2.02	−2.87	−2.83	−4.26**	−4.26**
Spain	−3.74**	−6.11**	−2.02	−2.87	−2.88	−4.35**	−4.1**
Switzerland	−2.21	−5.06**	−0.24	−3.18	−2.38	−4.4**	−3.25*
Thailand	−4.27**	−5.78**	−2.89	−3.01	−3.87**	−5.12**	−4.28**
Uganda	−3.26*	−4.83**	−2.59	−4.66**	−5.81**	−6.74**	−3.26*
USA	−4.41**	−5.94**	−2.46	−3.2*	−3.28*	−5.74**	−4.43**
Zambia	−2.58	−5.76**	−2.86	−3.65**	−2.58	−6.49**	−2.59
Zimbabwe	−3.24*	−6.6**	−2.86	−3.65**	−3.39*	−6.89**	−3.26*

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

one thousand times, each time with a different set of yield changes drawn from the historically observed distributions of yields preserving historically observed covariation between different crops and regions.

We illustrate the role of the insulating and tariff-correcting policies by running the model twice. In the first run, which represents our base scenario without any insulating policies, we vary the level of

Table 4
Integration and co-integration statistics for pork.[a]

Region	ADF test for tariff (log)	ADF test for change in tariff (log)	ADF test for global tariff (log)	ADF test for change in global tariff (log)	ADF test for border price (log)	ADF test for change in border price (log)	ADF test for cointegration factor
Australia	−2.6	−8.28**	−2.01	−3.38*	−2.33	−4.76**	−2.95
Canada	−3.17	−9.59**	−2.01	−3.38*	−2.66	−9.14**	−3.84**
Ecuador	−3.08	−5.8**	−2.87	−3.81**	−2.45	−7.37**	−3.26*
EU	−2.64	−6.65**	−2.01	−3.38*	−2.29	−5.37**	−2.88
Finland	−0.47	−4.95**	−2.33	−3.03	−2.52	−3.69**	−0.68
Iceland	−3.26*	−4.45**	−0.18	−3.3*	−2.27	−4.86**	−3.99**
Japan	−0.79	−6.34**	−2.01	−3.38*	−0.8	−5.66**	−1.56
New Zealand	−5.1**	−7.67**	−2.01	−3.38*	−3.09	−5.48**	−5.3**
Philippines	−2.38	−6.17**	−2.34	−3.57**	−2.47	−5.39**	−2.48
South Korea	−4.27**	−6.85**	−2.63	−3.99**	−0.65	−5.04**	−4.23**
Spain	−2.43	−3.16	−2	−2.9	−2.4	−2.61	−3.77**
Sweden	−2.41	−5.15**	−2.29	−3.17	−2.11	−4.27**	−2.82
Taiwan	−1.8	−6.16**	−2.55	−3.9**	−0.94	−4.88**	−2.25
Thailand	−4.08**	−8.72**	−2.15	−3.21	−4.52**	−6.94**	−4.03**
USA	−4.9**	−7.99**	−2.01	−3.38*	−2.33	−9.46**	−5.25**

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

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Table 5
Integration and co-integration statistics for poultry.[a]

Region	ADF test for tariff (log)	ADF test for change in tariff (log)	ADF test for global tariff (log)	ADF test for change in global tariff (log)	ADF test for border price (log)	ADF test for change in border price (log)	ADF test for cointegration factor
Australia	−2.57	−7.4**	−2.4	−3.36*	−2.57	−4.11**	−2.86
Brazil	−5.24**	−7.28**	−0.53	−2.96	−1.63	−3.94**	−5.24**
Canada	−2.43	−5.78**	−2.4	−3.36*	−2.12	−6.46**	−2.77
Dominican R.	−2.95	−7.47**	−2.65	−3.54**	−2.93	−7.12**	−3.13
Ecuador	−2.15	−4.84**	−3.03	−3.68**	−2.5	−4.59**	−3.37*
EU	−3.76**	−7.06**	−2.4	−3.36*	−2.59	−5.17**	−3.98**
France	−3.75**	−6.47**	−2.65	−3.54**	−2.73	−4.97**	−3.94**
Iceland	−3.62**	−4.56**	−0.42	−3.16	−3.07	−5.6**	−4.83**
Indonesia	−3.31*	−8.1**	−3.07	−3.1	−4.07**	−4.93**	−4.69**
Italy	−3.81**	−7.11**	−2.65	−3.54**	−2.85	−5.66**	−3.97**
Japan	−2.49	−5.75**	−2.4	−3.36*	−1.4	−4.88**	−2.88
New Zealand	−3.27*	−6.6**	−2.4	−3.36*	−2.28	−5.56**	−3.51**
Philippines	−3.65**	−5.14**	−2.71	−3.52**	−2.04	−3.86**	−3.86**
South Africa	−4.06**	−6.16**	−2.74	−3.84**	−1.8	−4**	−3.89**
South Korea	−3.78**	−4.39**	−2.68	−3.79**	−1.09	−3.69**	−3.73**
Sweden	−3.53*	−6.64**	−2.64	−3.21*	−2.21	−4.5**	−3.66**
Switzerland	−2.41	−5.02**	−0.42	−3.16	−2.6	−4.13**	−3.57*
Taiwan	−2.67	−5.24**	−2.62	−3.7**	−0.86	−3.96**	−3.55**
Thailand	−3.19	−4.59**	−3.07	−3.1	−2.24	−4.86**	−3.1
USA	−3.93**	−7.12**	−2.4	−3.36*	−2.35	−6.59**	−3.9**

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

Table 6
Integration and co-integration statistics for rice.[a]

Region	ADF test for tariff (log)	ADF test for change in tariff (log)	ADF test for global tariff (log)	ADF test for change in global tariff (log)	ADF test for border price (log)	ADF test for change in border price (log)	ADF test for cointegration factor
Australia	−4.53**	−7.46**	−2.42	−2.92	−5.28**	−4.67**	−5.29**
Bangladesh	−3.23*	−5.29**	−1.1	−4.08**	−3.41*	−5.58**	−3.66**
Brazil	−4.12**	−6.3**	−1.36	−3.36*	−2.15	−5.42**	−4.17**
Colombia	−2.73	−5.3**	−2.6	−2.96	−3.38*	−5.82**	−2.99
Dominican R.	−3.58**	−6.12**	−2.6	−2.96	−2.47	−5.33**	−4.53**
Ecuador	−3.76**	−6.88**	−3.19	−3.14	−2.34	−4.99**	−3.72**
France	−2.2	−5.71**	−2.42	−2.92	−3.05	−5.55**	−2.94
Indonesia	−2.24	−3.94**	−1.15	−3.24*	−2.16	−3.54*	−2.8
Italy	−1.9	−5.71**	−2.42	−2.92	−3.04	−5.52**	−2.42
Japan	−1.61	−5.29**	−2.42	−2.92	−2.72	−6.18**	−2.65
Malaysia	−2.05	−6.02**	−2.58	−2.98	−3.74**	−6.07**	−2.14
Philippines	−3.2*	−6.58**	−2.67	−2.98	−3.23*	−5.67**	−4.72**
Portugal	−1.91	−6.16**	−2.42	−2.92	−3.07	−5.58**	−2.75
South Korea	−5.47**	−7.49**	−2.86	−3.24*	−2.46	−5.53**	−5.85**
Spain	−2.12	−6.01**	−2.42	−2.92	−3.05	−5.56**	−2.97
Sri Lanka	−3.56**	−6.09**	−2.6	−2.96	−4.4**	−6.57**	−3.85**
Taiwan	−2.19	−5.44**	−2.78	−3.16	−2.01	−5.42**	−3.41*
Thailand	−4.61**	−7.68**	−2.63	−2.63	−3.46*	−4.61**	−4.86**
Uganda	−1.8	−4.98**	−2.97	−4.56**	−3.65**	−4.71**	−3.74**
USA	−3.23*	−4.67**	−2.42	−2.92	−2.74	−4.85**	−3.43*
Zambia	−2.25	−3.82**	−3.57**	−3.01	−3.31*	−5.07**	−2.59

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

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Table 7
Integration and co-integration statistics for soybeans.[a]

Region	ADF test for tariff (log)	ADF test for change in tariff (log)	ADF test for global tariff (log)	ADF test for change in global tariff (log)	ADF test for border price (log)	ADF test for change in border price (log)	ADF test for cointegration factor
Argentina	−2.22	−5.39**	−1.7	−3.16	−2.78	−4.57**	−2.95
Australia	−4.19**	−7.83**	−2.63	−3.64**	−5.43**	−7.24**	−4.41**
Brazil	−3.16	−6.17**	−2.67	−3.47*	−2.65	−5.57**	−3.24*
Canada	−4.42**	−10.28**	−2.63	−3.64**	−2.24	−4.87**	−4.43**
Colombia	−2.22	−5.15**	−2.73	−3.69**	−2.26	−5.97**	−2.13
Ecuador	−3.69**	−5.22**	−2.88	−3.55*	−2.46	−4.99**	−3.86**
France	−2.26	−5.04**	−2.16	−3.5*	−2.98	−6.47**	−2.29
Indonesia	−2.61	−4.58**	−2.99	−3.29*	−3.79**	−4.29**	−3.01
Italy	−2.26	−5.04**	−2.16	−3.5*	−2.81	−6.45**	−2.29
Japan	−1.41	−4.21**	−1.04	−3.05	−1.57	−4.43**	−1.95
South Korea	−4.74**	−7.53**	−2.7	−3.71**	−2.99	−5.12**	−4.88**
Spain	−2.57	−5.34**	−3.08	−3.23*	−2.63	−5.35**	−2.34
USA	−3.67**	−4.94**	−2.63	−3.64**	−2.38	−5.27**	−3.82**
Zambia	−2.63	−6.61**	−3.41*	−3.43*	−4.2**	−6.2**	−2.59
Zimbabwe	−1.02	−4.79**	−3.17	−3.46*	−2.89	−5.2**	−1.02

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

yields and hold protection constant. In the second run, we run the model for two periods and allow the countries to adjust their protection based on current price and the changes in their tariffs in the previous period. By comparing the volatility of prices between the two stochastic runs we are able to attribute any differences in volatility of world prices to price-insulating policies.

Table 8
Integration and co-integration statistics for sugar.[a]

Region	ADF test for tariff (log)	ADF test for change in tariff (log)	ADF test for global tariff (log)	ADF test for change in global tariff (log)	ADF test for border price (log)	ADF test for change in border price (log)	ADF test for cointegration factor
Australia	−2.66	−5.85**	−2.14	−3.34*	−2.28	−4.36**	−2.84
Austria	−3.59**	−6.07**	−2.44	−3.05	−2.83	−4.95**	−3.83**
Bangladesh	−3.52*	−5.09**	−1.53	−3.38*	−4.1**	−4.67**	−3.3*
Brazil	−4.94**	−5.01**	−1.92	−2.93	−4.25**	−4.8**	−4.95**
Chile	−4.04**	−6.06**	−2.87	−3.22	−3.84**	−7.67**	−4.39**
Colombia	−3.21*	−5.59**	−2.54	−3.57**	−2.75	−5.96**	−3.3*
Dominican R.	−1.83	−5.41**	−2.54	−3.57**	−3.87**	−6.78**	−1.83
Ecuador	−3.72**	−6.43**	−2.96	−3.72**	−3.99**	−5.85**	−3.99**
EU	−4.28**	−7.14**	−2.14	−3.34*	−3.39*	−5.54**	−4.52**
Finland	−3.12	−5.77**	−2.62	−2.92	−3.02	−5.2**	−4.44**
Indonesia	−3.07	−4.75**	−3.25*	−3.41*	−2.98	−4.45**	−3.65**
Japan	−4.45**	−4.5**	−0.6	−2.73	−1.88	−4.45**	−4.71**
Kenya	−3.58**	−5.05**	−2.53	−3.46*	−2.77	−5.04**	−3.58**
Mozambique	−2.53	−4.06**	−1.45	−3.15	−4.23**	−4.71**	−2.56
Pakistan	−3.3*	−5.64**	−2.56	−3.51*	−3.46*	−5.82**	−3.46*
Philippines	−3.36*	−5.6**	−2.56	−3.51*	−2.99	−5.68**	−3.34*
Portugal	−3.92**	−7.57**	−2.54	−3.57**	−2.74	−6.18**	−3.93**
South Africa	−3.15	−4.83**	−2.96	−3.81**	−2.86	−5.03**	−3.17
Spain	−4.6**	−7.12**	−2.54	−3.57**	−3.13	−5.49**	−4.75**
Sweden	−2.69	−5.87**	−2.86	−2.85	−2.88	−5.65**	−3.73**
Switzerland	−2.93	−6.3**	−0.6	−2.73	−2.34	−4.82**	−4.14**
Thailand	−3.04	−5.17**	−2.54	−3.11	−2.59	−3.96**	−3.15
Uganda	−4.81**	−6.45**	−2.72	−3.89**	−2.32	−4.66**	−4.91**
USA	−3.18	−5.03**	−2.14	−3.34*	−3.01	−5.62**	−3.27*

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

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Table 9
Integration and co-integration statistics for wheat.[a]

Region	ADF test for tariff (log)	ADF test for change in tariff (log)	ADF test for global tariff (log)	ADF test for change in global tariff (log)	ADF test for border price (log)	ADF test for change in border price (log)	ADF test for cointegration factor
Argentina	−2.33	−5.1**	−2.32	−3.3*	−3.54**	−6.18**	−2.62
Australia	−3.96**	−6.12**	−2.07	−2.93	−3.27*	−5.57**	−4.35**
Austria	−2.29	−6.19**	−1.88	−3.26*	−1.86	−5.09**	−2.58
Bangladesh	−4.55**	−6.21**	−1.43	−3.74**	−2.73	−4.83**	−4.87**
Brazil	−4.42**	−6.42**	−2.19	−2.78	−3.14	−5.93**	−4.58**
Canada	−3.14	−6.42**	−2.07	−2.93	−3.25*	−5.29**	−3.07
Chile	−4.23**	−6.6**	−2.19	−3.73**	−2.86	−5.57**	−4.3**
Colombia	−3.94**	−6.12**	−2.32	−3.3*	−3.13	−5.56**	−3.97**
EU	−2.86	−6.76**	−2.07	−2.93	−2.85	−6.6**	−3.04
Finland	−1.7	−5.38**	−2.15	−2.92	−1.98	−4.63**	−1.78
India	−3.43*	−7.98**	−1.66	−2.24	−2.08	−5.43**	−3.45*
Japan	−4.91**	−6.84**	−2.07	−2.93	−4.48**	−5.84**	−5.12**
New Zealand	−1.81	−4.32**	−2.07	−2.93	−3.33*	−4.8**	−2.8
Norway	−4.15**	−7.19**	−0.48	−2.7	−3.42*	−6.69**	−4.83**
Pakistan	−4.51**	−7.42**	−2.47	−3.32*	−3.38*	−6.58**	−5.24**
South Korea	−2.97	−6.05**	−2.14	−3.68**	−2.95	−5.07**	−3
Sweden	−2.04	−5.39**	−2.19	−3.48*	−1.95	−4.84**	−2.13
Switzerland	−1.72	−5.56**	−0.48	−2.7	−3.04	−5.05**	−3.04
USA	−3.37*	−5.82**	−2.07	−2.93	−4.33**	−6.61**	−3.28*
Zambia	−1.68	−4.28**	−2.75	−3.62**	−2.81	−5.6**	−1.79
Zimbabwe	−1.2	−4.66**	−2.14	−3.68**	−2.32	−5.42**	−1.25

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

Table 10
Beef coefficient estimates.[a]

Region	Short-run impact of border price (insulation)	Error correction	Long-run relationship with global protection	Trend coefficient for tariff	Intercept of long-run tariff	Number of observations	First-order autocorrelation
Argentina	−0.14**	−0.27**	−0.45	0.01	−11.94	48	−0.04
Australia	−0	−0.23*	−0.01	−0	0.51	49	0
Brazil	−0.13	−0.77***	−0.72	0.01	−11.67	30	−0.01
Canada	−0.03**	−0.38***	−0.03	0	−0.01	49	−0.25*
Chile	−0.18***	−0.33***	−1.29**	0.01***	−27.36***	44	−0.05
Colombia	−0.43***	−0.36***	−0.24	−0	8.24	48	0.06
Ecuador	−0.79***	−0.49***	−2.67***	0.02*	−30.26*	39	−0.15
Egypt	−1.25***	−0.29***	−4.08***	0.03***	−57.67***	48	−0.1
EU	−0.62***	−0.08	−0.79	−0.01	24.29	49	−0.11
Finland	−0.06	−0.03	−6.11	−0.03	62.54	37	−0.25
Iceland	−0.53***	−0.14	−0.15	−0.04	74.21	31	0.32*
Italy	−0.33***	−0.23**	1.07	−0.01	20.53	46	0.28
Japan	−0.52***	−0.21**	−1.76	−0.01	14.74	49	−0.1
New Zealand	−0.05*	−0.21**	−0.33*	−0	0.22	49	0.11
Philippines	−0.06*	−0.25**	0.25	0	−0.82	47	0.08
South Africa	−0.63***	−0.48***	−0.68	−0.01	11.64	44	0.28*
South Korea	−0.41**	−0.32***	−1.38	0.01	−25.51	43	0.21
Sweden	−0.64***	−0.13*	−0.48	−0.01	17.18	41	−0.14
Switzerland	−0.43**	−0.35*	0.13	−0.03*	56.71*	31	0.1
Taiwan	−0.71***	−0.36***	−1.11**	0.01**	−20**	41	0.23
USA	0	−0.47***	−0.02	−0**	0.81**	49	−0.2

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

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Table 11
Maize coefficient estimates.[a]

Region	Short-run impact of border price (insulation)	Error correction	Long-run relationship with global protection	Trend coefficient for tariff	Intercept of long-run tariff	Number of observations	First-order autocorrelation
Argentina	−0.26**	−0.47***	0.58	−0.01	12.43	48	−0
Australia	0	−1.04***	0	−0	0	46	−0
Austria	−0.54***	−0.26**	−0.44	−0.01	14.89	48	0.07
Belgium	−0.14***	−0.23**	−0.93	−0	3.31	45	0.1
Brazil	−0.25**	−0.4***	1.03	−0	2.88	44	−0.12
Canada	−0.09***	−0.29***	−0.05	0	−0.22	49	0.13
Chile	−0.42***	−0.6***	0.3	0	−2.62	44	−0.14
Colombia	−0.51***	−0.42***	−0.92*	0.01***	−25.09***	48	−0.08
Ecuador	−0.25	−0.66***	−1.08***	0.01	−10.33	39	0.08
EU	−0.51***	−0.18*	−1.33	−0	4.9	48	0.08
India	−0.36*	−0.55**	0.27	−0.01	17.97	38	−0.08
Indonesia	−0.37***	−0.53***	−0.32	0.01	−12.01	39	0.15
Netherlands	−0.09*	−0.2	−0.83	−0	6.52	31	0.19
New Zealand	0.01	−0.16	0	−0	3.28	31	−0.07
Pakistan	−0.54**	−0.59***	0.11	−0	3.32	48	0.02
Philippines	−0.68***	−0.55***	0.36	0	−9.05	47	−0.13
Portugal	−0.31***	−0.69***	−0.12	−0	4.23	30	−0.14
Spain	−0.37***	−0.5***	−0.48	−0	4.03	30	−0.02
Switzerland	−0.68***	−0.05	−2.85	−0.03	65.56	31	0.16
Thailand	−0.11	−0.72***	0.01	−0	3.84	37	0.06
Uganda	−0.13*	−0.56***	−0.26	−0	1.35	39	−0.12
USA	−0.12***	−0.27**	−0.03	0	−1.35	49	−0.04
Zambia	−0.81***	−0.37***	2.05	−0.01	15.2	44	0.09
Zimbabwe	−0.82***	−0.28***	−0.22	−0.03*	50.83*	44	−0.01

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

In Table 19, we compare the standard deviations of global prices for the eight focus commodities. Clearly, the introduction of price-insulating trade policy greatly increases price volatility in each case, especially in the cases of beef and pork, and sugar where the standard deviation of the global price trebles and in the case of rice where the standard deviation more than doubles as a result of the highest insulation coefficients.

Table 12
Pork coefficient estimates.[a]

Region	Short-run impact of border price (insulation)	Error correction	Long-run relationship with global protection	Trend coefficient for tariff	Intercept of long-run tariff	Number of observations	First-order autocorrelation
Australia	−0	−0.78***	0*	−0	0.01	49	−0.09
Canada	−0.04*	−0.62***	0.09	−0	1.3	49	−0.06
Ecuador	−0.49***	−0.33***	−1.31	0.01	−27.2	39	−0.05
EU	−0.49***	−0.32***	−0.45	−0.02***	35.68***	49	−0.07
Finland	−0.08*	0.05	0.85	−0	1.88	39	−0.01
Iceland	−0.75***	−0.26*	1.73	−0.04**	85.1**	31	0.21
Japan	−0.54***	−0.07	−1.87	0.03	−63.3	49	−0.1
New Zealand	0.26**	−0.75***	0.34	−0	2.7	49	−0.01
Philippines	−0.61***	−0.23***	−0.61	−0	1.1	47	−0.17
South Korea	−0.36	−0.38***	−0.54	−0	8.62	43	−0.08
Spain	−0.69***	−0.08	2.79	−0	1.44	33	−0.27
Sweden	−0.51***	−0.32***	−1.23***	−0.01***	29.3***	40	−0.05
Taiwan	−0.57***	−0.24**	1.58*	−0.01	21.57	41	0.03
Thailand	−0.78***	−0.49***	−0.14	−0	1.77	39	0.01
USA	−0	−0.56***	0.03	−0	0.2	49	0.14

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

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Table 13
Poultry coefficient estimates.[a]

Region	Short-run impact of border price (insulation)	Error correction	Long-run relationship with global protection	Trend coefficient for tariff	Intercept of long-run tariff	Number of observations	First-order autocorrelation
Australia	0	−0.78***	0*	−0	0	49	−0.08
Brazil	0.08	−1.12***	0.06	0.01	−11.39	30	−0.01
Canada	−0.41***	−0.2**	−0.48	−0	3.41	49	0.16
Dominican R.	−0.65***	−0.32***	1.98*	−0.02***	48.43***	48	−0.15
Ecuador	−0.82***	−0.16	−0.61	−0.05*	104.01*	39	0.14
EU	−0.48***	−0.41***	0.27	−0	6.29	49	0.09
France	−0.43***	−0.35***	0.17	−0	1.45	48	−0.13
Iceland	−0.35**	−0.31*	2.17*	−0.03**	69.35**	31	0.02
Indonesia	−0.64**	−0.7***	−1.53**	0.01**	−27.68*	38	0.08
Italy	−0.15**	−0.77***	0.46	−0	5.26	48	0.27*
Japan	−0.16	−0.27**	0.25	−0	3.73	49	0.07
New Zealand	0.08	−0.42***	0.19	−0	3.4	49	0.03
Philippines	−0.58***	−0.33***	−0.07	0	−6.97	47	0.15
South Africa	−0.07	−0.48***	0.13	0.01*	−14.83*	44	0.17
South Korea	−0.47**	−0.28***	0.47	0	0.3	43	0.2
Sweden	−0.15**	−0.75***	0.02	0**	−9.55**	44	−0.04
Switzerland	−0.17*	−0.25	0.74	−0.01	25.59	31	0.04
Taiwan	−0.59***	−0.31***	0.67	0.03***	−50.32***	41	0.09
Thailand	−0.69***	−0.37***	−1.16*	0.02***	−36.24***	38	0.18
USA	−0.11***	−0.59***	−0.09	0	−0.95	49	−0.02

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

Table 14
Rice coefficient estimates.[a]

Region	Short-run impact of border price (insulation)	Error correction	Long-run relationship with global protection	Trend coefficient for tariff	Intercept of long-run tariff	Number of observations	First-order autocorrelation
Australia	−0.02	−0.61***	−0.28**	−0**	5.88***	49	0.08
Bangladesh	−0.43**	−0.81***	0.71	−0.01	22.35	33	0.18
Brazil	−0.28***	−0.79***	0.18	0.01*	−19.98*	37	−0.16
Colombia	−0.54***	−0.16**	−0.9	0.02	−45.88	48	−0.1
Dominican R.	−0.61***	−0.35***	2.24**	−0	7.6	48	0.15
Ecuador	−0.77***	−0.45***	−0.35	0.01	−20.17	39	0.07
France	−0.76***	−0.03	−10.09	0.03	−56.87	49	−0.03
Indonesia	−0.17	−0.4**	−0.02	0	−0.9	34	−0.07
Italy	−0.7***	−0.34***	1.26	−0.01	13.88	49	0.43**
Japan	−0.82***	−0.01	−59.48	−0.12	267.81	49	0.23
Malaysia	−0.79***	−0.1**	−2.02	0.02	−30.06	46	0.04
Philippines	−0.69***	−0.27***	1.71**	−0	4.09	45	0
Portugal	−0.25**	−0.12	−1.58	0.01	−14.03	49	−0.17
South Korea	−0.89***	−0.28***	−0.41	0.04***	−74.9***	43	0.12
Spain	−0.8***	−0.03	−9.63	0.03	−54.37	49	−0.01
Sri Lanka	−0.38***	−0.52***	0.57	−0	6.26	48	0.02
Taiwan	−0.45***	−0.32***	1.03**	0.03***	−64.4***	41	0.15
Thailand	−0.28***	−0.82***	0.23	0.01*	−13**	37	0.07
Uganda	−0.09*	−0.76***	−1.96***	0.01***	−14.06***	40	0.02
USA	−0.27***	−0.12*	0.09	0	−4.31	49	−0.01
Zambia	−0.57**	−0.36**	−0.48	0.01	−19.78	35	0.12

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

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Table 15
Soybean coefficient estimates.[a]

Region	Short-run impact of border price (insulation)	Error correction	Long-run relationship with global protection	Trend coefficient for tariff	Intercept of long-run tariff	Number of observations	First-order autocorrelation
Argentina	−0.03	−0.49***	1.29**	−0.02**	31.44**	32	0.01
Australia	−0	−1.05***	0	0	−0	49	0
Brazil	−0.27***	−0.61***	0.56	0	−2.32	44	−0.06
Canada	−0.02*	−0.8***	−0.02	0	−0.47	49	−0.03
Colombia	−0.58***	−0.25***	−1.09*	−0	5.76	48	0.04
Ecuador	−0.74***	−0.72***	−0.39	−0.01	13.66	39	0
France	−0.33	−0.25**	−2.69	0.01	−26.28	37	−0.09
Indonesia	−0.44***	−0.45***	−0.71	0	−6.39	39	−0.04
Italy	−0.35	−0.25**	−2.72	0.01	−27.04	37	−0.1
Japan	−0.41***	−0.08	−2.68	−0.02	34.08	31	0.17
South Korea	−0.36**	−0.53***	−0.14	0.05***	−92.96***	44	0.01
Spain	−0.58**	−0.31***	−2.46	0.01	−12.22	35	−0.04
USA	−0.09***	−0.24**	0.18	0	−0.74	49	0.08
Zambia	−0.97***	−0.26**	3.98	−0.02	45.51	32	0.14
Zimbabwe	−0.81***	−0.62***	−0.08	−0.03**	55.85**	36	0.12

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

We also report the distributions of the domestic price volatility in the ten regions considered in our model in Fig. 4. As this figure shows, there is little change in domestic price volatility due to insulation. The figure therefore makes the very important point that even though countries may appear to insulate themselves from volatile global prices, in fact though their common action they drive up volatility

Table 16
Sugar coefficient estimates.[a]

Region	Short-run impact of border price (insulation)	Error correction	Long-run relationship with global protection	Trend coefficient for tariff	Intercept of long-run tariff	Number of observations	First-order autocorrelation
Australia	−0.12***	−0.21**	−0.13	−0	2.63	49	−0
Austria	−0.74***	−0.1	−6.28	0.01	−25.26	39	−0.18
Bangladesh	−0.83***	−0	−32.37	0.38	−738.94	33	−0.38**
Brazil	−0.57***	−0.31***	−0.4	0.05***	−92.6***	40	−0.17
Chile	−0.68***	−0.13	−0.16	−0.01	28.13	30	−0.14
Colombia	−0.85***	−0.13***	−2.17	0.02	−47.94	48	−0.09
Dominican R.	−0.49***	−0.15*	4.36	−0.02	47.8	48	−0.26*
Ecuador	−0.56***	−0.43***	−1.79	0.02*	−49.13*	39	−0.23
EU	−0.75***	−0.13**	−3.38	0.01	−18.75	49	−0.12
Finland	−0.37***	−0.12	−2.08	−0	2.68	38	0.17
Indonesia	−0.57***	−0.1	2.34	−0.03	66.53	34	0.26
Japan	−0.59***	0.01	19.66	0.41	−833.37	31	0.11
Kenya	−0.84***	−0.13*	−2.21	0.03	−65.8	45	0.05
Mozambique	0.1	−0.39**	3.66	0.05	−98.29	33	0.11
Pakistan	−0.82***	−0.1**	−1.29	0	−5.19	47	0.1
Philippines	−0.82***	−0.09	−4.18	0.03	−59.4	47	−0.06
Portugal	−0.79***	−0.13**	−4.46	0.04	−82.96	48	−0.03
South Africa	−0.68***	−0.22**	0.92	−0	5.23	40	−0.17
Spain	−0.78***	−0.18***	−0.99	−0.01	14.65	48	−0.2
Sweden	−0.26***	−0.18	−0.43	0	−0.05	38	0.05
Switzerland	−0.21**	−0.24*	3.78*	−0.07**	132.75**	31	0.29
Thailand	−0.17**	−0.34***	−0.31	0.01	−17.84	37	0.04
Uganda	0.19	−0.89***	−1.05	0.01	−24.82	39	−0.24
USA	−0.51***	−0.22***	0.5	−0	7.26	49	−0.01

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

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Table 17
Wheat coefficient estimates.[a]

Region	Short-run impact of border price (insulation)	Error correction	Long-run relationship with global protection	Trend coefficient for tariff	Intercept of long-run tariff	Number of observations	First-order autocorrelation
Argentina	−0.06	−0.41***	−0.06	−0	2.66	48	−0.09
Australia	−0.09***	−0.57***	0.11	−0**	4.09**	49	−0.03
Austria	−0.76***	−0.17***	0.8	0.02*	−45.88*	34	0.22
Bangladesh	−0.45**	−0.95***	0.3	−0.01	17.97	33	−0.37
Brazil	−0.22*	−0.52***	−0.27	−0.01	14.97	44	0.13
Canada	−0.02*	−0.35***	−0.2**	−0	0.24	49	0.17
Chile	−0.5***	−0.66***	−0.13	0.01	−13.3	44	−0.11
Colombia	−0.65***	−0.35***	−0.49	−0	1.9	48	0.1
EU	−0.6***	−0.16**	−0.66	−0.01	17.3	49	0.14
Finland	−0.65***	−0.08	−2.38	−0.01	29.9	39	−0.03
India	−0.86***	−0.2*	0.77	0	−5.97	34	0.17
Japan	−0.76***	−0.18**	−1.56	0.02	−47.05	49	−0.25*
New Zealand	0.01	−0.09	−0.01	−0**	6.51**	49	0.19
Norway	−0.76***	−0.15	−4.32	0.02	−28.83	31	−0.08
Pakistan	−1.22***	−0.39***	0.4	−0	4.52	47	0.04
South Korea	−0.6***	−0.2**	−1.5	0.02*	−46.33*	43	0.12
Sweden	−0.57***	−0.13*	−1.2	0.01	−25.85	40	0.03
Switzerland	−0.62***	−0.17	0.06	−0.05**	92.25**	31	0.02
USA	−0.18***	−0.19*	−0.19	−0	5.12	49	−0.16
Zambia	−1.13***	−0.16	1.74	0.03	−54.05	39	0.06
Zimbabwe	−0.26	−0.49***	−0.3	−0.04***	78.67***	43	−0.04

[a] *, **, *** imply significance of 0.10, 0.05 and 0.01.

enough to achieve nothing in lowering domestic price volatility—a point emphasized by Martin and Anderson (2012).

4. Discussion

The findings of our stochastic simulation suggest that the observed price-insulating policies raise the volatility of global prices from the levels that would apply in the absence of this policy response. While we are unable to observe global price volatility without the influence of the existing insulating policies, especially storage, to confirm the finding of our model, we can compare the model results with the empirically observed global price volatility, which we do in the last column of Table 19. This comparison serves as an important validation of our approach because it suggests that the introduction of the estimated endogenous insulation policies to a standard CGE model succeeds in bringing the modeled volatility, at least of the eight focus commodities, closer to the levels observed in the world. There are many other sources of volatility, such as those associated with stock dynamics, which are not captured in our model and which presumably raise global price volatility above the volatility captured in our very simple model.

Table 18
Global trade-weighted coefficients.

Commodity	Insulation	Correction
Beef	−0.35	−0.41
Maize	−0.2	−0.47
Pork	−0.59	−0.58
Poultry	−0.31	−0.81
Rice	−0.43	−0.68
Soybean	−0.2	−0.55
Sugar	−0.76	−0.16
Wheat	−0.3	−0.47

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Table 19
Global price volatility, standard deviation in percentage points.

Commodity	Simulated zero insulation	Simulated insulation	Observed
Rice	2.1	5.0	7.9
Wheat	3.8	6.7	8.2
Sugar	2.7	8.2	11.7
Beef	1.6	2.3	5.1
Maize	5.1	7.4	8.1
Soybeans	3.8	4.9	6.8
Pork	0.6	0.9	NA
Poultry	1.9	3.3	3.5

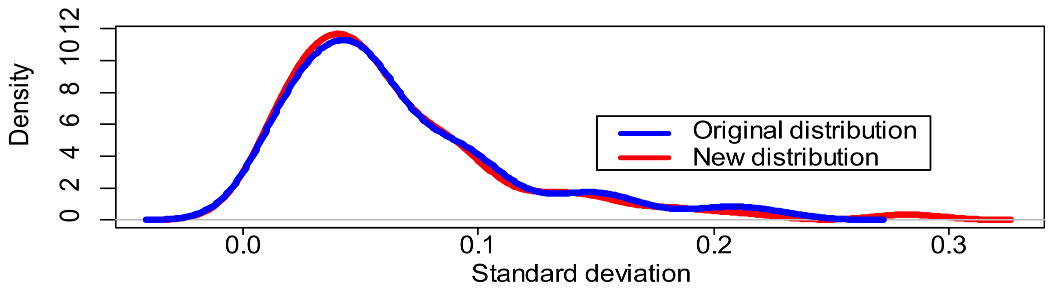

Fig. 4. Distribution of standard deviations of domestic prices (eight commodities, ten regions).

5. Conclusions

More volatile and generally higher food prices in recent years have raised genuine concerns whether food prices might be becoming less stable with potentially negative impacts on the growth and development of the poorest nations. Examination of domestic and world price data suggests that, in some cases, price transmission appears to be fairly direct, while in others seemingly quite complex patterns of price insulation emerge, with prices initially insulated but increases in world prices subsequently passed through to domestic prices. A key question examined in this paper is the extent to which these price changes contribute to changes in the volatility of world prices, and in their intertemporal correlations.

We find that the price and trade policy variables of most interest appear to be integrated series, making inferences based on OLS regression unreliable. After testing for the existence of cointegrating relationships we use a modified version of the Error-Correction Model that is widely used to estimate price transmission between border and domestic prices. The results of this analysis confirm that policies of price insulation are consistently applied to a wide range of commodities and that initial insulation against a shock tends to be followed by adjustments designed to reduce the gap between the previous level of protection and policy makers' target rates of protection. Simulating the impacts of the estimated policy response, we confirm that they greatly raise global price volatility—often doubling it.

Aside from destabilizing global prices, we also find that the observed policies do very little to stabilize domestic prices. Even though each country that attempts to insulate itself from the global market succeeds by having lower than global level of price volatility, the collective action of all insulating countries causes global price volatility to rise so much that the insulating countries end up with domestic prices nearly as volatile as they would have been without any insulation at all.

This paper reinforces the findings of earlier work, e.g. Anderson et al. (2012), Martin and Anderson (2012), that price insulating behavior by individual countries seeking to insulate their domestic markets from changes in world prices increases the volatility of world prices while being unsuccessful in reducing the volatility of domestic prices because of the collective action (or beggar-thy-neighbor) nature of this type of intervention. In addition, it shows that price-insulating behavior creates more complex intertemporal behavior in the world prices of key staple commodities such as wheat, rice and maize as countries seek to return their protection to the levels consistent with their domestic political-economy equilibrium.

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